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ABSTRACT

Contents of this volume of studies analyzing the causes of the alarming growth in illegitimacy, families lacking a father, and welfare caseloads, include the following studies: "The Family, Poverty, and Welfare Programs: An Introductory Essay on Problems of Analysis and Policy," Robert I. Lerman; "The Impact of Welfare Payment Levels on Family Stability," Marjorie Honig; "Income Supplements and the American Family," Phillips Cutright and John Scanzoni; "Illegitimacy and Income Supplements," Phillips Cutright; "Participation in the Aid to Families with Dependent Children Program (AFDC)," Barbara Boland; "Treatment of Families Under Income Transfer Programs," Irene Cox; "Poverty, Living Standards, and Family Well-Being," Lee Rainwater; "Child Welfare, Parental Responsibility, and the State," Harry D. Krause; "The Concept of Family in the Poor Black Community," Carol B. Stack and Herbert Semmel; "Black Family Structure: Myths and Realities," Andrew Billingsley; and, "Family and Community Life in the Working Class," Marc Fried and Ellen Fitzgerald. (JM)

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STUDIES IN PUBLIC WELFARE

PAPER No. 12 (Part I)

THE FAMILY, POVERTY, AND WELFARE PROGRAMS: FACTORS INFLUENCING FAMILY INSTABILITY

A VOLUME OF STUDIES

PREPARED FOR THE USE OF THE

SUBCOMMITTEE ON FISCAL POLICY

OF THE

JOINT ECONOMIC COMMITTEE
CONGRESS OF THE UNITED STATES

NOVEMBER 4, 1973

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LETTERS OF TRANSMITTAL

OCTOBER 25, 1973.

To the members of the Joint Economic Committee:

Transmitted herewith is a volume of studies entitled "The Family, Poverty, and Welfare Programs: Factors Influencing Family Instability." This is Paper No. 12 (Part I) in the series *Studies in Public Welfare*, prepared for the Subcommittee on Fiscal Policy as part of its comprehensive review of the Nation's welfare-related programs.

The views expressed in these studies are those of the authors and do not necessarily represent the views of the Subcommittee on Fiscal Policy, the Joint Economic Committee, or the committee staff.

WRIGHT PATMAN,
Chairman, Joint Economic Committee.

OCTOBER 23, 1973.

Hon. WRIGHT PATMAN,
Chairman, Joint Economic Committee,
U.S. Congress, Washington, D.C.

DEAR MR. CHAIRMAN: Transmitted herewith is a volume of studies entitled "The Family, Poverty, and Welfare Programs: Factors Influencing Family Instability," Paper No. 12 (Part I) in the subcommittee's review of public welfare programs.

The studies in this volume analyze the causes of the alarming growth in illegitimacy, families lacking a father, and welfare caseloads. The authors attempt to answer questions vitally important to our understanding of poverty and to the redesign of public welfare programs. Among them are:

Does the welfare system actually increase the breakup of families?

What is the effect of increased income on illegitimacy and disruption of families?

What factors account for the increase in illegitimacy and in families headed by mothers, and can we expect these increases to subside in the future?

How much of the recent increase in welfare caseload resulted from added participation by eligibles, from the increase in eligibility due to higher benefits, and from increases in family splitting?

What share of welfare-eligible families headed by women still do not receive welfare assistance?

These questions are controversial and difficult to answer with precision. But they are too important to ignore. The authors do an admirable job of carefully addressing these subjects. The papers represent the views of their authors and do not necessarily represent the views of the Subcommittee on Fiscal Policy, individual members thereof, or the subcommittee staff.

This volume was edited by Robert I. Lerman. Alair A. Townsend provided general direction and compiled many of the papers.

MARTHA W. GRIFFITHS,
Chairman, Subcommittee on Fiscal Policy.

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Part II

THE FAMILY, POVERTY, AND WELFARE PROGRAMS: AN INTRODUCTORY ESSAY ON PROBLEMS OF ANALYSIS AND POLICY

By ROBERT I. LERMAN*

I. INTRODUCTION

Problems of family breakdown jumped to the center stage of public debate upon publication of the Moynihan report[1] in 1965. The report drew emotional and extensive criticism, and in the years since political leaders and social science researchers have paid increasing attention to family structure of low-income groups. Their heightened interest arises chiefly from (*a*) the dramatic growth in family welfare recipients (the aid to families with dependent children program—AFDC) and (*b*) the intractability of poverty in families without a father, a rising proportion of America's families. Not only do broken families contribute to the slowdown in reducing poverty, but family disorganization is a key reason for the degree of income inequality between white families and those of minority races. Finally, some observers have concluded that poverty and growing numbers of broken families lead to increased crime. Directly and indirectly, then, family breakdown is seen as a problem related to poverty, racial inequality, and crime.

In addition, there is a growing awareness that government programs and family stability are interrelated. How people form and maintain family and household units affects government programs. In turn, government programs themselves influence family life, possibly to the point of encouraging fathers to desert their families. Nowhere is the interdependence between family structure and government programs more visible and more widely discussed than in the case of welfare programs and the problems of poor families. Some see welfare as *caused by* the irresponsibility of those who bear illegitimate children and of those who abandon their families. Others see welfare itself as the *cause* of broken and dependent families, as a barrier to work and family strength. Still others blame both welfare and family instability upon a third factor, the inadequate income and employment opportunities available to many fathers.

Daniel Moynihan raised these and other issues in his policy paper, *The Negro Family: The Case for National Action*. In summarizing sociological literature and adding his own analysis, Moynihan argued that family dissolution among low-income black families had reached a critical stage. He urged national policies to strengthen low-income

*The author is staff economist, Subcommittee on Fiscal Policy. The author wishes to thank Alair Townsend, Heather Ross, Vee Burke, Sharon Galm, Irene Cox, and James Storey for their useful comments.

[1] The entire text of the Moynihan report, officially titled *The Negro Family: The Case For National Action*, and an examination of reactions to the report appear in Lee Rainwater and William L. Yancy, *The Moynihan Report and the Politics of Controversy*, MIT Press, Cambridge, 1967.

black families. Coming soon after passage of major civil rights legislation, the report pointed out the largest remaining obstacles to racial equality were in the social and economic sphere. To make great headway in solving this problem, the report said, a major goal of government programs should be to help low-income black families stay together.

The report elicited widespread criticism and opposition. Primarily, the critics charged that Moynihan was blaming individual black families for their poverty instead of blaming racial discrimination and inadequate employment opportunities.[2] A second, more valid criticism was that there was insufficient evidence to prove Moynihan's assertion that families headed by women generally led to a "tangle of pathology."

The debate surrounding the report and discussions of family issues became highly charged emotionally. Anyone pointing to cultural factors as a cause of family disruption and to family disruption as a cause of poverty could expect to be called "racist," whether his opinions were based on analysis or whether he was simply using this line of reasoning as a rationale for strong action against racial discrimination. Emotional responses, in turn, gave many analysts an excuse for not dealing with the sound criticisms of the arguments.

In such an atmosphere, objective analysis is difficult. Nevertheless, issues of family structure are too crucial to social and economic policy to be evaded. Unfortunately, in spite of years of research and the publication of many books and articles, many questions remain unanswered because of the complex and sensitive nature of the dynamics of family behavior. It is essential that sound analysis continue on these matters if we are to achieve informed discussion about governmental policies to promote family strength.

To aid this discussion, the papers in these two volumes provide material which helps to answer the following questions, among others:

What are the primary causes and effects of illegitimacy, marital disruption, and female headship of families? What is the role of government transfer programs in influencing family structure? What government policies would increase family stability? Would vigorous pursuit of deserting fathers discourage desertion or increase the chances of disrupting second families? To what extent can extensive provision of birth control services reduce illegitimacy? What is the effect of various government transfer programs, both singly and in combination, on incentives for family stability and for various types of household arrangements? What is the appropriate role for government policy in influencing family arrangements?

These questions are relevant to the problems of existing transfer programs and the difficulties of overall reform of the system of transfers. Hence, the subcommittee staff obtained papers designed to improve basic knowledge and public understanding about the issues they raise. The 10 papers in these two volumes address a variety of issues and use a variety of analytical techniques. Phillips Cutright and John Scanzoni, and Marjorie Honig provide statistical analyses of the causes of high and increasing illegitimacy, marital instability,

[2] See, for example, William Ryan's charges that the Moynihan report "* * * seduces the reader into believing that it is not racism and discrimination but the weaknesses and defects of the Negro himself that account for the present status of inequality between Negro and white," in an article reprinted in Rainwater and Yancey, p. 458.

female-headed families, and participation in welfare programs. Barbara Boland also uses detailed statistical analysis to investigate the degree to which poor families eligible for AFDC actually receive benefits. Marc Fried and Ellen Fitzgerald, and Carol Stack and Herbert Semmel use participant-observer techniques to observe in actual cases how marital instability occurs and how poor families, particularly broken families, operate. Andrew Billingsley adds to this discussion by examining closely the complexities of family structure among low-and middle-income black families.

The policy focus of these papers varies from highly general to specific recommendations. Lee Rainwater makes the case that poverty means low relative income, not low absolute income, and that government policy should raise the income share going to people at the bottom of the income distribution. Papers by Harry Krause and by Stack and Semmel express different opinions on the more narrow issue of how strict the government enforcement policy should be in establishing paternity and obtaining support payments. Another important policy issue concerns which is the appropriate unit for purposes of determining eligibility and benefit levels under government income transfer programs: the individual, the family, the household, or some combination of these. Irene Cox examines closely the family unit treatment of eligibility and benefits in some existing programs; other authors make general recommendations about the issue.

This essay is an effort to pull together the findings from the 10 papers in these two volumes, to examine what these and other papers tell us about the causes and effects of family disorganization, about relative income and antipoverty policy, about problems in determining the appropriate recipient unit, and about the wisdom of vigorous efforts to collect child support payments. The essay and these volumes generally ignore family disoragnization problems among the middle and upper classes.

II. Perspectives on Family Disorganization

A. Some Problems of Methodology

One may legitimately ask whether it is possible to identify precisely the causal factors behind family instability. After all, how can we assess the quantitative importance of possible causes of desertion when the deserting parent does not know exactly why he or she left the family? This is a good question. Analysts should approach issues of family stability modestly. We should be aware of the limits to our ability to understand how behavior is affected by the intricate web of social, economic, and psychological processes.

A primary problem is the limitation of analytical techniques. Concern for privacy and sensitivity of subject matter impede asking some questions through large surveys. Yet these data may be essential for the analyst using statistical inference. Participant-observer studies allow a close look at family behavior but suffer from other restrictions, especially limited reliability of generalizations drawn from individual cases. Thus, to gain in-depth understanding of the operations of large numbers of families requires using the best of both techniques. Participant-observer studies suggest variables that should be included in any statistical analysis and on which data are required.

These case studies also provide knowledge on possible mechanisms of causation. For example, what is the process by which income affects family stability? Even the analyst using the best participant-observer studies to help formulate hypotheses must use great care in inferring causation.

Statistical inference can mislead as well as explain. Consider the fact that a negative, statistically significant correlation exists between income and illegitimacy. The percent of children that are illegitimate is highest in low-income families and is lowest in high-income families. One conclusion that might follow is that low income tends to cause high illegitimacy. However, one could also conclude that low family income is the result rather than the cause of illegitimacy, since having children without a husband leaves mothers without a primary income source. Another possible conclusion is that some third factor, say lack of education or an attitude of little concern for the future, causes both low income and high illegitimacy.

A second problem is defining what we mean by family disorganization. The common measures are the extent of illegitimacy, marital instability, and female headship of families. Many analysts have given little thought to what is the most appropriate measure because they assume that all three measures move together. Unfortunately, this assumption is often incorrect. The fact is that one measure may or may not move closely with the other two. High illegitimacy rates are consistent with low or high rates of marital stability, and with low or high rates of female headship of families.

Looking at one measure alone as an index can yield highly misleading conclusions. Suppose there is a decrease in the share of marriages that are disrupted. If the reason for this decline is that young unmarried pregnant women decide not to enter into a hasty marriage and decide instead to have more illegitimate children, then one could not conclude that family disorganization is declining. Similarly, larger numbers of female-headed families may not indicate growing family instability. The growth in female heads may occur because a larger number of mothers in disrupted marriages have enough income to establish and maintain their own households instead of living with other adult relatives. Thus, to understand trends in family disorganization, the analyst must at least examine measures of illegitimacy, marital disruption, and female headship of families.

A third problem in analyzing the implications of family disorganization is determining the appropriate unit by which to judge social and economic welfare. In some cases, it is the family; in others, the household; and in still others, groups of close relatives who live in separate households. From the standpoint of income adequacy, male discipline, and adult guidance, the child who lives only with his mother but who has grandparents or uncles living next door may be much better off than the child in a husband-wife family who has little access to his father's time or resources. Billingsley, Stack and Semmel, and Fried and Fitzgerald all point out the important role that close friends or relatives residing outside the household can play.

Considering the differences in how people associate and share resources is important in assessing economic welfare. If every poor *family* lived in a *household* with an above-poverty per-capita income that was shared equally by all household members, then the economic welfare of all *individuals* would exceed the poverty standard. Alternatively, the complete absence of poor *families* may leave some *individuals* poor because they have little access to their family's resources. Most discussions of poverty implicitly assume that the family unit is most relevant for judging the amount of income available to individuals. Although some family concept may be the single best measure, there remains the problem of precisely defining the family itself. Should it be the nuclear family or all household members related by blood, marriage, or adoption? Sometimes the most meaningful measure of income is household income, including that of nonrelatives.

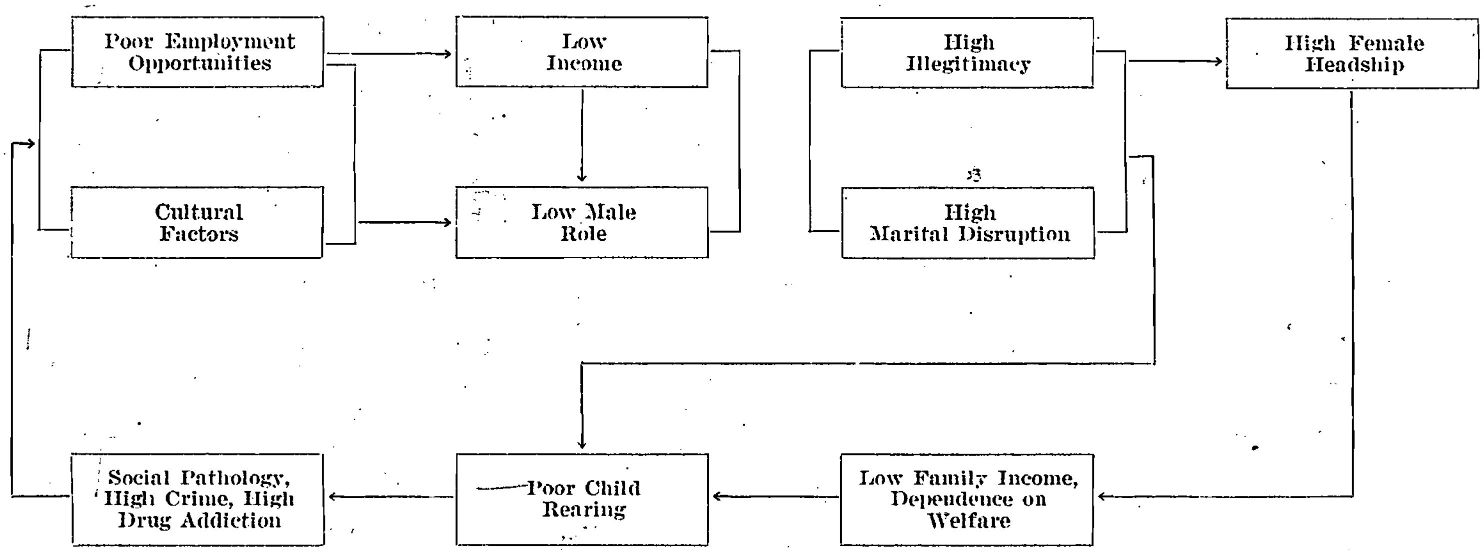
B. The Conventional Image of the Disruption Process

Like most broad social processes, family disorganization has many causes and effects. Any one marriage may end because of a mixture of money problems, infidelity, personality conflicts, and differences in cultural values concerning the sanctity of marriage. The process becomes increasingly difficult to understand when we consider large numbers of family breakups. Nevertheless, sociologists and other social scientists study such processes as family disorganization in an effort to be able to make valid generalizations about how various social forces influence behavior. The idea is that, although any given marriage becomes disrupted because the partners cannot get along, general factors such as income and education may account for the phenomenon that marriages among some groups (in some decades) seem to break down more often than marriages among other groups (in other decades).

Although differences of opinion do exist among analysts of family behavior, there appears to be a consensus among many authors about the general picture of how marital disruption occurs in low and lower middle class families and what its effects are. It is useful here to present a stylized description of this conventional image. Such a model necessarily oversimplifies. The model's usefulness for purposes of this essay depends on how well it captures key elements in the causal chain various authors describe.

The conventional image of family disorganization is a series of causal linkages. These appear in figure 1. Authors may differ significantly as to how much this image accords with their own views. For example, some authors may believe that poor employment opportunities reduce the male role in the family, but reject the notion that family disorganization is a prime cause of social pathology. The model in figure 1 provides a framework with which we can both highlight differences among various authors and examine what evidence exists to substantiate each linkage.

Figure 1.—The model of the disruption process.



Moynihan and many of his critics appear to agree that poor employment opportunities, low income, and the marginal male role all help cause high illegitimacy, marital instability, and female headship. The disagreement centers largely around the nature of family disruption and its effects. To Moynihan, illegitimacy, marital instability, and female headship are what family disorganization means. All three measures represent experiences of failure, bitterness, and alienation. To Moynihan, the result of this disorganization is that the rearing of children suffers from the absence of the adult male, which, in turn, causes such children to become poorly motivated and to engage in crime and drug addiction.[3] Critics of Moynihan object to his conclusion that female headship family patterns adopted by many poor blacks have a poor effect on children. For many of these critics, rejection of the link between female headship and poor child rearing grows out of their belief that illegitmacy, marital instability, and female headship do not necessarily constitute family disorganization but rather may be culturally superior family patterns or, at worst, healthy adaptations to problems of black people in a hostile society.

In spite of substantial scholarly debate on the nature of these family processes, the evidence marshaled to establish or to reject one or another of the linkages is often weak and inconclusive. This section examines the available evidence, much of which appears in the newly released papers in these volumes.

C. Causal Factors Behind High and Increasing Illegitimacy, Marital Disruption, and Female Headship

Basic facts on illegitimacy, marital disruption, and female headship are well known and can be described quickly. A few figures appear in tables 1 and 2. In comparing differences among population groups at one point in time, one finds that illegitimacy, marital disruption, and female headship are highest among blacks, the young, and those with the lowest income. The trend over time in the three indices is upward. Between 1950 and 1968, the share of births that were illegitimate more than doubled; the share of married women not living with their husbands went up slightly; and the share of families headed by women rose significantly in the case of blacks but has remained relatively constant in the case of whites.

The task of this section is to examine what light various papers shed on the causes of illegitimacy, marital disruption, and female headship. To what extent do specific factors influence differences among population groups and changes over time? Of particular interest is the evidence of these papers regarding the validity of the first half of the causal chain in figure 1. In attempting to explore the reasons for higher and growing family instability, we shall consider separately the effect of each causal factor, such as income or health, on all three measures of family instability.

1. DEMOGRAPHIC FACTORS

The role of demographic factors often receives far too little attention relative to behavioral factors, possibly because demography is lacking in drama. Cutright, and Cutright and Scanzoni help to correct

[3] See the discussion in the Moynihan report reprinted in Rainwater and Yancy, especially pp. 80–81 and 84–85.

this imbalance by explicitly calculating the quantitative significance of changes in various population components and in birth rates.[4]

TABLE 1.—*Trends in illegitimacy*

	1940	1950	1960	1965	1968
Illegitimacy ratios:[1]					
White	19.8	17.5	22.9	39.7	53.3
Nonwhite	166.4	179.5	215.8	263.2	312.0
Illegitimacy rates by age of mother:[2]					
White:					
Total, 15 to 44	3.9	6.1	9.3	11.5	13.0
20 to 24	6.0	10.0	18.5	21.7	22.6
25 to 29	4.3	8.8	17.1	23.8	21.5
30 to 34	2.6	6.0	10.8	16.4	1[illegible].9
35 to 44	1.3	2.1	3.9	4.9	4.7
Nonwhite:					
Total, 15 to 44	39.1	68.9	90.2	94.4	83.0
20 to 24	52.2	103.5	147.1	142.2	109.9
25 to 29	36.7	92.4	137.4	153.3	96.6
30 to 34	26.5	62.6	97.3	129.3	75.1
35 to 44	10.8	20.0	31.9	37.8	24.2

[1] Number of illegitimate births per 1,000 total births.
[2] Number of illegitimate births per 1,000 unmarried women.

Source: Phillips Cutright, "Illegitimacy and Income Supplements," this volume, tables 2 and 3.

TABLE 2.—*Selected trends in marital status and family types*

	1950	1955	1960	1965	1970
Marital status of women, age 14 and over (in percent):					
White:					
Single	19.8	17.9	18.6	20.4	21.3
Married, spouse present	63.6	63.9	63.7	61.5	60.2
Married, spouse absent	2.7	3.6	3.0	2.9	2.6
Widowed	11.5	12.4	11.9	12.4	12.4
Divorced	2.4	2.2	2.8	2.8	3.4
Nonwhite:					
Single	20.7	20.5	22.0	23.1	28.0
Married, spouse present	50.8	48.5	48.6	46.9	43.6
Married, spouse absent	11.2	13.6	12.0	12.9	11.3
Widowed	14.6	14.9	13.8	13.1	13.0
Divorced	2.7	2.5	3.6	4.0	4.2

[4] The discussions in the next two sections draw on the comprehensive work of Cutright, and Cutright and Scanzoni. See Phillips Cutright, "Illegitimacy and Income Supplements," and Phillips Cutright and John Scanzoni, "Income Supplements and the American Family," in this volume.

TABLE 2.—*Selected trends in marital status and family types*—Con.

	1950	1955	1960	1966	1968	1972
Type of family (in percent):						
White:						
Husband-wife	88. 0	87. 9	88. 7	88. 8	88. 9	88. 2
Other male head	3. 5	3. 0	2. 6	2. 3	2. 2	2. 3
Female head	8. 5	9. 0	8. 7	8. 9	8. 9	9. 4
Nonwhite:						
Husband-wife	77. 7	75. 3	73. 6	72. 7	69. 1	65. 7
Other male head	4. 7	4. 0	4. 0	3. 7	4. 5	4. 2
Female head	17. 6	20. 7	22. 4	23. 7	26. 4	30. 1

Source: U.S. Bureau of the Census, *Current Population Reports*, series P-23, No. 42, *The Social and Economic Status of the Black Population in the United States*, 1971, U.S. Government Printing Office, Washington, D.C. 1972, p. 100; and U.S. Bureau of the Census, *Current Population Reports*, series P-20, No. 144, No. 212; U.S. Census of Population: 1950, Vol. 4, special reports, pt. 2, ch. D, *Marital Status*: U.S. Census of Population: 1960, vol. 1, Characteristics of the Population, U.S. summary.

Understanding demographic influences on illegitimacy trends requires one to distinguish among measures of illegitimacy. Consider first the increase in the share of births that are illegitimate. This percentage, called the illegitimacy ratio, has more than doubled in the last three decades. From 1960 to 1968, the ratio has grown substantially among whites and nonwhites. A different picture of illegitimacy trends would appear if our interest were in the average propensity of an unmarried woman to have a child. The conventional measure of this propensity is the illegitimacy rate, the number of illegitimate births divided by the number of unmarried women of childbearing age. Illegitimacy rates have also risen dramatically between 1940 and 1970. However, since 1960 the rate actually declined among black women and the increase slowed among white women. Interestingly enough, the illegitimacy ratio and the illegitimacy rate can move in opposite directions.

The primary reasons cited for the differences in the 1960–68 trends are the general decline in the birth rate among married women and the rise in the share of unmarried women. As married women decide to bear fewer children, the share of all children who are illegitimate can rise in spite of no change in the illegitimacy rate. Note that the lower birth rate among married women has little to do with decisions of unmarried women concerning illegitimacy. Thus, the illegitimacy ratio rose partly because of behavioral changes among married women. The impact of the rise in unmarried women on illegitimacy trends is subject to differing interpretations. On the one hand, larger numbers of unmarried women mean a greater population at risk of conceiving out of wedlock, implying that even a constant propensity to bear illegitimate children leads to an increase in absolute numbers of illegitimate births. This interpretation makes the increase appear as a natural result of a larger population of unmarried women and assigns none of the rising illegitimacy to a change in the willingness of women to bear illegitimate children. Although this is plausible, such a view does presume that marriage and childbearing decisions are independent and that the sequence is the marriage choice first and the child-bearing choice second. An alternative view is that the larger share of women who are unmarried is itself partly related to a reduced fear of bearing an illegitimate child. If greater acceptance of illegitimacy influences women not to marry or remarry, then a constant illegitimacy rate could be consistent with a growing willingness to bear an illegitimate child.

Demographic factors also played an important role in the rapid increase in the numbers of women in disrupted marriages and heading families. In an interesting analysis of how much various components contributed to the growing numbers of female heads of families, Cutright and Scanzoni assign about 20 percent of the increase in white female heads and 13 percent of the increase in nonwhite female heads to the effect of population growth alone.[5] Although their results indicate a surprisingly small impact from rising marital disruptions, this finding becomes understandable when one separates two offsetting trends. Among black ever-married women, the share of women who are separated or divorced has increased sharply between 1940 and 1970 while the share of women who are widows has declined sharply. Counting the sum of these components as "women in disrupted marriages" leads one to conclude that marital disruptions have changed little in 30 years.

2. HEALTH FACTORS

The well-being of families is often related to family disorganization measures. To most analysts, high illegitimacy, marital disruption, and female headship are signs of family difficulties. However, increases in these indices could have been caused by higher rather than lower living standards. Cutright, and Cutright and Scanzoni bring this point out when demonstrating the role of improved health on family disorganization. It is interesting that this factor, which is more objective, less emotion-laden than most, is not mentioned by Moynihan or his critics.

Improved health can increase illegitimacy in a variety of ways. Cutright reports that better nutrition and health has apparently led to a decline in the age of menarch—age of first menstruation. This means a corresponding reduction in the age at which adolescents are fertile. A second way in which improved health is important is through declines in involuntary sterility. Third, spontaneous fetal losses have declined, again for health reasons. According to Cutright, these three health-related factors account for virtually all—88 percent—of the increased illegitimacy rates between 1940 and 1968 for nonwhites but much less—19 percent—of the increase among whites.[6]

One limitation of this important finding is the assumption that women's decisions regarding coitus, abortion, and marriage are independent of these health factors. This assumption may not have general validity. While a lower age of fecundity may increase chances of con-

[5] Cutright and Scanzoni, table 9.
[6] Cutright, table 9.

ception, the unmarried woman still makes decisions about birth control and abortion partly based on how strongly she resists bearing an illegitimate child. Further, a declining rate of spontaneous fetal losses does not exclude the role of a woman's decision to bear an illegitimate child. She may still be able to get an abortion. Thus, the influence of health improvements cannot be totally separated from behavioral decisions.

Health improvements may also affect the number of women in disrupted marriages who are mothers. Cutright and Scanzoni point out that childlessness used to be high among married women in large part because of disease and the effect of poor health on sterility and fetal loss. They report that childlessness declined significantly in the 1940–1970 period among women in disrupted marriages. They attribute most of the decline in childlessness among nonwhite women to improved health; for white women, both improved health and reductions in the use of birth control techniques contributed to the declining childlessness.

These health effects interact with decisions by individual married women, as in the case of the health effects on illegitimacy. Considerably more evidence than Cutright and Scanzoni present would be necessary to show the size of the independent health effect. Consider the facts Cutright and Scanzoni present in table 6 showing that the percent of all nonwhite women aged 15–44 in disrupted marriages who were childless fell from 30 percent to 18 percent between 1940 and 1970. They seem to argue that improved health allowed a greater share of married women who desired children while their husband was present to bear one. Then, with fewer of the married women childless, it follows statistically that there would also be a smaller share of childless women among those in disrupted marriages. An alternative explanation of the Cutright and Scanzoni data might run as follows. The presence of children normally reduces the chance of a marriage dissolving. The finding that fewer women in disrupted marriages are childless means that an increase has occurred in the chances for marital disruption among families with children.

Notwithstanding these limitations, Cutright and Scanzoni make important contributions, by highlighting the role of health factors on illegitimacy and marital disruption trends.

Cutright's related discussion of voluntary contraception and illegitimacy is also valuable. After noting the trends in effective use of contraceptive devices and their influence on illegitimacy trends, Cutright assesses the potential impact of birth control programs on illegitimacy. He points out that programs attempting to increase the use of birth control pills and IUD's will likely have only limited success in reducing illegitimacy. Since a large proportion of illegitimate births are first births among young, poor, unmarried women whose coital experience is infrequent and irregular, many women-at-risk will either not participate in the programs or encounter high failure rates. This is not to say that these young women want to bear an illegitimate child. Most do not and, where available, many have taken advantage of abortion services. Cutright argues convincingly that providing abortion services will be much more successful in reducing illegitimacy than will encouraging the use of contraceptive devices.

3. INCOME, EARNINGS, AND WELFARE PAYMENTS

Money is at the root of many conflicts, including some that lead to family disruption. One problem is caused by too little money. Moynihan and many of his critics agree that poor job and income opportunities of black males contribute significantly to marital disruption, that marriages are strained and the role of the father is made increasingly marginal when he cannot adequately support his family. (See Figure 1 for how these factors fit into an overall causal chain.) Although too little money is cited as the cause of trouble, the Government's attempts to provide money are said to make matters worse. The most dramatic charges against the welfare system are that it encourages desertion by fathers and that it encourages illegitimacy. Providing welfare payments and other benefits more generously to female-headed families (i.e., one-parent families) than to husband-wife families clearly reduces the economic importance of the male. Critics charge that welfare's preference for the broken family also stimulates illegitimacy, marital disruption, and female headship of families.

A number of papers in this volume examine the empirical basis for such hypotheses. In attempting to discover how low income, poor job opportunities, and high welfare payments actually influence behavior, the authors replace speculation with statistical analyses. Although the statistical analyses themselves have limitations which are discussed below, they reveal many new interesting results. The next three sections summarize the major findings concerning the effects of economic variables on illegitimacy, marital disruption, and female headship.

a. Illegitimacy

Low income may cause illegitimacy in a number of ways. First, and most direct, the pregnant mother may bear the child without getting married because the mother and/or father believe their income is insufficient for marriage. Second, her low income may prevent an unmarried women from getting a safe abortion. (This cause should diminish because of the Supreme Court's decision forcing States to legalize some abortions.[7]) Indirectly, poor income opportunities may increase illegitimacy by reducing the economic importance of the male to the family. Given the existence of welfare benefits paid upon his absence, the family's income with the father in the house might be little or no higher than its income without him. Further, as the reduced male role causes illegitimacy and marital disruption among some population groups, young people perceive that the chances of a stable marriage are low and that illegitimacy and female headship are acceptable to the community.[8]

[7] The Supreme Court has also ruled that States may elect not to cover the costs of abortion under the medicaid program.

[8] Each of these causal factors suggests the use of a different measure of income. The sufficiency of income to marry would seem to depend on family income opportunities relative to average family income. Purchasing a safe abortion is more closely dependent on absolute rather than relative income. The male economic role is most plausibly measured by expected family income, husband present, divided by expected family income, husband absent. In examining community effects, which would be highly complex, one would probably use neighborhood income and neighborhood family disorganization measures.

Cutright's paper examines the effect of socioeconomic status on illegitimacy in three ways. First, Cutright argues that racial differentials in poverty status account for at least 45 percent of racial differentials in illegitimacy rates. According to Cutright, counting all socioeconomic status factors would probably explain more than half of the differentials in illegitimacy. Second, Cutright maintains that in the absence of any narrowing over time in absolute income differences by race, one should not expect a narrowing in illegitimacy rate differentials by race. Finally, Cutright contends that income differences in the form of different AFDC payments do not affect State differentials in the levels or rates of change of illegitimacy rates.

In estimating the effect of income on racial differentials in illegitimacy, Cutright uses the fact that the share of nonwhite unmarried women, ages 15–44, who are poor or near-poor, is much larger than the corresponding share of white unmarried women. Unfortunately, this measure of income differentials has limitations. To the extent that poverty status of some unmarried women is measured after the woman has at least one illegitimate child and forms a new household, it is illegitimacy that is helping to cause poverty instead of the other way around. A second problem with this index occurs even when the poverty status of unmarried women is measured before they bear children. Their poverty may be irrelevant to their unwedded motherhood. A woman's family may be poor because she has many brothers and sisters. This poverty need not affect her marriage decision.

The theory of the illegitimacy process should play a large role in deciding on the most appropriate income variable. If low income causes illegitimacy by making marriage financially difficult, then the appropriate variable in comparing illegitimacy behavior of nonwhite with that of white unmarried women is what their relative *family* incomes would be if they married. High illegitimacy may also result from the fact that poor males would make only a small financial contribution to the family. If this is the causal factor at work, then the appropriate income variable is how well financially the mother and her child would fare with the father present relative to how they would fare in his absence.

Cutright does turn to family income measures in examining illegitimacy trends by race. Although Cutright argues that the income of nonwhite families relative to white family income is the important determinant of behavior, he does not use the traditional measure of relative income, the ratio of nonwhite to white median income. The relative income position of blacks shows improvement, according to this ratio. Instead, Cutright looks at changes in absolute dollar differences. Noting that nonwhites were further behind whites in 1970 than in 1950 in absolute dollars, Cutright has a justification for the absence of a significant narrowing in the racial differentials in illegitimacy. Unfortunately, the use of absolute dollar differences instead of ratios is highly suspect. To believe in absolute dollar differences as the appropriate measure of relative status, one would have to believe that nonwhites would feel more deprived relative to whites if their respective average incomes were \$95,000 and \$100,000 (a difference of \$5,000) rather than \$6,500 and \$10,000 (a difference of \$3,500).

Another problem with this measure is that the median family income figures for each race include all families, not simply husband-wife families. Since families with female heads make up a much larger and faster growing share of nonwhite families than of white families, this compositional effect alone accounts for part of the racial differential in median income and part of the slow narrowing of nonwhite-white income differentials. In assessing whether income differentials cause illegitimacy differentials, one would want to abstract from this compositional change. To avoid much of this problem, one could use either (a) changes in the nonwhite to white ratio of incomes of husband-wife families, or (b) changes in the nonwhite to white ratio of average family income of female-headed families relative to the same ratio for husband-wife families.

Pitfalls in this area of analysis can arise from an overly narrow view of causation. One must recognize that illegitimacy can be a cause of low income as well as an effect of low income. Caring for an illegitimate child hinders the job of earning above-poverty incomes. Also, some third factor may cause both low income and high illegitimacy. For example, Edward Banfield might argue that the lower class' lack of concern and inability to sacrifice for the future are the root causes of both low income and high illegitimacy.[9] Poor preparation by schools and by families is another possible cause.

Another highly publicized cause of high illegitimacy is the welfare system. AFDC payments and other benefits are said to reduce the cost of caring for illegitimate children. Cutright examined this hypothesis by comparing State AFDC payments with State illegitimacy rates. He found no discernible relationship. While his analysis is of interest, it does not consider the effects of many variables simultaneously. The advantage of the multivariate approach in analyzing the effects of AFDC on illegitimacy is that it allows one to examine all factors that may influence State illegitimacy differentials together, taking account of any simultaneities. The size of the AFDC payment would be treated as one of many important variables. Cutright's simple correlations showing that AFDC payments and illegitimacy rates are not positively related suggest that high welfare payments are not the overwhelming influence on illegitimacy. Estimates of more accurate AFDC effects on illegitimacy will come out of future research.

b. Marital instability

Do high private incomes help keep husbands and wives together? Do high welfare payments help break up marriages? Cutright and Scanzoni attempt to analyze these controversial questions empirically.

In examining the effect of private income, Cutright and Scanzoni focus primarily on the following question: To what extent do racial differentials in income account for racial differentials in marital stability? They report census data showing that marital disruption rates of adult males decrease as the male's income rises. This tendency holds for both white and black men. Although black men experience higher rates of separation and divorce than white men even at the same income level, the size of the total racial differential in marital stability is much larger than racial differentials within income classes. After

[9] See Edward Banfield, *The Unheavenly City*, Little, Brown, and Company, Boston, 1968, especially pp. 45–66.

expanding the discussion to consider the lower asset holdings, lower job satisfaction levels, and greater nonincome disadvantages of blacks as compared to whites, Cutright and Scanzoni conclude that economic status differences account for virtually all of the observed racial differentials in marital stability.

One may question their conclusion on a number of grounds. First, the fact that lower incomes are associated with higher marital instability does not necessarily mean that low income causes marital instability. It may be that marital stability is influencing the male's income. Separations, divorces, and delays in remarriage may reduce a man's need or his perceived need for income or may discourage him from working hard. A further possibility is that some third factor, such as lack of concern for the future or lack of regard for a stable family life, causes both low incomes and high disruption rates. Second, there is the unsolved problem of distinguishing nonincome disadvantages from income disadvantages suffered by blacks. Otherwise, one can count the same unfavorable factor twice. Finally, there is no evidence about the extent to which these added disadvantages induce high marital disruption rates.

Cutright and Scanzoni examine the impact of welfare payments on marital stability by relating State AFDC levels to State percentages of women living with their husbands. They report finding no relationship between the two variables and conclude that the size of the AFDC benefit does not influence marital disruptions. Because the procedure does not account for other influences of marital stability, these results are of limited value. Further discussion of AFDC effects appears in the next section.

c. *Female heads of families*

A primary concern about the rise in female-headed families is that the increase represents a growing breakdown of the family. Attracting most attention are the enormous increases occurring among already unstable minority families. The share of black families headed by women grew dramatically from 21 to 31 percent between 1960 and 1971. Accompanying this growth was an explosive 238 percent increase in welfare recipients in the aid to families with dependent children category. It is especially troubling that the rise in broken families accelerated during a decade of rapid gains in income.

What happened? Did the general prosperity fail to reach lower-class minority families? Did high welfare payments stimulate greater family splitting? What evidence is available to determine the effects of private income and welfare payments on the share of families headed by mothers? Those attempting answers to these questions examine the trends in female headship rates over time and differences among various groups in these rates at a single point in time.

In interpreting the puzzling events of the 1960's, one might first ask what the growth in the share of black female-headed families represents. Moynihan has argued that these increases indicate considerable and growing social distress among one segment of the black population.[10] Although this interpretation is the most obvious one, some authors disagree. Cutright and Scanzoni point out that the growing share of black female family heads did not result in any decline

[10] Daniel P. Moynihan, "The Schism in Black America," *The Public Interest*, No. 27, Spring 1972, pp. 4–9.

in the share of black women who were wives in intact families. If the share of intact black families remained constant, it would appear that family disorganization did not in fact become worse. What accounts for these differing views?

The Moynihan analysis does not recognize that there are three possible family positions of mothers with children. Such women may be wives in intact families, heads of broken families, or "other relatives" of the head. Examples of this third category would be a mother living with her children in a household headed by her father, uncle, or other relative. Mothers in this latter category are designated by the Census Bureau as "subfamily heads" but not as "family heads." Looking closely at the 1950 to 1972 trends, one finds that part of the increase in female family heads can be attributed to a declining share of subfamilies. That is, fewer mothers without husbands must double up with other relatives; more can form their own households. Among women living with their own children under 18 but not living with a husband, the share heading subfamilies fell from 33 percent in 1950 to 13 percent in 1972 with a corresponding rise in the share heading families from 67 percent to 87 percent. This factor alone contributed significantly to the observed rise in the number of female family heads with children. Of the absolute increase in female family heads between 1950 to 1972, 36 percent would not have occurred if the share of mothers without husbands who become subfamily heads had remained at the 1950 rate.[11] These and other figures suggest that some of the rising female headship could well be an indication of higher living standards from higher income including AFDC payments—rather than a sign of growing family disorganization. Mothers in disrupted marriages might simply have decided to spend part of their increased income to set up and maintain their own households.

Cutright and Scanzoni provide an overall look at the contributions of various components to the increase in female heads. Among black women, the largest component was the growing tendency for mothers in disrupted marriages to head their own families. Although a moderate increase in illegitimacy rates was evident, increasing marital disruption apparently accounted for little of the total increase in fatherless families. This is a surprising result. It would seem to flatly contradict the widespread view that black families have suffered increasingly severe breakdown.

[11] The data for these calculations come from U.S. Bureau of the Census, *Current Population Reports*, series P-20, Nos. 33, 106, 246, 251. The calculations were performed as follows. Let FH equal the number of female heads with children, SH equal the number of female subfamily heads, with subscripts 1 and 2 appling for years 1950 and 1972. Then, one can compute FH^*_2, an adjusted number of female family heads in 1972 that assumes for 1972 the 1950 behavior pattern regarding the formation of subfamilies and separate families:

$$FH^*_2=\left(\frac{FH_1}{FH_1+SH_1}\right)(FH_2+SH_2)$$

An estimate of the percent increase in FH atrributable to the fact that the share of subfamily heads did not remain at the 1950 rate is EST, where

$$EST=\frac{FH_2-FH^*_2}{FH_2-FH_1}=0.36$$

for the 1950–72 period. EST equals 47 percent for the 1950–60 period and 20 percent for the 1960–72 period. The comparable figures using current year rather than base year weights are 17 percent, 36 percent, and 13 percent for the 1950–72, 1950–60, and 1960–72 periods, respectively.

A closer look at the trends reveals a still more complex, but more accurate interpretation of the rising proportion of black families lacking fathers. Consider all women with children in one of three family status categories: (1) heads of families, (2) wives of family heads, or (3) other relatives of the family head. In both the 1940–60 and 1960–70 periods, the rise of female heads was almost completely offset by the declining share of women who were other relatives of the family head. However, the observed constancy in the share of wives occurred for different reasons in the two periods. While in the 1960–70 period the constant share of wives meant no significant behavioral changes, it was the result of two offsetting trends between 1940 and 1960. Over those two decades, the share of ever-married Negro women who were widowed declined from 25 percent to 18 percent.[12] By itself, such a decline would lead to an increase in the share of wives. Instead, the share of women who were separated and divorced rose, offsetting the expected increase in the share of wives. Thus, rising marital breakups played a significant role in the 1940–60 increases in black female headship but not in the 1960–70 changes.

What impact could the welfare system have exerted on these trends? One interesting potential effect is that high and growing AFDC payments might have allowed mothers in disrupted marriages to form and maintain their own households. Second, in spite of the constancy in the share of wives in husband-wife families during the 1960's, AFDC itself might have encouraged marital disruption, an effect offset by other factors, such as rising incomes. Boland indirectly looks at the potential effect of AFDC on the rise in female headship in the 1967–70 period. Her results show that the increase in AFDC recipients occurred as a result of higher participation by female heads and expanded economic eligibility rather than as a result of an increase in the number of low-income female family heads.[13] These findings are of major interest but they are not—and were not intended to be—good evidence on the question of AFDC's effect on family splitting.

Direct tests of the AFDC influence on female headship appear in papers by Honig and by Cutright and Scanzoni. These authors investigate AFDC effects by examining whether differences in AFDC payments among areas influence State or metropolitan area differences in female headship rates. As in their tests of AFDC's impact on illegitimacy and marital disruption, Cutright and Scanzoni do not simultaneously consider variables other than the two primary variables. Use of this procedure limits the value of their finding of no significant relationship between AFDC payments and incidence of female-headed families.

Honig provides the most careful study to date on the impact of AFDC payment levels. Her hypothesis is that, holding other things equal, higher AFDC payment levels in a metropolitan area will mean an increase in the area's percentage of women who are AFDC recipients—the recipient rate. The higher recipient rate will result from (*a*) larger shares of female family heads becoming welfare recipients—welfare-independent female heads—and (*b*) larger shares of women becoming both female family heads and welfare recipients—welfare-

[12] Reynolds Farley, "*Growth of the Black Population*," Markham Publishing Co., Chicago, 1970, p. 145.

[13] Barbara Boland, "Participation in the Aid to Families With Dependent Children Program (AFDC)," this volume.

induced female heads. An important feature of the Honig analysis is the inclusion of explanatory variables in addition to AFDC payment levels. For example, the level of male wage appears as a potential influence on female headship rates and recipient rates. This is critical to the analysis. High male wages should reduce the financial incentive for family splitting and thus discourage family breakup. High AFDC payments should have the opposite effect. Unless one takes account of each factor separately, the two factors could cancel each other out. The results would look as if there were no relationship when in fact there were two off-setting effects.

Honig's findings strongly support the notion that the size of the AFDC stipend influences the share of female-headed families. Tests using 1960 and 1970 data on 44 metropolitan areas provide estimates of the impact of higher AFDC payments. According to Honig's results for 1960, a 10-percent-higher area AFDC payment will induce, other things equal, a 3- to 4-percent-higher rate of female headship. This result holds for rates applying to nonwhite and white women. Honig's preliminary results for 1970 yield a similar conclusion but a slightly weaker relationship.

In considering Honig's findings, one should recall that high AFDC payments might induce female headship either by discouraging stable marriages or by helping mothers in disrupted marriages to head their own households rather than live with other relatives. Honig is unable to distinguish fully between these two types of AFDC effects. Analyzing these two effects separately as well as verifying Honig's findings with more detailed models are important tasks for future researchers. Such models should take account of the existence of other transfers, particularly food stamps and general assistance. These other transfers have an important effect on area differences in benefit packages available to female-headed relative to husband-wife families.

D. Effects of Family Disorganization

Many authors who agree on the causes of family disorganization disagree on the effects of such disorganization. According to Moynihan and E. Franklin Frazier, family disorganization causes a "Tangle of Pathology." Moynihan, in concluding his case for national action to strengthen black families, closes with the following quote from Frazier:

> As the result of family disorganization a large proportion of Negro children and youth have not undergone the socialization which only the family can provide. The disorganized families have failed to provide for their emotional needs and have not provided the discipline and habits which are necessary for personality development. Because the disorganized family has failed in its function as a socializing agency, it has handicapped the children in their relations to the institutions in the community. Moreover, family disorganization has been partially responsible for a large amount of juvenile delinquency and adult crime among Negroes. Since the widespread family disorganization among Negroes has resulted from the failure of the father to play the role in family life required by American society, the mitigation of this problem must await those changes in the Negro and American society which will enable the Negro father to play the role required of him.[14]

This view has come under a good deal of criticism. The most radical critics say that there is nothing wrong with illegitimacy, marital instability, and female headship, and that the family structure of the

[14] Rainwater and Yancy, p. 94.

poor is a healthy adaptation to difficult circumstances. Others point out that there is no necessary connection between family structure and social pathology. Herbert Gans reflects this attitude in the following quote:

> In addition, it must be stressed that at present, we do not even know whether the lower class Negro family structure is actually as pathological as the Moynihan report suggests. However much the picture of family life painted in that report may grate on middle-class moral sensibilities, it may well be that instability, illegitimacy, and matriarchy are the most positive adaptations possible to the conditions which Negroes must endure.
>
> Moynihan presents some data which show that children from broken homes do more poorly in school and are more likely to turn to delinquency and drugs. Preliminary findings of a study by Bernard Mackler of the Center for Urban Education show no relationship between school performance and broken families, and a massive study of mental health in Manhattan, reported by Thomas Langner and Stanley Michaels in "Life Stress and Mental Health," demonstrated that among whites at least, growing up in a broken family did not increase the likelihood of mental illness as much as did poverty and being of low status.[15]

Discussions of the effects of family disorganization occur in several papers in this volume. As noted earlier, the three measures used in this paper may differ in their effects. Readers will find surprising the result reported by Cutright and Scanzoni that whether a first birth is illegitimate or legitimate may have little effect on whether, in the long run, the mother is unmarried, is married with spouse absent, or is married with spouse present. This result holds for black and for white women. Of course, this does not mean that families with at least one illegitimate child are not fatherless for longer periods than other families. What it does indicate is that we should be cautious about the meaning and the effects of family disorganization.

Papers by Billingsley and by Stack and Semmel emphasize a point that requires continuing emphasis. It is that families technically classified as female-headed may take a variety of forms in practice. An adult male relative or nonrelative may live in the same household, may share some of his income with family members, and may help to bring up the children. Whether through a stable household composition or a relatively unstable one, children may have close near-familial relationships with adult men. Stack and Semmel, and Fried and Fitzgerald point out that domestic networks of close relatives and friends are important to the lower class family. In cases where the household consists only of a mother and her children, the mother may receive strong support in child rearing from other relationships. Stack and Semmel note that even the meaning of a household can differ from place to place. Children may commonly sleep and eat in households other than their primary residence.

The variety of household types has implications for research on the question of whether broken families help cause crime, drug addiction, and other social adjustment problems. Researchers must take account of the fact that the strict dichotomy of male- and female-headed families may be misleading in at least two ways. First, a study using such a dichotomy which finds no effect by sex of head does not necessarily mean that family structure does not influence child rearing. The researcher may have chosen the wrong family structure variable. Second, finding a significant effect by sex of head does not allow the

[15] Herbert Gans, "The Negro Family: Reflections on the Moynihan Report," as reprinted in Rainwater and Yancey, pp. 450–451

researcher to conclude that lack of male influence is causing the problems or successes. Many families headed by women may have more healthly male influences than some headed by poor men.

Billingsley cites a number of studies to support the following contention:

> Thus, working-class families, middle-class families, and upper-class families in the black community provide a progressively higher level of protection to their children than families in the lower class, and the under class. This is due not so much to the nature of the family structure as to the nature of the resources available to the family to help them care for their children. A husband and father is one, but not the only important figure and function necessary to insure the well-being of children. In the black community particularly, other family members, relatives and friends, neighbors, and other role models often provide some of the screens of opportunity which enable some families to function better than others.[16]

Child abuse seems to occur less among black families than among the general population. Billingsley's review indicates that many poor black families function well in child rearing while some function poorly. The results reported by Billingsley do not seem to show an independent effect from family structure apart from the impact of low income.

The problem of distinguishing between the effect of low income and the effect of a breakdown in the family is a difficult one. Still, one must cast doubt on those studies which find significant family breakup effects but do not separate these effects from the effect of income. Further, Gans points out that under some circumstances, such as personality differences between parents or mental illness, the child's environment may benefit from his parents' separation.

On this general issue it appears that the evidence still is inconclusive.

III. Antipoverty Policy and Relative Income

On the meaning of poverty reasonable people continue to disagree. Lee Rainwater argues for a definition based on the dollar distance between those with the lowest incomes and the average income in the country. Having defined poverty in relative terms, Rainwater goes on to advocate antipoverty policies that narrow the income distribution. He would prevent family incomes from falling below one-half of the median income. Only such a policy would alleviate Rainwater's notion of "poverty."

Many have argued that, as a matter of morality, the Nation should redistribute income on a vast and continuing basis in order to eliminate wide disparities among families. Rainwater attempts, by use of social science, to demonstrate that extensive redistribution is wise. In this effort, he seems to equate what people believe is economically necessary with what they think (or should think) is (or should be) a matter of right, perhaps even without work effort.

Rainwater's discussion of the causes and the cures of family instability goes to the crux of the matter. Rainwater does an excellent job of portraying the highly involved nature of the disruption process. A summary cannot be just to his discussion. As key elements, Rainwater identifies the economic marginality of the father, the lack of necessary monetary resources for proper family functioning, and a street scene that is especially attractive to those without extensive resources. This web of factors produces values and norms that lead to

[16] Andrew Billingsley, "The Black Family: Myths and Realities," in the second part of these two volumes.

high illegitimacy, marital disruption, and female-headed families. The behavior is seen as an adaptation by the lower class to low incomes. If inadequate income causes this behavior, then adequate income should produce a different set of adaptations, behavior similar to that of the working class. The only question remaining for Rainwater is whether "adequate" income means an absolute minimum or a relative minimum. Rainwater then uses opinion poll results to show that people believe a relative minimum—not an absolute minimum—is necessary for an adequate level of well-being. It follows that a minimum relative income must be provided to change the behavior patterns of the lower class.

Although Rainwater describes the disruption process as a complex matter of many elements, his discussion of a proposed solution fails to answer many questions. From Rainwater's point of view, the appropriate questions are: (1) to what extent is low relative income the root cause of family disorganization? and (2) would assuring a minimum relative income significantly reduce such disruptions? Lost are such subtleties as the economic importance of the male to the mother and her children, the problem of community influences that foster disruption, and the time required for families to change their behavior.

Rainwater and Cutright contend that low relative income causes family disorganization. Although this essay is not the place for a full-blown analysis of the issue, the subject merits further discussion beyond the remarks on page 13. The chain of reasoning behind the relative income hypothesis seems to run as follows: (1) Participant-observer studies and the data showing highest family disorganization among the lowest income groups confirm the significance of low income as a primary cause of illegitimacy and marital disruption, (2) absolute income has increased dramatically in the last 10–20 years with no accompanying decline in family disorganization, but (3) incomes of the bottom 20 percent have not improved relative to median family income, thus (4) the lack of relative income gains must be responsible for the fact that family disorganization has not declined.

There is good reason to question this conclusion. First, neither Cutright nor Rainwater demonstrates that the high correlation between low income and family disorganization means that low income is a cause of family disorganization. The causation can run in both directions or some third factor may be at work to cause both. Although it is difficult to demonstrate causation in social science analysis, a general model that takes simultaneous account of many forces would yield more convincing evidence.

Second, the direct evidence against the relative income hypothesis is at least as strong as the evidence in support of the hypothesis. Consider racial differentials over time and by region. The median income of black families relative to the median income of white families increased from 0.52 in 1959 and 0.55 in 1960 to 0.64 in 1970 and 0.63 in 1971. If we consider the more appropriate income variable, median income of intact families, the increase in black to white incomes is even larger; the ratio of such median incomes rose from 0.59 in 1959 to 0.73 in 1971.[17] In spite of these relative income gains by black

[17] U.S. Bureau of the Census, "Current Population Reports," series P–60, No. 85, "Money Income of Families and Persons in the United States," U.S. Government Printing Office, Washington, D.C., 1972, and U.S. Bureau of the Census, "Statistical Abstract of the United States," 1971 (92d ed.), Washington, D.C., 1971, pp. 316–320.

families, family disorganization was at least as high among blacks, and in many ways higher in 1971 than in 1959.[18] Looking at regional differences, one sees a similar picture. Incomes of blacks relative to whites are much higher outside the South than inside the South. According to Rainwater, this should imply higher family disorganization among southern black families than among northern ones. Yet, in fact, black families are no less stable in the South than elsewhere.

One should not interpret this evidence to mean that relative income plays no role in illegitimacy or in marital disruptions. Low relative incomes might still be one cause of family disorganization. But it is certainly not the only major cause. Rainwater is probably closer to the right track when he emphasizes the relatively low income contribution of the father. The male's earnings are financially less important to black than to white families. In 1971 median earnings of black males working full time were 68 percent of median family income of black husband-wife families; the comparable figure for whites was 77 percent.[19] While black men raised their earnings position relative to white men between 1959 and 1969, the improvement in relative earnings of black women was much larger. As a result, the ratio of black female to black male earnings increased from 0.45 to 0.55 in only 10 years.[20] Black male earnings have approximately kept pace with the large rise in average cash welfare payments. However, taking account of the increased availability of other transfer benefits such as food stamps and medicaid, one would probably find a decline in the ratio of average earnings of black males to transfer income available to female heads with children. These trends by themselves do not prove any particular relationship. But they suggest that diminished importance of the father's financial contribution to total family income may be the more significant economic explanation of broken families than is low combined income of some husbands and wives relative to average family income.

IV. Family Unit Policy

How to treat different types of families under income transfer programs is a difficult problem. Families differ not only by family income but also by sex of head, by number of children, and by presence or absence of parents and other related or unrelated adults. If Government policies are not to create inequities between families or to increase incentives for socially undesirable behavior, these family differences must be recognized in the design of benefit programs. Unfortunately, present Government programs suffer from poor design in this respect. Confusion is rampant because of the large number of different rules about family units. And the rules of some programs are obviously unfair to many families.

The most well-known and controversial family unit policy is the general exclusion of husband-wife families under the aid to families

[18] Rainwater's hypothesis might still hold if the income distribution among black families were more unequal in 1959 than in 1971. But the data show some slight improvement occurred for the lowest and second lowest fifth of black families.

[19] U.S. Bureau of the Census, Current Population Reports, series P-60, No. 85, pp. 45, 47, and 123.

[20] See Richard B. Freeman, "Changes in the Labor Market for Black Americans, 1948-72," *Brookings Papers on Economic Activity I: 1973*, The Brookings Institu-[illegible], Washington, 1973, p. 112.

with dependent children program. Although some two-parent families may qualify for AFDC because of disability or unemployment, most are ineligible. AFDC cash is prohibited to all families of fathers with full-time jobs who live at home, no matter how poor they are. This policy has been widely attacked as inequitable and socially disruptive. The policy treats families unequally, giving varying sums to families of the same size and income. Moreover, it encourages parents not to marry, to separate, and to avoid remarriage. What is not generally recognized is that welfare's discrimination against intact families falls under a general set of problems in family unit policy.

This section first provides a general framework for viewing the family unit problem. Next comes an examination of family unit policies under existing programs. These two sections and the findings of some papers in the volume lead to a discussion of the implications for structuring future Government programs.

A. The General Framework [21]

The family unit problem arises from the recognition that the sharing of income makes individual income a poor measure of potential living standards. At the same living standards, supporting a large family simply requires more income than supporting oneself. Because needs differ by family size, most people would regard as unfair rules that based tax payments and transfer benefits *only* on each individual's income. Common notions of equity suggest that the tax code and transfer program rules should take account of differences in family circumstances. Most people believe that a single individual should pay higher taxes than a father of five with the same income.

The tax code treats different family units differently by allowing deductions from taxable income and by using different tax rate schedules. The taxpayer may deduct a specific amount for each person he supports. The taxpayer's family relationships and income-sharing arrangements can also affect his tax rate schedule if he is married or qualifies as a household head. Similarly, in the case of transfer programs, the benefit an individual or group may receive depends on rules defining "countable" income and on rules prescribing the benefit schedule. As a result of these rules, taxes tend to be lower and transfer benefits higher in large family units than in small family units with the same income.

There is an important difference between family unit rules in the tax code and in transfer programs. While the tax code covers every individual or group of individuals qualifying on an income basis, transfer rules exclude some individuals or groups on grounds other than income. For the most part, this discussion considers the issue of eligibility as simply an added example of how benefit schedules differ for different family units.

Although family unit rules are intended primarily to provide equity among families, one cannot ignore their effect on incentives. As noted

[21] The author draws heavily on the excellent work of William A. Klein in writing this section. See his "Problems in Choosing Family Unit Rules for a Negative Income Tax" Institute for Research in Poverty. Discussion Paper 61–70, University of Wisconsin, Madison, Wis., 1970, and "Familial Relationships and Economic Well-Being: Family Unit Rules for a Negative Income Tax," vol. 8, *Harvard Journal of Legislation* 362, March 1971.

above, AFDC rules financially discourage marriage and encourage separation by making most intact families ineligible for benefits. In the case of the income tax code, the different tax schedules applying to married couples and to single individuals may cause a man and woman's total tax payments to rise or to fall if the couple marries. Unfortunately, one cannot always avoid influencing family formation incentives. The nature of the family unit problem is that efforts to promote equity sometimes cause undesirable incentive effects.

To determine the equity and incentive effects of transfer programs, one must examine the definition of countable income and the benefit schedule. Adding a person to the recipient unit may increase benefits because potential benefits rise with the number of members. On the other hand, counting the new member's income as available to the recipient unit tends to reduce benefits. The total effect on benefits will depend on which of the separate effects is larger.

A simple example illustrates the two separate effects. Consider the case of a man and woman living together and potentially eligible to receive benefits on an individual or married couple basis. That is, the filing unit consists of individuals or families, where a family is defined as persons living together related by blood, marriage, or adoption. The advantage or disadvantage of marriage depends on how benefits change with the size of the unit and on how much each individual earns separately. Suppose, first, that the incomes of the man and of the woman are zero. Then the advantage of remaining single and filing separately depends only on the benefit schedule. If the per person benefit is the same for the first and the second person in any unit, say $800 per person, then filing separately or together would not affect total benefits of $1,600. Alternatively, per person benefits may decline with unit size; for example, suppose the first person's benefit is $1,000 and each added person's benefit is $600. Then filing separately will yield $2,000 or $400 more than the $1,600 received by filing jointly.

Once we remove the assumption of zero income, advantages of filing separately may exist even with a benefit schedule whose per person amounts do not decline with family size. Suppose the woman earns $8,000 and the man earns zero. If they file jointly as a husband-wife unit, the $8,000 joint income may disqualify the couple from any benefits. But filing separately would allow the man to continue receiving benefits.

Although the benefit schedule and the income definition together influence the financial advantages of various family and household arrangements, the two types of rules ostensibly have different purposes. The benefit schedule, which specifies how benefits change with the size of the unit, should reflect the needs of units of various sizes. Larger units may require a smaller number of dollars per person to achieve the same living standard as smaller units. This is the usual justification for benefit schedules in which per person benefits decline with unit size. How the rules define the unit's income implicitly depends on how income is or should be shared. In the example noted above, the rules presume that husband and wife share their income but that single men and women living together do not.

B. Some Perspectives on Actual Family Unit Behavior

A good deal of work in these two volumes examines actual sharing patterns among the poor. Stack and Semmel, Fried and Fitzgerald, Rainwater, and Billingsley are among those who emphasize the importance for the poor of large units within which people share income and consume goods together. They observe that many families reside in households with relatives outside the immediate family and with nonrelatives. The sharing of resources by the poor often extends beyond the household. Stack and Semmel point to large domestic networks involving a number of households in which pooling of goods and income is extensive. The pooling may take a variety of forms. Contributions from one household to another may occur randomly or as a result of one household's temporary shortfall or temporary high point in income. Or the poor may do more sharing of such goods as dresses or cleaning equipment.

An application of economic theory provides a plausible explanation for these observations. Assume first that individuals gain satisfaction from normal goods and services and from privacy. Second, note that larger units can achieve the same living standards as smaller units at lower per-person costs. (Many see this tendency as analogous to economies of scale in production by a firm.) [22] According to economic theory, a rise in income should increase one's purchases of all normal goods. If privacy acts as a normal good, one would expect that as their incomes rise, people would buy somewhat more privacy (less sharing) and somewhat more of other goods. This reasoning would explain why we observe larger units more often among low-income than among moderate-income people. For example, statistics from another subcommittee study reveal that the percentage of sample households with nonnuclear households is highest in cities providing the lowest welfare benefits.[23]

The studies in this volume suggest that sharing may be an important way for low-income people to raise their living standards. To the extent that poor families gain access to such goods as a vacuum cleaner, a good television, and wide clothes selection, it is through extensive sharing. A second important advantage of large units is the insurance they provide against temporary shortfalls in income. Analytically, this arrangement is comparable to the pooling of risks in a large financial portfolio. Such an insurance-type advantage of large units may help allow poor families to consume a higher share of their income than nonpoor families do. But these benefits of large units may have associated costs. One is loss of privacy. Another is some loss of independence. If a particular good is shared among many,

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[22] Actually, the analogy falters when one looks closely at the household economies. One may be able to provide the same space per person at lower cost in large as opposed to small units. However, this does not imply an economy of scale because the two housing units are different goods by the very fact that they are shared by different numbers of people.

[23] Nonnuclear households are those which contain adults other than the head and spouse. James R. Storey, Alair A. Townsend, and Irene Cox, *How Public Welfare Benefits Are Distributed in Low-Income Areas*, Paper No. 6. Studies in Public Welfare. Prepared for the use of the Subcommittee on Fiscal Policy, Joint Economic Committee, U.S. Government Printing Office, Washington, D.C., 1973, p. 22.

the owner of the good may feel a necessity to consider other opinions before trading the good for another. A third cost is the risk that the unit may break down. After sharing your goods with others with the expectation of favors returned, the other parties may move or may decide simply to drop out of the sharing unit.

An interesting implication of this explanation is that a rise in incomes [24] among the poor is likely to lead to a less proportional rise in the consumption of goods and services. Some of the added income may be used to buy such intangibles as privacy, independence, and lower risks of extreme income inadequacy. Trends in living arrangements do show a statistical association between income and household formation. Rising incomes in the last two decades have occurred alongside large increases in household headship by women in disrupted marriages and by aged persons.[25] Ironically, increased income can add to observed poverty figures if the increase stimulates many low-income people to form their own households. That is, individuals who would be classified as nonpoor if they resided and shared income in a large family unit become classified as poor when they form an independent household.

Turning the analysis around leads to an alternate explanation of the facts and some different implications. One could argue that the taste for privacy influences income levels rather than the other way around. Consider the fact that leisure is a good. Then, one might expect those with a special desire for privacy and independence to give up more of another good, leisure, than do those with a taste for sharing in large units. We would again observe that those with moderate or high cash earnings tend to form smaller units and those with low earnings are in larger units. But, in this case, we would not necessarily expect added income to lead to smaller units. Although this latter hypothesis is plausible and may explain some behavior, it is probably not a good general explanation. It is not consistent with the observed fact that leisure is lowest among adult male workers with low wages and low earnings.

C. Family Unit Policy in Existing Transfer Programs

With the large number of Federal transfer programs have come many family unit policies. Cox examines some of these policies in a paper appearing in these volumes. Cox points out how family unit policies in different programs can affect a single family and how these program linkages can produce anomalous results. This section draws heavily on Cox's work to analyze some family unit policies in the context of the framework outlined above.

As in other areas of public policy, family unit treatment in existing programs has developed through a series of *ad hoc* adaptations by the President, by Congress, and by the Supreme Court rather than through conscious design. For example, the food stamp program initially used a household definition of the recipient unit. A Congressional amendment gave family status a role by requiring that household members be related in order to be eligible, but the Supreme Court recently ruled that this amendment was unconstitutional. The social

[24] We have not distinguished here between relative and absolute income effects. The actual stimulus to behavior could be a combination of both.
[25] See Cutright and Scanzoni, Table 4.

security program (Old Age, Survivors, and Disability Insurance, OASDI) focuses on individual contributions, workers' dependents, and their survivors to define the recipient unit. Over time, the definition of survivors and dependents has been extended from close relatives such as wives and children, to other relatives such as grandchildren who prove dependency, and now to a divorced wife who is currently not a dependent but who gained eligibility by spending many years as the wife of a covered worker. One result of these and other changes is that, in many programs, the working definition of a recipient unit is a combination of individual, family, and household concepts.

The AFDC program best illustrates the complexity of existing unit definitions. As a result of recent Supreme Court decisions sometimes the recipient unit has become the individual and other times it has become the family.[26] Looking at benefit schedules and income treatment separately may help simplify matters. The family is the recipient unit if one is establishing a benefit schedule. This clearly is the implication of *Dandridge* v. *Williams* (397 U.S. 471 (1970)). In this case the Supreme Court allowed States to use family maximums so that children born into families already receiving the maximum would not cause any change in the AFDC grant. The majority makes clear that benefits accrue to the family unit as a whole.[27] Somehow it does not seem to follow, according to the Court, that income accruing to one family member is available to all. In *Lewis* v. *Martin* 397 U.S. 552 (1970), the Supreme Court ruled that ". . . California may not consider the child's 'resources' to include . . . the income of a non-adopting stepfather who is not legally obligated to support the child as is a natural parent . . ."[28] This decision applies to virtually all States. At the same time, HEW regulations require that States consider the stepfather's income as available to the mother of AFDC children without proof of his actual contributions.[29] This is odd since it assumes that income which is available to the mother is not necessarily available to the children. In order for this case to make sense, one must consider the individual as the appropriate unit for the treatment of income. Admittedly, the case of the stepfather's income is an exception. In general, the income available to one family member is available to all. For example, earnings by the mother or by a child who is not a student reduce the total grant to the family, not simply that part of the grant intended for the earner.[30] On the other hand, income of a household member who has no legal obligation to support any of the recipients is not considered as available to any member of the unit.

[26] To some extent, household status is also relevant, in that the presence of other people in the household can affect grant levels. See the latter part of this section for further discussion of this point.

[27] Justice Stewart states, as part of the majority opinion, "It is not more accurate to say that the last child's grant is wholly taken away than to say that the grant of the first child is totally rescinded. In fact, it is the *family* grant that is affected." (Italic in original.) See pp. 476–477 for quotation and pp. 477–481 for elaboration. If the Court had prevented States from using family maximums, it would have been difficult to allow States to reduce per person benefits at all with increases in family size.

[28] See p. 560.

[29] See 45CFR233.20, (a), (3)(vi).

[30] Earnings of children who are students and, in some States, small amounts of income set aside for future educational needs of children, do not reduce AFDC payments. This case differs from the treatment of stepfather's income in that the income does not affect any member's AFDC grant.

To some extent, OASDI uses unit policies that are combinations of individual and family definitions. As with AFDC, the benefit schedule implicitly takes account of income sharing within the family by establishing family maximums. On the other hand, the treatment of income applies the individual definition. In contrast to AFDC, earnings by the surviving mother can reduce only her OASDI payment, not the payments intended for her children. Implicitly, a mother's income is viewed as available to her children if it comes in the form of a partial OASDI payment but is viewed as not available if it comes in the form of earnings. It is not easy to picture actual families dividing their total income into a pooled part and into individual parts on this basis.

Of primary importance in analyzing these family unit definitions is the question of how well existing policies achieve various social objectives. The objectives that most closely relate to family status are: (1) horizontal equity; that is, the notion that people in similarly needy situations should receive similar amounts of benefits, and (2) desirable family composition incentives; that is, the notion that the program should not itself provide financial inventives to split family or household units. Although family unit policies may also influence work incentives and childbearing incentives, the focus here is on family and household composition.

First, consider different family unit-policies with respect to marriage and separation. In the case of AFDC, the family-splitting incentive is well known. Poor families are not generally eligible for AFDC benefits unless one natural parent is absent.[31] This means that the income available to a mother and her children may rise if the father deserts the family. Although costs to the entire family (including the father) would increase as a result of his establishing a second household, the father's living standard might go up because his income can be used solely for his own support. Thus, both subunits of the original family can achieve a higher living standard than is available on an intact basis. This treatment leads to inequities between equally needy families and constitutes an incentive for parents to separate.

A less well-known aspect of AFDC policy is that the disencentive to marry is small or zero. A man and woman living together with their children are treated the same way whether or not they marry. They are generally ineligible.[32] Marriage might have an indirect effect on eligibility in helping to resolve the issue of paternity. However, where there is definite knowledge about paternity, marriage does not affect eligibility. A second important case is that of a mother living with her children and with a man other than their father. If she marries this man, the AFDC grant may only fall a small amount or not at all. The welfare agency may not assume that any of his income goes to support the children whether or not the man is married to the mother.

This means that marriage need not affect that part of the AFDC grant meant for the children. The marriage almost certainly would reduce the AFDC grant to the mother. The fall in the mother's grant would occur because, as the man became legally obligated for the support of the mother, the welfare agency could assume without

[31] See *Lewis* v. *Martin*, 397 U.S. 552 (1970), p. 560 and 45 C.F.R. 233.20, (a), (3),(vi).

[32] Families with two natural parents present may be eligible for AFDC if one parent is incapacited or, in some States, if the father is unemployed.

proof that he shared his income with her. Nevertheless, in some cases where the family's AFDC grant would equal the State's family maximum, whether or not the mother were included or possibly where the husband's income were zero, marriage would not legally affect the AFDC grant at all.

Food stamp program rules certainly do not discourage marriage. Currently, if a man and woman live together, with or without children, they receive exactly the same consideration whether or not they marry. They would generally qualify or not qualify as a single unit on the basis of their total income and assets. Congress apparently tried to change this policy by amending the food stamp law to require that all household members under 60 be related if any are to qualify for benefits. The effect of this amendment would have been to encourage marriage. However, the Supreme Court ruled this provision invalid, thereby allowing households whose members are not related to qualify.[33]

Old Age and Survivors Insurance rules encourage marriage in some cases and discourage marriage in others. Cox covers these rules in some detail. A surviving child beneficiary under age 18 or age 18–22 in school loses all benefits by marrying. On the other hand, a potential retiree living with a woman could increase the couple's benefits through marriage.[34] The widowed mother under 60 loses her survivor benefits if she remarries. However, this rule may have little effect for two reasons. First, her own earnings may easily reduce her benefits to zero even if she remains unmarried. Second, since benefits to her children are unaffected by her marriage and since her family may receive the family maximum whether or not she remains eligible for benefits, the marriage may have no effect on total family income.

Taken together, these three transfer programs do little to discourage marriage among adults living together. From the equity point of view, married couples do not receive harsh treatment under transfer programs relative to the treatment accorded unmarried couples living together. The well-known family-splitting incentive present in the AFDC program has to do largely with *household* composition. What is most at stake is the presence or absence of a natural parent in the house, not the marital status of the natural parents. Of course, these two factors are often closely related in practice. The emphasis on the household status of parents does not mean that the AFDC program relies on the household unit definition in other respects. To understand how AFDC food stamps and OASDI treat different types and sizes of households requires a separate analysis.

Another question is whether transfer programs alter the household composition incentives that would have existed in the absence of the programs. This is a difficult question to answer largely because of problems in determining the exact nature of economies of scale to household size. The effect is likely to vary substantially. Some groups may save a good deal by pooling resources and consuming many goods jointly but others may save little. Thus, providing the same benefit amount to two households of two as to one household of four may or may not encourage the formation of larger units. Although

[33] See *Moreno* v. *U.S. Department of Agriculture*, docket No. 72-534, June 25, 1973.

[34] The new wife of a retired man must wait 1 year after the marriage before receiving her benefits.

there is no unique standard of neutrality, one can examine whether transfer programs offer significant financial advantages to various types of household arrangements.

The case of AFDC is complex. As noted above, the program significantly discourages natural parents from living in the same household. What are the effects on other types of household groupings? Several factors are at work in determining the benefits to an AFDC family which shares a household with other persons. Essentially, the household members not part of the AFDC family are not included in the recipient unit. This means they receive no benefits, but it also means that their income does not reduce the AFDC grant. Were this all, one could say that AFDC does not at all discourage persons who combine to achieve economies of scale. However, AFDC interferes with reaping the savings associated with low housing costs per person in large units. The welfare agency generally reduces the family's grant to take account of the rent contribution presumably forthcoming from those not in the AFDC unit and sometimes to take account of lowered food costs. These rules reduce what is possibly the most important financial advantages to large household groupings.

On its face, the latter rule appears equitable in that similarly needy families receive similar benefits. After all, a family needing less than another because of household sharing arrangements should receive less. The problem with this reasoning is that household composition is partly a matter of choice. To some extent, one may view the decision to buy privacy by losing rent savings as an ordinary consumption decision. And just as needs do not differ because one person chooses good shirts over good shoes while the other person does the opposite, the AFDC family's needs should not be said to decline when it gives up some of one good (privacy) for more of another (say, high quality food).

In the case of food stamps, the household is the unit of primary importance in determining benefits. With some exceptions, the program takes account of the combined needs and combined income of all household members. The net effect on food stamp benefits of adding a person to the household depends on the person's income. If the added member has no income, food stamp benefits to the household rise. Household benefit levels increase with household size by nearly equal per person amounts. On the other hand, the added person may cause household food stamp benefits to fall (coupon prices to rise) if he has sufficient income. Thus, the food stamp rules may or may not discourage persons from saving money by combining into large household units.

The OASDI program is the easiest to assess with respect to household composition incentives. There is no financial discouragement at all to those who combine into large household units. The presence or absence of other household members does not affect the size of payments to OASDI beneficiaries.

Up to this point, the discussion has considered the extent to which AFDC, food stamps, and OASDI allow persons to gain higher benefits and/or living standards under some household arrangements than others. In some cases, it was found that differences in treatment under AFDC and food stamps can act as financial disincentives to recipients trying to save money by joining a large household unit.

In addition to these effects, the income from transfer programs could well influence people to choose smaller household units. Although the OASDI program does not directly discourage large household groupings, its benefit checks may enable people to live in smaller units than they could otherwise afford. Increased income alone may stimulate family heads to start their own household rather than share a household.

D. General Implications of Family Unit Policy

To improve family unit policies, one must understand not only what the policies are but also why they developed. What rationale exists for current policies? This section views current policies as the result of attempts to compromise conflicts among competing goals while recognizing some actual family patterns. Achieving a consensus as to goals is relatively easy. Most people would agree that family unit policies should not encourage the breaking up of families and should not discourage low-income people from living together to save money. Additional agreement is likely on the goal that people with similar needs should receive similar treatment under income transfer programs. Although people may disagree about the degree to which a family's needs change with size, poll results reported by Rainwater strongly suggest that people believe dollar needs per person fall significantly as family size rises.[35]

Unfortunately, agreement concerning these goals does not mean agreement on family-unit policies. The problem is that even these few goals conflict with each other. One can be achieved only at the expense of another. Thus, policymakers must choose among the goals, deciding which to stress and which to subordinate.

To illustrate the conflict among goals, let us consider defining the recipient unit in a variety of ways. Suppose the individual were the recipient unit. Benefits would be independent of family or household groupings. This has the important advantage of not discouraging marriage or large household groupings. However, use of an individual definition makes it difficult to base benefits accurately on needs. Under an individual definition, a family of six might receive six times the benefits available to an individual with the same private income. Such a policy would run counter to the widely held notion that per person needs decline as family size increases. Rainwater reports poll results indicating most people believe that total money amounts necessary for a given living standard rise slowly with family size. The Social Security Administration poverty lines suggest that a seven-person family needs only about 2.5 times as much money as a two-person family needs, while the people surveyed place the ratio at 1.46. Contrast these ratios with the 3.5 ratio that would result from use of the individual unit definition.

A second problem with the individual definition occurs if there is not a single proportional tax rate applying to transfer benefits as well as to taxable income. Currently, and under most reform proposals, the rate at which transfer benefits decline with the first dollar earned is higher (often it is set at 50–70 percent) than the rate at which taxes are paid on the first dollar of taxable income (14 percent). This factor, combined with an individual definition of recipient unit, would mean

[35] See Rainwater, table 5.

that a married couple in which husband and wife each earned $5,000 would end up with a lower after-tax, after-transfer income than a married couple in which the husband earned $10,000. Such a treatment would also present the difficult problem of how to avoid income-shifting between family members

Defining the family as recipient unit could solve these two problems of the individual definition. In addition, the family definition would not discourage savings attained through the formation of large households. The disadvantage is that the rules could create a financial disincentive for a man and a woman with children living together to marry. If the "family" included the man and counted his income, their benefits probably would be smaller than those of mother and children only. Another way of looking at it is that benefits could differ for two groupings of individuals merely because one couple is married and another is not.

A household definition would eliminate the adverse family formation incentives. People living together would receive the same treatment whether or not they married. Unfortunately, the household-unit rule could discourage savings through combining into a single household. A low-income family could lose its benefits by moving in with a moderate-income family.

One may view many of the current family unit policies as ad hoc procedures to attain compromise among competing goals. Surviving wives and children of workers covered by OASDI are eligible for benefits as individuals. Moreover, each recipient's income affects only that individual's OASDI benefit. However, the OASDI program uses a family maximum benefit, departing from the individual unit policy in recognition that individual needs drop as family size rises.

The AFDC program uses legal support status to help determine eligibility and benefits and to account for income. As noted above, the pure family definition can create a financial disincentive to marry and an inequity between married and unmarried parents. By defining the unit on the basis of legal support relationships, the AFDC program avoids these two problems.[36] This adaptation, however, gives vastly better treatment to families with a stepfather than to families with both natural parents present.

Since no family-unit policy can achieve simultaneously all desirable goals, the appropriate question is, How can Congress fashion rules that offer the best compromise among competing objectives? The first step is to recognize the impossibility of removing every bad incentive or equity feature. By taking account of this point, the Congress might show less inclination to adopt rules on an ad hoc basis.

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[36] A welfare program in England makes different types of ad hoc adjustments to the family definition of recipient unit. To avoid the disincentive to marry and the potential inequity between married and unmarried couples living together, the Government investigates on a case-by-case basis whether an unmarried man and woman living together should be considered as husband and wife for program purposes. See *Cohabitation, Report by the Supplementary Benefits Commission to the Secretary of State for Social Services*, Her Majesty's Stationery Office, London, 1971. If a transfer program in the United States used such a procedure, the Supreme Court might well declare the rules unconstitutional on grounds of invasion of privacy. Klein, "Problems in Choosing . . .," pp. 69–71, extrapolates this judgment from the logic of *Griswold* v. *Connecticut* 381 United States 479 (1965).

The second step is to avoid the worst distortions of family and household behavior. Most students of current law would agree that the AFDC policy of generally excluding intact families on eligibility grounds (regardless of income) constitutes the worst family-unit policy. According to Honig's results, this policy has a concrete effect on people's actions; it actually helps to disrupt families. Cutright and Scanzoni argue against this conclusion. However, whether or not this AFDC policy actually helps to break up families, the enormous inequity between families would remain. If the rationale for the current policy is that families with two natural parents present have better earnings opportunities than families with only one natural parent present, Congress might simply provide a less generous benefit schedule to the two-parent group.

The third step is to take account of actual family and household patterns, if only for administrative reasons. Stack and Semmel's findings of a substantial instability in household composition should warn against the difficulties of administering a program on the basis of a household unit definition. If Congress had known that a large number of persons to whom it wanted to give food stamps lived in households with nonrelatives, it might not have passed the recent amendment denying benefits to households in which not all persons were related.

The fourth step is to consider seriously how to adjust benefits by family size. Current and proposed policies may accurately reflect varying needs, but they may be far out of line. The poverty lines of the Social Security Administration, which have a good deal of influence on policy, probably overstate the needs of large relative to small families. The SSA poverty schedules take account only of savings in food costs achieved by large families.[37] If the savings possible through purchase of all other goods exceed the savings in the food area, as this author expects, then poverty lines overstate the rate at which money needs rise with family size.

Finally, rules about recipient units are under legal attack. Although the Supreme Court explicitly mentioned but did not rule on the constitutionality of using the household as the recipient unit,[38] recent decisions invalidating amendments to the food stamp law suggest possible legal trouble for the household definition. The Federal district court and Justice Douglas in a concurring opinion argued that the amendment making a group of related persons ineligible for food stamp benefits because they share a household with a person unrelated to the group was in conflict with fundamental personal freedoms, particularly the freedom of association.[39] By this logic, one could contend that any

[37] Mollie Orshansky, "Counting the Poor: Another Look at the Poverty Profile," *Social Security Bulletin*, January 1965.

[38] *Moreno* v. *U.S. Department of Agriculture*, U.S. Supreme Court No. 72–534, footnote 4, p. 5.

[39] *Moreno* v. *U.S. Department of Agriculture*, 345 F. Supp. 310 (1972), p. 314 and U.S. Supreme Court No. 72–534, June 25, 1973, concurring opinion of Justice Douglas, "The right of association, the right to invite the stranger into one's home is too basic in our constitutional régime to deal with roughshod. If there are abuses inherent in that pattern of living against which the Food Stamp program should be protected, the act must be 'narrowly drawn,' *Cantwell* v. *Connecticut*, 310 U.S. 296, 307, to meet the precise end. The method adopted and applied to these cases makes 3(e) of the act unconstitutional by reason of the invidious discrimination between the two classes of needy persons."

household unit definition results in a similar conflict since a poor family could lose all benefits by sharing a household with a middle-income family. This reasoning leads to another possible legal problem with the household definition—the assumption of income-sharing between persons who lack legal responsibility for each other. The Supreme Court recently declared unconstitutional a provision eliminating from eligibility for food stamps those households which include a person claimed as a tax deduction by a person who is himself ineligible for food stamps. To some extent, the Court objected to the presumption, in the absence of evidence of actual support, that the person claimed as a dependent is not needy.[40] However, the critical point is that the Court finds it irrational to judge the needs of some household member or members on the basis of the presence of a non-needy household member.[41] Would the Court also find it irrational to continue to disqualify many household members by counting income not actually available to them simply because it is earned by other household members? If the answer is yes, then the household unit definition itself would become unconstitutional.

If the household unit definition falls, then the Congress will probably increasingly rely on the family definition of recipient unit. If it allows the family definition, the Court will be saying that it is illegal to provide disincentives to household foundation, but legal to allow disincentives to family formation. Under these circumstances, it will be impossible to treat an unmarried couple like a married couple unless the transfer program moves to an individual definition of recipient unit.

V. The Pursuit of Absent Fathers

Fathers are absent from the overwhelming majority of AFDC families. In 1971, 31 percent of fathers of AFDC families had deserted or separated from their families; 14 percent had divorced the mother; and 28 percent left their children without marrying the mother. All told, about three-fourths of fathers of AFDC children had left their families. Only in 24 percent of ADFC families were the fathers dead, incapacitated, or unemployed.[42]

This problem of absent fathers has become a critical concern largely because of the dramatic growth in the AFDC caseload. Between 1961 and 1971, the number of AFDC recipients nearly tripled, growing from 3.6 million to 10.6 million persons. About 80 percent of this increase occurred among families affected by illegitimacy or marital instability. These families receive welfare payments not only because the father is absent but also because the father provides too little financial support. Given these facts and the public outcry against rising welfare costs, it is no wonder that pursuit of absent fathers has become a controversial issue. Many advocate vigorously pursuing the father and extracting more money from him in order to save money for the taxpayers. Opponents argue that vigorous pursuit could inter-

[40] *U.S. Department of Agriculture* v. *Murry*, U.S. Supreme Court No. 72-848, majority opinion of Justice Douglas, pp. 3–5.

[41] In Justice Douglas' words, "We conclude that the deduction taken for the benefit of the parent in the prior year is not a rational measure of the need of a different household with whom the child of the tax deducting parent lives and rests on an irrebuttable presumption often contrary to fact." Ibid., p. 6.

[42] Social and Rehabilitation Service, HEW *Findings of the 1971 AFDC Survey: pts. I and II*, NCSS reports AFDC-1(71) and AFDC-2(71), 1972.

fere with rights to privacy and in any case would cost the taxpayers a good deal more than it would save them.

Part 2 of this series contains papers on both sides of the issue. Krause favors the vigorous pursuit policy while Stack and Semmel oppose it. However, in spite of their basic policy disagreement, the authors appear to agree on some issues. This section attempts to pinpoint their areas of agreement and disagreement and to assess what important information gaps exist that are relevant to this issue.

A. Establishing Paternity

Establishing paternity is often a prerequisite for obtaining child support payments. It is generally only natural parents who are legally obligated to provide financial support for the child. Thus, one must begin any attempt to extract support payments by finding out who the natural parents are.

Krause points out that an equally important reason for establishing paternity is to help the illegitimate child gain the same rights as the legitimate child. He maintains that the child's right to know his father follows from the Supreme Court's rulings providing equal protection for illegitimate children. Krause argues, "Equal protection for the child born out of wedlock will remain an empty phrase if it is not combined with active efforts to find the man vis-a-vis whom the child is to have substantive rights." Identification of the father becomes especially important in establishing the child's rights to such social "entitlements" as social security, veteran's benefits, and health insurance. This suggests extending the vigorous pursuit policy to cover all families, not simply welfare families.

Is a decent program to establish paternity possible? Krause believes the answer is yes in spite of the current absence of humane and efficient State systems. He points to the experience of Sweden, where a mandatory paternity action takes place in cases of illegitimate births in which the father does not acknowledge the child. The threat and probable success of court action strongly encourages most fathers to acknowledge their children voluntarily. As a result Sweden is successful in determining the father in 95 percent of illegitimate births. Krause also cites the fact that modern medical techniques can greatly improve court procedures to establish paternity. Unfortunately, courts have done little to utilize the modern techniques, in large part because of the inadequacy of State procedures to make such evidence available. In fact, Krause calls current State procedures "scandalous" and in need of reform. A decent program to establish paternity is practical, not only because of scientific advances but also because of the number of fathers who voluntarily acknowledge their children. Krause and Stack and Semmel agree that a great many fathers voluntarily aid their illegitimate children directly and indirectly by providing access to financial help and friendship from the father's close relatives. The authors disagree about whether a significant decline in voluntary help might result from using involuntary methods of establishing paternity in combination with child support collection effects. Stack and Semmel warn that voluntary acknowledgment of fatherhood and voluntary contributions would decline significantly if extensive legal efforts were made to collect child support.

B. Vigorous State Efforts To Collect Child Support Payments

The authors stand on opposite sides of the question of whether or not the Government should devote substantial resources to collect child support payments from absent fathers of welfare children. Stack and Semmel argue against the vigorous pursuit policy.

First, they contend that the State would gain little if any cost savings. Collection efforts are costly and are likely to yield little added child support due to the fact that incomes of absent fathers are low. Second, Stack and Semmel point out that in most States added child support payments by the absent father would not financially aid the child. Each dollar of added child support usually causes a dollar reduction in the welfare grant and thus, no change in total income. The child may even lose financially as a result of the vigorous pursuit policy. Voluntary contributions by the father may dry up if he fears they will lead to large increases in his obligations. Further, the child may lose access to a whole network of the father's relatives as fewer fathers voluntarily acknowledge their children because of the fear of government pursuit.

Krause cites the alarming growth in the welfare caseload and in parental nonsupport as justification for action. He believes the continuing increase in child abandonment could break down the current system so as to require support for all children. Krause believes the vigorous effort to collect support would yield direct State benefits because many absent fathers have at least moderate incomes. However, even if direct government savings do not exceed collection costs, Krause believes the vigorous pursuit policy is important as a deterrent to family splitting.

Comparing the positions of Krause and of Stack and Semmel, one finds that some of their disagreement concerns three empirical questions. First, are the incomes of absent fathers large enough to yield cost savings that at least pay for collection costs? Krause says yes and Stack and Semmel say no, but reliable data to settle the issue are lacking. The second question is, could vigorous collection efforts deter illegitimacy and child abandonment? Krause says yes and, although Stack and Semmel do not deal directly with the question, they seem to say no. The third question is, is there a danger that child abandonment will continue to increase (with or without a vigorous policy) and potentially bankrupt existing programs? Stack and Semmel again do not comment, as their paper is not concerned with the question. Krause sees this danger as a live possibility and maintains that only a vigorous collection effort could prevent the enormous financial burdens that could result from further increases in family breakdown. Since settling these empirical questions would narrow the range of disagreement, these topics are important subjects for future research.

THE IMPACT OF WELFARE PAYMENT LEVELS ON FAMILY STABILITY

By MARJORIE HONIG *

SUMMARY

This paper examines the two most controversial questions surrounding the current welfare system. (1) Do high welfare benefits induce families to break up? and (2) do high welfare benefits influence families to choose welfare over work? It is generally recognized that the AFDC program, largest of the current public assistance programs, contains financial incentives favoring family splitting and discouraging work. Poor families with male heads are not generally eligible to receive AFDC assistance; furthermore, female heads of families may receive incomes from the AFDC program equal to what they could earn if they entered the labor force. It is clear that these incentives are undesirable from several points of view. The purpose of this study is to estimate empirically the extent of the impact of these incentives on family behavior. Can the proportion of female-headed families in the largest metropolitan areas be related to, among other things, the size of the AFDC monthly payment available to a typical female-headed family? and secondly, can the proportion of female-headed families who are actually AFDC recipients similarly be related to the amount of income available from the AFDC program?

These questions have played an important role in public discussion of the merits of the present welfare system and the desirability of finding new approaches to dealing with poverty. There has been however relatively little effort toward obtaining empirical evidence of the public's response to the incentives inherent in the present system. This study attempts to provide some tentative answers to these important questions.

According to the findings reported in this paper, *high welfare payments do help to cause family splitting and do influence women heading families to become welfare recipients.* These findings are based on empirical tests of a model designed to examine the impact of differences in AFDC payment levels in 44 metropolitan areas in 1960 and 1970. The model postulates that there are several influences, including AFDC payment levels, on interurban differences in the share of women heading families with children and on differences in the share of such families receiving AFDC payments. Specifically, the model

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seeks to determine whether the size of the average AFDC payment in metropolitan areas exerts an independent, positive effect on both these proportions. The data suggest positive answers to both questions.

Although AFDC payment levels were found to exert a discernible impact on behavior, the overwhelming share of AFDC families became one-parent families for reasons other than AFDC. Nevertheless, the magnitude of the AFDC impact is impressive. The 1960 figures indicate that the independent effect of a 10 percent higher AFDC stipend was to raise by 3–4 percent the share of families headed by women. This result occurred among both white and nonwhite families. The relationships were weaker in the preliminary test using 1970 data, but they remained statistically significant.

While high AFDC payments encouraged family splitting and welfare dependency, high male wages and low unemployment rates did the opposite. Holding other factors constant, those areas in which male wages were high relative to AFDC payments had lower shares of females heading families with children. The effect of low unemployment rates was also favorable, though considerably less significant than the wage effects. These favorable effects of high male wage rates relative to AFDC payments and of low unemployment rates took place among both whites and nonwhites.

The implications of these findings for public welfare policy are straightforward. The findings provide empirical support for proposals to broaden public assistance programs in order to help working and nonworking two-parent families. If high AFDC support for female-headed families and low male wage rates cause family splitting and greater degrees of welfare dependency, it would appear advisable to reverse the policy emphasis. Without worsening the position of current recipients, the policy shift would entail: (1) improving wage and employment opportunities of low-income males, or providing the same benefits to two-parent families as to one-parent (female-headed) families; and (2) continuing to improve work incentive features in public assistance programs. Allowing a support level for two-parent families that is comparable to amounts available under the current AFDC program would reduce the existing family-splitting incentives and reduce the incidence of family splitting. The improvement in work incentives requires moderate income guarantees to families with no income while increasing the amounts of retainable earnings of welfare recipients. Other methods for raising work incentives and low incomes may be to use wage subsidies, earnings subsidies, or public employment programs.

I. Introduction

Much of the current controversy surrounding the welfare system is concerned with the question of whether welfare itself causes workers to leave jobs and influences fathers to desert their families. While attempting to raise the incomes of the poor, government programs have created undesirable disincentives to work and to maintain family unity. Providing government grants to those with little income is one way to eliminate poverty, if poverty is defined—as it usually is—as lack of income. The difficulty arises when the availability of a government grant encourages the poor and near poor to forego working for income they can receive without work. If, in addition, welfare income is available only to single-parent families, work disincentives are augmented by incentives to dissolve families for purposes of welfare eligibility.

The importance of work incentive aspects of public assistance programs is evident in the following figures.

In 1970 the United States spent about $10 billion in cash transfers under various assistance programs. At estimates of the poverty population at the time, expenditures would have increased another $11.4 billion if the incomes of all families then estimated to be in poverty had been brought up to the "poverty line"—defined at $3,968 per year for a family of four. A further increase in costs would have occurred, however, if all wage-earners in families under the poverty line had withdrawn from the labor force to receive their total income in the form of transfer payments at the gain of increased leisure. Total cash transfers for public assistance would have reached $28 billion in this case. This figure would be larger if wage-earners with incomes *above* the poverty line reduced their work hours because of strong preferences for leisure.[1]

The question of work disincentives is not new, of course. It has emerged once again, however, as a possible explanation of two dramatic—and paradoxical—patterns which have developed in the economy in the last decade, both of which have required substantial increases in expenditures in the public assistance sector. The first of these is the significant growth of the public assistance population in a period of increasing incomes and employment opportunities. In 1960, 5.2 million persons were receiving welfare assistance in the United States; by 1970, this number had increased to 12.5 million, with the largest increases beginning in mid-decade, the period of the greatest expansion of the economy. Although the recession in the latter part of the period could be expected to swell the welfare rolls, the rapid growth in the numbers receiving public assistance was established several years previous.

The second pattern to emerge in this period was the significant increase in the proportion of families headed by females, especially among the nonwhite population. In 1960, 22.4 percent of all nonwhite families were headed by females; by 1970 this figure was 26.7, a 20-percent increase in the 10-year interval. The change was smaller for the white population but consistent: In 1960, 8.7 percent of all white families were headed by females, by 1970, 9.1 percent.[2] This pattern as well is remarkable in that it occurred in a period of increasing prosperity. The breakdown of family life predictably would occur more readily in periods of depressed economic opportunities, where frustration and continued inability to provide for the family would produce increased desertions. Although increased incomes in this period may have allowed some women to establish separate households, it is unlikely that this factor alone could account for the sizable increase in female-headed households, especially in the nonwhite population.

In short, it is difficult to account for these trends on the basis of either macroeconomic conditions or changes in underlying social patterns during these years. The existence of increasing work disin-

[1] Policy Research Division, Office of Economic Opportunity, and "Characteristics of the Low-Income Population," *Current Population Reports*, series P-60, No. 81, U.S. Bureau of the Census.

[2] U.S. Bureau of the Census, *Current Population Reports*, series P-20.

centives from the public welfare sector provides a plausible hypothesis for both patterns, however. During this period the average welfare payment rose relative to market earnings. From 1960 to 1970, the average payment in the Aid to Families with Dependent Children (AFDC) program, the largest of the welfare programs, rose 72 percent while the average wage in manufacturing, for example, rose by 49 percent.[3] There was in this period therefore an increased monetary incentive for consumers to leave the labor force and to apply for public assistance.[4]

The increasing payment levels may have accounted as well for the increase in the proportion of families headed by females since AFDC provides assistance primarily to families with female heads.[5] There was, in other words, a built-in incentive for family dissolution in the public assistance system during this period. This may have taken two forms. The male supporter may in fact have deserted the family since its financial needs could be met by public assistance (or the mother may have encouraged the father to leave since she now had a steady source of income). Alternatively, the father may have appeared to desert, a case of what might be called "statistical desertion," although he remained in the vicinity to lend additional support to the family.

The AFDC caseload may therefore have experienced increases in this period from two sources related to increases in the size of the AFDC payment: (1) an increased propensity to apply for AFDC assistance on the part of women who became female heads of families for reasons unrelated to the level of welfare income, and (2) increases in the population of those nominally eligible for the program—female heads of families with children less than 18 years of age—due to desertions and reductions in remarriages. If the disincentive impact

[3] Annual Statistical Supplements, 1960 and 1970, Social Security Bulletin; U.S. Bureau of the Census, *Statistical Abstract of the United States: 1971*, p. 232. The increase in the average wage in manufacturing during this period was probably larger than the increase in earnings for the potential welfare population due to adverse structural changes in demand for low-skill labor during this period.

[4] In addition to increases in payment levels, two other changes in the 1960's may have added to the welfare rolls. Amendments to the Social Security Act in 1962 and 1967 lowered the reductions in welfare payments associated with increased earnings. Instead of reducing payments $1 for each dollar of earnings the new formulas allowed deductions for work expenses and for the first $30, plus one-third of monthly earnings above $30, before applying the dollar-for-dollar reductions. While encouraging work effort by allowing working recipients to keep a higher share of their earnings, the amendments also added to the attractiveness of welfare for those with some earnings. A further effect was to reduce the income loss from cutting back on work hours suffered by those with moderate earnings. Consider a State paying its full standard of $240 per month. A mother of three (father absent) working 150 hours per month at $3 per hour earns $450 per month in gross earnings and about $415 per month after taxes. In 1960, cutting her work hours from 150 to 100 would cause her net earnings to fall to $284, a drop of $131, but still too high for welfare. By 1970, the decline in hours would mean only a $45 drop in total income since, with $284 in net earnings, she could receive $86 in welfare payments.

[5] During the 1960's various States adopted programs providing assistance to families with unemployed male heads of households (AFDC-Unemployed Father Program). However, by 1969, only 4 percent of all AFDC cases were male-headed households (*Welfare in Review*, U.S. Department of Health, Education, and Welfare, September/October 1969). The AFDC program experienced the most rapid growth during the 1960's of all public assistance programs. By 1969 it accounted for nearly one-third of the public assistance budget with expenditures totaling $3.6 billion and assistance received by 1,876,000 families.

of welfare income took the latter form as well as the former, the expected market earnings of males may have been as important a factor in determining the size of the welfare population as the earnings of females. This factor has not previously been considered among the determinants of the size of the AFDC caseload. This study attempts to distinguish between the two forms of the disincentive impact of the AFDC program, and in so doing, to provide some estimates of the impact of AFDC on family stability.

Since the AFDC program is the largest public assistance program, the implications of policy alternatives such as changes in the level of assistance provided, as discussed in this study, are of interest in themselves to policymakers. In addition, the AFDC program lends itself to convenient statistical analysis of consumer response to the incentives inherent in all programs which provide a guaranteed minimum income and the opportunity for consumers to exercise choice regarding the alternatives of public assistance income and labor market activity.[6]

State welfare policies vary considerably. Variations exist regarding not only the amount of money provided to poor families but also the eligibility qualifications for aid. Many States in the past required applicants to prove long-term residency in the State while others did not. Some States allow two-parent families to qualify for assistance while others do not. Differences also have existed or continue to exist with respect to inclusion of employables, maximum age for child eligibility, and provision for poor childless couples. The widest variations in State eligibility criteria occur in the State and local programs, often general assistance programs.

These variations must be held constant when estimating the incentives regarding labor supply and family stability. If not, the size of the eligible population will vary across areas; for example, new residents will be eligible in some areas, not in others, and estimation biases may result. One difficulty with analyzing public assistance programs is that it is often impossible to identify the rules regarding eligibility since the programs vary considerably even in different localities within the same State.[7]

This is less true of AFDC, a Federal-State program. Intrastate variations in eligibility conditions tend to be smaller, and interstate variations are well documented. In addition, in contrast to the general assistance programs where payment levels tend to be uniformly low, average monthly payments in the AFDC program vary considerably

[6] There is no reason to assume that the magnitude of the disincentive effect will be the same in the AFDC program as in more general income maintenance programs such as the family assistance plan (FAP) once advocated by the President or a negative income tax scheme. The eligible populations are different and the factors which influence consumer choices in one program may differ from those in other programs. Nevertheless, since most public discussion concerning public assistance has been concerned not with the precise size of the work disincentive, but whether in fact it is present at all, the appearance of the disincentive in the AFDC program establishes a presumption that it may be present in such other income maintenance programs as a negative income tax.

[7] Previous attempts to measure the size of the work disincentive have been limited by the use of programs which tend to maximize these variations. See especially C. T. Brehm and T. R. Saving, "The Demand for General Assistance Payments," *American Economic Review*, vol. 54, December 1964, and Hirschel Kasper, "Welfare Payments and Work Incentives," *Journal of Human Resources*, vol. III, No. 1, Winter, 1968.

between States so that their influence on family behavior, if present, should be discernible.

Moreover, the AFDC program offers the consumer a choice between labor market activity and public assistance income, and is therefore akin to the proposals for more general income maintenance programs such as the FAP. This is in contrast to many State and local programs which do not offer assistance if there is an employable person present in the household. AFDC was developed in the 1930's to provide aid to children when one or both parents were either disabled or absent. The program was designed to permit the mothers of children to remain in the home. This remains a basic tenet of the program, although there have been some shifts in attitude due in large part to the growing tendency in the population at large for women with children to enter the labor force. Still, mothers are not required to work while receiving assistance from this program in most States. In the few States which so stipulate, the program specifies that work should be undertaken only if adequate child care is available. The opportunity for consumer evaluation of the relative returns from employment and welfare participation is therefore written into the program (it is always possible of course that in practice this opportunity may vary somewhat across areas due to different interpretations of the basic provisions of the program on the part of local AFDC administrators, but these differences can be presumed to be relatively minor).

Furthermore, the AFDC program is designed so that it is an easy matter to identify the nominally eligible population—female heads of families with children less than 18 years of age. It is therefore possible to identify the separate effects of the availability of welfare income on both the existing population of female heads of families and also on the population of women who become female heads of families for purposes of receiving AFDC assistance.

The method of analysis is to assess whether area differences in AFDC payment levels produce area differences in the share of women heading families and in the share of female heads receiving welfare payments. To perform this analysis, one must account for effects other than the AFDC payment level. These include State program features, alternative sources of income, and such area factors as the local unemployment rate, for example. One can then examine the independent effect on behavior of the size of the AFDC payment. Those expecting a significant impact would predict that high AFDC payments imply high shares of females heading families and receiving welfare. If AFDC does not affect behavior, then area payment variations should show no statistically discernible effect on female headship or welfare participation.

It should be recognized that the analysis does not provide a direct answer to the question as to whether increases in AFDC payments between 1960 and 1970 caused the observed increases in female headship or AFDC recipient rates. Instead, the paper examines whether variations in the level of AFDC payments were related to variations in the proportions of female headed families or the proportions of such families receiving AFDC assistance *in a given year*. It is, however, plausible to expect that if differences in AFDC payments across metropolitan areas are seen to influence behavior, increases in payments over time should have similar effects.[8]

The units of observation are the largest standard metropolitan statistical areas SMSA's with census data on both white and nonwhite populations). Of the 66 such SMSA's enumerated in the census, 44 are used in this study. Lack of data or reliable sample estimates on the racial breakdown of the caseloads in the smaller SMSA's, especially in the non-South, reduced the feasible sample set. Since suspected racial differences in the response to welfare income play a significant role in public discussion, it seemed advisable to maintain the racial breakdown at the cost of sample size. Cross-section analyses on SMSA's are presented for 1960, with some preliminary results for 1970. In addition to the 1960 and 1970 censuses, data were collected from State welfare agencies on the number of AFDC cases in each SMSA, the racial proportion of the caseload, and the average AFDC payment in the SMSA. The specific eligibility constraints such as residence are established at the State level but are presumed to be effective in each SMSA in the State. The SMSA provides a better estimate of a labor market than the State, and the problem of urban-rural differences in attitudes to public welfare, which may confound estimates of welfare response at the State level, is avoided.

II. The Welfare Model

This section describes the model used to estimate the effect of the size of the AFDC payment on family behavior. The model consists of two functions:

$$FH/F = \alpha WL + \gamma X \qquad (1)$$

$$C/F = \beta WL + \gamma X \qquad (2)$$

where

WL is the average AFDC stipend,

FH is the number of female heads of families with children less than 18 years of age,

C is the number of AFDC cases (families),

F is the number of females aged 14 to 54, excluding childless single women,

X is a vector of standardizing variables, and

γ is the coefficient vector.

The first equation states that the share of females aged 14 to 54 who are female heads of families with children less than 18 years of age is predicted to be a function of the size of the AFDC payment, as well as a function of several other factors (to be discussed below). The second equation states that the same factors are predicted to influence the proportion of females aged 14 to 54 who are AFDC recipients.

Coefficient α in equation (1) is an estimate of the response to the level of welfare income (measured here by the mean AFDC stipend) on the part of women who, in the absence of the welfare alternative, would

[8] There may of course be changes in conditions over time which are not evident in a cross-section analysis. An intercensal study is now in progress to analyze these changes.

be members of families which contained a supporting male member. For want of a better term, these women will be called "welfare-induced" female heads of families, and the coefficient α is an estimate of the size of welfare-induced desertion; that is, of the higher proportion of families headed by females which is directly related to larger AFDC payments.

In equation (2), the dependent variable is the ratio of AFDC cases to the female population, hereafter referred to as the recipient rate. Coefficient β in equation (2) is an estimate of the response of the recipient rate to the level of AFDC payments. If in fact larger AFDC payments lead to larger AFDC caseloads, the object is to break down the caseload response into increased AFDC participation (a) by women who were female heads of families for reasons unrelated to the level of AFDC income, or "welfare-independent" female heads of families, and (b) by women who were female heads as a result of high AFDC income levels, or "welfare-induced" female heads of families. To calculate the welfare-induced component requires adjusting coefficient α. Coefficient α represents the impact of a percent difference in AFDC payments on the percent of women who are female heads of families. On the assumption that these women, whose headship is a response to high AFDC payments, actually become AFDC recipients, it is possible to calculate the component of the higher recipient rate due to welfare-induced female heads. Coefficient α is too small since a 1-percent increase in the share of women who are female heads (FH/F) implies a larger than 1-percent increase in the recipient rate (C/F), since the mean proportion of female heads of families is larger than the mean recipient rate. Adjusting for the different base rates yields the desired figure, the recipient rate response to higher payment levels that is due to larger numbers of welfare-induced female heads.[9]

Subtracting this component from the total recipient rate response (β), leaves the other component, that part of the increase due to welfare-independent female heads.[10]

The dependent variables in equations (1) and (2) are standardized by a set of variables which, like the average AFDC stipend, are likely to affect the decision to apply for welfare assistance. These are: female earnings opportunities in the form of a measure of the female wage and the female unemployment rate; male earnings opportunities; the amount of income held by females from sources other than earnings or public assistance; and the most important requirements limiting eligibility within the population of female heads of families.

[9] This is done by multiplying α by the reciprocal of the mean of the caseload as a share of female heads (C/FH).

[10] Using AFDC caseload as a percent of female heads as the dependent variable (C/FH) would yield misleading results. To the extent that the level of welfare income induces desertion, either actual or "statistical," the coefficient of the welfare payment in a single-function model where the dependent variable is the AFDC caseload as a proportion of female heads of families will be a downward biased estimate of the total response to welfare income. Increases in the level of welfare income in this case will be related to increases in the denominator as well as the numerator of the dependent variable.

These added variables are necessary in order to isolate the effect of the AFDC stipend level from the effects of other factors. The decision to apply for public assistance clearly depends on other factors as well as the income available from the AFDC program. For example, what concerns many potential recipients is the amount of income available from AFDC *relative* to the amount which could be earned in the marketplace. Many potential recipients may be unwilling to sacrifice large amounts of money income even at the gain of increased leisure, in those areas where the AFDC payment is low, relative to potential earnings. The sacrifice is clearly less in areas where the AFDC payment compares favorably with the average female wage. It is, therefore, predicted that the larger the AFDC stipend relative to the anticipated earnings of female heads with children, the more likely consumers will be to opt for AFDC income. Including the mean wage of females by SMSA in the equation makes it possible to look at the increase in the AFDC recipient rate (equation 2), or the proportion of female heads of families with children (equation 1) *related to* increases in the AFDC stipend, holding constant expected market earnings. Putting it somewhat differently, the prediction is that SMSA's with larger AFDC stipends relative to female earnings would have larger AFDC recipient rates and larger proportions of female heads of families. The coefficient of the AFDC stipend is, therefore, predicted to have a positive sign, and the coefficient of the female wage rate a negative sign (holding constant the AFDC payment in the latter case, SMSA's with high female earnings are likely to have relatively small AFDC recipient rates, for example).[11]

[11] The average AFDC stipend is defined as the average payment in the SMSA to families receiving AFDC income. To the extent that the average stipend does not reflect variations in nonassistance income of welfare families, the figure does not fully reflect differences in the attractiveness of AFDC. Although the data were not readily available to test the importance of this bias, a number of factors indicate any such bias is small. Another measure of the relative attractiveness of AFDC is differences in the average total income AFDC families actually attain. In tests using selected State data from 1961 and 1969 surveys, one finds an almost perfect correlation between average total income and average AFDC stipends.

The female wage rate for 1960 is the estimated mean earnings for all females employed 50–52 weeks in the SMSA. This figure was derived from the average earnings of females who worked any number of weeks in several occupational categories and the proportion of females who worked 50–52 weeks in those occupations, weighted by the occupational mix. The median earnings of females who worked 50–52 weeks is provided by the 1970 census. To the extent that those females who make the welfare/employment choice incur child care costs while employed, the relevant wage for their consideration is their market wage net of child care costs. Data are not available on these costs, however, but biases will not result unless there is a systematic negative relationship between costs of child care and the mean market wage (in which case SMSA's with large mean wages would in fact have small net wages, and the observed negative relationship between the mean wage and the AFDC recipient rate will be a downward biased estimate of the true relationship). There is no reason to expect that a negative relationship holds between these two variables; it is more likely that there is a positive relationship, or that there is no relation at all. Furthermore, since many of the females considering welfare as an alternative to market activity are undoubtedly ghetto-dwellers, the extended family patterns which exist among the low-income population are likely to provide low-cost or free child care

Other factors in addition to the AFDC stipend and the average female wage also may influence decisions concerning public assistance. Female earnings, as measured above, represent the average earnings which females in the SMSA could earn if employed 50–52 weeks per year. However, it may be difficult for women to attain that wage on average, if there are few employment opportunities. Therefore, the female unemployment rate in the SMSA is included in the equation as a measure of the difficulty of obtaining employment in the local labor market. The sign of the coefficient of this variable is predicted to be positive. A positive sign indicates that high female unemployment rates increase the likelihood of choosing welfare.

Furthermore, it has been suggested above that a portion of the AFDC caseload may consist of women who became female heads of families for the purpose of receiving AFDC income, those females defined as "welfare-induced" female heads of families. In this case, male earnings as well as female may be a factor in consumer decisions: the larger the male head of household's earnings related to the AFDC stipend, the less incentive he will have to desert his family. The mean earnings of males in the SMSA are therefore included in the equation, and the sign of the coefficient is expected to be negative. Holding the AFDC stipend constant, one would expect that the larger male earnings are, the smaller will be the proportion of female heads of families in the SMSA (equation 1), and the smaller, therefore, will be the AFDC recipient rate (equation 2).[12]

The amount of income held by females other than earnings or AFDC income—alimony, child support, property income—is also likely to affect consumer decisions regarding public assistance. Women with sufficient alimony or property income may be able to refrain from working without the necessity of going on welfare. It is therefore predicted that the coefficient of this variable will be negative: The larger the average amount of this type of income in the SMSA, the smaller should be the AFDC recipient rate or the proportion of female

[12] The earnings of males who worked any number of weeks are included rather than the full-year (50–52 weeks) earnings plus the male unemployment rate, since the male and female unemployment rates are highly correlated and therefore cannot both be included. Since the female rate may influence both groups of females, "welfare-induced" and "welfare-independent" female heads of families, it was included separately rather than the male rate.

heads of families.[13] A negative relationship between this variable and the AFDC recipient rate is predicted for yet another reason. Female heads of families with income of this type above a certain level are ineligible for the AFDC program.

The dependent variables in both equations were also standardized for the more important rules regarding eligibility in the AFDC program within the population of female heads of families with children. These are, for 1960, a residence constraint requiring at least a year's residence in the State (this requirement was voided by the Supreme Court in 1969 and therefore does not apply to the 1970 data); an age maximum on child eligibility of less than 18 years (in 1970, the standard age maximum was 21 years for children still in school or in training programs, and the eligibility constraint applicable for various States is a maximum of less than 21 years); and a stipulation that the mother must work if adequate child care facilities are available (applicable also in 1970). During the 1960's various States initiated a program whereby families with male heads were eligible for the AFDC program if the father was unemployed (AFDC-unemployed parent program). Although the numbers on this newer program have remained small relative to the number of female-headed families receiving AFDC income (see footnote 5), a variable was included in the 1970 equations to account for this program. These eligibility constraints are included in the equations since, if they are enforced, fewer females will be eligible for the AFDC program or will find AFDC attractive in SMSA's where these rules are in effect.

All of the above variables are defined separately by race, with the exception of the eligibility constraints.[14]

III. Estimation Results

This section examines how well the model fits the data for 1960 and 1970. Using the statistical technique of multiple regression analysis allows one to isolate the independent effects of many influences on female headship rates and on AFDC recipient rates.

[13] Male-headed families where the female has substantial amounts of income of this type are less likely to require AFDC income; therefore, the coefficient is predicted to be negative in equation (1) as well. However, there may be a positive correlation between the amounts of this income and the proportion of female heads of families insofar as large average amounts of this type of income may in itself indicate the presence of relatively large numbers of widows or divorced women with alimony or insurance income. This variable may therefore have an insignificant or positive coefficient in equation (1), depending on which influence is the stronger. (The coefficient was negative and significant for the nonwhite population in equation (2) for both 1960 and 1970, and negative but insignificant for the white population. The coefficients were positive and insignificant for both populations in equation (1)). The amount of this income was calculated by subtracting total earnings and total welfare income from the amount of all income received by females in the SMSA, and dividing by the number of females who received any income.

[14] Dummy variables were also included to account for differences both in the intercept between South and non-South (definition of "South" is that of the Bureau of the Census) and in the coefficients of the standardizing variables, since it is possible that both the supply and demand for public assistance may vary by region. Dummies to account for regional differences were included when coefficients or intercepts proved to be significantly different in separate South, non-South regressions. The coefficients of the AFDC stipend were not significantly different between the two regions.

Table 1 presents the logarithmic coefficients of the main independent variables by race from the regressions using 1960 data.[15] The interpretation of these coefficients is straightforward. Consider the coefficient of the AFDC stipend in the first column of table 1, .3854, which represents the effect of the AFDC stipend on the share of nonwhite women who are heads of families with children. This coefficient implies that, holding constant the effect of other differences among SMSA's, a 1-percent-higher AFDC stipend resulted in a .3854 percent higher share of females heading families in the area. As predicted by the welfare model developed above, the coefficients of the AFDC stipend are all positive and statistically significant. In other words, in 1960, the size of the AFDC stipend itself was an important determi-

[15] All regressions were run in double-logarithmic form on the assumption that the earnings distributions are log-normal. In this case, values of the AFDC stipend lying closer to the mean of the earnings distribution involve proportionately larger numbers of wage earners. This is relevant in particular in comparing responses from the white and nonwhite populations since the differential between the mean nonwhite wage and the mean AFDC stipend is smaller than the differential for whites. The coefficients, therefore, indicate elasticities.

The proportions of widows and divorced female heads of families with children less than 18 are also included in both equations as standardizing variables. While the independent variables cited above should account for the major influences on the dependent variable in equation (2), they may not constitute the entire set of factors likely to determine the dependent variable in equation (1). Specifically, they may not account for the proportion of "welfare-independent" female heads of families. Although there are no reasons on theoretical grounds for expecting a correlation between the factors determining "welfare-independent" family dissolution and the AFDC stipend, it is obviously desirable to reduce the possibility of bias in the coefficient of the AFDC stipend. This may be accomplished by standardizing both dependent variables for the proportions of those female heads of families most likely to be welfare-independent. (There may be bias in β as well as α to the extent that "welfare-independent" female heads of families become AFDC recipients.) Among the five subsets of female heads of families with children less than 18, those who are widowed or divorced are less likely to have attained that status for purposes of AFDC eligibility—widows for the obvious reason and divorcees because the costs of divorce are considerable and the status is not required for the AFDC program. The remaining categories are more likely to contain "welfare-induced" female heads of families; separated females since this category contains those reported as deserted; single females since the AFDC program provides an incentive for the single female with children to refrain from marriage; and "others", to the extent that the male's whereabouts is unknown, but he is not listed as having deserted. When each of these subsets is regressed on the AFDC payment and the standardizing variables cited above, the coefficient of the AFDC payment, as predicted, is significant or nearly so for separated, single, and "other" women for both populations. However, the coefficient is also significant for white widows, and is larger than its standard error for black divorced females, both coefficients with a positive sign. These latter coefficients indicate that there exists some positive relationship between the AFDC stipend and these groups in the population which may bias the coefficient of the AFDC stipend. This positive relationship may be explained by at least two hypotheses however: (1) that the factors determining welfare-independent family dissolution happen to be positively correlated with the AFDC payment for some unknown reason; or, (2) that some "welfare-induced" female heads of families report themselves as widowed or divorced to the census (or, for example, that women who became widowed or divorced for reasons unrelated to the AFDC payment may refrain from remarrying for purposes of AFDC eligibility, or may migrate to cities with relatively higher AFDC stipends). If the first hypothesis is correct, standardizing the dependent variables in equations (1) and (2) for the proportions of female heads of families who are widowed or divorced will produce an unbiased coefficient for the AFDC payment. If the latter is correct, the coefficient as an estimate of the response from the total population will be biased downward. When the proportions of widows and divorced females are omitted from the regressions, the coefficients of the AFDC payment are larger, as predicted, but not significantly different from the coefficients in table 1.

nant of both the proportion of females heading families (indicating family dissolution in response to the availability of AFDC income), and the proportion of the female population receiving AFDC income, for both the white and nonwhite populations.

TABLE 1.—*Determinants of the AFDC recipient rate and the proportion of families headed by females, 1960*

[Regression coefficients (N=44)]

	Dependent variables			
	Nonwhite population		White population	
Independent variables	Percent female heads (Equation 1)	Recipient rate (Equation 2)	Percent female heads (Equation 1)	Recipient rate (Equation 2)
AFDC stipend	[1] +0.3854	[1] +1.4241	[1] +0.2641	[1] +1.0881
Female wage	+0.0591	−0.4989	+0.1611	−0.4060
Male wage	[2] −0.8537	[2] −1.8978	[1] −1.6403	[1] −10.7409
Unemployment rate	+0.0700	+0.3116	[1] +0.2773	[1] +0.8761
$\bar{R}^2$	.7644	.7538	.7026	.6317

[1] Significant at .01 level.
[2] Significant at .10 level.

The coefficients of the AFDC stipend in equation (2) indicate the *total* recipient rate response to changes in the level of welfare income, that of both "welfare-independent" and "welfare-induced" female heads of families. Table 2 indicates the relative role of each of these groups. In the case of nonwhites, an AFDC stipend 1 percent higher than average meant a 1.42 percent higher share of women receiving AFDC. Of the total coefficient of 1.42 for the nonwhite population, .80 can be attributed to increases in the number of "welfare-independent" female heads of families in the AFDC caseload, and .62 to increases in the number of "welfare-induced" female heads of families in the caseload.[16]

TABLE 2.—*Composition of disincentive coefficients*

	Coefficients of AFDC recipient rate with respect to AFDC stipend	
	Nonwhite population	White population
Total coefficient	+1.42	[1] +1.9
Contribution of:		
Welfare-induced female heads of families	+.62	+1.53
Welfare-independent female heads of families	+.80	[1] +.37

[1] Estimated. See text.

[16] See the discussion on p. 44 and footnote 9 for the derivation of welfare-induced and welfare-independent effects. The coefficient in equation (1) indicates the percent change in the proportion of female heads of families from a 1-percent change in the AFDC payment. The change in the proportion of female heads is then calculated as a percent change in the AFDC recipient rate. The increase of .38 percent in the share of nonwhite females heading families implies an increase of .60 percent in the share of nonwhite females on AFDC since the mean proportion of nonwhite females heading families is somewhat less than twice the mean proportion of nonwhite females on AFDC.

A similar calculation for the white population requires an additional adjustment since the percentage increase in the AFDC recipient rate due to "welfare-induced" female heads of families (equal to 1.53, calculated from the coefficient in equation (1)), is in itself larger than the coefficient for the total effect (1.09).[17] Assuming that the composition of the total disincentive effect does not vary by race, then the total coefficient and that portion due to increased numbers of "welfare-independent" female heads of families can be estimated for the white population by setting them in the same relationship to the coefficient in equation (1) as the relationship which holds among the three coefficients for the black population. The contribution of nonwhite "welfare-independent" female heads of families to the total disincentive coefficient is 56 percent (0.8 is slightly more than one-half of 1.42). Nonwhite "welfare-independent" female heads of families constitute approximately 12 percent of the total nonwhite female population.[18] Since white female heads of families constitute only about 4 percent of the total white female population, the relative contribution of white "welfare-independent" female heads of families could be expected to be roughly one-third of 56 percent, nearly 20 percent. The contribution of white "welfare-induced" female heads of families would therefore be 80 percent, and the total disincentive coefficient for the white population roughly 1.9.

It should be noted at this point that the coefficients measuring the effect of the AFDC stipend in equation (1) do not differ significantly by race. Although the observed coefficient for the white population in equation (2) is considerably smaller than that of the black population, it is clear that this coefficient is too small relative to the coefficient in equation (1). The estimated coefficient of 1.9 for the white population, admittedly a crude approximation, is not highly different from that of the observed coefficient for the nonwhite population. Regardless, neither the observed coefficient of 1.09 nor the estimated coefficient of 1.9 warrants the conclusion that the total responses to the availability of AFDC income vary greatly by race.[19]

Reading from table 2, a 10 percent higher AFDC stipend, holding all other factors constant, would result in an increase in the AFDC recipient rate of approximately 14 percent for the nonwhite population, slightly less than half of which would be due to desertions of supporting males. The estimated impact of a 10 percent higher AFDC stipend on the white recipient rate is 19 percent, with slightly more than three-quarters due to desertions. The contributions of "welfare-induced" female heads of families, as a proportion of the total disin-

[17] An increase of .26 in the proportion of white female heads of families is equivalent to a 1.53 increase in the white recipient rate since the mean recipient rate (.007) is slightly less than one-sixth the size of the mean proportion of female heads (.041). The small total coefficient (equation 2) for whites may have resulted from downward bias due to the inclusion of the proportion of widows as a separate independent variable and the high degree of correlation between that variable, the AFDC stipend, and the AFDC recipient rate. Changes in variable or model specification, however, do not remove the discrepancy in the sizes of the two coefficients, and it is as yet impossible to determine the source of the bias. The bias does not appear, for example, in the coefficients of the male wage rate. The coefficient of the male wage in equation (2) is slightly over six times the size of the male wage coefficient in equation (1) for the white population, as expected.

[18] This is approximate since the number of "welfare-independent" female heads is not known; 12 percent is the proportion of *all* female heads of families.

[19] This finding is consistent with a recent analysis of work orientations and attitudes regarding welfare income. See Leonard Goodwin, *Do the Poor Want to Work?* Brookings Institution, 1972.

centive coefficient, is considerable. This does not, however, indicate that the proportion of families which dissolve for purposes of AFDC eligibility is greater than the proportion of female heads of families who become AFDC recipients. This result would be surprising in light of the heavy costs involved in family dissolution. Actually, as proportions of their respective populations, *seven times as many "welfare-independent" female heads of families become AFDC recipients as do families which dissolve for AFDC eligibility* in the case of the nonwhite population, for example. The relatively large impact of the AFDC caseload from desertions arises from the larger base population for family dissolution.

There is additional evidence which suggests that high AFDC stipends influenced family splitting. If income from AFDC relative to other income possibilities affects behavior, then the expected earnings of males, as well as those of females, should influence the proportion of female-headed families and consequently the AFDC recipient rate. The results in table 1 confirm these predictions. As suggested in the welfare model, the expected earnings of male exert a negative and statistically significant effect on both dependent variables for both populations. That is, low male earnings relative to AFDC stipends are related to high female headship rates and high AFDC recipient rates.

The coefficients of the AFDC stipend in equation (1) for both racial groups indicate that a *10 percent higher AFDC stipend resulted in a 3- to 4-percent-higher proportion of females heading families*. These effects are statistically significant at the 1-percent level of confidence, which strongly suggest that the findings are not statistical quirks. The data imply, in other words, that independent of other factors, the size of the AFDC payment itself was an important determinant of family dissolution.[20]

[20] In addition to desertions of fathers, high AFDC payments provide an incentive for mothers living with relatives to establish independent households. Although this effect may be operative for once-married women as well, it is likely to be strongest for single mothers. However, if the subgroup of single female heads of families is regressed on the independent variables, the coefficient of the AFDC payment is statistically less significant for both populations than the coefficient for separated female heads. In other words, the response to AFDC income is not very strong for the group of mothers most likely to be living with relatives. This suggests that the coefficient of the AFDC payment in equation (1) is primarily the response of mothers already maintaining separate households, and that the main response to AFDC income is desertion by fathers. Additional evidence is provided by the male wage rate, whose coefficient is negative and statistically significant in equations (1) and (2) (see discussion below). There is reason to believe that the coefficient of the male wage should be less significant for the group of mothers living with relatives because of offsetting influences of the male wage. For example, for mothers who prefer to remain with relatives, high wages of male relatives may enable them to resist the incentives to set up independent households resulting from high AFDC payments. In this case the coefficient of the male wage would be negative, that is, high male wages would be related to lower proportions of female heads of families. On the other hand, high wages of male relatives may enable women who prefer to establish separate households to do so even if low AFDC payments do not cover additional costs. In this case the coefficient of the male wage would be positive. There is evidence that these factors may be operative in the regression of single female heads of families. The coefficient of the male wage, although negative, is statistically insignificant (in fact, less than its standard error) for both the white and nonwhite populations. This should be compared to the strong statistical significance of the coefficient in the regression of separated female heads (easily significant at the 1% level for both populations). Thus, the fact that the male wage is negative and significant in equations (1) and (2) suggests that the major impact of the availability of AFDC income on family stability is desertion of fathers.

The coefficient of the female wage is of the correct sign but is not statistically significant in equation (2) for either population. However, this result is not surprising in light of the sizable portion of the total disincentive coefficient attributable to "welfare-induced" female heads of families, and the high correlation between the two earnings figures.

Female employment opportunities, in the form of the female unemployment rate, were found to have a significant effect on the white AFDC recipient rate but the coefficient was not significant for the nonwhite population.[21]

Table 3 presents preliminary results for the welfare model using 1970 data. With the exception of equation (1) for the white population, all coefficients of the AFDC stipend are statistically significant.[22] The predicted negative and significant relationship between male expected earnings and the dependent variables is also present in the 1970 data.

Unlike the earlier results the female wage is positive and statistically significant in these preliminary results. It is possible, but as yet untested, that this may be due to the omission of allowable earnings retained by recipients in the AFDC program in 1970. In general it is the case that the more affluent States tend to be more liberal in the administration of welfare programs. On the other hand, low payment States often use accounting formulas that result in lower benefit reductions than is the case in high payment States. If high wage areas tend to allow about the same proportion of earnings to be retained, the AFDC program will be more attractive to consumers and, holding the male wage constant, there may be higher proportions of female headed families and AFDC cases. These speculations remain to be tested fully when SMSA data on retained earnings become available. As in the 1960 results, the female unemployment rate is a significant factor for the white population but not for the nonwhite population.

[21] The coefficient for the nonwhite population, however, was larger than its standard error. The significant coefficients for the female unemployment rates in equation (1) for both populations may be reflecting the importance of the male unemployment rate on the decision to desert since the male rate itself is not included as a separate variable. Much of the discussion in previous attempts to estimate the determinants of the size of the welfare population has revolved around the question of whether the level of welfare income or labor market conditions such as the unemployment rate was the stronger and more consistent determinant of the proportion of the population receiving public assistance. (See, for example, Kasper, op. cit.). These studies, however, used statewide unemployment rates, which are not accurate descriptors of conditions in particular labor markets. In this analysis the SMSA unemployment rate is used, and the coefficients are less consistently significant than those of the AFDC stipend.

[22] The lack of a significant relationship between the AFDC stipend and the proportion of white women with children who are heads of families appears to be a southern phenomenon. If south and nonsouth SMSA's are handled separately, the coefficient of the AFDC payment is almost significant at a 10-percent level, with only 12 degrees of freedom, for nonsouthern SMSA's.

Table 3.—Determinants of the AFDC recipient rate and the proportion of families headed by females, 1970

[Regression coefficients (N=44)]

Independent variables	Dependent variables: Nonwhite population: Percent female heads (Equation 1)	Nonwhite population: Recipient rate (Equation 2)	White population: Percent female heads (Equation 1)	White population: Recipient rate (Equation 2)
AFDC stipend	[1] +0.2276	[1] +0.6928	+0.0518	[1] +1.5013
Female wage	[2] +0.7318	[2] +3.5875	[1] +1.3059	[2] +5.1738
Male wage	[1] −0.8223	[1] −2.1787	[1] −1.4439	[2] −4.1553
Unemployment rate	+0.0406	−0.3904	+0.0586	[3] −0.5946
$\bar{R}^2$	.6122	.7551	.8027	.7524

[1] Significant at .01 level.
[2] Significant at .05 level.
[3] Significant at .10 level.

In general, the coefficients of the AFDC stipend are smaller in these latter results than for the 1960 data, with the exception of equation (2) for the white population which is in the range estimated for 1960. Since this latter coefficient is statistically the most significant of the four coefficients in the 1970 results, it is possible that the lower coefficients may be the result of imperfect specification of the 1970 model. There were changes both in the administration of public welfare, and possibly in the public's response to the availability of public welfare in the decade following 1960—the allowance of higher retained earnings, the development of a movement emphasizing welfare as a right, possible changes in the attitudes of middle-class consumers toward the receipt of welfare income, for example—which may be creating statistical "noise" in the estimated coefficients, creating smaller coefficients and in general weaker relationships between the AFDC stipend in particular and the dependent variables. It is also possible of course that consumer response to the availability of AFDC income may have weakened over time. Better specification of the 1970 cross-section model, and more important, an analysis of changes from 1960 to 1970 in the variables specified in the model, should shed some light on these questions. Even the preliminary result, however, suggest that the basic relationships present in 1960 are still in evidence in the more recent data.

INCOME SUPPLEMENTS AND THE AMERICAN FAMILY

By Phillips Cutright *and* John Scanzoni*

Introduction and Conclusions

To assess the future of marriage and the family it is useful to understand the past. The first part of this paper summarizes trends in marriage and divorce rates and changes in the marital status of white and nonwhite Americans. The living arrangements and family status of children over the past several decades also provide clues to changes in American families.

Because much of the concern with the current and future status of the American family is generated by the dependent female family head with children, we measure the impact of different factors that explain the increase in the number of female family heads with children over the past three decades.

Direct analysis of the impact of the level of AFDC benefits on several measures of marriage and the family is possible when the variation among the states in AFDC benefits is used as a quasi-experimental design. If, for example, the level of AFDC benefits is a powerful stimulus affecting entry to marriage, "family splitting", or marital instability, one would expect to find that the marital status of women or the family status of children would vary among the States with high or low benefits. Presumably, States paying large benefits would have more never married women, more children living in families other than husband-wife families, and high marital instability than would States paying low benefits.

Comparison of the marital status of women among States with higher or lower AFDC benefits allows a test, albeit gross, for possible effects of the AFDC benefit levels on rates of entry to marriage by younger women. If high benefits deter marriage, we should be able to detect such effects. If the level of past AFDC benefits for female family heads with children is not related to the family status of children or the marital status of women, then the record of the past may be used, albeit cautiously, as a likely guide to predict the effect of income supplements on marriage and families in the future.

We compare the commitment to marriage and the family of adult white and nonwhite men, and examine the marital status of men as a function of their economic status. This allows a judgment of the strength of commitment to marriage and the family among white and black men, as well as providing clues to the likely effect of income supplements in undermining the institutions of legal marriage and legitimate childbearing.

Differences in the percentage of white and nonwhite ever-married women in disrupted marital status are viewed in terms of a

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general model of marital instability. The effect of economic factors and racial discrimination in producing the observed difference among white and black women is discussed. The probable impact of income supplements on each major cause of marital instability allows an informed judgment of the effect of future supplements on the American family.

Conclusions

1. The marital status of the adult population has shown no steady trend since 1890 in either the white or black population. Marriage and divorce rates fluctuate with economic and demographic conditions. The recent increase in the age at marriage among younger women is a response to current economic conditions that delay marriages.

2. Stability of marriages, as measured by the percentage of ever-married women 15–44 who are separated, widowed, or divorced has scarcely changed in either the white or black population since 1940.

3. Between 1940 and 1970 the number of ever-married white mothers aged 15–44 heading a family increased from 552,000 to 1,609,000—a rise of 1,057,000. Among comparable nonwhite mothers the increase numbered 509,000 female family heads.

4. The percentage of mothers aged 15–44 in disrupted marriages (separated, widowed, or divorced) who were family heads increased from 43 to 79 percent between 1940 and 1970. This change in the propensity to form separate households (rather than living as a child or other relative of the family head) is similar to shifts toward separate housing and living arrangements among younger married couples, aged couples, and widowed persons.

5. Four causes of the increase in the number of white and nonwhite families headed by an ever-married woman aged 15–44 were examined: (1) increasing numbers of ever-married women—or population increase; (2) declining marital stability; (3) declining childlessness among women in disrupted marital statuses; and (4) the increasing propensity to live in separate households.

6. The change in propensity to live as a separate household accounts for about 38 percent of the total increase of 1,566,000 ever-married female family heads aged 15–44. About 18 percent of the increase is due to the increase in the number of ever-married women, independent of other changes, while 71 percent of the total increase is due to both these factors in combination. Seven percent of the total increase is related to declining childlessness alone. Ninety percent of the total increase was related to the combined effects of changes in propensity, population increase, and declining childlessness, leaving a maximum of 10 percent that is related to declining marital stability alone and in combination with the first three factors. There was very little difference between the white and nonwhite population in the contribution that the four components make to the increase in numbers of female family heads.

7. When the increase in female family heads aged 15–44 from never-married mothers was included, the total 1940–70 increase moved up by 183,000. Of the new total increase of 1,749,000 female family heads, a maximum of 6 percent was related to rising illegitimacy.

Of the increase of 1,749,000 female family heads in all marital statuses, 36 percent was related to changing propensity to form separate households alone. Population increase and the change in propensity combined accounted for about 68 percent of the total increase. Less than 13 percent of the total increase was due to declining childlessness among the ever-married, or increasing illegitimacy among the never-married, independent of other changes. About 3 percent of the total increase was due to declining marital stability alone.

8. Illegitimate childbearing, bridal pregnancy and other measures of fertility do not account for the large difference in the proportion of white and nonwhite women who are female family heads. The main reason a higher proportion of nonwhite than white women with children are heads of a family is that a higher proportion of nonwhite women are in a disrupted marital status. A higher rate of breakup of first marriages among nonwhites and a longer period between separation, divorce, and remarriage for those whose first marriage fails, accounts for the major portion of the racial difference in the percent of women in a disrupted marital status.

9. If AFDC benefit levels are a cause of the increase in the number of female family heads, the program could have such an effect only through its effect on one or another of the four components of change discussed under conclusion 5, above. The program cannot affect the portion of change attributed to population growth; nor is it likely that the program has affected that component that is related to changing fertility patterns. Analysis of benefit levels and the marital status of women in 1950, 1960, and 1970 did not discover a relationship of the program data to age at marriage, stability of marriages, or the proportion of women married in either the white or nonwhite population. A possible AFDC program effect through the marital status or marital stability factor is, therefore, rejected. This process of elimination leaves only the propensity factor as a possible route through which the program could have increased the number of female-headed families.

10. If the AFDC program has powerfully contributed to the rise of the female-headed family this effect should be reflected by a relationship between the level of AFDC benefits and the family status of children among the States. Analysis of State AFDC benefit levels in 1960 and 1970 found no negative effect of program benefit evels on two measures of the family status of white children. In fact, white children were somewhat more likely to live in husband-wife families in States paying high rather than low benefits.

11. A similar analysis of State AFDC benefit levels and the family status of nonwhite and black children found no evidence indicating that the AFDC program encourages the formation of female-headed families. In fact, urban black children in high benefit States were somewhat more likely to live in husband-wife families in 1970 than were urban black children in low benefit States.

12. By 1970 the AFDC program had been operating for over 30 years. After three decades of program activity we found no impact of high or low program benefits on the family status of children, the formation of female-headed families, age of marriage, marital status, or the stability of marriages in either the white or nonwhite populations. A separate paper in this volume found no AFDC effects on illegitimacy rates between 1940 and 1970.

13. The increasing propensity to form separate households by mothers at risk of female headship is a trend shared by other Americans. A similar trend was observed among older persons and among

young husband-wife families. The decline in the percentage of aged persons or young intact families who live with relatives is not a function of changes in public assistance benefits. In the case of young couples, the increased propensity to live in separate housing cannot be linked to any type of income support program. Groups that do and those that do not have access to public assistance or social insurance have moved in the same direction.

14. There is substantial evidence that blacks and whites do not differ in their commitment to the institution of legal marriage. Within both populations the rate of entry to first marriages and to remarriage is similar when men with similar incomes are compared.

15. First marriages are more likely to be disrupted in the black than white population. About one third of the racial difference in stability of first marriages is related to differences in the distribution of income (measured in a single year) between the two populations of men.

16. When remarriages are included, 96 percent of white and 87 percent of black ever-married men aged 35–44 in 1960 were currently married and living with their wives. If nonwhite men had incomes similar to those of whites this 9-percent difference would probably reduce to 5 or 6 percent. Since many variables other than male income in the year prior to census are not accounted for, we judge the expected difference of 5 or 6 percent to be trivial, and conclude that the marriage behavior of nonwhite males demonstrates a commitment to marriage and the family similar to that shown by white men.

17. Racial differences in the proportion of women in a disrupted marriage are the result of different levels of constraints against marital instability and different levels of marital satisfaction between the two populations. High constraints against hasty marital dissolution were related to high levels of income as well as to social factors. Low incomes result in a lower level of constraint. Marital satisfaction and social constraints are the main factors that differentiate couples that do from those that do not remain together. Satisfaction with marriage is a net result of the balance of positive and negative inputs to the marriage. Racial discrimination in labor and housing markets lowers the level of constraints related to income among nonwhites, and also depresses the balance of positive over negative inputs to nonwhite marriages. Therefore, nonwhite couples have a lower level of economic constraints and are also less satisfied with their marriages than white couples. Therefore, marital instability among nonwhites is higher. Higher rates of marital disruption result in a higher proportion of nonwhite than white mothers living in a disrupted marital status and being the head of a family. The racial difference in the risk of female headship is not a function of differences between the races in the value placed on legal marriage and husband-wife families.

18. A program that would provide husband-wife as well as broken low-income families with an income supplement might increase marital instability or depress rates of entry to legal marriage if the economic benefits of family splitting outweighed the social and other benefits of legal marriage. There is no evidence that the past supplements provided primarily to broken families by AFDC stimulated marital dissolution, or caused low income persons to avoid marriage. If the past AFDC program has not produced a clear-cut effect on marital behaviors after three decades, it seems unreasonable to believe that a future program that would not directly penalize intact families would stimulate family splitting.

I. Trends in Marriage, Divorce, and Marital Status

It is important to have some idea of the past if we are to have a perspective on future behavior. Restricting an overview of American marriage patterns to recent data would unnecessarily limit the perspective one should have on very recent changes in the population. If, for example, the recent increase in the percent never married were a continuation of a long-run increase, then the significance of this change would be heightened. On the other hand, if the recent change is not an extension of a long-run trend, then it may merely represent an accommodation to recent socioeconomic events, rather than being the wave of the future.

Trends From Vital Statistics

Because of inadequacies in marriage and divorce registration in the United States, color-specific rates from vital statistics are not available. The trend in national rates reflects the white trend, due to the numerical superiority of the white over the nonwhite population.

Vital statistics on marriage rates per 1,000 unmarried women aged 15 and older are available from 1920. (Carter and Glick 1970: table 3-2 and NCHS, 1971: table 1). The rate in 1920 was 92 per 1,000. By 1930 the rate had declined to 67.6, and it remained in the low seventies during that decade. The rate moved up to about 83 in 1940, peaked at 118 in 1946, receded to about 80 by 1954, and declined to the low seventies during the 1958–63 period. The marriage rate has increased each year since 1963, and was about 82 in 1971.

As with marriages, the national divorce rate reflects most accurately the trend among whites. The divorce rate per 1,000 married women aged 15 and older was 8 in 1920. It declined to a low of 6.1 in 1932 and 1933, and then moved upward and peaked at 17.9 in 1946. The rate declined to 8.9 in 1958 and then increased reaching 13.4 in 1968. The divorce rate for the 12 months ending in January 1972 was about 16—not far below the post-World War II record high. (Rates prior to 1968 from Carter and Glick 1970: table 3–9; 1968 from *Statistical Abstract of the United States, 1972*, table 79; 1972 rate estimated by the authors.)

The rise in the divorce rate after 1958 has been accompanied by an increase in the number of children *per divorce decree*—from 92 children per 100 divorces in 1955 to 134 children per 100 divorces in 1968. (*Statistical Abstract of the United States*: Ibid.). The number of children involved in divorce cases rose from 347,000 in 1955 to 782,000 in 1968. With no further increase in the number of children per decree, the number of divorces in 1971 (773,000) would imply that 1,035,820 children were involved.

The general conclusion from this brief review is that both marriage and divorce rates have varied in response to economic and other conditions in the past. One should expect fluctuations in the future. A change from one year to the next does not herald the beginning of a new era. This point can be documented with Census Bureau materials which also allow analysis of long run trends in marital status, as well as recent changes by age, sex, and color.

Trends in the Percent of Adults Who Are Single

The 1890 to 1969 trend in the percent of men and women aged 15 and older who were single has been reported by Farley and Hermalin

(1971: table 1). Among nonwhite men, the percent single (26 percent) was the same in 1969 as it had been in 1890. Among white males, the percent single was stable from 1890 through 1940, and it then declined to a low in 1969. We conclude that the data show no shift away from entry to legal marriages in either male population. Men appear as willing now as they have been in the past to enter a legal marriage.

Among women the percent single among whites aged 15 and older follows a pattern of stability from 1890 through 1940, with declines in the percent single from 1940–1960, and no change since 1960. Among nonwhite women there was an increase of 1 percent single in 1969 over 1900—hardly evidence of a significant trend away from entry to marriage.

Percent Single by Color and Age, 1960–71

The percent never married among the total population of persons 15 and older will not reflect recent changes among younger age groups. Thus, while the 1960 and 1969 trend for persons aged 15 and older shows little change, the trend toward later age at marriage among young persons does indicate substantial change. The percent single among women aged 14–34 in 1960 and 1971 is shown in table 1.

TABLE 1.—*Percentage of women aged 14–34 never married, by color: United States, 1960 and 1971*

Age of women	Color					
	White			Nonwhite		
	1960	1971	Change	1960	1971	Change
14–17	94.6	97.3	2.7	93.9	97.4	3.5
18–19	67.7	76.8	9.1	69.1	81.3	12.2
20–24	27.4	35.1	7.7	35.4	47.7	12.3
25–29	9.8	10.3	0.5	15.7	24.4	8.7
30–34	6.6	5.9	−0.7	9.6	14.2	4.6

Source: U.S. Bureau of the Census (1963). *PC (1)–1D:* Table 176: U.S. Bureau of the Census, *Current Population Reports, Series P–20, No. 225:* table 1.

For whites aged 18–24, notable increase in the percent single occurred. Among whites aged 25–34, little change can be found. Among nonwhite women, the 1960–71 increase in the percent single is larger than it has been for whites at the same age. In 1971 a higher proportion of nonwhite than white women were single at all ages. The difference in the percentage single between white and nonwhite women at the same age is understandable, given the different level of income in the respective male populations (table 17, below).

Recent changes in the percent single are related to economic and demographic changes that affect the propensity of men and women to marry—changes that were, in part, responsible for the temporary decline in marriage rates during the depression. Since the economy is not now at depression levels, one might claim that economic factors could not be responsible for the recent change in marital status. However, careful analysis by Easterlin (1972) had documented the relationship between changing economic conditions among younger persons and the change in marital status. Also, one of the delayed demographic effects of the postwar baby boom is the recent "marriage squeeze" (Akers, 1967) which causes a deficiency of numbers of men of marriageable ages relative to women in the prime years for first marriage. Demographic and economic changes explain the recent trend toward later age of first marriage.

A shift toward later marriage has occurred before in response to economic events; the current shift is a response to recent changes that affect young persons, and is no evidence of rejection of marriage by them. The shift toward later marriage should have positive effects. Fertility is reduced by later marriage, and the high rates of marital instability among those who marry young will now affect a smaller proportion of the population than in the past, a change that may help to stabilize the rising divorce rate.

The Percentage of Ever-Married Women in a Disrupted Marital Status: 1940–70

A major portion of this report is concerned with the 1940–70 change in the number of female family heads aged 15–44, a trend that is related to the changing number of ever-married women in disrupted marital status categories. Table 2 shows the number ever married and the percent of ever-married women aged 15–44 in a disrupted marital status (widowed, divorced or husband absent).

TABLE 2.—*Number of women 15–44 ever married, and percentage whose spouse is not present, by color, United States, 1940–70*

[Numbers in thousands]

Year	Number ever married		Percent whose spouse is not present	
	White	Nonwhite	White	Nonwhite
1970	25, 775	3, 376	9. 9	28. 0
1960	24, 166	3, 068	9. 0	28. 2
1950	22, 768	2, 852	9. 5	26. 5
1940	18, 405	2, 424	9. 3	25. 3

NOTE.—The numbers of women whose spouse is not present are shown in table 6, below.

Source: Marital status from *U.S. Census of Population, 1940, 1950 and 1960*, and *Current Population Reports*, Series P-20, No. 212, table 1.

About 9.3 percent of white ever-married women aged 15–44 were in a disrupted marital status in 1940, compared to 9.9 percent in 1970. Among comparable nonwhites, the percentage increased from 25.3 to 28 percent over the 30-year period. This measure of the impact of changing marital stability on the current marital status of women shows little change.

II. TRENDS IN LIVING ARRANGEMENTS OF ADULTS AND CHILDREN

Trends Among Adult Women Aged 25–64

Living arrangements of adults have changed over the past several decades. The causes of changes in the distribution of the population among various household status categories have not yet been explained. We can, nonetheless, document the change, and then examine its impact on increases in the number of female heads of families.

Table 3 shows the distribution of the adult female population by household status in 1950 and 1969. Among nonwhite women there was an increase of 2 percent living as the wife of the head of a family between 1950 and 1969; over the same period, there was an increase of 9 percent in the proportion heading their own family. These two changes occur, in part, because of the large decline in the percentage of women living in families as a child or other relative of the family head. Fewer nonwhite women are now "doubled up" in another family. They are more likely than in the past to be the wife of a family head or to be the family head themselves.

TABLE 3.—*Trends in household status of adult women aged 25-64: United States 1950-69*

[In percent]

Household status	Color: Nonwhite 1950	Nonwhite 1969	Nonwhite 1950-1969 change	White 1950	White 1969	White 1950-1969 change
Wives of head	55	57	2	74	79	5
Head of family	13	22	9	6	7	1
Child or relative of family head	16	9	−7	12	5	−7
Primary or secondary individual	12	11	−1	6	8	2
Not in household	4	1	−3	2	1	−1
Total	100	100		100	100	

Source: Derived from Farley and Hermalin 1971: Table 4. Age standardized on 1960 total population 25-64.

Among whites, the percentage of women who were the wives of the family head increased by 5 percent between 1950 and 1969, while the proportion who were female heads of families increased by 1 percent. The 7-percent decline in the proportion living as a child or other relative of the family head generally was absorbed by an increase in the percentage who were, by 1969, the wife of the family head.

Other Measures of Trends in Household Status of Adults: 1940-70

One can examine change in living arrangements for different types of adults. In table 4, we have displayed the 1940-65 or 1970 trend toward separate housing arrangements for a variety of types of people. The first column shows the percentage of mothers aged 15-44 in a disrupted marital status who are family heads. (Not all of these female family heads are living in separate housing. In 1960, about 23 percent of all female heads aged 15-44 compared to 14.6 percent in 1970 were in subfamilies—see table 5, below. As the percent in subfamilies declines, one may assume the percentage of all female-headed families in separate housing is increasing.) The trend from 1940 through 1970 shows a clear upward swing in the probability that mothers at risk of heading a family will be doing so, and this change is a major cause of the increasing numbers of female-headed families.

TABLE 4.—*Percentage of ever-married mothers in disrupted marriages, currently married couples and widowed persons living as heads of families or in separate households: United States, 1940–70*

Year	Mothers aged 15–44 in disrupted marital statuses	Couples, wife aged 24 or less	Couples, all ages: White	Couples, all ages: Nonwhite	Couples, wife aged 65 or over	Widowed women	Widowed men
1940	43	81	94	89	93	19	21
1950	39	83	na	na	93	27	28
1960	69	92	98	95	97	36	38
1965	na	94	98	96	98	na	na
1970	79	na	na	na	na	50	47
Effective change ratio	63	68	67	64	71	38	33

Source: Mothers aged 15–44 from table 7, below. Widowed persons from Chevan and Korson, 1972: Table 1; Couples from Beresford and Rivlin, 1966, table 2 and Carter and Glick, 1970, table 6.2. Missing data indicated by "na."

Mothers at risk of being a family head are not the only group of Americans who are shifting from joint family housing arrangements to separate housing. In table 4 we find shifts toward separate housing among widowed men and women, among couples aged 65 and older, among couples with the wife aged 24 or younger, and among married couples of all ages. An appropriate measure with which one may compare the 1940-70 change in separate family status among these disparate types of people must consider the 1940 level, in addition to the absolute percentage change between 1940 and 1970. Taking the 1940-70 percent change as the numerator, and the difference between 100 percent and the 1940 percentage figure as the denominator, we have an "effective change ratio." This ratio measures the extent to which the observed change equals or is less than the possible maximum change.

The bottom row of table 4 compares this ratio for different types of persons or couples. Among women at risk of female family headship, the ratio is 63, a figure slightly lower than that for all other groups, except widowed persons. Although the absolute change in percentage points among widowed women (31 percent) is close to the percent increase among women 15-44 at risk of headship (36 percent), the lower percentage of the widowed in separate living quarters in 1940 results in a lower effective change ratio. Table 4 demonstrates that the 1940-70 period was one of decline in "doubling up" living arrangements among a wide variety of Americans. Individuals and couples, old and young, those with and those without public assistance or social insurance shared a common move toward separate housing and separate household status.

Trends in the Living Arrangements of Children: 1910-68

Since 1910 the percentage of children under age 5 who are living with their mothers (with or without the father present) has not declined. The rate in 1910 was 97 percent for white and 87 percent for nonwhite children. In 1960 the rate was 99 percent for white and 86 percent for nonwhite children under age 5 (Farley and Hermalin, 1971: p. 13). When we ask what percentage of children under age 6 live with *both* parents a trend emerges during the 1960's. Farley and Hermalin (1971: table 8) report that in 1960 about 69 percent of nonwhite children under 6 lived with *both* parents, 19 percent were with the mother alone, 1 percent with the father alone, and 11 percent with neither parent. By 1968 only 61 percent were living with both parents, 25 percent with the mother, 1 percent with the father, and 13 percent with neither parent. The decline in the percentage living with both parents is closely related to the increase in those living with the mother alone.

Among whites, 93 percent of children under age 6 lived with both parents in 1960, 5 percent with the mother alone, 1 percent with the father alone, and 1 percent with neither parent. In 1963, the percent living with the mother alone was up one point to 6 percent; the percent living with neither parent was down one point, and the percent living with both parents or with the father alone was unchanged.

III. Analysis of Change in the Number of Female Family Heads: 1940–70

Analysis of the trend in the number of ever-married female family heads proceeds in two stages. We first comment on different measures of trends in the risk of becoming a female family head. Having settled on the measure we then examine the causes of change in the number of ever-married women aged 15–44 who are heading a family. Following this analysis we include single women, and then examine one common explanation of the increase in the number of female family heads—the alleged impact of the AFDC program.

Changes in the Numbers of Families Headed by an Ever-Married Woman Aged 15–44

Table 5 shows the number of families headed by an ever-married woman aged 15–44 with a child at the time of the census. The number excluding subfamilies in 1960 and 1970 is also shown. The number of female-headed families in 1940 and 1950 is thought to use definitions that make the 1960 and 1970 count which included subfamilies appropriate to trend analysis.

Table 5.—*Number of ever-married female family heads aged 15–44 with children, 1940–70, by color*

[Numbers in thousands]

Year	Color	
	White	Nonwhite
1970—all	1,609	681
1970—excluding sub-families	(1,369)	(586)
1960—all	1,210	451
1960—excluding sub-families	(923)	(351)
1950—all	618	217
1940—all	552	172
1940 to 1970 change	1,057	509

Source: 16th Census of the United States 1940: *Population, Families, Types of Families*, tables 3 and 7; Census of Population 1950, *P-E No. 2A*, tables 4, 5; Census of Population 1960; *Families, PC (2)-A-4*: Tables 5 and 21; *Current Population Reports*, series P-20: 218: Table 4. Never-married family heads are excluded.

Between 1940 and 1970 the number of white families headed by an ever-married woman aged 15–44 increased by 1,057,000 while the increase among nonwhites was 509,000. Neither increase suggests that American families are less stable than they were in the past. Changes in marital stability are, in fact, just one of several causes of the increase in the number of families headed by a woman.

Declines in the stability of marriages which increase the number of ever-married women in disrupted marital statuses, is one cause behind the increased numbers of female family heads. A second cause is found in population increase. Third, the number of female family heads may increase if the proportion of ever-married women in disrupted marital statuses who are childless declines. Finally, the number of female heads will increase if the population of women at risk of headship (those with a child in disrupted marital statuses) changes its propensity to form a separate family.

Changes in the Percent of Disrupted Marriages

Our analysis of the effect of changing marital stability is taken from table 2, above. Only small changes were found between 1940 and 1970 in the proportion of ever-married women 15–44 in disrupted marital statuses.

Change Due to Population Increase

With other factors under control the effect of population increase on the increase in the number of women aged 15–44 who are head of a family is taken from the counts of ever-married women aged 15–44—see table 2, above.

Change in the Propensity To Be Head of a Family

In table 4 we noted an increase in the propensity to head a family among those women at risk of female headship. This change is shown by color in table 7 below.

Change in Fertility Patterns Affecting the Risk of Female Headship

Trends in fertility affect the probability that a woman of a given age will become a female family head with own child. Changes in the percentage of women in an age group that is childless may increase the number of female heads with children.

Table 6 shows the trend in childlessness among women at risk of becoming a female family head. The trend and level of childlessness is similar among both whites and nonwhites, although the causes of the shared decline are different. The high levels of childlessness among nonwhites in the earlier years were, primarily, the result of disease and other health conditions that increased sterility and fetal loss. As health conditions improved, childlessness declined (Cutright, 1972; Farley, 1970).

Among whites a similar involuntary set of conditions partially explains the high early level of childlessness, but increasing bridal pregnancy and a decline in the use of birth control early in marriage among whites after 1940 is probably more important (Cutright, 1972: table 9). In sum, the decline in childlessness among nonwhites is largely due to changes in involuntary controls over fertility, while the decline among whites is primarily due to a shift away from voluntary controls over fertility shortly before and shortly after marriage. This shift occurred among all strata of the white population.

TABLE 6.—*Number of women aged 15–44 in disrupted marital status categories including and excluding childless women, by color: United States, 1940–70*

[Numbers in thousands]

Year	Number ever-married in disrupted marriage categories		Number who have borne children		Percent childless	
	White	Nonwhite	White	Nonwhite	White	Nonwhite
1970	2,551	945	2,096	803	17.8	18.3
1960	2,170	866	1,718	691	20.8	20.2
1950	2,157	756	1,610	508	25.4	32.8
1940	1,719	614	1,237	428	28.0	30.3
1940 to 1970 change	832	331	859	375	−10.2	−12.0

Source: Table 2 Childlessness from U.S. Bureau of the Census, "*Differential Fertility 1940 and 1910: Women by Number of Children Ever-born*" 1950 from "*Fertility*," vol. IV Pt. 5, chapter C, tables 16 and 17; 1960 from "*Women by Number of Children Ever-born*," PC (2)-3A, tables 16 and 17; 1970 from *Current Population Reports*, series P-20, No. 218, partially estimated.

Changes in the Risk of Being a Female Family Head With Children: 1940–70

Table 7 shows the change in the risk of being a female family head aged 15–44 with children using three different denominators for the population at risk. Never-married women are excluded from all numerators and denominators. Table 7 indicates an increase in the risk of being a female family head for both white and nonwhite women after 1940, regardless of the denominator chosen. With the exception of the slight decline in the risk among whites between 1940 and 1950 (see note to Table 7) all these measures show an increasing risk of female headship between 1940 and 1970. The major change in the risk occurred between 1950 and 1960—with an added but smaller increase between 1960 and 1970. For nonwhites the increase in risk was large and nearly equal from 1950 to 1960 and from 1960 to 1970.

TABLE 7.—*Measures of risk of female family headship for women aged 15–44: 1940–70*

[In percent]

Year	Percent ever married aged 15–44 who are female heads of family with own child		Percent of women in disrupted marital status that are female family heads with own child		Percent of women in disrupted marital status who have ever born children that are female family heads with own child	
	White	Nonwhite	White	Nonwhite	White	Nonwhite
1970	6.2	20.2	63.1	72.0	76.8	84.8
1960	6.3	14.7	55.7	52.1	70.4	65.3
1950	2.7	7.6	28.6	28.8	38.4	42.8
1940	3.0	7.1	31.2	28.0	44.6	40.2

NOTE.—1960 and 1970 data include female subfamily heads. 1950 data include both secondary and primary families. 1950, 1960, and 1970 data use own children while 1940 includes related as well as own children. The difference in definition between 1940 and 1950 may account for the slight decline in risk for whites and moderate the increase in risk among nonwhites. The effect of this change in definition in the subsequent analysis is trivial.

The large differences in risk of female family headship between white and nonwhite women using all ever-married women generally disappears when the more appropriate population at risk is chosen. Thus, the major explanation in any year of the higher percent of nonwhite than white ever-married women being female family heads is that a higher proportion of nonwhite than white women are in disrupted marital statuses. As noted in Table 6 there is little difference in childlessness levels of white and nonwhite women in disrupted statuses in the same year. Therefore, the larger percentage of ever-married nonwhite women heading their own families shown in columns 1 and 2 of table 7 is caused, for the most part, by the higher proportion of nonwhites than whites in disrupted marital statuses.

Among whites the trend in the risk of female family headship in the population at risk (defined as women in disrupted marital statuses with children ever-born) moves from about 45 percent in 1940 to 77 percent in 1970; among nonwhites the percentage increases from 40 to about 85 percent. We refer to this trend as one of changing *propensity* to form *separate* families in which the woman is head. The effect of changing propensity on the increase in the number of female family heads, independent of the increase in the population at risk, is considered in the following section.

Components of Change in the Number of Female Family Heads, 1940–70

In table 8 we display the absolute and relative impact of changing propensities to form separate households, population increase, declining childlessness and marital stability on the 1940 through 1970 increase in the number of families headed by an ever-married mother aged 15–44.

TABLE 8.—*Components of 1940–70 change in the number of ever-married female family heads aged 15–44 with children, by color*

[Numbers in thousands]

Components of 1940–70 increase in the number of female family heads	Color: White Number	Color: White Percent	Color: Nonwhite Number	Color: Nonwhite Percent	Total Number	Total Percent
(1) Higher propensity alone	398	37.6	191	37.5	589	37.6
(2) Population increase alone	221	20.9	67	13.2	288	18.4
(3) Components 1 and 2 combined	779	73.7	333	65.4	1,112	71.0
(4) Declining childlessness alone	78	7.4	38	7.5	116	7.4
(5) Components 1, 2, and 4 combined	966	91.4	444	87.2	1,410	90.0
(6) Marital stability alone	33	3.1	18	3.5	51	3.3
Total increase	1,057	100.0	509	100.0	1,566	100.0

NOTE.—Formula for computing the numbers in rows 1 through 6: (1) $\left(\frac{H_2}{M_2}-\frac{H_1}{M_1}\right)M_1$; (2) $\frac{H_1}{P_1}\left(P_2-P_1\right)$; (3) $\left(\frac{H_2}{M_2}\cdot\frac{M_1}{D_1}\cdot\frac{D_1}{P_1}\cdot P_2\right)-H_1$; (4) $\frac{H_1}{M_1}\left(\frac{M_2}{D_2}-\frac{M_1}{D_1}\right)D_1$; (5) $\left(\frac{H_2}{M_2}\cdot\frac{M_2}{D_2}\cdot\frac{D_1}{P_1}\cdot P_2\right)-H_1$; (6) $\frac{H_1}{D_1}\left(\frac{D_2}{P_2}-\frac{D_1}{P_1}\right)P_1$; where subscripts 1 and 2 refer to 1940 and 1970, respectively and H=number of female family heads with a child in disrupted marital status; M=number of mothers in disrupted marital status; D=number of ever-married women in a disrupted marital status, and P=number of ever-married women.

Source: Tables 2, 5, and 6, above.

Rows 1, 2, 4, and 6 show the increase in numbers expected from the observed change in each component, had there been *no* change in any of the remaining three components. Rows 3 and 5 show the contribution of various components in combination with other components to the increase.

Among whites, for example, the change in propensity alone accounts for 398,000 of the total 1,057,000 increase—or about 38 percent. The increase in the number of ever-married women alone would, had no other components changed, resulted in a gain of 221,000 female family heads. The combination effect of change in both propensity and population is larger than the sum of the change from each component considered in isolation—row 3 shows about 74 percent of the total increase is a result of joint changes in propensities and population increase. By itself alone, the decline in childlessness is related to an increase of 78,000 heads; when considered jointly with the first two components 91 percent of the total increase is attributed to factors other than changing stability of marriages. The contribution of the decline in marital stability alone is about 3 percent of the total increase.

Among nonwhites the center columns in table 8 indicate results quite similar to those detailed above for whites. The combination effect of changing propensities and population increase accounts for 65 percent of the increase in numbers of nonwhite female family heads, while the addition of declining childlessness to the first two components accounts for 87 percent of the gain. The decline in nonwhite marital stability alone is related to 3.5 percent of the increase.

The right hand columns of table 8 add the numbers resulting from separate analysis of white and nonwhite change. We see that 71 percent of the increase in all ever-married female family heads aged 15–44 is related to changing propensities to form separate households and the increase in the number of ever-married women. When the decline in childlessness is added to the first two components we find that 90 percent of the increase can be allocated to the first three components. The role of marital stability is relatively small—it accounts for just 3 percent of the increase when considered alone.[1]

Adding in Never-Married Female Family Heads

The census of 1940 reported 28,000 white and 18,000 nonwhite never-married mothers aged 15–44 as female heads. In 1970 the

[1] The size of the components in table 8 will change if 1970 rather than 1940 measures are substituted at various points. For example, among nonwhites formula 6 results is an estimate of 18,000 added heads due to change in marital status alone. One might ask whether this component would change had 1970 rather than 1940 data been used. The first term in formula 6 gives the percentage of ever-married women with children in disrupted marital statuses that were heads. In 1940 the figure was 28 percent while in 1970 it was 72 percent. Substituting the 1970 for the 1940 statistic will increase the marital status component to 46.3 thousand. This increase comes, in part, from reducing the actual effect of rising propensities on the increase in numbers of heads and then allocating this change to marital status. Substantive interpretation of results obtained after substituting 1970 for 1940 inputs is unclear.

respective numbers were 63,000 and 166,000. (See table 5, above, for sources). Among whites the total increase was 35,000; among nonwhites it was 148,000. The upper panel of table 9 breaks down the total increase for never-married female family heads into the three relevant components behind the increase. (Changing marital status is not relevant.) Because no direct measure of changing propensity is available for the never-married, we assume that the change in propensity observed for the ever-married is equal to the change for never-married mothers. This assumption can tolerate considerable error without serious consequences on the following analysis. The data for the population increase component are the counts of never-married women 15–44 in 1940 and 1970. Allocation of change to these two components considered separately and then in combination is identical to that detailed above in table 8. Allocation of change to declining childlessness is a residual term—the difference between the change in the number of never-married heads allocated to changing propensities and population increase in combination, and the total increase. We are unable to calculate the contribution of rising illegitimacy (which is analogous to declining childlessness among the ever-married) because the necessary data are not available. However, the method selected to represent the contribution of rising illegitimacy is constrained by the previously deducted contribution from the first two components. Finally, the relatively less precise calculations used for the never-married can have little impact on our final assessment of the components of change, since the contribution of rising numbers of never-married heads to the increase from all marital statuses is just 10 percent.

TABLE 9.—*Components of 1940–70 change in the number of never-married female family heads aged 15–44 with children and female heads with children in all marital statuses, by color*

[Numbers in thousands]

Components of 1940–70 increase in the number of female family heads	White		Nonwhite		Total	
	Number	Percent	Number	Percent	Number	Percent
NEVER-MARRIED FAMILY HEADS						
(1) Higher propensity alone	20	57.1	20	13.5	40	21.9
(2) Population increase alone	2	5.7	15	10.1	17	9.9
(3) Components 1 and 2 combined	24	68.6	52	35.1	76	41.5
(4) Increased illegitimacy in combination with 1 and 2	11	31.4	96	64.9	107	58.5
Total increase	35	100.0	148	100.0	183	100.0

ALL MARITAL STATUSES

(1) Higher propensity alone	418	38. 3	211	32. 1	629	36. 0
(2) Population increase alone	223	20. 4	82	12. 5	305	17. 4
(3) Components 1 and 2 combined	803	73. 5	385	58. 6	1, 188	67. 9
(4) Declining childlessness alone [1]	89	8. 1	134	20. 4	223	12. 7
(5) Components 1, 2, and 3 combined	977	89. 5	592	90. 1	1, 569	89. 7
(6) Marital stability alone	33	3. 0	18	2. 7	51	2. 9
Total increase	1, 092	100. 0	657	100. 0	1, 749	100. 0

[1] The numbers in this line represent the change for ever-married from declining childlessness alone, and the illegitimacy component. The latter numbers include a few thousand never married women who, strictly speaking, should not be included in this line.

NOTE.—The lower panel is the sum of the numbers for ever-married women from table 8 and the appropriate numbers for never-married women in the upper panel.

The upper panel of table 9 indicates that the change in propensity accounts for 57 percent of the white but only 13.5 percent of the nonwhite increase. Population increase is a relatively minor factor when considered alone. In combination the change in propensity and population increase account for 69 percent of the white and just 35 percent of the nonwhite increase. The difference between the total gain and that which has been related to the first two components is allocated to the effect of rising illegitimacy. For white and nonwhite never-married women together about 58 percent of the total increase is related to rising illegitimacy. Some portion of this figure is actually coming from combination effects of illegitimacy with the first two components. Even so, just 6 percent (107/1749) of the total increase can be attributed to rising illegitimacy.

The lower panel of table 9 summarizes the data for women all marital statuses. For whites and nonwhites together, over two thirds of the increase in female family heads with children aged 15–44 stems from the combined effects of changing propensities to form separate households and population increase. Less than 13 percent of the total increase can be allocated to changes in fertility patterns among the never-married and the ever-married. About 90 percent of the increase is a function of changes in propensities, population increase and fertility. Changing marital stability alone accounts for about 3 percent of the increase.

Expected Change in Numbers of Female Family Heads, 1970–90

Estimates of the change in the number of female heads over the next 30 years can be made by considering the likely change in each of the four components. Of these four components the changing number of women aged 15–44 can be estimated with a high degree of accuracy, since many are already born, and changes in fertility will have only moderate effects. In 1990 the number of women aged 15–44 will be about 57.6 million—about 38 percent more than the count for 1970. (U.S. Bureau of the Census, Statistical Abstract of the United States, 1970: Table 7, using series D projection). Thus, if there were no further change in marital status, propensities or fertility patterns we would expect about 3,476,000 female heads in all marital statuses aged 15–44 in 1990—a gain of 957,000—or 38 percent—over the number in 1970. This gain from population growth alone will underestimate the actual change unless the net effects of changes from the remaining three components are zero or negative. Is it likely that the other components will have zero or negative impacts?

By 1970 the level of propensity to form separate households was high, and while it probably will not decline, there are constraints related to youthful age and low parity (Sweet, 1972) that should prevent it from undergoing much further increase. If the proportion of women who are single increases this may reduce the number of female heads, net of other changes, because the risk of headship is so different between ever-married and never-married women. Given the small change over the past three decades in the marital status of the ever-married one might assume little further increase in the percentage of ever-married women in disrupted marital statuses.

The last component represents the effects of illegitimacy and childlessness among the ever-married. It is likely that a favorable change (i.e. one that will reduce the number of female heads) will occur. The recent decline in illegitimacy rates (Cutright, 1973: Table 3) should continue. The enormous decline in marital fertility among both lower and upper income women (Jaffe, 1972) should be accompanied by increasing childlessness among ever-married women in the *early* years of marriage. This latter change would increase childlessness among women at high risk of marital disruption in, say, the first 3 years of marriage. There is, however, little reason to think that childlessness will return to its 1940 level, and given the relatively small impact of the fertility factor under the stimulus of enormous change between 1940 and 1970, the likely impact on female headship of a moderate reversal of past trends should not be overestimated.

In short, this speculative application of our method suggests that the components responsible for change in the past will not assume the same order of importance in explaining changes in the future. From population growth alone, an increase of about 957,000 female heads aged 15–44 with own children is expected by 1990. The actual increase may be smaller or larger, but given the likely course of the components measuring the propensity factor, marital status, and fertility patterns, it seems likely that population increase will be the major factor behind changes in the number of female family heads over the next 30 years.

The Effect of Timing the First Birth on Female Headship

Because illegitimacy rates are higher among nonwhite than white women, it can easily be assumed that one factor that explains why a higher proportion of nonwhite than white women are female family heads must be related to this difference in the timing of first births in the two populations.

The 1967 Survey of Economic Opportunity, conducted by the U.S. Bureau of the Census, provides data that can be used to assess the impact of illegitimacy and the timing of legitimate first births in relation to the data of marriage on the probability of female headship. The average mother in this 1967 sample was about 40 years of age, so we are looking for a *long run* impact on illegitimacy or the timing of first legitimate births. Does timing of the first birth affect the chances that a mother will be a female family head some years after her first birth?

Table 10 shows the percentage of American mothers aged 59 years or less who were family heads, according to a classification that groups mothers according to marital status and the time of the first birth in relation to their marriage, if any. In the first row are mothers of illegitimate children who were never married at the time of interview. (Seven percent of white and 19 percent of nonwhite mothers whose first births were illegitimate were never married.) The third and fourth columns show the percent of mothers who were heads of a family. Overall, 11 percent of white and 33 percent of nonwhite mothers were family heads. The group with the highest risk of headship in either population are mothers of illegitimate children who have never married. However, these never-married mothers are a small minority (6 percent of nonwhite and under 1 percent of white) of all mothers.

TABLE 10.—*The percentage of mothers aged 59 and under who were female family heads, by fertility history of first births and marital status, by color: United States, 1967*

Fertility history and marital status	Number of mothers		Percent female family heads		Adjusted percent family heads	
	White	Black	White	Black	White	Black
Never married: Illegitimate	30	375	88	98	NA	NA
Ever married:						
Illegitimate	386	1,589	13	27	13	25
Pregnant brides	956	1,034	11	29	11	29
8 to 14 months	3,491	1,071	10	28	11	29
15 to 24 months	2,325	674	10	26	10	27
More than 25 months	3,430	975	10	28	10	28
Others, including never married	318	701	21	74	20	73
Total	10,906	6,044	11	33	11	33

Source: 1967 Survey of Economic Opportunity. Tabulations provided by James Sweet, Department of Sociology, University of Wisconsin, Madison. Timing from marriage to first birth defines births within 7 months as representing premarital pregnancy, hence the category of "pregnant brides." Adjusted columns statistically control for differences among women in different fertility history and marital status categories (within each color group) on years of education, rural or urban residence and age at interview. Adjusted differences for never married women alone were not computed. Never married women are included with "others," a group that includes women with missing data that prevented their being placed in a specific fertility history code. Because never married women are included with "others" the column total is greater than the actual total count.

If we ask whether the risk of headship among women who ever marry is greater among those who have had a first illegitimate rather than a first legitimate birth, we have a clearcut answer. Among whites, there is a difference of only 1 or 2 percent between unwed mothers who later marry and other types of mothers. Among nonwhite mothers, the unwed mothers who later marry are no more likely than are nonwhite mothers of legitimate firstborn children to be a female family head. This lack of an impact of an illegitimate first birth for ever-married unwed mothers holds up with statistical controls for education, place of residence, and age at interview. Therefore, for unwed mothers who later marry, illegitimacy does *not* heighten the risk of later female family headship in comparison with the risk run by mothers of legitimate children. Because 93 percent of white and 81 percent of nonwhite mothers whose first birth was illegitimate have married, the effect of illegitimacy in raising the level of female family headship in either population can only be relatively small.

The presence of unwed mothers not yet married inflates the level of female family headship for nonwhites by 4.2 percent. The white level is inflated by 0.2 percent. Although the effect of illegitimacy on the level of female family headship among blacks is sizable, one should note that the white-nonwhite difference in the level of female headship is reduced by only one-fifth when the effect of higher nonwhite illegitimacy is removed.[2] The basic cause of higher levels of female family headship among nonwhites is the higher proportion of nonwhite ever-married women in a disrupted marital status—not illegitimacy.

Income Supplements and the Increase in Female Family Heads

What does this analysis of the components of change behind the rise in the number of female family heads suggest about the role of past income maintenance policies and family structure? Have past policies been a cause of the increase in female family heads? What might future income policies do to moderate the impact of each component of change in the future?

One component responsible for a large share of the past increase in female family heads was population increase alone and in combination with changing propensities to form separate households. It is difficult, if not impossible, to empirically link past income supplement policies to the population component. The component of change related to the decline in birth control use early in marriage cannot be linked to past income supplement policies, any more than income supplements can be held responsible for rising illegitimate births. Nor has solid evidence emerged that would show a causal connection between income supplement programs and the small changes in marital status that also play a small role in the rise in the number of female family heads.

Outside of the income programs, we know of no social programs whose intent was to increase population growth, shift the pattern of fertility control downward, increase illegitimacy, or decrease marital stability. If so, then there are no easily identifiable programs that one might seek to revoke, and thus reduce the number of female family heads in the future.

One major cause of change, not yet discussed, is the propensity component. Change in this component might be considered to be a function of past income supplement programs which have reduced economic constraints that may have depressed the propensity to form female headed families in the past. Why has the propensity among women at risk of headship to form separate households changed?

[2] This impact can be assessed by statistically removing unwed mothers who never marry from the numerator and denominator used to calculate the percent of women who are family heads. If no nonwhite unwed mothers were found in the category of never married, the percent of the remaining nonwhites that would be female heads would be 28.8 percent. A more detailed analysis of these data is reported in Cutright (1973c).

Is this change a function of increasing benefits under the AFDC program?

AFDC as a Cause of Increasing Propensity To Form Separate Families

We now ask whether it is likely that changes in welfare benefits are responsible for the increase in the propensity of women at risk to establish a separate household. Since the AFDC program is a major source of income to recipients aged 15–44 with children and no husband present, we ask whether variation among the States in the level of AFDC benefits [3] is a cause of (1) differences by State in the living arrangements of children or (2) differences by State in marital status. We can also ask whether changes in benefit levels in various States are related to changes in marital status and the living arrangements of children. Available 1970 census data do not allow direct study of the propensity to establish separate households on a State-by-State basis. However, data on marital status and family status of children, by color, do provide an indirect index of State and color variations in the level of female headed families in 1960 and 1970. If, for example, most children live in husband-wife families, the level of female headed families will be low. If the AFDC program has contributed to increasing the propensity of women with children to establish separate households, then differences in the family status of children by State should be related to different levels of AFDC benefits. Monthly benefit levels for the average AFDC family for specific years were adjusted to 1967 consumer prices. We recognize deficiencies in our analysis. For example, in many high benefit States there are programs for male headed families that emerged in the 1960's. Also, many program related characteristics other than benefit levels are not controlled. Finally, possible effects from nonprogram factors such as urbanization levels, etc., are not adjusted. Still, it is usually the case that strong direct effects emerge, if they exist, under relatively simple analysis. Given the lack of previous work in this area the following sections may still, for all their limitations, constitute a useful exploratory analysis.

The small differences in family size among the States produce only minor average family benefit effects, and these were not adjusted. States are grouped into four strata of nearly equal size in the analysis of whites, and three strata in the analysis of nonwhite or black characteristics. The smaller number of nonwhite strata is required because many States have small nonwhite populations and some States have few blacks among their nonwhite population.

AFDC and the Family Status of White Children

Table 11 ranks the four white strata by the monthly family benefit in 1960 and 1970. Most States remained in the same stratum in the two periods. In 1960 the mean family benefit in the stratum paying the

[3] Utilization of AFDC by the population defined as eligible under State regulations tends to increase as benefit levels increase. Not surprisingly, States with high benefit levels also tend to be States in which a larger proportion of female headed families receive AFDC benefits. Therefore, our tables contrasting the level of benefits with the family status of children tend to maximize program effects because both utilization and the benefit level covary and we have made no attempt to control for utilization among benefit strata.

highest benefits was $186 per month, and this rose to $208 in 1970. In contrast, the States in the lowest stratum paid an average benefit of $80 a month in both 1960 and 1970.

TABLE 11.—*Percentage of families with own children that are husband-wife families, 1960 and 1970: whites*

AFDC benefit stratum	Monthly average family benefit		Percentage of families with children that are husband-wife families		
	1960	1970	1960	1970	Change
Highest	$186	$208	93.1	91.2	−1.9
2d	152	166	93.7	90.3	−3.4
3d	120	121	92.8	90.4	−2.4
Lowest	80	80	92.6	90.5	−2.1

Source: 1960 family status of children by State from 1960 Census, *PC (1) 2d*: Table 111. 1970 family status of children by state from 1970 Census, *PC-1B*: Table 22. *Social Security Bulletin*, various years for December 1960 and 1970 average monthly benefits per family. All 1960 and 1970 benefits expressed in 1967 dollars.

1970 Strata:

Highest: New York; Massachusetts; New Jersey; Pennsylvania; Connecticut; Illinois; Minnesota; Rhode Island; New Hampshire; Alaska; Wisconsin; North Dakota.

2d; Vermont; Michigan; Washington; Kansas; California; South Dakota; Iowa; Colorado; Idaho; Virginia; Oregon.

3d; Maryland; Montana; Utah; Nebraska; Maine; Wyoming; Indiana; Oklahoma; Delaware; Arizona; New Mexico; Texas;

Lowest: North Carolina; Kentucky; Missouri; West Virginia; Tennessee; Georgia; Nevada; Arkansas; Florida; Louisiana; South Carolina; Alabama; Mississippi.

1960 Strata:

Highest; New York; Wisconsin; Connecticut; California; Illinois; New Jersey; Massachusetts; Washington; Minnesota; New Hampshire; Idaho; North Dakota.

2d; Iowa; Oregon; Utah; Rhode Island; Wyoming; Kansas; Colorado; Michigan; Maryland; Montana; Pennsylvania; Ohio

3d; New Mexico; Nebraska; Arizona; Alaska; Vermont; South Dakota; Indiana; Oklahoma; Maine; West Virginia; Louisiana; Virginia.

Lowest; Nevada; Missouri; Delaware; Georgia; Kentucky; North Carolina; Texas; Tennessee; Florida, Arkansas; South Carolina; Alabama; Mississippi.

The right-hand side of table 11 shows the percentage of families with own children present that were husband-wife families. For example, in the high benefit stratum in 1970 we find that 93.1 percent of all white families with own children were husband-wife families. The remaining 6.9 percent of families with own children were, predominantly, female headed families. Comparing the percentage of families with own children that were husband-wife families across benefit strata in 1960 we find no significant pattern. States paying the lowest benefits were nearly equal to States paying the highest benefits in this measure of the family status of children. Thus, the proportion of white families with children that were headed by women does not seem to have been related to the level of AFDC benefits in 1960. Inspection of the next column allows a similar conclusion for 1970. The right hand column indicates that the 1960 to 1970 decline in the family status of white children was about equal in high and low benefit States.

Table 12 groups States by the change in monthly AFDC benefits between 1960 and 1970, and relates these changes to the change in the percentage of families with children that are husband-wife families. The top stratum of States shows a mean AFDC benefit gain of $45, while the lowest stratum shows an average decline of $21. States with the largest benefit increases show a decline of −2.2 percent and States with the greatest benefit losses show a decline of −2.7 percent. The difference between the two extreme strata is minor and it is opposite to the pattern that would be observed if rising benefits were a cause of the rise in white female headed families.

TABLE 12.—*Change of monthly average AFDC benefit and change in percentage of families with own children that are husband-wife families; 1960–70: whites*

AFDC benefit change stratum	Monthly average family benefit change [1]	Change in percent of families with own children that have husband wife families
Highest	$45	−2.2
2d	15	−2.5
3d	−1	−2.4
Lowest	−21	−2.7

[1] AFDC benefit in 1967 dollars.

Source: See table 1. It should be noted here that changes in average family benefits do not reflect only changes in AFDC payment levels. Average benefits can rise or fall as families gain or lose income from other sources such as alimony, veterans pensions, or social security.

Highest stratum: Pennsylvania, Alaska, Vermont, New York, Virginia, South Dakota, Rhode Island, Massachussetts, Michigan, New Jersey, Oklahoma, Minnesota.

2d stratum: North Dakota, Texas, Illinois, Maine, Delaware, New Hampshire, Connecticut, Tennessee, Kansas, Florida, Arkansas, North Carolina, Colorado.

3d stratum: Alabama, Kentucky, South Carolina, Wisconsin, Maryland, Washington, Indiana, Missouri, Iowa, Nebraska.

Lowest stratum: Mississippi, Georgia, West Virginia, Idaho, Oregon, Nevada, California, Arizona, Utah, New Mexico, Wyoming, Louisiana.

A different measure of the family status of white children is available for 1970. In table 13 we group States by the level of AFDC benefits in 1970 and show the percentage of white children under 18 that were living in husband-wife families. Census also provides a tabulation for persons in urban areas. Among urban children 87.5 percent in the highest benefit stratum were living in husband-wife families, compared to 84.1 percent of urban children in the lowest benefit stratum. This finding, of course does not mean that high AFDC benefits encourages husband-wife families. It does tend to support the conclusion that the AFDC program does not encourage the formation of female headed families. If the AFDC program paying high benefits encourages formation of female-headed families by decreasing marital stability, by decreasing the propensity to marry, by encouraging illegitimacy, or by increasing the establishment of separate households, the effects of one or more of these alleged effects of high AFDC benefits should be manifest in the living arrangements of children. We conclude that this evidence does not support the view that the AFDC program affects the formation of white female headed families with children.

TABLE 13.—*Percentage of children under 18 living in husband-wife families in 1970, by monthly average AFDC benefit and place of residence: whites*

AFDC benefit stratum	Average monthly benefit [1]	Percent in husband-wife families	
		State total	Urban areas
Highest	$208	88.2	87.5
2d	166	86.6	85.1
3d	121	86.3	85.4
Lowest	80	84.9	84.1

[1] Benefits in 1967 dollars.

Source: See table 11 for States by benefit stratum in 1970.

AFDC and the Family Status of Nonwhite and Black Children

Table 14 groups States in which 90 percent or more of nonwhites are Negro into three groups. In both 1960 and 1970, low-benefit States show a lead of about 3.5 percent over the highest benefit States in the percent of nonwhite families with children that are husband-wife families.

TABLE 14.—*Percentage of families with own children that are husband-wife families, 1960 and 1970: nonwhites* [1]

AFDC benefit stratum	Average monthly family benefit [2]		Percent of families with children that are husband-wife families		
	1960	1970	1960	1970	Change
Highest	$171	$201	74.8	65.8	−9.0
Middle	109	108	76.7	68.0	−8.7
Lowest	67	68	78.4	69.2	−9.0

[1] Excludes States in which nonwhite population is less than 90 percent Negro, and States with few nonwhite families.
[2] Benefits in 1967 dollars.

1960:
Highest stratum: New York, Connecticut, Illinois, New Jersey, District of Columbia, Michigan, Maryland, Pennsylvania.
Middle: Ohio, Indiana, Louisiana, Virginia, Missouri, Georgia, Kentucky.
Lowest: North Carolina, Texas, Tennessee, Florida, Arkansas, South Carolina, Alabama, Mississippi.
1970:
Highest stratum: New York, New Jersey, Pennsylvania, Connecticut, Illinois, Michigan, District of Columbia, Virginia.
Middle: Maryland, Indiana, Delaware, Texas, North Carolina, Kentucky, Missouri, Tennessee.
Lowest: Georgia, Arkansas, Florida, Louisiana, South Carolina, Alabama, Mississippi.

Source: See table 11.

Between 1960 and 1970, the share of husband-wife families with children declined sharply: this decline was of equal magnitude in the highest and the lowest benefit stratum.

When States are grouped by the amount of AFDC benefit change between 1960 and 1970 we find, in table 15, that the decline in family status of children was slightly higher in the stratum with a $10 benefit gain than in the stratum with a $36 gain, while States with a $9 loss showed a slightly lower decline.

TABLE 15.—*Change in monthly average AFDC benefit and change in the percentage of families with own children that are husband-wife families, 1960–70: nonwhites*

AFDC benefit change stratum	Average monthly family benefit change [1]	Change in percentage of families with own children that are husband-wife families
Highest	$36	−9.8
Middle	10	−10.0
Lowest	−9	−8.1

[1] Benefits in 1967 dollars.

NOTE.—Stratum: Highest: Pennsylvania; New York; Virginia; Michigan; New Jersey; Texas; Illinois; Connecticut. Middle: Tennessee; Florida; Arkansas; North Carolina; District of Columbia; Alabama; Kentucky. Lowest: South Carolina; Maryland; Indiana; Missouri; Mississippi; Georgia; Louisiana.

Source: See table 11.

The data in tables 14 and 15 might be used as weak support for the hypothesis that AFDC was influencing female family formation among nonwhites if it were true that other characteristics that vary with benefit levels were constant. Up to this point, no other characteristics related to family structure have been controlled. Table 16 partially corrects this omission.

TABLE 16.—*Percentage of children under 18 living in husband-wife families in 1970, by AFDC benefit and place of residence: blacks*

Benefit level stratum	Average monthly benefit	Percent in husband wife families	
		State total	Urban areas
Highest	$205	55.4	55.3
Middle	132	56.8	55.6
Lowest	76	55.7	53.0

NOTE.—High stratum: New York, Massachusetts, New Jersey, Pennsylvania, Connecticut, Illinois, Wisconsin, Michigan, Washington, District of Columbia. Middle: Kansas, California, Virginia, Maryland, Indiana, Oklahoma, Delaware, Texas, North Carolina. Low: Kentucky, Missouri, Tennesse, Georgia, Arkansas, Florida, Louisiana, South Carolina, Alabama, Mississippi.

A few States excluded from the 1970 nonwhite analysis (table 14) are included in this table, because census reported black children separately from other nonwhites.

Source: See table 11.

In table 16 we have the percentage of black children under 18 living in husband-wife families in 1970. Children in urban areas are separated from children in the State as a whole. The black population in States paying the lowest benefits is more likely to be in rural than urban areas, and studies of family characteristics in both white and black populations indicate lower rates of marital instability and fewer female headed families in rural than in urban places (Cutright 1971b: table 2). Thus, table 16 allows control of a demographic characteristic that helps produce the weak pattern of effects related to AFDC benefits in table 14 and 15. We find, in table 16, that the percentage of black children living in husband-wife families in 1970 is 55.7 percent in the lowest AFDC stratum, compared to 55.4 percent in the high benefit stratum. The small advantage to children in low AFDC benefit States is removed when the effects of place of residence are controlled. Among urban black children (over 75 percent of the black population is urban) 55.3 percent in high benefit and 53 percent in low benefit States are in husband-wife families. Thus we conclude, as we did for whites, that this preliminary analysis yields no evidence indicating that the AFDC program encourages the formation of female headed families. Rather, the level of benefits appears to have no effect on those events that generate different types of families.

AFDC and Marital Status: 1950–70

If the level of AFDC benefits does not effect the family status of children, it is unlikely that it has large effects in increasing the percentage of women who are married but separated from the husband. Direct examination of the relationship of benefit levels to the marital status of women was conducted on a State by State basis for 1950, 1960, and 1970. One question was whether the level of AFDC benefits depressed the percentage of white or nonwhite women who were living with a husband at ages 15–19, 20–24, or 25–34.

Among white women in these three age groups we found no relationship between the level of average monthly AFDC benefits and changes in marital status after 1950. States paying benefits in the lowest quartile were similar to higher benefit States in the decline in the percentage of women who were married and spouse present. Most of this decline, which is limited to the two youngest age groups, is a function of the increase in the percentage of women remaining single after 1960, rather than being a result of declining marital stability among the ever-married. Between 1950 and 1970 there was virtually no change in the percentage of whites aged 25–34 living with a spouse in any of the four AFDC benefit strata. Direct analysis of changes in benefits and changes in marital status found no relationships between the two measures.

States were grouped into three AFDC benefit strata, and a similar analysis for nonwhite women of comparable ages was conducted. We discerned no pattern of effects indicating that the benefit level discourages early marriage or disrupts marriages in the older age groups.

We conclude that the evidence from these exploratory analyses does not indicate that white or nonwhite women in States with relatively high or low AFDC benefits respond to these real or potential benefits by changing their marriage pattern or by disrupting their marriages because some measure of economic security is provided by the AFDC program.

AFDC and the Propensity Factor

We used the available data in a State by State analysis of the family status of white and nonwhite children to see whether the percentage of children in husband-wife families, or the percent of families with children that were husband-wife families, was related to the level or to the change in the level of AFDC benefits. Neither whites nor nonwhites displayed a relationship between the family status of children and the level of AFDC benefits. Our measure of the family status of children in 1960 allowed the AFDC program to operate for over 20 years to differentiate the family status of children by States. At the end of this period of program operations we found no differentiation by State that was related to the AFDC program. In 1970 we had two measures of the family status of children that might have been affected by the operation of the AFDC program. Again we found no relationship between the level of benefits and the family status of children for either whites or nonwhites. We concluded that the AFDC program has not been a factor behind past changes in white or nonwhite marital status or family structure. It seems impossible to attribute the change in the propensity of white and nonwhite mothers in disrupted marital status to form separate families to the effects of the AFDC program.

If the AFDC program is n t a cause of change over time in the propensity to form separate female-headed households we may ask "What is the cause of this change in propensity?" This is similar to a question posed by Beresford and Rivlin (1966) in their discussion of the trend toward separate living arrangements among older Americans. Although those writers attributed some of the increase in the propensity of the aged to form separate households to social security benefits, this cause was not empirically demonstrated. The authors did note a number of empirically-based contradictions to the thesis that increases in income of the aged were a basic cause of changing propensity, and concluded that other factors—which they called "tastes" for separate housing—were responsible. The shift to separate housing may often involve a tradeoff in which a gain in privacy leads to a worsening of economic position, but the value of privacy is higher than the cost of nonprivacy incurred by living with relatives. Many aged persons now trade off higher living standards for privacy. At the moment no one appears to know why older people have done this, any more than we really understand why younger and economically hard-pressed husband-wife families are far less likely to be doubled up than was the case in the past (table 4, above). Many young husband-wife families now also trade off privacy for lower economic status. These changes among the aged and among young married adults may allow us some perspective in viewing a similar change in propensity among mothers in disrupted marital statuses or never-married mothers. These women are not the only Americans whose propensity for separate living arrangements has changed. Groups that do and groups that do not have access to public assistance or social insurance programs have moved in the same direction. The growth of transfer programs for the aged and for the dependent mother has tended to make things a bit easier for those who receive the benefits but the transfer programs are not in themselves a cause of the increase in the propensity to form separate households or families.[4] Finally, the level of headship among those at risk is now so high that further changes in this component are not likely to inflate the number of female family heads greatly.

IV. The Income of Men and Racial Differences in Marital Status

The higher proportion of nonwhite than white mothers in a disrupted marital status is the major explanation of higher rates of female family headship among nonwhite women. Why is the marital status of mothers in the two populations so different? Does the difference in marital status indicate a lower commitment to the institution of marriage among nonwhites? Are the differences due to cultural effects, or can they be linked to more concrete casual factors? These and similar questions can only be subjected to partial empirical tests because the research required to provide fully satisfactory answers has yet to be

[4] Using 1960 census materials James Sweet (1972) has examined characteristics of mothers aged 59 and under in disrupted marital status that are related to the probability that they will be a female family head. Neither race nor education level is related to the probability of headship. The factors that are related to the probability of headship (e.g., age of youngest child, number of children, the age of the mother, and the type of martial disruption she has experienced) do not appear to provide a clue as to the causes of the trend in the propensity to form separate families. While Sweet also finds that the probability of headship is related to the level of earnings and to the level of "other" income, some 52 percent of 1960 mothers at risk had no other income. Within this group with no other income the probability of headship was 74 percent, compared to 84 percent among women with other income. The level of headship among women with no income other than earnings in 1960 was nearly double the 1940 and 1950 levels of headship for all women at risk. This provides some added weight to the view that changes in public assistance have not been a significant cause of changing propensities to female-headed families.

undertaken. Nonetheless we can provide a relatively clear answer to the question of whether significant differences exist between whites and nonwhites in their commitment to the institution of legal marriage. We can also consider the effect of economic differences on the marital status of the white and nonwhite population.

Male Income and the Marital Status of Men: 1960

Detailed tabulations from the 1960 census allow comparison of three measures of the marital status of men. Only males aged 35–44 are shown, although similar results are obtained using younger and older age groups. Table 17 allows one to compare white and nonwhite men according to their 1959 income. Within each color group we can answer the following questions: "How does income affect the probability that a man will still be never married at this age? Among those who are ever married, how is 1959 income related to the stability of first marriages? Finally, if remarriages are allowed among ever married men, what relationship does income have with this measure of male marital status?"

The first two columns of table 17 display the percentage of men never married. Among white and nonwhite males sharp declines in the percentage never married accompany rising income. The major variation in the percent never married occurs below the median income level for ever married men in each population. There is little change in the percentage never married above that median income. If we contrast white and nonwhite men with no control on income differences we find 11 percent of nonwhite and 8 percent of whites never married. This 3 percent difference provides no support to the view that nonwhite men are more reluctant to enter legal marriages than are whites. Comparing the percent never married within the income interval in which the relevant white median income ($3,913) is found, we observe no difference between the two populations. In sum, we find no support for the view that nonwhite men are less likely to enter legal marriages than are whites at comparable income levels. Indeed, it is somewhat remarkable that lower income nonwhite men are as likely to be ever married as we find them to be.

TABLE 17.—*The relationship of the marital status of men 35–44 to their 1959 income, by color: United States 1960*

1959 male income	Measures of marital status								
	Never-married			Ever-married with first wife			Ever-married not disrupted		
	White	Nonwhite	White less nonwhite	White	Nonwhite	White less nonwhite	White	Nonwhite	White less nonwhite
Percent if income is:									
None	42	33	9	55	34	21	82	70	12
$1 to $999	29	19	10	65	55	10	85	80	5
$1,000 to $2,999	15	11	4	75	62	13	92	86	6
$3,000 to $4,999	8	8	0	81	68	13	96	90	6
$5,000 to $6,999	5	6	−1	85	72	13	97	92	5
$7,000 to $9,999	3	5	−2	87	77	10	98	94	4
$10,000 or more	3	6	−3	89	79	10	98	95	3
Total	8	11	−3	83	64	19	96	87	9
Median income of never-married	$3,913	$2,322	$1,591	na	na	na	na	na	na
Median income of ever-married	na	[illegible]	na	$5,865	$3,337	$2,528	$5,865	$3,337	$2,528
Color difference at white median interval			0			13			5

Source: Derived from 1960 Census. See Cutright, 1970: Table 5, and Cutright, 1971: Tables 4 and 5 for references.

Among ever-married men white median income was $2,528 higher in 1959 than that of ever-married nonwhite men. This higher level of white income affects racial differences in the stability of first marriages, as we observe in the center columns of table 17. The percentage of ever-married men still living with their first wife increases as incomes increase in both populations. Overall there is a 19-percent difference favoring whites; this difference declines to 13 percent in the income interval containing the white median income, and is only 10 percent above $7,000. The percentage of men still living with the first spouse is a poor measure of commitment to the institution of marriage because it does not index the motivation of the male population to marry and live in a husband-wife family. The percentage of ever-married men not currently divorced or separated due to marital discord allows for remarriages, and therefore provides a superior measure of the strength of commitment to the institution of marriage as a means of fulfilling adult roles. The data are shown in the righthand columns of table 17.

We find that 96 percent of ever-married whites and 87 percent of ever-married nonwhite men aged 35–44 are not divorced, or permanently separated from their current spouse. In both populations increasing income is related to more favorable marital status. The overall racial difference of 9 percentage points declines to just 5 percent when nonwhites with incomes similar to the white median income are compared.

The control on 1959 income of the man does not adequately control for economic factors that affect marital stability. At the same level of income in a given year, per member consumption in the nonwhite family will be lower due to larger family size; assets will also be lower; the cumulative past income and the prospect for future income growth will be lower; the larger number of dependent other adults aggravates crowding in the nonwhite household; the quality of neighborhood is lower; the physical quality of housing per dollar expended is lower; and health of family members is poorer [Cutright, 1971a]. None of these factors is fully controlled with a control on last year's income, although all these differences will tend to increase racial differences in marital stability. Given these considerations we conclude that the level of commitment to the institution of legal and permanent marriage is similar in the two populations of men. While this appears likely, we still have the question of why instability of first marriages is so high among nonwhites and why, in spite of small differences in the marital status of white and nonwhite men, we have found large differences in the marital status of ever-married women.

Why Marital Status of Ever-Married Women Differs Between Whites and Nonwhites

Although this review of 1960 census materials has found strong effects of male income on male marital status, a comparisom of white and nonwhite ever-married men at the same 1959 income level did not completely eliminate differences in marital status. And, while a measure of marital status that allowed for remarriages showed small differences, this observation should not blind one to the fact of higher rate of disruptions of first marriages. It is the higher rate of first marriage disruption, in conjuction with a longer period from separation divorce and divorce to remarriage that accounts for the large difference in the percentage of nonwhite as compared to white women who are in a disrupted marital status (Glick and Norton 1971: table 3). We now ask why the rate of first marriage breakup is higher among nonwhites.

INCOME, CONSTRAINTS, AND SATISFACTION

The decision to terminate a marriage is the outcome of the level of satisfaction with the marriage, and the level of constraints that inhibit dissolution of the marriage. At the same level of satisfaction a high level of constraints may result in a stable marriage, while a lower level may allow the marriage to dissolve. At the same level of constraint a high level of satisfaction may result in stable marriage while a lower level of satisfaction may allow the marriage to break up.

Constraints explain, in part, why increases in income are accompanied by higher levels of stable marriages. The higher the income the higher the level of economic and social constraints that inhibit hasty marital dissolutions. At higher income levels the loss of social status to husbands and wives steming from dissolution of the marriage may be greater than that experienced by couples further down the income distribution. The decline in ecomonic well being as a consequence of family breakup may also be greater for the upper income couple. Cumulated assets increase with income, and this factor may also inhibit hasty decisions to terminate the marriage. Since a higher proportion of nonwhites are in the lower income intervals, the constraints against nonwhite marital breakup are lower. Economic constraints can be viewed as a negative type of control over marital disruption that varies with past, current and expected future income levels.

A positive control over marital dissolution comes from feelings of happiness with the marriage. Couples that are very happy with their marriages do not part company. At the same income level white and nonwhite couples are not equally pleased with their marriages (Renne, 1970: tables 6 and 7). Why are nonwhites less satisfied? To understand this finding requires understanding of how a high or low level of marital satisfaction is achieved. Careful analysis has shown (Orden and Bradburn 1969; Bradburn 1969; Bird 1970) that satisfaction with marriage is not the simple result of the feeling one has about one's spouse. Rather, the level of satisfaction is determined by the balance of positive and negative inputs to the marriage. Further, the measures used to index the level of positive or negative inputs to the marriage are not associated; the level of negative inputs is not a function of the level of positive imputs. Therefore, both dimensions of marital satisfaction operate to determine the level of satisfaction each spouse feels toward their marriage.

The negative dimension has been measured by the number of disagreements the couple has had recently about children, spending money, and so forth. This measure is called the "marital tension index" (Bradburn, 1969). The measure of the positive dimension for both husbands and wives is usually derived from questions on companionship and sociability of the spouses. However, Bird (1970) has shown that the husband's positive feelings about the marriage are closely related to his satisfaction with his job. This finding was expected because, in addition to the role of husband and father, work is the major role through which men fail or succeed in establishing themselves as successful and worthwhile adults. The experience the man has at work should affect his marriage, and table 18 indicates that this expectation can be documented. In table 18 we tabulate the marital tension level against the level of job satisfaction, and find how these positive and negative determinates of marital happiness are related to the probability that different types of men will say that they are "very happy" in their marriages.

TABLE 18.—*Percentage of husbands reporting very happy marriages, by marital tension and the level of job satisfaction*

Marital tension index (negative input)	Job satisfaction (positive input)			Total
	High	Middle	Low	
Low	89	69	65	74
Middle	74	50	36	56
High	56	36	27	39
Total	72	50	39	54

Source: James H. Bird, "A Statistical Study of Marital Happiness," Master's paper, Department of Sociology, Vanderbilt University, 1970. Data courtesy of National Opinion Research Center, University of Chicago.

NOTE.—N=310 whites.

In this sample of white couples we find the percent "very happy" increases from 39 to 74 percent as we move from high to low levels of marital tension. We also note that the percentage "very happy" declines from 72 to 39 percent as job satisfaction declines. Because job satisfaction and marital tensions are not related, each variable affects the level of marital happiness when the other variable is being held constant. For example, among persons low on marital tensions, the percentage "very happy" is 89 among those with high job satisfaction, but only 65 percent among those low on job satisfaction. The percent high on marital happiness varies from 89 to 56 percent, as a function of marital tension. These two variables, representing the positive and negative inputs to marriage, clearly show the effect the world of work on male satisfaction with marriage.[5]

High job satisfaction is caused by on-the-job mobility, income growth (Bradburn, 1969: chapter 10), and the feelings that one's talents are being used. The difference between white and nonwhite men in experiencing job mobility and income growth is severe, even after other differences between the two groups are statistically con-

[5] A comparable table for blacks is not available. Renne (1970: table 13), using different measures of job satisfaction, compared the level of marital dissatisfaction among husbands and wives by the level of job satisfaction among those currently employed. Among white husbands and wives not satisfied with their jobs, 24 and 25 percent respectively were dissatisfied with their marriages. Among whites saying they were "very satisfied" with their jobs only 14 percent of the husbands and 20 percent of the wives reported dissatisfaction with the marriage. Among black husbands and wives [illegible] and 55 percent not satisfied with their work were not satisfied with their marriage in comparison to 22 and 28 percent of black husbands and wives respectively who were "very satisfied" with their jobs. Within each population the level of job satisfaction is related to marital dissatisfaction, and the percentage effect is larger among blacks than among whites. Finally, at the same level of job satisfaction nonwhites are less satisfied with marriage than are whites. This difference would be reduced had income and other factors been controlled.

trolled (Cutright, 1973; Duncan, 1969: table 4-4). A higher proportion of Negro than white men are denied job mobility and income growth, events that reduce nonwhite males' feelings of satisfaction from work. Feelings of dissatisfaction with work "spill over" and lower the positive inputs to marriage, thus depressing the husband's level of marital satisfaction.

What accounts for the lower levels of job mobility and income growth that depress job satisfaction and thus limit the positive input to nonwhite marriages? The causes of racial differences in marital happiness can be broken down into two components—the part of the total difference that is a function of racial discrimination in the labor market, and the part that is not.

Racial discrimination affects the level of current income, the level of permanent income, the growth of income with age, and occupational mobility. Therefore, racial discrimination affects both marital constraints and marital satisfactions. Recent work by Duncan (1969: table 4-4) breaks up the difference in 1964 earnings of white and Negro men aged 25-34 years. Duncan estimates that $660 of the $3,030 difference in 1964 earnings can be attributed to the nonwhite male's lower socioeconomic origins—his disadvantaged class background. An additional $720 is a function of his lower level of academic achievement in school and fewer years of schooling. The remaining $1,650 of the white-nonwhite earnings difference (or 54 percent of the total difference) is independent of all the preceding factors, and may be thought of as the cost of discrimination borne by a Negro male aged 25-34 in 1964.[6]

This "cost" is cumulative over time. We estimate (from U.S. Bureau of the Census, 1968) that the lifetime earnings of white and nonwhite males will differ by some $200,000 of which some 54 percent (or $108,000) is a function of racial discrimination, net of class background, education, or academic ability effects.

The difference in the marital status of ever-married white and nonwhite women is viewed as one of the consequences of racial discrimination. Racial discrimination depresses positive inputs to the marriage that stem from satisfaction with work and performance of marital roles. Low positive input reduces satisfaction with marriage and leads to higher rates of disruption in the nonwhite population. The level of constraints against marital instability is also related to the effect of racial discrimination on the level of income, past, present, and future.

After a marriage is disrupted, the time to divorce and the time to remarriage is longer among nonwhites than among whites (Glick and Norton, 1971: tables 3, 5, and 7). This difference is related to lower incomes of the nonwhite population which block rapid reentry to marriage after one marriage dissolves. Both the higher level of marital disruption and the length of time to remarry inflate the difference in the percent of white and nonwhite women in a disrupted marital status. The basic cause of a higher percentage of nonwhite women in disrupted marital statuses can be traced to the consequences of

[6] Duncan notes that attributing the class background difference as a "cause" of lower Negro male incomes ignores the point that this difference is itself a result of racial discrimination in the previous generation. Therefore, the "effect of discrimination" measured accounts only for discrimination against the sons, but fails to include discrimination against their fathers.

racial discrimination, rather than to hypothetical "cultural" differences between blacks and whites.

V. The Likely Impact of Future Income Supplements

If future income supplements to intact or broken low-income families are to change the marital status or family status of the adult population, they must change the rate of family formation through marriage or change the rate of marital instability and remarriage.

To depress the marriage rate, the supplements must overcome the value placed on the institution of legal marriage by both men and women. We have found no evidence of a racial difference in the commitment of men and women to legal marriage. Recent upward shifts in the age of marriage are the result of economic and demographic factors, and do not signal a movement away from legal marriage. The swing to delayed marriage should be welcome, because it will reduce completed family size and may increase marital stability. A study of State-by-State variations in average monthly AFDC benefits and marital status did not uncover any effect of this income supplement program on the recent shift to a later age at marriage among either white or nonwhite women. It is unlikely that a future income supplement program for intact or broken families will depress rates of family formation through legal marriage.

Among persons already married, variation in average monthly AFDC benefits among the States was unrelated to the level of disrupted marriages. Income supplements to intact families in the future might increase the rate of marital disolution if they lowered the level of social and economic constraints or depressed marital satisfaction. However, the level of constraints increases with income, and it is difficult to think of the way in which supplements would reduce constraints among married couples receiving supplements unless both husband and wife were willing to dissolve the union simply to increase family income through "family splitting." This step would be unlikely unless the supplement program was unable to cope with such fraud and the judicial system was unable to extract court-ordered support payments from the husband. Moreover, the constraint factor includes both social and economic factors, and the social costs of marital dissolution (e.g., loss of the status of being legally married) are ignored if one considers only the possible economic advantage of family splitting. Further, it appears reasonable that of the two major factors that determine whether or not a marriage dissolves (constraints and marital satisfaction), the latter is a major determinant of marriage disruption among persons that would become eligible for income supplements. Will supplements to intact families decrease satisfaction with the marriage?

Supplements could encourage marital instability if they altered the balance of positive and negative inputs to marriage that determine marital happiness. To do this, the supplements would have to increase the level of negative inputs flowing from the number of marital disagreements and/or decrease the level of positive inputs. We find it difficult to imagine how the supplements would increase marital tensions, particularly since one major source of marital tension has to do with allocation of a limited amount of money. Supplements might increase marital instability if they reduced the positive inputs to marriage that come, for example, from job satisfaction, or husband-wife interaction. It is unlikely that a supplement to the family income would depress the job satisfaction of either the husband or the wife, or depress positive inputs to marriage from husband-wife interaction.

It is unlikely, then, that a supplement to an intact family would cause a decline in social constraints that inhibit marital dissolution among those eligible for the benefits. It is also unlikely that additional income will depress the level of positive inputs to the marriage. Nor is it reasonable to expect that added income will increase negative inputs to marriage. In terms of our model of the general causal factors that determine whether couples remain together or part company, we conclude that there exists no reason to expect that supplements will increase marital instability among persons already married, or among those who will marry in the future.

The vast majority of persons in our society still prefer and gravitate toward monogamous marriage because it is uniquely able to provide a variety of rewards. Persons marry with a commitment toward permanence, but they also believe that if certain rewards do not flow from the union they have the right to end it. Thus, Americans have desires for marriage and for permanence, and they will marry and stay married if the marriage turns out to be sufficiently rewarding (Scanzoni, 1970, 1971, 1972).

While most persons think of marital rewards solely in interpersonal terms, it is evident that these rewards often hinge on the level of economic resources that are available to husbands and wives. Thus, marital stability increases with corresponding increases in income of the husband. Providing additional funds to low-income husband-wife families should increase the probability of their remaining stable, rather than decrease it. Both Farley (1971) and Cutright (1971b) estimate moderate *positive* effects on nonwhite marital stability from various policies to shift the distribution of nonwhite males across income intervals to equal that of white males.[7]

Would the infusion of transfer income from an income supplement program decrease commitment to marriage and the family? We have found nothing in our review of the past or recent level of commitment to marriage and the family to indicate that supplements would have a negative impact. We conclude that income supplements to intact families pose no threat to legal marriage, legitimate child-bearing within marriage, or the stability of marriage.

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[7] The moderate size of these effects is constrained by the static nature of the data used to estimate the impact of higher nonwhite incomes. The actual process that would have to precede a change in the distribution of nonwhite men to higher income intervals through a change in the labor market would be expected to increase the effect, because the process would involve the elimination of racial discrimination. As noted in the text, the effects of discrimination on marital stability are, in part, independent of the effects of income levels.

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ILLEGITIMACY AND INCOME SUPPLEMENTS

By PHILLIPS CUTRIGHT*

SUMMARY

In comparative international perspective the U.S. illegitimacy rate around 1960 ranked 23rd in a list of 46 countries. The rate for U.S. nonwhites ranked 17th and the rate for U.S. whites ranked 37th in a list of 48 populations. The U.S. nonwhite rate is comparable to the rate for Iceland and below that of 15 Latin American nations—largely because U.S. nonwhites have rejected the institution of consensual or common law marriage, and have better control over fertility than do populations in Latin nations.

For both whites and nonwhites illegitimacy rates were relatively stable from 1920 through 1940. After 1940 the white illegitimacy rate gradually increased and reached a high in 1968—the last year for which national data are available. The nonwhite rate rose from 1940 through 1965, but declined between 1965 and 1968. All age groups participated in the increase of illegitimacy rates between 1940 and 1965; older nonwhite women have shown declines of 30–35 percent between 1965 and 1968. Illegitimacy rates among teenage nonwhite and white women continued to increase through 1968.

Illegitimacy rates are determined by the degree to which a population of unmarried women in the childbearing years is subject to involuntary controls over conception and gestation, as well as by the amount of sexual activity, and use of effective contraception and induced abortion.

The rise in U.S. illegitimacy rates after 1940 was not caused by decline in the use of effective contraception.

Induced abortion by pregnant unmarried women probably declined between 1940 and 1950, and then rose slowly during the 1954–65 period among both whites and nonwhites. The rise in abortion after 1954 could not keep up with the increase in the white and nonwhite "illicit" pregnancy rate—therefore, illegitimacy rates continued upward during a period of increasing abortion use. Induced abortion of illegitimate pregnancies was about the same in the 1960's as it had been in 1940.

Improvements in control of venereal and other diseases since 1940 have greatly reduced sterility and spontaneous fetal loss among nonwhites and, to a lesser extent, among whites. Improvements in nutrition in both populations have contributed to an increase in fecundity among girls aged 17 and under. We estimate that some 88 percent of nonwhite increases in illegitimacy to women aged 15–44 between 1940 and 1968 and 19 percent of the white increase can be accounted for by improvements in health conditions that reduce involuntary hindrances to both gestation and conception. From 1940 through 1968 *all* of the increase in nonwhite teenage illegitimacy

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rates and a quarter of the increase in white teenage rates was related to improved health, rather than increasing sexual activity.

Small increases in sexual activity ma, account for that portion of the rise in illegitimacy not already accounted for by improved health conditions.

About 60 percent of white unwed mothers and 80 percent of nonwhite are below 125 percent of the poverty level at the time of birth. The illegitimacy rates of low-income women are about eight times higher than are the rates of nonpoor women in the white population; the rates of low-income nonwhites are three times greater than those of nonpoor nonwhites. About 45 percent of the difference between white and nonwhite illegitimacy rates is related to the higher proportion of nonwhites below the low income line.

Unmarried mothers have lower utilization of medical service before, during, and after pregnancy than married women of similar economic status. This partially accounts for their higher rates of fetal loss and maternal mortality in childbirth, and the higher infant mortality rate of illegitimate children.

Since some 90 percent of illegitimate children are not wanted by the mothers, the prevention of pregnancy or birth can be achieved by a shift in, access to, and utilization of medical services from the present pattern. Subsidized contraception programs were evaluated for their potential to reduce illegitimacy. The characteristics of unwed mothers and of the present subsidized contraception-only program suggest that illicit pregnancies will be not reduced by more than 13 to 27 percent. However, if this effect were to be reinforced by the addition of legal abortion to program services, we would expect a reduction of perhaps 50 percent in the total U.S. illegitimacy rates within 6 to 8 years.

The effect of different types of government "welfare" programs on illegitimacy were reviewed. Illegitimacy rates outside the United States are not related to the benefit level of family allowance programs.

In the United States illegitimacy plays an increasing role in determining total AFDC expenditures, and currently accounts for about 30 percent of all AFDC benefits. However, we cannot conclude that AFDC benefits are a cause of illegitimacy. Analysis of trends in numbers of families on AFDC and change in illegitimacy rates revealed several examples of periods of declining or stable numbers of families on the rolls accompanied by rising illegitimacy. Further, the great explosion of AFDC in the 1965–70 period was accompanied by declining nonwhite illegitimacy rates for women aged 20 and older, and by nearly stable rates for older white women. During this period teenage rates rose, as they had been rising since 1940.

Gross comparisions of State benefit and State illegitimacy rate changes from 1940 to 1960 found no relationship between AFDC benefit change and illegitimacy rate change for either whites or nonwhites. In 1960 there was no difference in illegitimacy rates for whites between States paying high benefits and States paying low benefits. In 1970 there was no systematic association between white or nonwhite illegitimacy rates and AFDC benefit levels. It is likely that attempts to control illegitimacy by lowering benefits or restricting access to benefits will have no more effect on illegitimacy in the future than similar efforts have had on past illegitimacy.

Economic theories of fertility may not apply to illegitimacy because these theories assume rational deliberate calculation by parents and ess to perfect means of fertility control.

We compared birth control, adoption, and income supplements as methods of reducing or eliminating social and material punishments related to illegitimate childbearing. Income supplements have little impact in contrast to a wide range of benefits from birth control or adoption. It should be noted that punishments, while formidable, have not eliminated illegitimacy in the past. The option to reduce the sanction of sole responsibility for child care through adoption is far more available to whites than nonwhites. In spite of having this option available, the white illegitimacy rate is far lower than the nonwhite.

We examined the possibility that income supplements might reduce use of birth control or increase the number of illegitimate children living with their mothers—and thus increase the public dependency burden. We concluded that income supplements would affect none of the factors that determine the living arrangements of illegitimate children and their mothers. Furthermore, all available data on recent fertility trends among low-income women suggest increasing utilization of contraception and abortion during a period of rapid expansion of the AFDC program. Existing barriers to birth control, not the present or projected income supplement program, are the main hindrance to voluntary control over illegitimate fertility.

I. Measurement and the Immediate Causes of Illegitimacy

Measures of Illegitimacy

The most common measures of illegitimacy are: (1) The number of illegitimate births, (2) the ratio of illegitimate to all births, and (3) the illegitimacy rate—the number of illegitimate births per 1,000 unmarried women of childbearing age. Unmarried women include never-married, widowed, and divorced women. For analytic purposes the number of illegitimate births is not useful and the illegitimacy ratio is so powerfully affected by changes in marital fertility and marital status that it has only limited utility. The illegitimacy rate per 1,000 unmarried women in their childbearing years (defined in the United States as 15 to 44) provides a measure of the birth rate to the unmarried population. Since most studies of levels or changes in illegitimacy want to measure the probability of an illegitimate birth to the population at risk of such a birth (unmarried women for the most part), it is necessary to use the illegitimacy rate.

The accuracy of an illegitimacy rate depends on: (1) Valid and complete registration of illegitimate births; (2) accurate counts of the unmarried female population by census; and (3) minimal distortion of the resulting rate due to the infusion of illegitimate births from women who are married, but nonetheless deliver children that are registered as illegitimate. In general, analysis of these problems has indicated that U.S. illegitimacy rates (when adjusted for known errors) provide a reasonably accurate measure of the level as well as the trend of illegitimacy. (Cutright, 1972c: appendix: Measuring Illegitimacy in the United States.)

The Immediate Causes

In the absence of measurement error stemming from one or another of the above factors, the immediate causes of an illegitimacy rate are similar to the immediate causes of fertility rates in a population. Following Davis and Blake (1956) we have three immediate causes.

1. Exposure to risk of pregnancy, which is determined by: (a) the proportion of unmarried women having sexual relations; (b) the age at which sexual activity begins; and (c) the frequency of coitus.
2. Control over conception among those unmarried women exposed to risk of pregnancy, a factor influenced by: (a) involuntary control due to inability to conceive; and (b) voluntary control through the use of contraception or sterilization by women able to conceive.
3. Control over gestation, involving: (a) involuntary fetal loss from spontaneous abortion or stillbirth; or (b) voluntary fetal loss from legal or illegal abortion.

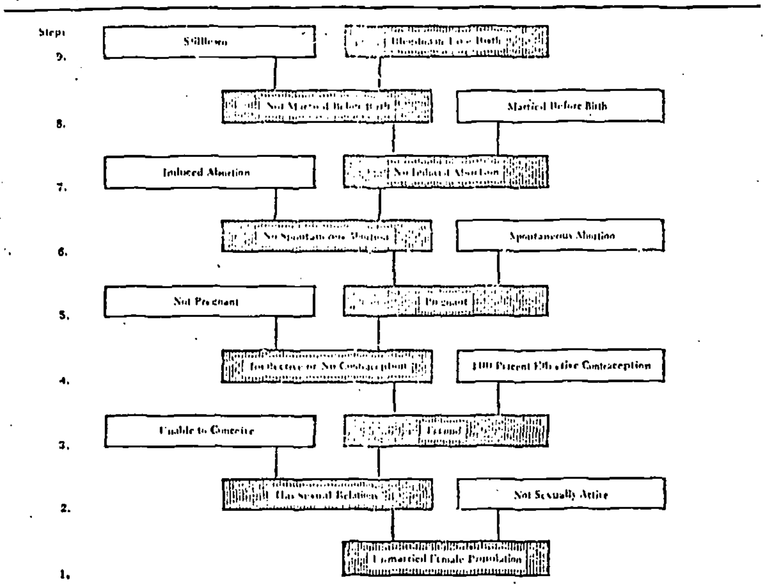
The number of illegitimate births and the rate will also be affected by the percentage of pregnant unmarried women who carry the fetus to term but marry before the birth. These women are pregnant brides—the legitimate child was conceived before marriage. The percentage of women pregnant outside of marriage who marry prior to the birth varies among populations and among subgroups within the same population. Both cross-national comparisons and analysis of subgroups within the United States indicate that differences in the probability of marriage before birth among populations is a function of such factors as the number of previous births to the mother, the age of the unmarried woman, her relationship with the putative father, and the marital status of the putative father. Differences in the probability of marriage prior to birth exist, and these differences are related to the illegitimacy rate.

A population with a high out-of-wedlock conceived birth rate (legitimated *plus* illegitimate births per 1,000 unmarried women in the childbearing years) will have a low probability of legitimation (Cutright, 1972b; 1972c). A population with a low out-of-wedlock conceived birth rate will have a low illegitimacy rate *and* a high probability of legitimation. Women who become pregnant brides are by and large women who would have become brides with or without the pregnancy, since their sexual activity leading to pregnancy is with a man they want to marry and is restricted to a man who wants to marry his sexual partner. Women who become unwed mothers rather than pregnant brides have sexual relations with men they do not care to marry, men ineligible to marry, or with men they know do not want to marry them.

In the United States nearly 70 percent of white compared to 24 percent of nonwhite pregnant unmarried women marry before birth, thus legitimating the out-of-wedlock conceived birth. This difference in the probability of legitimation is a function of the higher rate of out-of-wedlock conceived births in the nonwhite population. It is the rate of out-of-wedlock conceived births, not the probability of legitimation, that is the main immediate cause of racial differences in illegitimacy rates.

Figure 1 illustrates the stages that precede an illegitimate birth. The analysis begins with the unmarried female population, some of which is sexually active and therefore is at risk. These women are retained as we move to stage 3. Some sexually active women are sterile and are unable to conceive. Among women who are fecund (able to conceive) only those not using contraception effectively remain at risk of pregnancy. Only the sexually active fecund women not using effective contraception will become pregnant. Between 20 and 40 percent of pregnancies of 4 or more weeks gestation in different populations terminate in spontaneous fetal loss (Shapiro, et al., 1971; James, 1970). A pregnancy not terminated by spontaneous fetal loss can be terminated by induced abortion.

Figure 1. Steps to Unwed Motherhood



Source: P. Cutright, 1971a: 27.

The pregnant unmarried women who carries the fetus to term may marry or not marry before the birth. Those not marrying will have an illegitimate birth if they avoid stillbirth. Stillbirths are infrequent—under 2 percent of all deliveries. In the United States only live born children are counted as legitimate or illegitimate births. Unmarried women with an illegitimate live birth become unwed mothers.

Similarly, married women who have adulterous or incestuous intercourse can be followed through the same stages, except that legitimation through marriage is not an option for them.

Thus, the level and the trend in levels of illegitimacy is the end result of patterns of sexual behavior, fecundity differences, contraception, involuntary and voluntary fetal loss. These immediate causes prior to legitimation by marriage interact to affect the probability of legitimation which, in recent decades at least, is not believed to be a true immediate cause of illegitimacy.

II. Historical and Comparative Perspectives

Trends Since 1750

Historical data on illegitimate births exist for a number of European states and nations since around 1750. Analysis of trends in Europe allows one to divide the years since 1750 into three distinct stages. The first period extends from around 1750 to around 1870, the second from around 1880 to 1940, and the third includes the post-World War II years (Cutright, 1972b).

The first explosion of illegitimacy in Western nations occured after 1750. All across Europe the rates drove upward, peaking between 1860 and 1880 in most nations (Shorter, 1971: 265–272). Recent work (Shorter, 1972) attributes this long-run rise to social, demographic and economic changes that resulted in the diffusion of modern ideas of self-expression and individualism among the lower classes who, for the first time, had moved from a life situation that repressed nonmarital intercourse (and perhaps premarital sex with the future spouse) to one in which family and community authorities were no longer able to exercise control. Thus, rising sexual activity among couples that would not marry brought with it increasing illegitimacy rates.

After about 1880 illegitimacy rates all across Europe receded (Cutright, 1971a; Shorter, Knodel and van de Walle, 1971). In nation after nation the rates began a decline that continued through the 1930's. What explains the decline of illegitimacy after 1880?

The decline in illegitimacy was accompanied by a common change in nearly all European nations—the decline of marital fertility rates. Declining marital fertility was not caused by a decline in coital activity; rather the decline was due to increasing use of abortion and male methods of contraception—*coitus interruptus* and condom. Increasing use of birth control by the married population during this period indicated a set of conditions that allowed birth control among the unmarried sexually active population to increase as well. It seems likely that illegitimacy declined in most nations because birth control increased. In some nations a decline in common law marriages whose issue were defined as illegitimate may also account for some part of the declining rate to older women. The decline in the rate was less pronounced among teenage girls, a fact that may be accounted for by a dramatic rise in fecundity among the young after 1880 (Tanner, 1968) as well as by improvements in other health conditions that decreased sterility and spontaneous fetal loss (Cutright. 1972a). Improvements in health conditions must have moderated the decline in illegitimacy after 1880, but no measures of this effect are available.

The third era in the history of illegitimacy begins around 1940. Illegitimacy rates in Europe remained low during World War II; after the War some nations experienced stable, others declining and still others rising rates. We can statistically account for most of these different post-War patterns, by examining differences among nations in the control of marital fertility and changes in the age at marriage and legitimate childbearing (Cutright, 1971a: table 2). That analysis of change in post-World War II rates included the United States, Canada, Australia and New Zealand, as well as Japan and all European nations that do not allow legal abortion on demand. The results of that study of 23 populations clearly showed that post-World War II illegitimacy rates tended to increase when marital fertility rates indicated weak efforts to control legitimate childbearing; also, illegitimacy rates tended to rise when the age of marriage was going down, and when, therefore, the age of legitimate childbearing was declining. The 1950–60 increase in illegitimacy in the United States is not unusual. Other populations with similar behaviors (e.g. Canada, England and Wales, Scotland, New Zealand, and Australia) also experienced a similar rise in illegitimacy.

Awareness of these historical trends and the changes that appear to explain them do not support traditional explanations of illegitimacy as resulting from secularization or social disorganization. Nor do the fluctuations over time give support to some psychological explanations of illegitimacy which treat unwed mothers as typically "disturbed," "neurotic," "psychotic," or "acting out" various needs. [Pauker, 1969, has provided an excellent review of this literature. Also, analysis of illegitimacy rates among women with the same years of birth [Cutright, 1970: Ch. 6] indicates that the same cohort that had a very high illegitimacy rate at one age may have a low rate in later years; the same unmarried women may have a very low illegitimacy rate in their early years of childbearing but then have a high rate in the later years of childbearing. These findings, when considered along with the cyclical nature of illegitimacy rates indicates that excessive reliance on psychological characteristics of unwed mothers to explain varying illegitimacy rates is no more useful than would be an effort to explain unemployment rates with psychological variables.

U.S. Illegitimacy in Comparative Perspective

The United States obviously is not the only nation in which illegitimate children are born. Table 1 shows illegitimacy rates in 46 nations around 1960. The note to table 1 indicates differences in measurement of these rates—primarily between Latin American nations and Iceland, compared to the other populations. In this table the treatment of the consensually married in high illegitimacy nations allows a somewhat similar treatment of separated women in the U.S. population. One can apply either the rate of 90 or 64 to the U.S. non-white population and emerge with little difference in the comparative picture. The illegitimacy rates of European populations or U.S. whites are not changed much by including or excluding separated women, because few women aged 15–44 are separated.

Among these 46 nations the United States ranks 23d. Nationally we are about average, although the total U.S. rate is below that of only five non-Latin populations. The total U.S. rate is far below the rates for Uruguay and Argentina, which also have just 10–12 percent nonwhite populations.

TABLE 1.—*Illegitimacy rates per thousand unmarried women in 48 populations around 1960*

Nation	Rank	Rate	Nation	Rank	Rate
Dominican Republic	1	218	Sweden	25	20
Nicaragua	2	199	Denmark	26	18
Honduras	3	193	Australia	27	17
El Salvador	4	189	Canada	28	17
Venezuela	5	188	Poland	29	16
Guatemala	6	169	France	30	16
Panama	7	168	England and Wales	31	15
Peru	8	167	Yugoslavia	32	15
Ecuador	9	137	Hungary	33	13
Paraguay	10	135	West Germany	34	12
Costa Rica	11	106	Scotland	35	11
Cuba	12	104	Czechoslovakia	36	11
Mexico	13	100	United States, white	37	9
Colombia	14	78	Norway	38	9
Iceland	15	76	Finland	39	8
Uruguay	16	66	Switzerland	40	7
United States, nonwhite	17	64	Luxembourg	41	7
Chile	18	48	Belgium	42	6
Argentina	19	43	Italy	43	5
Austria	20	27	Spain	44	5
New Zealand	21	26	Ireland	45	4
Portugal	22	24	Netherlands	46	4
Bulgaria	23	24	Greece	47	2
United States	24	22	Japan	48	2

NOTE.—Latin American nations, Bulgaria and Yugoslavia use women 15 to 49 while other populations use women 15 to 44. The addition of unmarried women 45 to 49 produces rates slightly *lower* than would be found had only women 15 to 44 been used. Rates in Latin American nations and Iceland include consensually married with single, widowed and divorced women in the denominator of the rate. Illegitimate births, regardless of the marital status of the mother, are in the numerator of all rates. Color specific rates in the United States included separated women, a change which has substantial effects on nonwhite rates (they move from 90 to 64) but little effect on white rates (they are 9.3 and move to 8.9). The total U.S. rate declines from 21.7 to 20.1 by adding separated women. Illegitimacy rates in Latin nations and Iceland would be lower had illegitimate births to consenually married women been removed from the numerator and this group of women also removed from the denominator. Latin American and U.S. births adjusted for under reporting; U.S. population adjusted for census undercount. Rate for New Zealand excludes the Maori population.

Source: Latin American nations from Johnson and Cutright (1973: table 13.4); Iceland from Cutright, 1970: Table 1.15; remaining nations from Cutright, 1971b: Appendix table I, United States rates from Cutright, 1972c: Table 2 and appendix table 7.

What may come as a surprise is the relatively low illegitimacy rate of the U.S. nonwhite population. Its rate is below that of all but two Latin Nations—both of which have large white populations—and is lower than that of all-white Iceland. This suggests that (1) the level of fertility control among U.S. nonwhites at risk of illegitimacy may be higher than that of other populations in which coital activity is also common among the unmarried, and (2) U.S. nonwhites reject an alternative to legal marriage widely adopted by other depressed populations—that of consensual marriage. The percentage of U.S. nonwhite women in consensual marriages is no greater, and perhaps smaller than that of the Nordic population of Iceland—3 to 4 percent of nonwhites compared to 5.6 percent of Iceland's in 1960, (Cutright 1970: Table 1.12, and Beasley and Frankowski, 1970: Table 5). Since consensual or common law marriages are recognized as legal in a number of U.S. States, the few nonwhites in such living arrangements should, in practice, register births as legitimate, although no information on this subject exists.

III. Trends in U.S. Illegitimacy: 1920–68

Trends in Numbers of Births and the Illegitimacy Ratio

Table 2 shows the trend in the number and the ratio of illegitimate births, by color, from 1920 through 1968. These measures of illegitimacy differ from the figures published over the years by various agencies within the U.S. Government. First, the data shown here are corrected for underregistration of births. Therefore, the numbers of illegitimate births are higher than estimates previously published. Second, as a result of the correction for underregistration of births, the national illegitimacy ratio—the number of illegitimate per 1,000 total births—is higher than previosuly published estimates because adjustment for underregistration of white and nonwhite births adds relatively more illegitimate than legitimate births to the adjusted national birth estimates.

TABLE 2.—*Estimated number of illegitimate births and illegitimacy ratios: United States, 1920–68*

Year	Illegitimate births			Ratio per 1,000 births		
	Total	White	Nonwhite	Total	White	Nonwhite
1920	86,365	38,490	47,875	29.3	15.0	125.0
1930	90,800	42,296	48,504	34.7	18.6	141.1
1940	102,996	43,473	59,523	40.3	19.8	166.4
1945	128,190	58,670	69,520	44.8	23.6	179.3
1950	148,372	54,353	94,019	40.9	17.5	179.5
1955	189,733	64,812	124,921	46.2	18.6	202.4
1960	230,428	83,333	147,095	53.4	22.9	215.8
1965	297,055	124,196	172,859	78.6	39.7	263.2
1968	343,815	155,200	188,615	97.8	53.3	312.0

Source: P. Cutright 1972c: table 1.

The total number of illegitimate births was about 86,000 in 1920. From this low point the number rose to 103,000 in 1940, 148,000 in 1950, 230,000 in 1960, and 344,000 in 1968. Projections by the National Center for Health Statistics (1968) of the expected number of illegitimate births for 1980 are 403,000, under the assumption that the 1965 rate will continue, and marital status of women aged 15–44 does not change.

With few exceptions the trend in the number of illegitimate births has been accompanied by a rising illegitimacy ratio for the Nation and for the white and nonwhite population. From a low of 15 illegitimate per 1,000 total births in 1920, the white ratio rose to 53 in 1968. The illegitimacy ratio for nonwhites was 125 in 1920 and it also rose over the years reaching 313 in 1968.[1]

Although the illegitimacy ratio is frequently used as a measure of the level or changes in the level of illegitimacy, it is a measure of doubtful utility, because it is heavily influenced by marital births and by marital status of the population. Changes in marital fertility and/or changes in the marital status of women 15–44 will affect the illegiti-

[1] Although the proportion of births that are illegitimate has increased, and is much larger in the nonwhite than white population, demographic analysis finds virtually no impact of illegitimacy on the completed fertility of whites and nonwhites. Nor is the racial difference in fertility increased by the racial difference in illegitimacy. Finally, illegitimacy had no significant impact on the growth of the white, nonwhite, or total population of the United States over recent decades. (Cutright, 1973a.)

macy ratio. These problems are discussed in the appendix of Cutright (1972c). For analytical purposes the number of illegitimate births per 1,000 unmarried women of childbearing age—the illegitimacy rate—is preferred, because it is not directly affected by changes in marital fertility or marital status.

Trends in Illegitimacy Rates

Table 3 shows the trend in illegitimacy rates, by age and color—where available data permit age-specific calculations. These rates differ from those published in the U.S. Government reports (National Center for Health Statistics, 1968) because the number of illegitimate births is larger after births were corrected for underregistration and the denominators of the rates were adjusted to account for under-enumeration of the population of unmarried women by Census (Siegel, 1968). The effect of the correction for underregistration of births diminishes as registration improves, and becomes negligible around 1960. The effect of correcting for underenumeration o the population, however, remains rather steady over time. Adjustment for underenumeration always has a much larger effect on nonwhite than white rates, because nonwhites are more likely to be missed by Census than are whites. When this adjustment is not made nonwhite rates are inflated because the count of women at risk of an illegitimate birth is artificially low.

TABLE 3.—*Adjusted age-specific illegitimacy rates, all women and by color: United States, 1920–68*

Year and color	Age of mother						Age standardized, 15 to 44
	15 to 19	20 to 24	25 to 29	30 to 34	35 to 44	15 to 44	
All women:							
1920						8.7	
1930						7.8	
1940	8.7	10.6	8.1	5.8	2.8	8.0	7.9
1945	10.1	16.2	13.0	7.7	3.3	10.5	10.1
1950	13.6	21.7	20.5	13.7	4.7	14.5	14.1
1955	16.1	33.6	33.9	22.4	6.4	19.5	19.3
1960	16.4	39.9	41.7	27.8	8.8	21.7	21.7
1965	17.5	39.3	48.4	37.2	10.3	23.4	23.7
1968	20.7	36.4	37.6	28.0	8.6	24.1	23.2
White:							
1920						4.4	
1930						4.3	
1940	3.6	6.0	4.3	2.6	1.3	3.9	3.6
1945	4.4	10.1	7.6	3.9	1.5	5.5	5.1
1950	5.3	10.0	8.8	6.0	2.1	6.1	5.8
1955	6.2	14.9	13.3	8.7	2.8	7.9	7.7
1960	6.9	18.5	17.1	10.8	3.9	9.3	9.3
1965	8.0	21.7	23.8	16.4	4.9	11.5	11.4
1968	9.9	22.6	21.5	14.9	4.7	13.0	12.4
Nonwhite:							
1920						41.5	
1930						31.6	
1940	48.4	52.2	36.7	26.5	10.8	39.1	39.7
1945	51.5	64.8	49.1	31.8	13.0	45.4	45.3
1950	69.8	103.5	92.4	62.6	20.0	68.9	69.1
1955	77.6	127.4	120.5	98.2	24.4	83.2	83.5
1960	78.5	147.1	137.4	97.3	31.9	90.2	90.2
1965	79.6	142.2	153.3	129.3	37.8	94.4	95.3
1968	86.5	109.0	96.6	75.1	24.2	83.0	80.2

Source: P. Cutright 1972c: table 2. Age standardization used the distribution of unmarried women to age groups in 1960 as the standard population.

Looking first at the rate for both white and nonwhite women combined, we find a slight decline in illegitimacy from 1920 through 1940. From 1940–65, illegitimacy rates for each age group increased; the age-standardized rate in 1965 was triple the rate of 1940. Between 1965 and 1968, the age-standardized rate declined slightly, in spite of a further gain in the teenage illegitimacy rate. The 1965–68 decline in the age-standardized rate was the result of declines in each age group 20 and older. In general, the trend using all women also applies to both the white and the nonwhite populations. Both populations participated in the upward trend from 1940 to 1965, and the decline among older women between 1965 and 1968. Whether measured by percentage change or by absolute change in rates, older nonwhites experienced a sharper decline than did whites between 1965 and 1968. In percentage terms, the 1965–68 increase among teenagers was larger among whites than nonwhites, but the absolute increase was considerably greater among nonwhite than among white teenagers.

IV. Changes in the Immediate Causes of Illegitimacy: 1940–68

The following sections review changes in the immediate causes of illegitimacy. As with any fertility rate, illegitimacy will be affected by the degree of voluntary or involuntary control over conception and voluntary or involuntary control over gestation. We first review what is known about changes in sexual activity, contraception, and sterility—the controls over conception. Then we discuss changes in spontaneous fetal loss and induced abortion—the controls over gestation. The impact of improved health conditions on illegitimacy rates are estimated. After removing the change in illegitimacy rates caused by improved health, we arrive at estimates of change due to rising sexual activity.

Voluntary Controls Over Conception: Coital Experience

A 1971 national probability sample of teenage unmarried girls (Kantner and Zelnik, 1972; Zelnik and Kantner 1973) provides the only available measures of coital experience among girls at risk of an illegitimate birth. No national data for older unmarried women are available. The 1971 study cannot be compared to earlier Kinsey data (Kinsey et al., 1953; Gebhard et al., 1958) because the Kinsey reports were not drawn from representative samples of the population.

The first column in the first row of the top panel of table 4 refers to the percentage of white girls aged 15 years in 1971 who reported ever having intercourse. The next figure in that row is the percent of single girls aged 15–19 years who reported intercourse by age 15 or earlier. The difference between the two figures may reflect errors in memory by older girls, a trend toward earlier intercourse within the sample, the attrition from the sample of teenage girls who marry and are thus lost to this sample, failure to understand the question, or other factors such as sampling error.

TABLE 4.—*Percentage of never-married women reporting coitus at age 15, age 19, and ages 15–19, by race: United States, 1971*

Race and coitus by age	Age at interview (percent)		
	Age 15	Age 19	Ages 15 to 19
White:			
15	11	(1)	7
19	(1)	40	(1)
15 to 19	(1)	(1)	23
Black:			
15	32	(1)	24
19	(1)	81	(1)
15 to 19	24	(1)	54

[1] Statistic cannot be computed.

Source: Kantner and Zelnik, 1972: Table 1; Zelnik and Kantner, 1972.

The second row in the top panel reports that 40 percent of whites aged 19 at interview claimed to have had coitus at some time, while the third row notes that 23 percent of unmarried whites aged 15–19 reported coitus at some time prior to interview.

Among black teenagers, the percentage with coitus by age 15 and age 19 was 32 and 81 percent, respectively. The percent reporting coitus by age 15 is lower among girls aged 15–19 than among those aged 15 alone; about 54 percent of nonwhite teenagers reported coitus.

At age 15, about three times as many nonwhites as whites have had coitus; using either the reports for girls age 19 at interview, or the data for all teenagers, the percentage of nonwhites with coital experience is about double the figure for whites. These differentials in coital experience do not adequately reflect racial differences in out-of-wedlock conceived births among teenagers. For example, the teenage nonwhite illegitimacy rate is nearly 10 times the white teenage rate. When both illegitimate and legitimated (by marriage) out-of-wedlock conceived births are combined, the nonwhite teenage rate is about four times higher than the white rate (102 births per 1,000 unmarried nonwhite teenagers compared to 27 births per 1,000 unmarried white teenagers during the 1964–66 period—the only available years; Cutright, 1974: Table 1). Thus, the racial difference in coital experience is not large enough to account for racial differences in births conceived by unmarried teenagers.

Voluntary Controls Over Conception: Contraception

Illegitimacy rates may increase if effective use of contraception declines. There is substantial evidence that the rise in illegitimacy after 1940 was not caused by a decline in effective contraception. Table 5 shows the available data from large studies of both whites and blacks from the 1930's through 1971.

TABLE 5.—*Measures of contraceptive use by unmarried teenagers and unwed mothers, by race: 1930, 1960, and 1971*

Race	1971		1960, percent of unwed mothers using every time	1930's, percent of unwed mothers ever using contraception
	Percent always using, aged 15 to 19	Percent of currently pregnant regular user		
White	21	24	10	12
Black	15	17	7	7

Source: Zelnik and Kantner, 1972. Cutright, 1972c: Table 8. See Kantner and Zelnik (1973), for analysis of contraceptive use among unmarried sexually active teenagers.

Among girls 15 to 19 in the 1971 sample 21 percent of white girls who ever had intercourse and 15 percent of similar black girls said they always used contraception. The 6-percent advantage to the white sexually active vanishes when we compare the percent (not reported in table 5) never using—some 15 or 16 percent of both racial groups never used anything. Knowing that a young girl reports always using contraception provides no measure of the effectiveness of her contraceptive effort—a point documented by the second column of the table, where we find that 24 percent of white currently pregnant-out-of-wedlock girls and 17 percent of black girls similarly pregnant report that they were regular users. One might conclude that contraceptive use was higher among the pregnant than among the nonpregnant sexually active teenager. Rather, we suggest that survey data on contraceptive use indicate very low levels of effectiveness among unmarried young girls.

The 1960 and the 1930 data also shown in table 5 refer to unwed mothers of all ages—most of whom are young women. After the birth of the child these unwed mothers were asked about contraceptive use. In the 1960 sample the percent reporting use every time was 10 percent for whites and 7 percent for black unwed mothers; In the 1930's study the percent reporting use at some time was 12 and 7 percent—a figure that would have to be lowered to make it comparable with the 1960 measure.

The 1930's studies also found no difference in pregnancy rates between black women reporting use or no use of contraception. For all practical purposes the level of effective use by blacks before 1940 was zero (Pearl, 1936; Farley, 1970). The only way for effectiveness to go was up, although the 1971 data indicate that little improvement among young unmarried women had occurred. Similarly, the 1971 data show nothing that would indicate notable levels of effective contraception among young white unmarried girls. These observations are supported by other evidence on the efficacy of contraception and trends in use of various methods reviewed elsewhere (Cutright 1972a). We conclude first, that the illegitimacy rate did not increase because of a decline in contraceptive use and second, that the present level of effective contraceptive practice among young unmarried sexually active couples is little greater than it was in the past.

STERILIZATION

Although temporary contraception is still the major method of voluntary conception control among the sexually active, sterilization operations also control conception. Changes since 1920 in the percentage of unmarried women exercising voluntary control over fecundity through contraceptive sterilization must be very very small. Both males and females electing this form of permanent contraception are overwhelmingly married and older. Increases in male and female sterilization among married couples (Phillips, 1971) cannot change the illegitimacy rate.

VOLUNTARY CONTROLS OVER GESTATION

Voluntary fetal loss from legal therapeutic induced abortion is not an important factor in accounting for changes in U.S. illegitimacy rates. Tietze (1968:784) reports a decline in the ratio of therapeutic abortions per 1,000 live births from 5.1 to 3.5 and then to 1.8 for New York City, between the middle 1940's and the early 1960 period. Other national estimates are compatible with both the level and trend in New York City. Since the majority of therapeutic abortions are to white married women, the possible impact of a decline in legal abortion on a rising white or nonwhite illegitimacy rate can be discounted.

Illegal induced abortion

A review of trends in induced illegal abortion among whites and nonwhites since 1940 (Cutright, 1972a: 1972c) using data on maternal death by cause concluded that for whites unmarried women were more likely than married women to resort to illegal abortion; this was also true for nonwhite women, though the difference in utilization of illegal abortion between married and unmarried nonwhites was smaller than in the case of whites.

Since we know that there has been a vast decline in spontaneous abortion, a stable or increasing abortion death ratio (the ratio of maternal death from abortion to maternal death from other causes) from 1940 to 1965 could not have been caused by increases in spontaneous abortion. Unfortunately, the only period for which the abortion death ratio can be calculated by color and marital status is 1949–51.

Abortion death ratios for all white and for all nonwhite women are shown in table 6 from around 1940 to 1965. For both whites and nonwhites, the abortion death ratio for all women was much higher in 1940 than it was 10 years later.

TABLE 6.—*Abortion death ratios by color and marital status: United States, 1940–65*

Year	White		Nonwhite	
	All women	Unmarried only	All women	Unmarried only
1939 to 1941	22.7	98.3	23.5	31.9
1949 to 1951	11.7	50.7	14.6	19.8
1950 to 1953	12.0	52.0	14.6	19.8
1954 to 1957	14.5	62.8	17.8	24.1
1959 to 1961	19.2	83.1	29.4	39.9
1962 to 1965	21.3	92.2	27.9	37.8

Source: Cutright, 1972c: Table 27. Ratios for unmarried women in 1949–51 from direct observation; ratios for unmarried women in other years are estimated from the observed ratio of unmarried to all women abortion death ratios in 1949–51.

This large decline in the abortion death ratio probably indicates a decline in the induction of abortion between 1940 and 1950. This interpretation is consistent with the increase in general fertility rates over this 10-year period. The abortion death ratios increased from 1950 to 1960 and they remained high for nonwhites or increased for whites after 1960. The leveling off of the general fertility rate began in the mid-1950's, and it declined after 1958. This association between trends in abortion death ratios and general fertility rates therefore may provide some support for our use of the abortion death ratio as an indicator of trends in induced abortions for married and unmarried women. For 1951–62 the proportion of maternal deaths from abortion in New York City increased from 26 to 42 percent. Among whites it went from 14 to 25 percent; among nonwhites it moved from 36 to 49 percent (Omran, 1971:505). This trend provides added support for the validity of the measures used here.

Because we have only one point in time for which abortion death ratios may be calculated for unmarried women, a fixed multiplier of any kind will provide a trend in abortion death ratios for unmarried women that will mirror the trend for all women. If one can accept the idea that trends in induced abortion among married women may be accompanied by a similar trend among unmarried women, then the application of a fixed multiplier (see table 6) to estimate the abortion death ratio for unmarried women would be appropriate.

How do these estimates of the trend of induced abortion among the unmarried fit with the trends in illegitimacy rates? For nonwhites, illegitimacy rate increases after 1940 can, in part, be explained in terms of improved health conditions—see tables 8 and 9 below. After 1955 the increase in the nonwhite illegitimacy rate slowed down. It is impossible to say that the decline in the rate of increase after 1955 was caused by increasing induced abortion among nonwhites, rather than by some other change. Still, the likely increase in induced abortion was accompanied by a leveling off of the increase in the nonwhite illegitimacy rate.

Among whites the rate of increase in illegitimacy was slower between 1940 and 1950 than it was after 1950. Some part of the 1940–50 increase in illegitimacy may be due to declines in induced abortion, but the change in the illegitimacy rate was quite small—just 2.2 births per 1,000 women 15–44 over that 10-year period. During the period of rising abortion death ratios from around 1955 to 1965, the white illegitimacy rate also increased—by 3.7 births per 1,000. Small changes over a decade may easily be overinterpreted, since few women are involved. Still, we have the question of why the white rate should increase at all, if induced abortion among whites was increasing at the same time.

From studies of other populations with illegitimacy rates similar to white U.S. levels (Cutright, 1970: ch. 4) we know that rising illegitimacy often accompanies a rising abortion death ratio. Also, for the same post-World War II time periods, illegitimacy rates tend to be stable or decline when the abortion death ratio is also stable or declining. A plausible explanation of why both abortion and illegitimacy increase or are stable or declining at the same time, can be approached by thinking about the pregnancy rate that underlies an illegitimacy rate. When the pregnancy rate increases by, say, 10 per 1,000, every one of these added pregnancies will have to be aborted (or legitimated) if the illegitimacy rate is to remain stable. When pregnancy rates are increasing, the induced abortion rate (whether figured in terms of the rate to a population of women, or the likelihood that a pregnant woman will abort) may increase. But the abortion rate will not gain enough to do more than slow down the rising illegitimacy rate. There must be a net gain in abortion over and above the net gain in the pregnancy rate if an increase in induced abortion is to produce a decline in the illegitimacy rate. Large net gains were not likely when abortion remained illegal.

THE EFFECT OF INDUCED ABORTION ON DIFFERENCES BETWEEN WHITE AND NONWHITE ILLEGITIMACY RATES

With or without a control on age, the abortion death ratio is higher for unmarried whites than for unmarried nonwhites. Since a larger proportion of white unmarried maternal deaths are from abortion, this may indicate a higher use of induced abortion among pregnant unmarried whites than among pregnant unmarried nonwhites. Therefore, one might conclude that some portion of the difference between white and nonwhite illegitimacy rates is due to the greater likelihood that the pregnant unmarried white than the pregnant unmarried nonwhite would have an induced abortion.

This conclusion can be true at the same time that the induced abortion rate per 1,000 unmarried women is higher among nonwhites than whites. Assume two unmarried female populations aged 15–44, one with a pregnancy rate of 40 and the other with a pregnancy rate of 140. If 50 percent of the pregnancies in the first population are voluntarily aborted, while only 25 percent are aborted in the second, the induced abortion rate per 1,000 women in the first population will be 20, while it will be 35 in the second. Because pregnancy rates per 1,000 women in the white and nonwhite unmarried populations are so different (Cutright, 1972c: table 31 estimates 39 and 183 per 1,000 unmarried women aged 15–44 in 1964–66 respectively for whites and nonwhites), one cannot conclude that the absolute effect of induced abortion on depressing the illegitimacy rate is less for nonwhites than for whites, in spite of the conclusion that pregnant unmarried whites are more likely than pregnant unmarried nonwhites to abort.

Finally, under legal abortion in New York City 1971 data (Tietze, 1973) show a legal abortion rate per 1,000 women 15–44 at 32 among whites and 72 among nonwhites.

Involuntary Controls Over Conception

There are two involuntary controls over conception that may change and thus alter the trend of illegitimacy. The first type of involuntary control occurs among the population of unmarried younger women, and pertains to those factors that affect the age at which they will be able to conceive a child if they have sexual intercourse. The second involuntary control occurs among older women who, although past the age of adolescent sterility, never become fecund, or become sterile through no choice of their own.

CHANGES IN INVOLUNTARY STERILITY AMONG YOUNGER WOMEN

Although a decline in the age of menarche (age at first menses) among Western populations has been documented for many years, the possible effects of such changes on illegitimacy rates among younger women are rarely discussed. It is worthwhile investigating the decline in the age of menarche for its possible effects on illegitimacy rates among young women because such a decline should affect two factors that may, in turn, increase illegitimacy. First, a decline in the age at menarche will increase the percent of women fecund (able to conceive) at a given age. Second, a decline in the age of menarche may also tend to increase the percent of women in a given age group having sexual intercourse. Data are available to allow us to construct some limits on which age groups could plausibly have their illegitimacy rates affected by a decline of adolescent sterility.

THE TREND IN AGE AT MENARCHE

Many writers (Novak, 1921; Gould and Gould, 1932; Mills, 1937; Tanner, 1968; Damon *et al*, 1969; Zacharias *et al*, 1969) have documented and attempted to explain the long-run decline in the age of menarche in Western populations. There is general agreement that one major factor responsible for the decline is improved nutrition and health during preadolescent years. Recent work suggests that improved health conditions increase the rate of physical growth which, in turn, decreases the age of menarche (Frish and Revelle, 1969 and 1970).

A large study in England (Wrey, 1971) found no class differences among English girls in the 1950's, although studies in earlier years reported later age of menarche among lower class girls. Among broad social classes it may well be that in modern industrial nations only small differences if any in age at menarche exist. The 1965 national fertility study found no difference in age at menarche between white and Negro women born after 1910 (Ryder and Westoff, 1971).

It is impossible to state, with certainty, the extent to which age at menarche changed from around 1940 through 1968 in the United States.

One large sample of the U.S. population that may also be fairly representative of income groups provides an estimate of age at menarche for the population of white women who were around age 15 in 1960 (Zacharias et al, 1970). No study of comparable size and representativeness for earlier years exists. However, the various small studies of white women between 1920 and 1940 yield a consistent pattern; women in earlier decades achieved menarche at a later age than did women around 1960. The available data suggest that teenage girls in 1940 had a mean age of menarche around 13.5 years, while teenage girls in the 1960's had a mean age at menarche of 12.5 years.

ADOLESCENT STERILITY

Virtually no woman is fecund before first menses, and most women do not become fully fecund for some time after first menses. Recent data indicate that the period of partial sterility following menses may be around 2.5 years. This period is suggested by the finding that the difference between the mean age at menarche and "regular menses" is about 2 years 3 months (Zacharias et al., 1970: table II).

ESTIMATING CHANGES IN FECUNDITY FOR YOUNG WOMEN: 1940-68

The impact of a 1-year decline in the age of menarche on fecundity has been reported by Cutright (1972a). A decline in age at menarche will result in a net increase in the percentage fully fecund at a given age. For example, at age 15.5 years and assuming a 2.5-year period of sterility, about 69 percent of girls were fully fecund in 1968 compared to just 37 percent in 1940—a net gain of 32 percent.

If we arbitrarily take an increase of 15 percent fully fecund as a change that might affect illegitimacy rates, with the hypothesized 1-year decline in the age of menarche from 1940 to 1968, and the 2½-year period of sterility, then pregnancies among women below the age of 14 would not be affected. Therefore, illegitimacy rates of women 14 and under should not increase much due to increasing fecundity.

Under the assumption of a 2½-year sterile period, pregnancy rates among unmarried women at ages 17 or 17.5 years and older should not be greatly affected by 1-year decline in the age at menarche. Therefore, any sizable increases in illegitimate births to women 18 or older could not be due primarily to increasing fecundity caused by the decline in the age at menarche. On the other hand, illegitimacy rates at ages 15, 16, and perhaps 17 might increase because the fecundity of women at these ages has increased substantially.

CHANGES IN INVOLUNTARY FIRST CHILD STERILITY AMONG OLDER WOMEN: DISEASES AND GENERAL HEALTH

Changes in involuntary first child sterility have been measured by Farley (1970: 109-111). Using census data Farley argues that since very few women who marry wish to remain childless, changes in the percent of ever-married women who reach age 45 without bearing a child can be used as a measure of change in involuntary sterility.

Trends in childlessness among ever-married women show that only 6 to 8 percent of ever-married white and nonwhite women attaining age 45 in the 1880's were childless. Thereafter the childless percentage among nonwhites increased more rapidly than it did among whites, and this measure of involuntary sterility reached its peak among women completing childbearing in the 1950's. The early rise in childlessness, and its decline after 1940 was due to the rise and eventual decline of untreated venereal and other diseases. Improved living conditions after 1940 may also have contributed to the decline of childlessness among women entering their childbearing years after 1940. (See Grove and Hetzel, 1968 and Farley, 1970, ch. 9 for trends in diseases related to childlessness.)

The association of the childless percentage with illegitimacy rates and bridal pregnancy from 1920-39 is quite striking. The level of childlessness among cohorts reaching age 20 from 1920 to 1939 remained both high and stable—while illegitimacy rates and bridal pregnancy among whites and nonwhites were also low and also relatively stable. The decline in involuntary sterility after about 1940 was accompanied by a rise in white and nonwhite illegitimacy and bridal pregnancy. (Cutright 1972c: tables 9 and 10.) Because the decline in involuntary sterility among nonwhites was greater than it was among whites entering their childbearing years after 1930, the impact of this decline on illegitimacy should be larger for nonwhites than for whites. Numerical estimates of the effect of the decline of involuntary sterility on illegitimacy rates are discussed in a later section.

EFFECT OF INVOLUNTARY FECUNDITY LOSS DUE TO AGE: 1960

Childlessness does not measure involuntary fecundity losses that are related to aging in a "healthy" population. The impact of fecundity loss largely due to age, rather than to diseases related to poor public health or severe poverty, can be assumed to be constant over time. While consideration of normal loss of fecundity with age does not explain trends in illegitimacy rates, it does help understand the impact of this factor on the level of illegitimacy in a healthy population, and also the different levels of illegitimacy between younger and older age groups in a given year.

In table 7 we see that the fecund percentage peaks in the 20–24 age group, and then gradually declines. In the second column we give the illegitimacy rates for 1960. The third column shows the expected illegitimacy rate assuming that all women had been fecund, and assuming that sexual activity, contraception, abortion, and legitimation of out-of-wedlock-conceived births were no different among the fecund and the subfecund. By mathematically transforming all subfecund women into fecund women, we can, under the above assumptions, estimate the effect of fecundity loss on the illegitimacy rate of each age-color group. This effect is shown in the last column of table 7.

TABLE 7.—*Percentage fecund, observed, and "fecund only" illegitimacy rates by age and color: United States, 1960*

Age and color	Percent fecund	Observed illegitimacy rate, 1960	Fecund only rate	Effect of fecundity less than 100 percent
White:				
15 to 19	76.8	6.9	9.0	−2.1
20 to 24	87.5	18.5	21.1	−2.6
25 to 29	79.0	17.1	21.6	−4.5
30 to 34	64.0	10.8	16.8	−6.0
35 to 39	53.0	6.2	11.7	−5.5
40 to 44	35.0	1.9	5.4	−3.5
15 to 44	73.0	9.3	12.7	−3.4
Nonwhite:				
15 to 19	78.6	78.5	99.9	−21.4
20 to 24	87.5	147.1	168.1	−21.0
25 to 29	79.0	137.4	173.9	−36.5
30 to 34	64.0	97.3	152.0	−54.7
35 to 39	53.0	50.2	94.7	−44.5
40 to 44	35.0	13.7	39.1	−25.4
15 to 44	73.0	90.2	123.6	−32.4

Source: P. Cutright, 1972c: Table 11.

Among whites and nonwhites the largest absolute effect of fecundity loss occurs among women 30–34. At all age levels the absolute effect of fecundity loss is much smaller on white than on nonwhite rates. However, because fecundity loss with age is the same in both populations, the percentage effect of fecundity loss on the illegitimacy rate of each population is the same. For each group the "fecund only" rate is

37 percent higher than the observed rate. Thus, the effect of an estimate of the total impact of fecundity loss, under 1930 health conditions, is to reduce illegitimacy rates by 37 percent.

A second useful result of computing illegitimacy rates under the assumption that all women in all age groups are fecund is to test alternative hypotheses about why illegitimacy rates among women 30 and older are lower than the rates of women 20–29. One explanation is that the older the women the less likely they are to be able to conceive. While true, this factor has been removed from the "fecund only" rate: we see that the "fecund only" illegitimacy rate for white and nonwhite women 30–34 is lower than the "fecund only" rate at ages 25–29; the fecundity-adjusted rates for women 35–39 are below those at ages 30–34, while the adjusted rates to women 40–44 are lower still. Therefore, fecundity loss does not account for the steady decline in observed illegitimacy rates above age 29.

An alternative explanation of the decline may be that sexual activity decreases. This hypothesis, however, is not supported by the only data pertaining to the issue (Kinsey et al., 1953: tables 76 and 168). The Kinsey figures suggest that the likelihood of illicit coitus among both never married and previously married women increases from age 15 through age 29 and, with the exception of previously married women, does not decline through age 30–45. At all age levels, the active incidence of coitus (percentage exposed to risk) is higher among previously married than among never married women. Also, the weekly frequency of coitus among sexually active women is higher among those previously married than among those never married. Because the proportion of previously married among the total unmarried increases with age, it is clear that the decline of illegitimate birth rates at age 30 and above cannot be ascribed to a decline in coital activity after age 30.

The decline in illegitimacy rates among older age groups cannot be explained by an increase in legitimation of births conceived out of wedlock, since the percentage of out of wedlock conceived births legitimated by marriage declines with age (Cutright, 1972c, table 23). We have now eliminated fecundity loss, declining sexual activity and increasing legitimation as explanations for declining illegitimacy rates after age 29 with increases in age. We conclude that the declining illegitimacy rate across older white age groups is largely caused by spontaneous or induced abortion and/or contraception.

Estimated fetal loss at all gestation periods

Estimated fetal loss per 1,000 pregnancies by age and color, are available for a population of married women enrolled in the Health Insurance Program (HIP) of Greater New York—a prepaid medical care program. Suspected cases of induced abortion have been removed (Shapiro, *et al.*, 1971). About 187 white and 315 nonwhite spontaneous fetal deaths per 1,000 pregnancies of 4 or later weeks' gestation are estimated. These estimates understate total spontaneous fetal loss because no fetal loss occurring before the fourth week is included. Erhardt (1963), for example, has estimated that of every 1,000 pregnancies, 112 abort *before* the beginning of the fourth week of gestation. James (1970) estimates a third of fertilized ova are lost before the first month.

ESTIMATING CHANGING SPONTANEOUS FETAL LOSS, 1940–60

Estimates of change in fetal loss of 4 or more weeks' gestation from 1940 to 1960 may be made by first calculating the true level of fetal loss in 1960 (Shapiro, *et al.*, 1971) per 1,000 live births, by color. These data are then adjusted to account for the higher level of fetal loss among unmarried than married white and nonwhite women in 1960. The ratio of the true level of fetal loss, by color and marital status, in 1960, to the registered level of late fetal loss is computed for 1960. The multiplier for unmarried women is then applied to registered late fetal loss to unmarried women, by color, in earlier years. This calculation yields the estimated true level of fetal loss of 4 or more weeks' gestation in 1940. The difference between the estimated true level in 1940 and 1960 represents the change in spontaneous fetal loss over time. Declining spontaneous fetal loss after 1960 is not taken into account, an omission that places a conservative bias on our estimates of declining fetal loss over time because we actually apply the data to the 1940–68 period. Our procedures yield an estimated decline of 316 spontaneous fetal deaths per 1,000 white illegitimate births, and a decline of 639 spontaneous fetal deaths per 1,000 nonwhite illegitimate births between 1940 and 1960. (See Cutright, 1972b and 1972c for details.)

The Impact of Improved Health on Illegitimacy Rates

We can measure the likely impact of improved health on the 1940–68 change in illegitimacy rates by asking this question: What would the 1940 illegitimacy rate have been had 1940 health conditions been equal to those of 1968? If we find that the 1940 rate to be expected under 1960's health conditions would be equal to the 1968 illegitimacy rate, then one could readily claim that all the increase in the illegitimacy rate between 1940 and 1968 was caused by improved health effects on involuntary fertility controls. If only 50 percent of the observed 1940–68 change in rates is a function of improved health, then the remaining change might be allocated to increased sexual activity, or possibly errors in the measurement of the effects of improved health.

TABLE 8.—*Effect of 1940 health conditions on depressing 1940 illegitimacy rate, by age and color*

Color	Age	Health condition: Higher spontaneous fetal loss	Lower teenage fecundity	Higher first child sterility	Total
Nonwhite	15 to 19	−23.2	−9.6	−4.6	−37.4
White	15 to 19	−.8	−.6	−.2	−1.6
Nonwhite	15 to 44	−24.9	−3.9	−10.8	−38.6
White	15 to 44	−1.2	−.2	−.3	−1.7

Source: Data for girls 15 to 19 from Cutright 1972a: table 3. Here we have added the decline in first child sterility. Change in illegitimacy rates to girls 15 to 19 excludes births to girls 14 and younger. See Cutright 1972a: table 1 for these exact rates. Data for women 15 to 44 from Cutright, 1972c: table 10.

Table 8 shows the probable impact of poorer health conditions in 1940 as compared with those in the 1960's on the 1940 illegitimacy rate of teenagers, and of women aged 15–44. Among nonwhite teenagers, for example, higher levels of spontaneous fetal loss suppressed the 1940 rate by 23.2 births per 1,000; lower teenage fecundity (due to a higher age at menarche) suppressed the rate by another 9.6 births per 1,000, while higher first child sterility had an additional impact of 4.6 fewer births per 1,000 unmarried nonwhite teenagers. The total impact of all health conditions in this group is −37.4 births per 1,000 in 1940.

Among white teenagers, identical procedures do not result in health effects comparable to those among nonwhites. The total impact on the 1940 white teenage rate was to depress it by 1.6 births per 1,000.

The illegitimacy rate of women 15–44 shows effects similar to those indicated for teenagers. Among nonwhites, the 15–44 rates was depressed by 38.6 births per 1,000 in 1940; among whites, the 1940 rate of women 15–44 was depressed by 1.7 births per 1,000.

The combined health effects shown in table 8 can be compared to the observed 1940–1968 change in the appropriate illegitimacy rates. This step allows calculations of the percent of the observed change that may be allocated to improved health rather than other causes. Table 9 shows the 1940–1968 observed change in illegitimacy rates, the change due to improved health conditions, and the change due to other causes. Among nonwhite teenagers, 102 percent of the change is related to health conditions—a finding indicating that changes in other immediate causes (e.g., sexual activity), may have declined.

TABLE 9.—*Change in 1940–68 illegitimacy rates related to improved health and other causes: United States, by age and color*

Color	Age	Total observed change	Health related change[1]	Other causes	Percentage of change related to health
Nonwhite	15 to 19[2]	36.5	37.4	−0.9	102
White	15 to 19	6.2	1.6	4.6	26
Nonwhite	15 to 44[3]	43.9	38.6	5.3	88
White	15 to 44	9.1	1.7	7.4	19

[1] From table 8 above.
[2] 15 to 19 rates exclude births to mothers age 14 and under, see Cutright 1972a, table 1.
[3] 15 to 44 rates from table 3, above.

Among teenage whites, however, only 26 percent of the observed change is related to improved health: the change in the white illegitimacy rate not related to improved health amounts to 4.6 births per 1,000 girls 15–19.

Among women aged 15–44, improved health accounted for 88 percent of the increase in the nonwhite rate, and 19 percent of the white increase. The change in the illegitimacy rate of unmarried women 15–44 not related to improved health is 5.3 and 7.4 births per 1,000 among nonwhites and whites, respectively.

If *all* of the increase in illegitimacy rates not allocated to improved health is a function of increasing coital activity, the minimum estimate of the increase in the percent of women having coitus is found by converting the change in the rate per 1,000 due to "other causes" (in table 9) to percentage terms. This yields an estimate of a minimum increase in the percent with coitus of around one-half of 1 percent for 3 of the 4 groups. Since not all women having coitus become pregnant, and some of those who do become pregnant fail to carry to term, this minimum estimate must be multiplied by an estimate of the percent of the sexually active who carry to term. A reasonable estimate may be 20 percent (see Cutright, 1972c, for details) although this will differ for women in groups with varying rates of fetal loss, contraception use, and coital activity. If this multiplier is doubled (to 40 percent), the alleged increase in the proportion of sexually active would be estimated at only 5 to 7 percent.

V. Economic Variables and Illegitimacy

Economic Status of Unwed Mothers

Although a national survey of the economic status of unwed mothers at the time of birth does not exist, it is possible to provide rough estimates. We then can calculate approximate illegitimacy rates by color and economic status, and thus test the hypothesis that differences in illegitimacy rates among color groups are a function of the economic status of these subsets of the U.S. unmarried female population.

A study of 1967 births in California for which medical expenses were paid from public funds provides the first set of data. Berkov (1971: Table 4) reports that 52 percent of white illegitimate—but only 10 percent of white legitimate births—had medical costs paid from public funds. Among blacks, 76 percent of illegitimate and 40 percent of legitimate births were paid for from public funds. Because public financing of births is means-tested and restricted to the low-income population, it is clear that unwed mothers are more likely than wed mothers to be poor. A second interesting pattern is that among both color groups, the younger unwed mother is about as likely as her older counterpart to be poor.

These California data may understate the percentage of unwed mothers in poverty. First, it is unlikely that every poor unwed mother delivering in California in 1967 had her birth paid from public funds, while it is improbable that any appreciable number of nonpoor unwed mothers had their births paid from public funds. Second, the economic status of California is well above that of the rest of the Nation, and one might expect that a smaller proportion of California's unwed mothers would be poor than is the case in the Nation as a whole. Third, eligibility levels for publicly subsidized births may be lower than the near-poverty line.

Thus, we would expect that a correct estimate of the poor among the Nation's unwed mothers would be somewhat higher. Campbell's estimates for 1960–65 births shown in table 10 lack a direct empirical foundation but appear, nonetheless, to be quite plausible.

TABLE 10.—*Distribution of births by legitimacy status, color, and poverty status: United States, 1960–65*

Legitimacy status and color	Poverty status: Poor and near poor	Poverty status: Not poor	Total
Total births	29	71	100
White	23	77	100
Nonwhite	65	35	100
Legitimate births	26	74	100
White	21	79	100
Nonwhite	59	41	100
Illegitimate births	74	26	100
White	62	38	100
Nonwhite	82	18	100

Source: Arthur Campbell, 1968, derived from tables A-1 and B-1, using median estimates of illegitimate births.

In the following analysis we estimated (regardless of age of the mother), that 60 percent of white and 80 percent of nonwhite illegitimate children born in the 1964–1966 period were to women below the near-poor poverty line. From the 1967 Survey of Economic Opportunity we tabulated the number of unmarried women, by age, color and 1966 poverty status. The percentage of unmarried women, who were low-income or nonpoor, when applied to our estimates of the number unmarried in 1965 provides the appropriate denominator from which illegitimacy rates can be estimated for each racial, age, and economic group. Table 11 shows the estimated illegitimacy rates by poverty status, age, and color for the years 1964–66.

TABLE 11.—*Estimated annual average illegitimacy rates by color, age, and poverty status, United States, 1964–66.*

[Illegitimacy rate per 1,000 unmarried]

Color and age	Poverty status: Poor and near poor	Poverty status: Not poor	Total
White:			
15 to 19	33. 5	3. 6	7. 8
20 to 24	97. 0	9. 9	21. 5
25 to 29	70. 4	11. 7	23. 4
30 to 34	32. 3	9. 1	16. 1
35 to 44	12. 0	2. 5	4. 8
15 to 44	42. 1	5. 4	11. 3
Nonwhite:			
15 to 19	101. 9	33. 6	72. 4
20 to 24	223. 1	54. 0	137. 3
25 to 29	258. 2	53. 8	146. 2
30 to 34	183. 0	52. 2	121. 9
35 to 44	46. 3	17. 7	34. 9
15 to 44	128. 8	39. 0	88. 2

NOTE.—If illegitimacy was a cause of poverty status the rates to the poor would be biased upward. Illegitimacy should not affect the poverty status estimates for mothers having a first birth, because poverty status is estimated prior to that birth. A first illegitimate child will not affect the poverty status of higher birth order unwed mothers if they have adopted or released the first or prior child. Prior illegitimacy will not affect poverty status of most mothers already below poverty at the time of the first birth. Only a small number of women nonpoor at first birth who retain the child and are thus moved below the poverty line prior to a second or higher order birth could be counted as mothers whose illegitimacy affected their poverty status. Thus, the direction of possible causality is from poverty status to illegitimacy, not from illegitimacy to poverty status. The negligible effects of illegitimacy on poverty status some years after the first birth are reviewed in Cutright, 1973b.

Source: P. Cutright 1972c: table 30.

Among whites in all age groups, poor and near-poor women are more likely than nonpoor women to have an illegitimate child. For example, the illegitimacy rate for low-income white teenagers is over 33 per 1,000 while the rate for the nonpoor is less than 4. The poor white teenager is nearly 10 times as likely as the nonpoor white teenager to have an illegitimate child. A similar differential exists among whites aged 20–24. Among whites 25 and older there is somewhat less difference in the illegitimacy rates of poor and nonpoor women, but it is still quite substantial. Among whites of all age groups, the illegitimacy rate among the poverty group was nearly eight times the rate among the nonpoor.

Among nonwhites, illegitimacy rates of the poverty group are about triple those of the nonpoor. This difference is fairly stable across age groups.

From data on the economic status of unmarried women from the 1967 Survey of Economic Opportunity we found that 16 percent of white unmarried women aged 15–44 and nearly 55 percent of similar nonwhites were below the near-poverty line. Because the illegitimacy rates of both whites and nonwhites vary by economic status, it is clear that some part of the white-nonwhite difference may be related to the higher rate of poverty among nonwhites. We can take the nonwhite population in 1965 and ask, What would the illegitimacy rate of the nonwhite population be had unmarried nonwhite women been no more likely than whites to be poor?

To answer this question we first apply the percentage poor among whites (16 percent) to the total number of unmarried nonwhite women aged 15–44 in 1965. We then apply the observed nonwhite illegitimacy rates from table 11 to this hypothetical population of poor and nonpoor nonwhites. Had these rates been maintained, but the population of nonwhites been distributed to economic groups in the same proportion as whites, the resulting nonwhite illegitimacy rate would have been 53.4—rather than the observed 88.2 per 1,000. Under conditions of similar white and nonwhite economic status, instead of the observed annual average number of nonwhite illegitimate births of 161,500, there would have been about 98,000. The expected illegitimacy rate of 53.4 represents a decline of 34.8 from the observed rate of 88.2 Thus, one might conclude that some 40 percent of the observed nonwhite rate is related to factors resulting from the different distribution of white and nonwhite women to economic classes.

If one imagines the host of noneconomic as well as economic changes that would have to occur before the poverty status of the nonwhite and white populations would become equal, then our calculation of a hypothetical nonwhite illegitimacy rate under equal economic conditions certainly must understate that expected rate. With all these qualifications, one can still estimate a minimum "effect of a higher risk of poverty" on the difference between the white and nonwhite illegitimacy rates in the 1964–66 period. The observed color difference is 76.9 per 1,000. Under conditions of equal poverty status this difference would be reduced to 42.1. Thus, about 45 percent of the white-nonwhite illegitimacy rate difference during 1964–66 can be allocated to the higher proportion of the nonwhite population in poverty.

Income, Income Distribution, and Illegitimacy

The temptation to investigate the influences of economic variables on illegitimacy rates has been stoutly resisted by most researchers. Given that some fairly reliable data on illegitimacy and economic status are now available, what kind of speculations may be made? Remembering that the small change in illegitimacy rates we are trying to explain with economic variables is about the same for blacks and whites (after correcting for change due to health), we first ask whether the patterns of changing economic status within and between the white and black population are related changing illegitimacy rates.

TRENDS IN FIXED MEASURES OF INCOME OR POVERTY: 1950–70

If we measure trends in economic status with a fixed poverty line (defined here as the percent of families with less than $3,000 a year, corrected for price changes), we find 20 percent of white families in 1950, but only slightly over 7 percent in 1970 had that little income. The proportion of nonwhite families below this fixed poverty line declined from 50 percent in 1950 to 20 percent in 1970. (Economic Report of the President, 1972: Table B–20.) Both white and nonwhite illegitimacy rates increased over this time period. Thus, there is little reason to believe that illegitimacy will disappear with further declines in the percent of white or nonwhite families living below a fixed poverty line. An alternative measure of changing economic status is median family income. Corrected for price changes the median nonwhite family income was $6,516 in 1970, compared to $3,014 in 1950—an increase of $3,502. Median white family income in 1970 was $10,236, compared to $5,601 in 1950—an increase of $4,635. (Economic Report of the President, 1972: Table B–20.) These large gains in median family income were not accompanied by declining illegitimacy in either population.

MEASURES OF RELATIVE INCOMES

Between 1950 and 1970, nonwhite median family income more than doubled, but the gap between the purchasing power of the median black and white family increased—whites had $2,587 more than nonwhites in 1950; by 1970, the dollar gap had increased to $3,720. Thus, one can conclude that the economic position of both whites and nonwhites has dramatically improved, or that the position of nonwhites compared to whites has become relatively worse. If nonwhites view income in about the same way as do whites, it would be reasonable for nonwhites to view their rising purchasing power in a relative perspective and consider their relative economic position as becoming poorer. If illegitimacy rates are in some way related to feelings about relative economic position, then the relative worsening in nonwhite economic status since 1950 might account for the nonhealth-related increase in nonwhite illegitimacy since 1950. This simple illustration would not, however, explain rising white rates, a problem we will return to shortly, after comment on whether people view income in relative rather than fixed terms.

Since World War II, Gallup polls have repeatedly asked adults to estimate "the smallest amount of money a family of four needs to get along" in the respondent's community. White and blacks in similar communities give similar responses, a finding that should not surprise anyone. What is, perhaps, surprising is that the "smallest amount of money" estimated is constantly going up. The public does not define poverty in terms of a fixed measure of poverty. Rather, the public definition of poverty moves with the trend in median family incomes. If the median family income is $5,000, the smallest amount of money needed to get along will be estimated at around $3,000 by the public; if the median family income is $10,000, the "smallest amount of money" needed to get along will be about $6,000. Whatever the median, the smallest amount of money needed to avoid poverty will be about 60 percent of the median (Rainwater, 1973). This finding is of potential importance in understanding the relationship of poverty, level of income, income distribution, and race to illegitimacy rates.

A measure of relative economic status, suggested by Gallup poll data, would calculate the percentage of families that lived below 60 percent of median family income in each year. This would measure the percent of families defined by the public (and themselves) as being poor. Although this measure is not at hand, a measure of the percent of families with less than 50 percent of median family income is available since 1947 (President's Commission on Income Maintenance Programs, 1969: Table 3–4). Since 1947, the percent of families with less than half the median income has not changed—in each year it is about 20 percent of all families, or one-fifth of all families. The lowest fifth of all families consistently are found to share about 5 percent of all money income in a given year (Ibid.: Table 3–2). The income of the lowest fifth of families has continually fallen further behind the purchasing power of the average family. This is true within the white and nonwhite population (Merriam, 1968). The lowest fifth of whites and nonwhites in 1970 are further behind the median income than was the lowest fifth of whites and nonwhites in 1950. Thus, the percentage of families defined as poor by themselves and others has probably not changed, but their economic position relative to that of average families has become worse.

If illegitimacy rates are, in some way, linked to the relative economic position, then the trend in relative incomes might suggest that both white and nonwhite rates should increase, rather than decline, at least among the daughters of the families in the lower depths of the income distribution. Why the rates would increase (if they have) among girls above a relative poverty line is a question that is probably best answered by abandoning a purely economic theory of illegitimacy. Indeed, given the lack of real evidence to link the recent U.S. trends in income distribution with illegitimacy trends; it would be premature to claim that relative economic position, if changed, would have a large and immediate impact on the illegitimacy rates of the poor.

The Health of Unwed Mothers and Their Children

Given that unwed mothers are so heavily concentrated in the lower income groups, we could expect that before and after their birth they would be less likely than married women to receive health care from physicians or fee-charging clinics. Unfortunately, there has been no national study of the extent to which unmarried women do or do not receive adequate medical service before, during, and after pregnancy. Similarly, no national study of the medical care received by illegitimate children is available. In spite of these difficulties, some scanty but direct information on contact with health services prior to delivery is available. When coupled with indirect indicators, it is possible to reach general conclusions regarding medical care and health of unwed mothers and their children. Kovar (1968a) has reported the number of visits to physicians and medical facilities during the 12 months preceding births occuring in 1963. The average white unwed mother had 7.6 such visits. The average white wed mother with a family income under $3,000 had 10.4 visits, and wed mothers with higher incomes had even more visits. Thus, the unwed mother was less likely than the wed mother in the lowest income interval to seek or receive (or some combination of these events) health care prior to her delivery. Among nonwhite women, a similar pattern emerges—the higher the family income, the more frequent the visits for health care. Also, nonwhite unwed mothers are less likely than wed mothers to receive health care. Further, at the same level of family income, nonwhites are less likely than whites to receive care prior to a birth.

Of the visits to physicians or medical facilities that do occur, Kovar (1968a) reports that the unwed mother is more likely than the wed mother to use medical facilities rather than private physicians. Controlling legitimacy and income, the white is more likely than the nonwhite to visit a private physician. Only 35 percent of visits for medical care among nonwhite unwed mothers were to private physicians, while 64 percent of white unwed mothers' visits were to private physicians. Among married women in the lowest income interval, 47 percent of nonwhite and 71 percent of white visits were to private physicians, while in the $5,000–$6,999 income interval, the figures were 64 and 80 percent for nonwhites and whites, respectively. We have no evidence that the source of the patient's care has any effect on the health of the mother or her child. We cite the statistics on place at which service is received primarily because it documents the fact that money is not the only reason nonwhite women are less likely than whites to have equal health care during pregnancy. These figures are relevant to our later discussion of possible steps that may reduce unwanted illegitimate births.

If it were the case that nonwhite women were less in need of medical services than white women, or that unwed mothers were less in need than married women, then one might argue that no unmet need exists. Clearly this is not the case, since maternal mortality is higher among nonwhites than whites, fetal loss at all gestation periods is higher among nonwhites, and prematurity is higher among nonwhite than white live-born children (Kovar, 1968b: Tables 1, 2, 5, 6).

Differences in U.S. infant mortality by legitimacy and color have recently been measured (National Center for Health Statistics, 1971: Table 1). The differential risk of infant death to legitimate and illegitimate children, and by color, follows a pattern similar to that found for late-registered fetal deaths. For example, the infant mortality rate (IMR) for white illegitimate children is 64 percent above the IMR for legitimate children. Among nonwhites the difference is less—13 percent. These results follow a pattern observed for white and nonwhite legitimate and illegitimate infant deaths in North Carolina for the periods 1957–61 and 1962–66 (Scurletis et al., 1969: Table 2).

We conclude that a relatively high risk of infrequent medical care—a risk above the level that could be predicted from economic status alone—differentiates white unwed from white wed mothers prior to delivery. This effect is less obvious among nonwhites. Since the available indicators of the health of the fetus, the live-born child, and the mother after delivery all indicate poorer health among the unwed and the illegitimate than among the wed and the legitimate, it appears that unmarried pregnant women and their children could use considerably more medical services than they now receive. Utilization of preventive medical care to reduce the risks of illicit pregnancy is also less frequent among the unmarried than the married. (Kantner and Zelnik, 1972.) To remedy this problem subsidized contraception programs have been introduced over the past several years. The current status and likely impact of these programs in reducing illegitimacy is considered in the following section.

VI. The Limits of Contraceptive Programs in Reducing U.S. Illegitimacy [1]

We can assess the likely future of illegitimacy in the United States by considering its immediate causes, the characteristics of unwed mothers, and the limitations of the present subsidized contraception program that now is attempting to reduce unwanted pregnancy among low income women. First, which of the immediate causes of illegitimacy can be changed by public programs?

Immediate causes subject to change.—No one has the vaguest idea how to decrease sexual activity that results in illegitimate births. The proportion of out-of-wedlock conceived births that are legitimated by marriage could only be increased by measures that would force unwilling couples to marry. This is not a desirable goal of public policy. A deliberate reduction of public health programs to increase spontaneous abortion and involuntary sterility is unthinkable. This process of elimination leaves only two remaining immediate causes—voluntary control over conception and voluntary control over gestation—that can be affected by public programs. We first consider the view that contraception-only programs can have a significant impact on illegitimacy, and then briefly consider the likely impact of adding abortion on request to public programs.

Are Illegitimate Births Wanted Births?

Before any voluntary family planning program can effectively reduce illegitimacy it must enlist the cooperative efforts of sexually active unmarried women. If it were the case that most illegitimate births were deliberate, then the prospects for a successful program would indeed be dim.

Both direct and indirect indicators of the wanted status of illegitimate births have been reviewed elsewhere (Cutright 1971a). That review suggested that no more than 10 percent of illegitimate births were deliberate wanted births. Of further interest is the fact that among couples trying to avoid pregnancy in the Bowerman study, 86 percent of whites and 84 percent of Negroes reported that only condom or withdrawal methods were used, while an additional 11

[1] For detail and documentation not included in this section see Cutright, 1974.

percent of whites and 12 percent of Negroes reported male and female methods combined. Only about 3 percent of white and less than 5 percent of black unwed mothers reporting contraceptive use were using exclusively female methods. (Bowerman et al. 1966: 408–409.) It is clear that these unmarried mothers were dependent on the male for protection. The resulting pregnancies therefore can hardly be seen as a deliberate result of the woman's nonuse of contraception; rather they are also the result of nonuse or ineffective use by men. The pattern of male contraceptive use resulting in illicit pregnancies provides no evidence that the women wanted them.

Because it is unlikely that efforts to increase male contraceptive use will work (Cutright 1971a), programs to increase voluntary control over illegitimate conception focus on unmarried women. Effective contraception for women (pill, I.U.D. and diaphragm) is controlled by physicians. What characteristics of unwed mothers will affect the reduction of illegitimacy by a physician-oriented female contraception program? What are the limits of a contraception program in reducing illegitimate pregnancy and births?

Characteristics of Unwed Mothers and Contraceptive Programs

Poverty.—Table 11 demonstrated that low-income unmarried women are much more likely than the nonpoor to become unwed mothers. Studies of fertility control problems among married couples indicate that the low-income wife is twice as likely as others to report an unwanted birth (Bumpass and Westoff 1970: Table 4). Although it is undoubtably true that some of this difference in effective birth planning is related to financial and location barriers that diminish contraceptive effectiveness among the married poor, one may still conclude that effective use of contraception is lower among the poor than the nonpoor using the same methods of contraception. The fact that unwed mothers are concentrated heavily among the poor population means that efforts must be made to overcome this pattern of ineffective use of contraception by poor people. In the case of the unmarried, three other characteristics also indicate further difficulties for a contraception program.

Low birth order.—The previous fertility of unwed mothers differs from that of wed mothers; nearly 73 percent of white and 54 percent of nonwhite unwed mothers in 1968 were having their first child. (U.S. Bureau of the Census, Current Population Reports, Series P–23, No. 36, 1970: Table 28.) About 63 percent of all illegitimate births were first births. This finding is radically different from what one finds in studies of unwanted legitimate births—less than 2 percent are first births. Contraceptive programs to prevent the bulk of unwanted legitimate births can utilize the postpartum period in maternity wards. But postpartum programs cannot possibly reach nearly two-thirds of potential unwed mothers. To date, most of the contraceptive programs that reach a substantial portion of the target population are postpartum programs. These programs can have only a limited impact on illegitimacy rates.

Age.—The third characteristic of unwed mothers that bodes ill for a contraceptive program is their youth. In 1968, for example, 10 percent of white and 18 percent of nonwhite illegitimate births occurred to girls aged 16 or under. (National Center for Health Statistics, unpublished data.) Put another way, nearly 10 percent of white and 18 percent of nonwhite illegitimate births resulted from pregnancies to girls aged 15 and under. Nearly 30 percent of white and 40 percent of nonwhite illegitimate births were to girls who became pregnant before they were 18. Under the laws of most States these minors cannot legally be treated by government programs without parental consent—nor can they legally receive the services of private physicians without it. (Pilpel and Wechsler, 1971.) Even if legal barriers were removed, the youth of this large number of unwed mothers will work against effective contraceptive practice, since they are less likely than older women to be able to cope with moral confusions surrounding sex and contraception.

Coital activity.—The median number of coital acts per year for sexually active unmarried women is about 12–16 (Zelnik and Kantner, 1972: table 7). About half of all sexually active women will have less and half more than the median number. In contrast, the average number of coital acts for married women is about 80 (Westoff and Westoff 1971: table 1). The sexual activity of unmarried women having sex at some time during the year not only is much less frequent, but also tends to be less predictable. The fact that the majority of sexually active unmarried women do not experience regular and frequent intercourse over an extended period of time should depress the self-perceived need for protection. For, unlike married women, many women who will be sexually active at some time during the year are unable to look ahead and see a future of frequent and regular intercourse. However, the group of sexually active women most likely to get pregnant are those having frequent intercourse. This group should be more likely than the less active to perceive the need for contraception and to come into the clinics to get help (if not already pregnant when the need is perceived).

At the present time both the pill and IUD are defined by most women as methods to be used during long-run periods of exposure to risk. For many sexually active unmarried women, actual sexual behavior does not fit this definition, and they cannot be expected to accept this type of contraception. This should, in theory, be true of poor and nonpoor, young and old, low and higher-parity women. Until alternative methods are developed, this segment of sexually active unmarried women are unlikely to receive an acceptable female method of contraception from either public or private sources. In the absence of a breakthrough in technology (and in use when it comes to pass) the distribution of coital acts among sexually active unmarried women suggests that a successful program using currently available methods is likely to attract only 40 percent of low-income women at risk of pregnancy. (See discussion around table 12, below).

In sum, the characteristics of unmarried women most likely to become unwed mothers imply low use of a contraceptive program. We now consider the characteristics of the current family planning program that make it unlikely that it will do a great deal to reduce U.S. illegitimacy.

Family Planning Program Characteristics

At the present time the family planning program is limited to providing services to women below the near-poverty line (Scheyer, 1970). This excludes the 84 percent of white and 46 percent of nonwhite unmarried women aged 15–44 who are *above* the near-poverty line. (Data from the 1967 Survey of Economic Opportunity.) Some 40 percent of white and 20 percent of nonwhite women having illegitimate children are excluded because they fall above the near-poverty line—see table 10, above.

A second program characteristic that restricts reduction of illegitimacy is its limitation to contraception, and inadequate patient care that results in high patient contraceptive failure rates. For example, Tietze and Lewit (1971) report that 24 percent of "pill" patients in one large program were pregnant within 12 months, while another program had a failure rate of only 8 percent. Pill patients under age 20 had an annual failure rate of 27 percent, while patients 30 and older had a rate of just 9 percent. In contrast, IUD failure rates were two to three times lower than pill failure rates for all age groups, and were lower in all programs. Clearly the choice by physicians of whether to push the pill or IUD will have an effect, but other characteristics of the program will affect failure rates as well. Still, even the best programs have failure rates of around 10 percent among their patients.

Contraceptive failure rates among clinic patients should be contrasted with those for married persons before one concludes that programs cannot affect marital fertility or even illegitimate fertility of the poor. The 1965 National Fertility Study, for example (Westoff and Westoff, 1971:69), reports 12-month failure rates among the married population of 4 and 7 percent for pill and IUD users respectively—rates clearly better than the 16-, 18-, and 21-percent failure rates experienced by users of condom, diaphragm and withdrawal, respectively. Failure rates for users of rhythm and foam were 28 and 29 percent. These data illustrate the importance of method—but they also remind us that unmarried women are not alone in having problems in conception control with present contraceptive methods. If contraceptive programs among unmarried women could promote only the level of fertility control of married women, one would still expect considerable numbers of illegitimate pregnancies and births to occur.

Probable Impact of Contraceptive Programs

A systematic assessment of the likely impact of varying levels of patient program participation on reducing illicit pregnancy is given in table 12. In this table we also show three different failure rates in the population of sexually active unmarried women *before* they do or do not become patients. We vary the percent of sexually active women who are patients in the program from 10 to 100 percent, and have alternate annual contraceptive failure rates of 10 and 20 per 100 patients in the program.

TABLE 12.—*Percentage reduction in numbers of illicit pregnancies under alternative failure rates before and after a program: by percent of sexually active unmarried women in the program*

Contraceptive failure rate before program	Percent sexually active women in program	Percent reduction of illicit pregnancies by patient failure rate	
		10 percent contraceptive failure	20 percent contraceptive failure
50	10	−8	−6
	40	−32	−24
	70	−56	−42
	100	−80	−60
30	10	−7	−4
	40	−27	−13
	70	−46	−23
	100	−67	−33
20	10	−5	0
	40	−20	0
	70	−35	0
	100	−50	0

Table 12 shows that a contraception program with a failure rate of 20 could reduce illicit pregnancies by as much as 60 percent, so long as 100 percent of women were in the program and the failure rate had been 50 prior to the beginning of the program. While a failure rate of 50 may be appropriate for the American population of married indigent women, it is too high for the population of sexually active unmarried women, because the coital activity over a year's time of the unmarried is well below that of married women. However, if it were true that half of all unmarried women having one or more coital acts over a 12 month period did become pregnant, the 50 percent failure rate would be appropriate. It can be demonstrated that this is unreasonable (Cutright, 1972c).

If we take the 20 percent preprogram failure rate and see what this rate implies for the percentage of poverty level women who are having coitus, this also leads to unreasonable estimates.

If we see what a failure rate of 30 implies about coital activity among poverty level sexually active women, we would estimate 48 percent of whites and 69 percent of unmarried nonwhites aged 15–44 had coitus at some time during the year. These estimates are consistent with survey reports cited above (table 4). The 30 percent failure rate seems more reasonable than either the 50 or 20 percent failure rates.

Taking this failure rate of 30, we see in table 12 that a program that could reduce the failure rate to 10 percent could, theoretically, reduce illicit pregnancies by 67 percent—if 100 percent of sexually active women were in the program. If the patient failure rate could only be moved to 20 percent, the program could—even with 100 percent of the women in it—only reduce illicit pregnancies by 33 percent. However, implications of the distribution of coital activity among unmarried women yield a more realistic level of program utilization of 40 percent (rather than 100 percent). The 40 percent estimate of program participation may be optimistic, for not only do the majority of potential unwed mothers have infrequent coitus, but they also tend to have never been pregnant, and they are very young and poor. To each of these characteristics associated with ineffective contraceptive use we should add a multiplier effect for the fact of being unmarried. That is, each of the above characteristics has a greater *impact in depressing effective* contraceptive use among the unmarried than among the married.

For the low-income population having sex at some time during a year, table 12 indicates that the present program realistically could expect to reduce unwanted illegitimate pregnancies by some 13 to 27 percent, depending on whether patient failures rates are 20 or 10 percent. Because the program does not reach the nonpoor population, its effects on the total number of illegitimate pregnancies and births will be even less. We conclude that, if unwanted illegitimate pregnancies and births are to be greatly reduced through public programs, the programs will have to be changed. One major barrier to increased impact is the contraceptive failure rate. However, the programs can only prescribe those contraceptives which exist and even with stronger followup efforts to switch patients who have discontinued one method, the problems inherent in contraception-only programs will remain. (Potter, 1971, and Potter and Sakoda, 1966.) For some years to come, all available contraceptive methods will leave large numbers of patients with unwanted pregnancies.

Changing the Program

This assessment is supported by evaluation indicating no impact of the contraception program on illegitimacy in Georgia and Tennessee (Cutright, 1972c). In the immediate future the goal of reducing illegitimacy through direct intervention by public programs is unlikely to succeed unless new methods of birth control become available. If the goal of preventing pregnancy is shifted to include that of preventing unwanted births, the inclusion of abortion as a backstop method for contraceptive failure would lead to a reduction of illegitimate births.

The women least likely to use effective contraception are most likely to use legal abortion when it is available. In New York City, for example, the ratio of legal induced abortions per 1,000 live births is higher among the poor than the nonpoor; it is higher among nonwhites than among whites; it is higher among the never-pregnant than mothers; and it is higher among women under 20 than it is for women 20 and older. In short, abortion is voluntarily used to prevent unwanted births when the woman cannot prevent pregnancy. (Pakter and Nelson, 1971: 6–7; Cutright, 1972c: table 35.) The evidence from other populations as well indicates that in the first few months of legal abortion their number is about equal to that of illegal abortions previously performed. After about a year of legal abortion their number begins to exceed that of previously hidden illegal abortions with a decline in illegitimacy rates resulting. In a period of 6 to 8 years, illegitimacy rates decline by up to 50 percent in populations when no restriction is placed on access to this method of birth control. (Cutright, 1972c: table 36.) There is no evidence that any contraception program anywhere has had a comparable effect. (Furstenburg, et al., 1972.)

The January 22, 1973 Supreme Court decision on abortion will have little impact unless abortion services are readily available across the Nation. From Tietze (1973) the likely short run impact of readily available legal abortion on illegitimacy can be assessed in a crude way.

Prior to legalization the number of illegitimate births to New York City residents had been increasing by about 10 percent each year. In 1971—the first full year after legalization—the number of illegitimate births was 12 percent below the number for 1970—the first decline since 1954, when illegitimate births were first recorded in New York City. This decline is similar to that of other populations the first year after legalization of abortion, and it should continue. Tietze (1973: 39) estimates that about half of pregnancies not spontaneously aborted by unmarried women were terminated by legal abortion in New York City in 1971. The 1970 to 1971 decline in the number of illegitimate births (3,800) when contrasted with the number of legal abortions (42,150) to unmarried women in 1971 indicates that legal abortion has replaced abortions that were previously illegal and not counted—with only a small net gain in the total number of induced abortions in the first year of the new law.

VII. Government Income Support Programs and Illegitimacy

Income Support Programs as a Cause of Illegitimacy

If total program expenditures are the dependent variable, and benefits are directly tied to births, (and if illegitimate children are eligible and receive benefits) illegitimacy will be a cause of Government spending. But one can ask whether the level of benefits affects illegitimacy, thus reversing the direction of causal effects. Critics of "welfare" programs often see their benefits as a cause of illegitimate births, because the economic sanction against the unwed mother has been replaced by what they assume is a reward.

The assumption that illegitimacy rates are dependent on Government programs rests on the belief that the level of support acts to stimulate nonmarital fertility—that an illegitimate birth is perceived as a net benefit by the mother due to an improvement in her economic position caused by the resultant grant. This view assumes that most unwed mothers and their children will have access to government funds, and that the costs to the mother of obtaining them will be relatively low. According to this model, illegitimacy rates should be most responsive to Government programs providing a high benefit level, no means testing or other barriers to access to the program and 100 percent coverage of illegitimate children. Family allowance programs in developed nations do not always meet all three criteria, but several programs fit one or more of the program characteristics that might stimulate illegitimacy.

Family Allowance Benefits and Illegitimacy

One might not expect a tiny grant to stimulate illegitimacy directly. Also, small grants could not be expected to have an indirect effect on illegitimacy by nurturing a pronatalist mood in the population. On the other hand, it might be argued that a large allowance relative to average wages would not only encourage high marital fertility but would also here a "spillover" effect on illegitimacy because such a program would be seen by the population as evidence of pronatalist orientation by Government. Therefore one might expect higher illegitimacy, as well as higher marital fertility.

Table 13 shows the value of allowance for families with three children as a percent of average wages and the illegitimacy rates to the relevant age group (women 30–34 in these populations are the group having a third or higher order birth). For comparison the U.S. rate is included, along with teenage rates in all nations.

TABLE 13.—*1966 family allowances for 3 children as a percent of average monthly wages and illegitimacy rates to women aged 15 to 19 and 30 to 34, circa 1965*

Nation	Allowance as percent of wages	Illegitimacy rates circa 1965, age of mother	
		30 to 34	15 to 19
France	41	30	6
Belgium	34	13	2
Italy	29	10	2
West Germany	20	19	7
Netherlands	15	10	3
Sweden	14	22	28
Canada	7	27	13
United Kingdom	4	44	11
United States	0	37	18

Source: Illegitimacy rates from Cutright 1971b: App. 1; family allowance relative to wages are maximum levels (which are generally close to minimum benefits) and are taken from President's Commission on Income Maintenance Programs (1970), table 2, 1-1: 415.

France pays the highest relative benefit and the United Kingdom the lowest. The illegitimacy rate at age 30 to 34 in the United Kingdom was 44 compared to a rate of 30 in France. Inspection of the rank order of benefits against illegitimacy rates to older women reveals no association, and we conclude that illegitimacy rates of older women are not affected by the relative size of family allowance benefits. There is no "spillover" effect.

Inspection of teenage rates allows a similar conclusion. The rate is 6 in France—with high benefits, and 28 in Sweden with its low relative benefits. Both Canada and the United Kingdom have very low relative benefits and teenage illegitimacy rates two to six times greater than the five nations paying the highest relative benefits. Clearly study of factors other than the relative size of the family allowance program are needed before national differences in illegitimacy rates can be understood.

AFDC and Illegitimacy in the United States

Comparing the number of AFDC families with the illegitimacy rate is misleading because the illegitimacy rate measures one population at risk, while the total count of AFDC families is not specific to the same population. Still, several examples of the absence of the expected relationship between changes in illegitimacy and total AFDC families can be shown. For instance, one finds the illegitimacy rate unchanged from 1930 to 1940, although the number of families on AFDC was zero in 1930 and 372,000 in 1940. The illegitimacy rate increased from 8 to 11 between 1940 and 1945, while the number of AFDC families declined from 372,000 to 274,000. Between 1950 and 1955 the number of AFDC families declined from 651,000 to 602,000, while the illegitimacy rate increased from 15 to 20. The number of families on AFDC increased from 1,054,000 in 1965 to 2.5 million in 1970, while illegitimacy rates did not change. Clearly, then, there are numerous examples of time periods when illegitimacy rates increased while AFDC rolls were stable or in decline; also, we have examples of periods in which illegitimacy rates were stable while the number of AFDC families was rapidly increasing. If one used increases in number of client families over short-run periods as a measure of change in access to and utilization of AFDC by eligible female-headed families, the historical record will provide data that would either support or disprove the hypothesis that increased access will increase illegitimacy rates.

TABLE 14.—*Number of illegitimate children 17 and under on AFDC in 1961 and 1969, by color*

[Numbers in thousands]

	1961			1969		
	White	Nonwhite	Total	White	Nonwhite	Total
Number surviving	1,111	1,734	2,845	1,522	2,281	3,803
Number after adoption and legitimation	333	1,438	1,771	457	1,939	2,396
Number on AFDC	179	470	649	[1] 397	[1] 1,171	1,568
Percent on AFDC	54	33	37	[1] 87	[1] 60	65

[1] Estimate.

NOTE.—Data on survivors estimated by assuming 97 percent of white and 94 percent of nonwhite births during the 17 preceding years survived. Of these survivors, 70 percent of white and 15 percent of nonwhite children were estimated to have been either adopted or legitimated by the marriage of their natural parents, thus removing them from the population of survivors at risk of being counted on AFDC roles as illegitimate.

Source: P. Cutright 1972c: table 24.

THE EFFECT OF ILLEGITIMACY ON AFDC: 1961–69

Table 14 allows us to test, although somewhat crudely, some alternative explanations of the rise in the numbers of illegitimate children on AFDC between 1961 and 1969. In 1961 there were 649,000 illegitimate children on AFDC: The number rose to about 1.6 million in 1969. Of the increase of 919,000, about 26 percent is due simply to the increase in the number of illegitimate children under 18 years of age between 1961 and 1969, and 74 percent is due to higher utilization of AFDC in 1969.

In May 1969, there were 454,900 AFDC families in which the father was absent and not married to the mothers. All 1.1 million children in this type of family were illegitimate—the average number of children was 2.4 per family. Another 478,000 illegitimate AFDC children were in other types of families. A minimum of about 2 million AFDC recipients therefore were illegitimate children or unmarried mothers. In families with the father not married to the mother, the monthly AFDC benefit per recipient in May 1969 was about $45.16.

Using this figure to calculate the annual cost of AFDC benefits to unwed mothers and illegitimate children yields an estimate of about $1.1 billion in 1969. This was about one third the total cost of the AFDC program in 1969.

THE EFFECT OF AFDC ON ILLEGITIMACY: 1940–70

To find that illegitimacy affects the cost of the AFDC program does not mean that the program is a cause of illegitimacy. Although explaining the rise in illegitimacy since 1940 by the AFDC program is quite popular, little systematic effort to test the hypothesis has been made.

In testing the AFDC program as a cause of illegitimacy we cannot simply use national trend data to relate benefit levesl and changes to illegitimacy rates. This is because the AFDC benefit varies greatly among the States. This complication, however, provides the analyst with the means of testing program effects, because the variation in State benefits is quite large. Viewed in this way the State differences in AFDC payments provide a quasi-experimental set of historical data to test the main hypothesis, while also allowing assessment of the probably future effects of minimum income supplement programs.

AFDC benefits per recipient are little different in families with or without illegitimate children. However, family benefits are higher in families with legitimate children because family size is somewhat larger. When States are grouped into high-to-low benefit strata and family benefit data are used (as in tables 16 and 17, below) the average family benefit will be higher than the average to families with illegitimate children. This fact does not alter the relative position of AFDC payments to families with illegitimate children among the States, because the rank order of per recipient benefits or family benefits using all families or only families with illegitimate children is virtually identical. States will be in the same high- or low-benefit stratum whether we do or do not specify benefit levels by the presence or absence of illegitimate children. (U.S. Welfare Administration, 1963: table 56).

BENEFIT CHANGE AND ILLEGITIMACY RATE CHANGE: 1940–60

Eighteen States (see table 15) have large numbers of nonwhites and whites as well as illegitimacy data for 1940, 1950, and 1960. Table 15 treats changes in AFDC benefits (in 1958 dollars) in terms of the pattern of small or large changes between 1940 and 1960. In this table, States are grouped according to whether they maintain a consistent pattern of large benefit increases, whether they varied from decade to decade, or whether they had a consistent policy of small benefit increases.

TABLE 15.—*Pattern of change in annual AFDC benefits and change in illegitimacy rates, by color, 1940–60*

Increase in annual AFDC benefit per recipient		Change in illegitimacy rate, 1940–60	
1940–50	1950–60	White	Nonwhite
High	High [1]	4.3	61.9
Low	do.[2]	5.5	58.1
High	Low [3]	5.9	63.9
Low	do.[4]	2.7	64.0

[1] Virginia, West Virginia, Illinois, New Jersey.
[2] Kentucky, Louisiana, North Carolina, Ohio, District of Columbia.
[3] Florida, Tennessee, Texas, Pennsylvania, Michigan.
[4] Alabama, Missouri, Mississippi, South Carolina.

NOTE.—"Large" benefit gain between both periods was $51 or higher, with the mean gain in "large gain" States of $92 (compared to $23 in small gain States) between 1940 and 1950. Mean gain in large gain States between 1950 and 1960 was $89 compared to $19 in the small gain States. A change in the annual benefit per recipient of, say, $51 means that the annual benefit at the end of the decade period was $51 a year higher than it was at the beginning of the decade.

Source: Benefit data by States from U.S. *Social Security Bulletin*, various years, adjusted to 1958 dollars. State illegitimacy rates are age standardized, and were calculated from State birth data adjusted for under registration of births, and Census counts for unmarried women 15 to 44.

We find that States with consistently high benefit increases had smaller white illegitimacy rate increases than States with inconsistent benefit changes; States with consistently small benefit increases had smaller white illegitimacy increases than any of the other three types.

There is no relationship between changing benefit levels and changes in nonwhite illegitimacy rates. For example, nonwhite illegitimacy rates in States with consistently high benefit increases jumped by about 62, while those in States with consistently low benefit increases went up by 64. Changing AFDC benefits are not related to changing illegitimacy rates in the 1940–60 period.

AFDC BENEFITS AND ILLEGITIMACY: 1960 AND 1970

Illegitimacy rates in 1960 and 1970 may be compared after including married but separated women in the denominator along with single, widowed, and divorced women. Counts of 1960 illegitimate births by State are somewhat more refined than 1970 data, because they include illegitimate births to State residents who delivered in another State that reports illegitimacy, while 1970 data generally exclude out-of-State births. The number of such births, however, is small.

It is necessary to continue to confine an analysis of nonwhite rates to States with predominantly black populations, because illegitimacy rates among various nonwhite groups are different. For example, in Oklahoma in 1970 the rate was 77.8 per 1,000 unmarried nonwhites aged 15–44; among American Indians the rate was 48.3 while among blacks it was 91.8. Although Indians formed only one-third of the nonwhite population at risk, their presence in the nonwhite population lowers the rate from an expected level per thousand of 91.8 (under the common assumption that virtually all nonwhites are blacks) to 77.8—a large drop solely attributable to the composition of the nonwhite population in this State. Similarly, the illegitimacy rate to nonwhites in Hawaii was 27.2 per thousand in 1970—this low nonwhite rate is due to the different ethnic composition of the Hawaiian nonwhite population compared to the United States as a whole. For these reasons those

States with few blacks in the nonwhite population are omitted from the following comparisons of illegitimacy rates and AFDC benefits in 1960 and 1970—a procedure that alters the States included in various benefit strata in some cases.

AFDC benefits and illegitimacy rates in 1960

Table 16 compares the illegitimacy rate for white and nonwhite women aged 15–44 in high-, middle-, and low-benefit States. For whites we find a mean illegitimacy rate of 9 births per 1,000 in the States paying the highest benefits, a rate of 10 births per 1,000 in the middle stratum, and a rate of 9 per 1,000 in the lowest benefit stratum.

TABLE 16.—*Illegitimacy rates and AFDC monthly family benefits, by color: 1960*

1960 AFDC monthly family benefit	Mean benefit	Number of States	Illegitimacy rate per 1,000 unmarried aged 15 to 44
White:			
High	$180	8	9.0
Middle	116	9	10.2
Low	74	9	8.9
Nonwhite:			
High	159	6	71.6
Middle	102	7	90.1
	(103)	(6)	(82.3)
Low	67	7	89.9
Extreme States:			
White:			
Highest	189	(1)	8.3
Lowest	44	(2)	5.8
Nonwhite:			
Highest	189	(1)	94.2
Lowest	44	(2)	98.7

[1] Illinois.
[2] Mississippi.

NOTE.—Monthly benefits in 1967 dollars.
For whites, high benefit States are: District of Columbia, Iowa, Illinois, New Jersey, Minnesota, Oregon, Washington, and Wisconsin.
For whites, middle benefit States are: Michigan, Pennsylvania, South Dakota, Indiana, Louisiana, West Virginia, Virginia, Missouri, and Delaware.
For whites, low benefit States are: Georgia, Kentucky, North Carolina, South Carolina, Tennessee, Texas, Florida, Alabama, and Mississippi.
For nonwhites, high benefit States are: District of Columbia, Illinois, New Jersey, Michigan, Pennsylvania, and Indiana.
For nonwhites, middle benefit States are: Louisiana, West Virginia, Virginia, Missouri, Delaware, Georgia, Kentucky.
For nonwhites low benefit States are: North Carolina, South Carolina, Tennessee, Texas, Florida, Alabama, and Mississippi.

Among nonwhites the illegitimacy rate in the highest benefit stratum was about 72 births per 1,000, in the middle stratum it was 90 (82 when Delaware, with its possibly unreliable rate is omitted) births per 1,000, while the nonwhite rate in the lowest benefit stratum was 90.

Comparison of the two States with the highest and the lowest AFDC benefit show a slightly higher rate to whites in the highest benefit State; among nonwhites the State with the lowest benefits had the higher rate. We conclude that, for 1960, this analysis yields no evidence that the level of illegitimacy is a function of the level of AFDC benefits.

AFDC benefits and illegitimacy rates in 1970

The level of AFDC benefits and illegitimacy rates for women 15 to 44 in 1970 are shown in table 17. Among whites the average rate in the States paying the highest benefits is 13.5; in the next lower benefit stratum the average illegitimacy rate is 15, while it is 11 in the lowest benefit stratum.

TABLE 17.—*Illegitimacy rates by AFDC monthly family benefits, by color: 1970*

1970 AFDC monthly family benefit	Mean benefit	Number of States	Illegitimacy rate per 1,000 unmarried aged 15 to 44
White:			
High	$193	9	13.5
Middle	120	14	15.1
Low	71	8	10.9
Nonwhite:			
High	189	7	82.9
Middle	101	9	80.4
Low	68	7	80.9
Extreme States:			
White:			
Highest	218	(1)	9.4
Lowest	40	(2)	9.6
Nonwhite:			
Highest	218	(1)	81.7
Lowest	40	(2)	107.6

[1] New Jersey.
[2] Mississippi.

NOTE.—Benefits in 1967 dollars.
For whites the States are: High benefits: New Jersey, Illinois, Minnesota, New Hampshire, Wisconsin, Michigan, District of Columbia, South Dakota, Pennsylvania, and Washington. Middle benefit white State are Iowa, Virginia, West Virginia, Oregon, Delaware, Arizona, Texas, North Carolina, Kentucky, Colorado, Nebraska, Indiana, Oklahoma, and Missouri. Low benefit white stratum is Tennessee, Gerogia, Arkansas, Florida, Louisiana, South Carolina, Alabama, Mississippi.
Nonwhite high benefit States are New Jersey, Illinois, Wisconsin, Michigan, District of Columbia, Virginia, Pennsylvania.
Nonwhite middle benefit States are Delaware, Texas, North Carolina, Kentucky, West Virginia, Tennessee, Nebraska, Indiana, and Missouri. Nonwhite low benefit States are Alabama, Arkansas, Georgia, Florida, South Carolina, Mississippi.

Among nonwhites the average illegitimacy rate in the highest benefit stratum is 83 births per 1,000, in the middle benefit stratum it is 80, while the rate is 81 in the lowest benefit stratum.

The lower panels of table 17 show the illegitimacy rates to unmarried women aged 15 to 44 in the State paying the largest and the State paying the smallest benefit. We find that the benefit in New Jersey is five times greater than is the benefit in Mississippi. This difference has no apparent effect in boosting either the white or nonwhite illegitimacy rate in New Jersey: for both racial groups the illegitimacy rate in Mississippi is higher than is the rate in New Jersey.

THE RATIO OF AFDC BENEFITS TO MALE EARNINGS

There is a strong positive relationship between the ratio of AFDC benefits to male earnings and the level of AFDC benefits (President's

Commission on Income Maintenance Programs, 1970: table 3.3–1). This variable, which some claim to be a cause of illegitimacy because a high ratio may discourage marriage thus encouraging illegitimacy, might obscure the "true" impact of high benefits on illegitimacy if its relationship to benefits levels were negative rather than positive. However, the States in which the ratio of benefits to male earnings is highest are also the high benefit States. In spite of having both factors at work, these States have illegitimacy rates no higher than lower benefit States.

The view that yet-to-be measured economic variables mask the "true" effect of benefits on illegitimacy may be supported by further work. Such an economic factor must be negatively related to benefits and positively related to illegitimacy. The search for such a variable should proceed with historical as well as current measures that can be directly linked to one or another immediate cause of an illegitimacy rate. At the present time. I have no evidence that such variables eixst, nor do I have a model that says that they should exist.

This review of the relationship of AFDC benefits and illegitimacy rates failed to find an association between the level of benefits and the level of illegitimacy in either 1960 or 1970. Further, analysis of change in illegitimacy rates and change in benefit levels between 1940 and 1960 failed to find that States with consistently high- or low-benefit changes had high or low changes in illegitimacy rates. We conclude that variation among the States in AFDC benefit levels is not related to illegitimacy rates. Although illegitimate children are a major cause of AFDC expenditures, the AFDC program is not a cause of illegitimacy. This conclusion is counter to recent work by economists whose theories suggest that fertility should respond to economic factors, like income support programs. Whatever the possible impact of income supports on the marital fertility of lower income groups, it is unlikely that income support to unwed mothers will affect illegitimacy rates. This conclusion is developed in the following sections.

WHY BENEFITS DO NOT CAUSE ILLEGITIMATE BIRTHS

Recent work on economic explanations of fertility (Robinson and Horlacher, 1971) develops the theory that income (measured in a variety of ways) is positively related to fertility, because children cost money. Since a program like AFDC meets the most obvious costs of childrearing, a naive argument might be that reducing the costs of illegitimate children stimulates illegitimacy. Since our analysis does not indicate that this is the case, we must ask why the economic theory fails.

The economic theory of fertility is qualified with the provision that it will work only when childbearing is rational and deliberate—that is, under conditions of perfect fertility control. This precondition is not met in the case of most illegitimate births. Thus, an effort to "control" illegitimacy by "cracking down" on welfare is unlikely to have any more effect in future than similar efforts have had in the past. It is also important to point out that increasing AFDC benefits probably will not reduce illegitimacy rates either.

THE EFFECT OF INCOME SUPPLEMENTS ON REDUCING SANCTIONS AGAINST ILLEGITIMACY

The possible effects of guaranteed income supplements on illegitimate childbearing by the next generation may be clarified by comparing the effect of different public programs and private efforts that reduce the negative sanctions resulting from illegitimacy. Table 18 lists eight types of punishment—both social and economic. The degree to which each punishment will be experienced by the pregnant unmarried women or the unwed mother will vary according to her economic status at the time, and other characteristics.

TABLE 18.—*Effect of birth control, income supplements and release of illegitimate child on punishments of the unwed mother*

Types of punishments	Birth control or abstinence	Income supplements	Adoption or release
Negative effects during pregnancy and childbirth	—	0	0
Sole responsibility for 16 to 18 years childcare	—	0	—
Social sanctions related to unwed mother status	—	0	0
Declining chance for marriage	—	0	—
High risk of female family head status	—	0	—
Decline in economic status during pregnancy	—	0	0
Lower economic status after birth	varies	varies	varies
Increase in poverty gap related to added children	—	varies	—

NOTE.—A "0" indicates no effect on the punishment while "—" indicates strong relief from the punishment. The term "varies" means, for example, that the risk of lower economic status after birth is a risk dependent on the economic status prior to the birth, the likely future economic status had the birth not occured and the size of child benefits offered by the income supplement program.

If the woman does not experience pregnancy and illegitimate childbirth she will not be sanctioned unless the economic reward for illegitimate childbearing exceeds the reward for avoiding it. Even if economic status were improved for a poor young woman, a glance at the other consequences of illegitimacy suggests that the net effect would still be negative, if she retained the child.

The three columns in table 18 express the effect of each program and individual response type in decreasing the specific form of punishment. There is at present no way to assess the relative importance of one or another punishment. For example, use of birth control (or abstinence) means that illegitimacy is avoided. Thus the consequence is to eliminate the effects of illegitimacy (expressed by the negative sign). The same elimination of effects of some punishments can be obtained if the woman releases the child for adoption—see column 3. This course of action (primarily used by whites rather than nonwhites) allows the unwed mother to avoid the sanction of 16 to 18 years of sole responsibility for child care, puts her back into the marriage market, and allows her to avoid the status of female head of a family or subfamily. In most cases release of the child for adoption also allows the unwed mother to avoid further increase in the poverty gap, if she is already at or below the poverty line at the time of birth.

Avoiding childbirth or giving up the child allows women above the income supplement line to maintain their economic position. Income supplements cannot affect this group.

Giving up a child after birth has no effect on social sanctions related to illegitimacy, the negative feelings and material affects experienced during pregnancy and childbirth, or likely wage loss during pregnancy.

In contrast, an income supplement program has no effect on any punishment except those involving the decline in economic position for women whose prepregnancy income was low enough to allow them to improve their economic status by bearing a child. Even so, the income supplement for this group of unwed mothers does not remove the sanction of sole responsibility for child care over a 16-year period of time or longer; it does nothing to improve the woman's chances for marriage and normal family life; nor does it affect the other negative consequences of unwed motherhood.

From this perspective, an income supplement program can only alleviate some immediate and longer term economic consequences of bearing an illegitimate child. This does not suggest that future amelioration of the economic consequences for low-income women will stimulate illegitimacy, any more than that their past alleviation through the AFDC program has stimulated it.

Possible effects of income supplements on living arrangements and birth control

We can turn the main factors in table 18 around and ask whether it is likely that the introduction of an income supplement program will influence the spread of birth control, adoption, or other factors that reduce the number of illegitimate children living without the father.

The living arrangements of illegitimate children would be affected if an income maintenance program changed adoption levels, legitimation by marriage of illicit pregnancies before or after the birth, or provided funds to allow establishment of separate living arrangements by unwed mothers, or changed the likelihood that other illegitimate children would be born to the mother.

Adoption

If income supplements reduced adoption levels and thus increased the number of female headed households with illegitimate children, such a change would affect whites rather than blacks. Present adoption levels among blacks can hardly be depressed further, although informal temporary shifts in child care might change. Available data indicate that the trend toward separate living arrangements among women with children in a disrupted marital status (see Cutright and Scanzoni, 1973) has neared a maximum level; further, illegitimate childbearing occurs most often among young women who must live with their mothers because of their young age—a constraint on separate living arrangements that will not change in the future.

It is possible that the recent alleged decline in white adoption is simply a function of changes in the age and class characteristics of white unwed mothers. In any case there seems to be no reason to believe that adoption rates will change in response to income supplements. Historically, the unwed mothers most likely to release the child were white and tended to be middle class. This group will not be directly affected by income supplements.

Legitimation before or after marriage

Somewhere between 10 to 20 percent of white and nonwhite unwed mothers have, in the past, married the alleged father after the birth of the illegitimate child. These are couples with strong bonds to each other. This tiny minority can hardly be much reduced by income supplement disincentives to marriage.

About 24 percent of nonwhite pregnant out-of-wedlock women and two-thirds of whites marry before the birth. Analysis of the characteristics of the women who marry prior to the birth and those who do not indicates little if any economic difference (Cutright, 1972c). The decision is not determined by economic condition; rather, it is determined by the commitment to marry between the couple prior to or shortly after pregnancy. Couples with little interest in marriage do not marry because they do not want to—not because they are constrained from marriage by low income. Thus, income supplements should not reduce the rate of legitimation prior to birth.

Effects on use of birth control

Income supplements might increase illegitimacy by depressing birth control use. Is such a change likely?

Recent trends among both white and nonwhite women indicate record low fertility. Census reports (Current Population Reports, P-20, No. 240, September 1972) that young wives aged 24 and under have the lowest recorded expected family size (2.3 children per wife.) Use or expected use of reliable methods of contraception by married women increased during the 1960's (Westoff, 1972). By 1970, there was a difference of only 3 percent in the percentage of white or black wives "just not using" contraception (Westoff, 1972: 10). Contraceptive effort in the two populations is nearly equal, but its effectiveness is lower in the black population. This difference is expected, given the earlier adoption of contraception by the white population. Legal abortion is used twice as frequently to terminate pregnancy by nonwhites as whites in New York City (Tietze, 1973).

Fertility rates per 1,000 white women aged 15-44 with family incomes below 125 percent of the poverty line, declined from 141 in 1960-65 to 114 in 1966-70; over the same years, the fertility rate per 1,000 nonwhite women in the low-income group declined from 180 to 136 (Jaffe, 1972: Table 3). These enormous declines in the fertility of low-income women during a period of massive increase in utilization of AFDC hardly indicate that adopting a more uniform national income supplement program would be accompanied by a decline in contraceptive or abortion use.

In sum, the behavior of the white and nonwhite low-income women in recent years document a decline in desired fertility and an increase in use of abortion and effective contraceptive methods. Although it is the fertility of married women that has been responsible for the decline in white and nonwhite general fertility rates, notable declines in illegitimacy rates among older nonwhite women have been documented (see table 3). The interrelated factors that have depressed the fertility of low-income women should continue to operate. None of these facts suggest that income supplements will result in abandonment by unmarried low-income women of voluntary methods of birth control.

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PARTICIPATION IN THE AID TO FAMILIES WITH DEPENDENT CHILDREN PROGRAM (AFDC)

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SUMMARY AND CONCLUSIONS

Poor female-headed families with children, generally considered an approximate measure of potential eligibles for AFDC, numbered about 1.7 million in 1970. In contrast, close to 2.4 million female-headed families received AFDC benefits sometime during 1970. That this discrepancy is anything more than statistical is not evident, since the current rules of AFDC eligibility include the potential of extending benefits to many families that are nonpoor (according to the official U.S. Government definition of poverty). The methodology for estimating AFDC eligibles developed in this paper shows that, indeed, more families are eligible for AFDC than are now participating. The 1970 estimate of eligible female-headed families of 2.7 million, more than accounts for the 2.4 million on the rolls—illustrating the inadequacy of poverty statistics as a measure of AFDC eligibility.

To date, the estimating methodology has been specified for data years 1967 and 1970. These years were chosen to accord with a period of rapid caseload growth so that the change in participation could be evaluated. The resulting estimates show that while the eligible pool increased by 24 percent between 1967 and 1970, the caseload doubled, suggesting that the rate of participation increased substantially. Comparing the estimates of eligibles to annualized estimates of the participating caseload shows that in the aggregate participation increased from 56 to 78 percent from 1967 to 1970. Similarly, for the female-headed portion of the caseload, participation increased from 63 to 91 percent. These aggregate figures reflect a general and widespread increase in participation which seems to have affected most geographic and demographic groups. The relative pattern of participation, however, exhibits no radical change between 1967 and 1970. Female-headed families, nonwhites, central city residents, and families located in the northeastern and western census divisions maintained a higher probability of being on AFDC than their counterparts.

The high level of participation for female-headed families suggests that in the future any substantial growth in caseloads for this group will have to come from the creation of new eligibles. To a limited extent, new eligibles had to account for some of the growth of the female-headed portion of the caseload between 1967 and 1970. Although the magnitude of the numbers involved is not overwhelming, the issue is one of significant policy concern. It is tempting to attribute the growth in female-headed families eligible for AFDC to AFDC

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itself, since the program offers incentives for intact families to split and for single females to have illegitimate children. Data at the national level suggest, however, that over the 1967 to 1970 period the dominant phenomenon was expansion of income eligibility rather than creation of female-headed families. The analysis presented here does not prove the absence of a behavioral link between the growth in AFDC caseloads and the growth in female-headed families—but simply suggests that over the 1967 to 1970 period the dominant phenomenon appears to have been a large and widespread increase in participation augmented by expansion of eligibility up the income scale.

I. Introduction

By the end of 1971, almost 3 million families were receiving money payments from the Aid to Families with Dependent Children (AFDC) program. The number of recipient families had tripled in a decade and more than doubled since 1967. Although the general trend in recipients has been upward since the inception of the program in 1935—interrupted by declines only during World War II and the Korean war—the growth of the latter 1960's and early 1970's was unprecedented. By 1970 the annual rate of growth reached a peak of 36 percent representing an addition of 677,000 families to the rolls in that year. Since then, the rate of growth has begun to moderate even though the absolute rise continues.[1]

The past decade of growth has been particularly perplexing because at the same time most of the Nation has been enjoying continuous prosperity. Between 1959 and 1971 close to 14 million Americans moved out of poverty.[2] The conventional scenario of the financially pressed father deserting his family so they may receive AFDC does not seem adequate to explain what has been happening. A more complex economic explanation is required and/or other factors must be taken into account. Indeed, a number of other economic and noneconomic phenomena, with potentially significant affects on AFDC, were occuring coincident with the rise in caseloads.

Although a general decline in poverty was observed over the decade, the process was selective—those who left were primarily male-headed families. Census figures show a slow but steady rise in the number of female-headed families with children in poverty—the primary population at risk for AFDC. Between 1959 and 1971 the number of such families increased by 20 percent. It is tempting to attribute this rise to the increasing financial attractiveness of AFDC itself. In many areas of the country, AFDC now provides a regular flow of cash benefits which may be supplemented by earnings, automatic eligibility for food stamps and medicaid, and public housing. But still a causal relationship is far from clear. The increase in all families over the

[1] U.S. Department of Health, Education, and Welfare, Social and Rehabilitation Service, National Center for Social Statistics. "Public Assistance—Annual Statistical Data, Calendar Year 1969" (NCSS Report A-7) and "Public Assistance Statistics," December 1970 and 1971 (NCSS Report A-2).

[2] U.S. Bureau of the Census, *Current Population Reports*, P-60, No. 86, "Characteristics of the Low-Income Populations, 1971," U.S. Government Printing Office, Washington, D.C. 1972. All references to poverty or the poor in this paper refer to official U.S. Government definitions and statistics.

same period was 18 percent and nonpoor female-headed families with children increased by an astounding 120 percent.[3]

It is clear, however, that program rules have been liberalized over the decade. Between 1961 and 1970 State standards of need for a family of four rose from an annual average of $2,286 to $3,300,[4] expanding eligibility up the income scale. This expansion was further heightened by the work incentive features of the 1961 and 1967 amendments to the Social Security Act. By 1962 States were required to deduct work-related expenses from earnings in calculating benefits and determining eligibility; and by July 1969 States were required to disregard the first $30 and one-third of the remainder of monthly earnings in calculating benefits for AFDC units already on the rolls. Such a work incentive feature has the effect of extending eligibility to units with substantial levels of income. In a generous welfare State like New York the "30 and ⅓" provision raises the 1970 annual breakeven point to $5,688 for a family of four with earnings well above the annual guarantee of $4,032.[5] The breakeven point is even higher if deductions for work-related expenses are taken into account.

Noneconomic rules of eligibility have also been liberalized. The 1961 amendments permitted States to extend benefits to families headed by unemployed males, and by 1967, 21 States had such programs. In 1969 the Supreme Court struck down the residency requirement which States could impose to deny payments to residents of less than 1 year, and in 1968 abolished the "man-in-house-rule." Prior to being outlawed, this rule deemed a family ineligible for assistance if there was a man present regardless of his legal requirement to support. Now any male may be present as long as he is not the father of the children.

In addition to liberalization of eligibility, an increase in the rate of participation is generally thought to have occurred as a result of increases in "recruiting" for the welfare rolls. Great Society programs like Community Action, Model Cities, and OEO legal services as well as private organizations like the National Welfare Rights Organization appear to have acted as advertisers and casefinders for AFDC. New York City welfare workers testified before congressional hearings that welfare clients ". . . can quote your procedure better than you can." ". . . it is the client action groups and someone working with them that tells them that they are entitled to this, that, and the other thing, and they inform the client, and they come in and they seem to know, and if you check it through you find out they were correct."[6]

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[3] Ibid.

[4] Derived from "Monthly Cost Standards for Basic Needs Used by States for Specified Types of Old-Age Assistance Cases and Families Receiving Aid to Families With Dependent Children," January 1961. Bureau of Family Services, Department of Health, Education, and Welfare, December 1962, table 5, and "OAA and AFDC: Standards for Basic Needs for Specified Types of Assistance Groups," July 1970 (NCSS Report D-2, July 1970), U.S. Department of Health, Education, and Welfare, Social and Rehabilitation Service, National Center for Social Statistics, table 3.

[5] Ibid. The guarantee refers to the dollar amount of benefits a family of a given size would receive if it had no other income. The break-even point refers to the income level at which a family is no longer eligible for benefits.

[6] *Problems in Administration of Public Welfare Programs.* Hearings before the Subcommittee on Fiscal Policy of the Joint Economic Committee, Congress of the United States, Part I, March 20; April 11, 12, 13, 1972, U.S. Government Printing Office, Washington, D.C., 1972.

To understand what has been happening to ADFC—or more basically, to measure what has been happening—requires sorting out a complex set of interwoven factors. One aspect of this task is the development of a more precise measure of AFDC eligibles than the number of female-headed families in poverty. In 1970 poor female-headed families with children numbered approximately 1.7 million, while an estimated 3.1 million families received AFDC benefits at some time during the year.[7] That this discrepancy is anything more than statistical is not evident, since the rules of AFDC eligibility include the potential of extending benefits to many families that are neither poor nor female-headed. The purpose of this paper is to report on the development of a methodology for estimating AFDC eligibles from census data, and the subsequent analysis of participation which is possible once the eligible pool has been estimated.

II. Methodology

To estimate the number of families eligible for AFDC, the eligibility criteria of the program are applied to the data files of the Current Population Survey (CPS)—the Census Bureau's recurring sample survey of 50,000 households. Once the crucial economic and demographic rules of eligibility are determined, they are translated into a set of decision rules and computational formulas suitable for screening each family in the CPS file for the presence of an eligible AFDC assistance unit. Such a procedure, in effect, counts the number of eligible units in the CPS sample and is thus subject to the same problems of sampling error as other CPS estimates.[8] To date, the estimating methodology has been specified for data years 1967 and 1970, but the general approach could easily be adapted to other years for which CPS files are available. The detail of this methodology is outlined in appendix A—only an overview is presented here.

The decision rules and computational formulas, which approximate actual rules of eligibility, are specified on a State-by-State basis as much as possible to reflect the variation in AFDC program characteristics across States. Although Federal law sets down general guidelines which States must follow to be eligible for Federal financial participation, a great deal of discretion in interpreting and implementing the Federal rules is left to the States. It is not surprising, therefore, that wide variations in program rules between States have resulted. The most dramatic example is the economic standards of need which are set by each State individually. In 1970 State standards for a family of four varied from a low of $2,112 per year in Arkansas to a high of $5,184 per year in California.[9] Administrative data on these standards are published periodically and are fully implemented in the computational procedures described below. Unfortunately, published information on other aspects of eligibility is not always available, necessitating more imprecise approximations. The general approach in

[7] Op. cit., U.S. Bureau of the Census for the poverty counts. See Technical Appendix B, pp. 2–5 for derivation of the estimate of annual AFDC recipients.

[8] For a discussion of the data base and its limitations see U.S. Bureau of the Census, *Current Population Reports*, Op. Cit.

[9] Op. cit., NCSS Report D-2, 7/70.

deriving these approximations has been to make liberal, but practicable assumptions so the resulting estimates represent an outside count of the pool of eligibles.

Categorical Eligibility

Since the AFDC program covers only certain demographic types or categories of families, the first step is to define the characteristics that determine categorical eligibility. The original intent of the 1935 Social Security Act, which initiated Federal participation in public assistance programs, was to confine assistance to categories of individuals who were unable to work because of age, blindness or absence of wage earner.[10] Consequently, AFDC payments were initially confined to needy children under 16 who were deprived of parental support because of death, incapacity, or absence from the home of a parent.[11] Over the years the Federal definition of categorical eligibility has been extended to include: such needy children under 18, and 18 to 20 if in school; the needy parent or caretaker relative of these children; both parents in families headed by incapacitated fathers; and in States with programs for unemployed parents, both parents in families headed by unemployed fathers.

Between 1967 and 1970 there were no major changes in the Federal law regarding categorical eligibility; therefore, the same set of decision rules is applicable for both years. The first rule identifies the universe of families with eligible children, defined in most States as never-married children under 18, and 18 to 20 if in school, following the Federal guideline. States may be more restrictive than the Federal guidelines, in this case, but cannot receive Federal aid for any child outside the Federal definition, e.g., children 18 to 20 not in school. Consequently the major deviation from the Federal standard is a few States that restrict eligibility to children under 18.[12]

Within this universe of families with eligible children, the next most important determinant of eligibility is the status of the family head. Thus a family is considered categorically eligible if:

(1) The head is a female. The female-headed family is the most typical AFDC assistance family, accounting for approximately 78 percent of the total caseload in 1970. Fortunately, these families are the easiest to identify in the CPS file.

(2) The head is an unemployed father and the family lives in a State with an unemployed parent segment of the AFDC program. In terms of CPS variables, an unemployed father is defined as a male head who did not work for at least 3 months because of the inability to find a job. The program definition of unemployment is considerably more complex

[10] Not until 1950 was the category for persons totally and permanently disabled added.

[11] Originally the program was entitled Aid to Dependent Children (ADC) reflecting the limitation of payments to children only. It was not until 1950 that a Federal contribution was provided for the caretaker of the children. The name of the program was changed to AFDC in 1962.

[12] See app. A for the definition of eligible children on a State-by-State basis.

than this but cannot be replicated with the information in CPS file.[13]

(3) The head is an incapacitated father. In terms of CPS variables an incapacitated father is defined as a male head who did not work at least 3 months because of illness. Although definitions of incapacity vary by State, this rule represents the extent of the disability information in the CPS. Even if State-by-State definitions of incapacity were determined this would still be the best approximation.

(4) The head is a male and no spouse is present, i.e. a male-headed-single-parent family. The wording of the Federal law "absence from the home of parent" allows eligibility for this type of family. Such families are relatively rare in the population as a whole and thus account for a fraction of 1 percent of the participating caseload.

For each of these four types of categorical families, only the head, spouse (if present), and eligible children are considered to be part of the AFDC assistance filing unit. All other persons in the household and their income are excluded from the assistance unit for the purpose of determining economic eligibility. This does not necessarily represent actual practice. A number of States have relative support requirements which legally obligate relatives present in the same household (as well as those not present) to lend support to the eligible assistance unit. Also, welfare officials presumably pro-rate certain shared expenses, such as rent, in determining benefits to the AFDC family when other individuals in the household are obviously contributing to the household budget. In estimating eligibles these considerations are ignored because no definitive information regarding actual treatment of such issues is available.

In addition to the above, one other type of categorical unit is considered. According to the AFDC studies of 1961, 1967, and 1971,[14]

[13] "* * * AFDC benefits may be made available to a needy child whose father is unemployed only in a State which has elected to cover such fathers and only when the father (1) has been unemployed for at least 30 days prior to the receipt of such benefits, (2) has not without good cause refused a bona fide offer of employment or training for employment, (3) has six or more quarters of work in any 13-calendar-quarter period ending within 1 year prior to the application for such benefits (or the father received, or was qualified to receive, unemployment compensation within 1 year prior to the application for such benefits), (4) is registered with the State employment offices, and (5) is not receiving unemployment compensation for the week in which he receives AFDC assistance.

A State's definition of an unemployed father must include any father who is employed less than 100 hours a month, or whose employment exceeds that standard for a particular month only if his work is intermittent and the excess is of a temporary nature as evidenced by the fact that he was under the 100-hour standard for the 2 prior months and is expected to be under the standard during the next month." Quoted from Studies in Public Welfare, Paper No. 2, *Handbook of Public Income Transfer Programs*, a staff study prepared by Irene Cox for the use of the Subcommittee on Fiscal Policy of the Joint Economic Committee, Congress of the United States, October 16, 1972, U.S. Government Printing Office, Washington: 1972.

[14] U.S. Department of Health, Education, and Welfare, Social and Rehabilitation Service, National Center for Social Statistics, "Findings of the 1967 AFDC Study," July 1970, and "Findings of the 1971 AFDC Study," Dec. 22, 1971. The 1961 study was prepared by the Welfare Administration, Division of Program Statistics and Analysis, "Characteristics of Families Receiving Aid to Families with Dependent Children, Nov.-Dec. 1961," April, 1963.

a substantial number of assistance units contain no adults—in 1971 almost 10 percent of the caseload. The earlier studies suggest that many of these units are children living with relatives other than parents—such as grandparents, aunts and uncles, or older brothers and sisters. Strict interpretation of the Federal law generally allows payments to these children regardless of the income of their relatives.[15] Thus, if a family contains at least one related but not own child, the child by himself (or themselves if more than one) constitutes a categorically eligible AFDC unit.

These five types of AFDC units cover the major categories outlined in the Federal law and account for about 97 percent of the current participating caseload.[16] The minor exclusions are not expected to have a significant effect on the estimates of eligibles and the subsequent analysis of participation.

Economic Eligibility

Once a CPS sample family is identified as categorically eligible, the next step is to compute economic eligibility. While in reality economic eligibility involves a review of both assets and current resources, the computational formulas are limited to current income since the CPS has no asset information.[17] The omission of the asset test, theoretically, biases upward the estimates of eligibles. However, it seems reasonable to assume that assets are not important for the population in question.

In terms of current income, all States determine economic eligibility and payments in the same general manner—by comparing a family's resources to the State's standard of need. Federal law requires that each State set a standard of need (S) which represents the State's assessment of what is required by a family of a given size, with no other income, to meet its basic needs. Because a State may not be fiscally capable of paying its standard of need (or simply chooses not to provide benefits at that level), several methods may be employed to lower payments and/or limit eligibility.

(1) The need standard (S) may be reduced by a percentage (p) to obtain a reduced payment standard (pS). Current resources are then compared to the reduced standard to determine eligibility and payments.

(2) The difference between the full standard or the reduced standard, where applicable, and a family's resources is termed the budget deficit and in most States is equal to the amount of the payment. But some elect to reduce benefits by paying only a certain percentage (r^d)—termed a ratable reduction—of the budget deficit.

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15 If the relatives are poor, one of them may be included in the assistance unit as the "needy caretake relative." Such a category has been omitted here because of lack of program information regarding the definition of a needy relative. The omission results in a slight underestimate of eligible recipients but does not affect the estimates of eligible units which are the primary focus of this analysis.

16 The largest group excluded here is the stepfather category which numbered 66,000 in Jan. of 1971, accounting for 2.6 percent of the total caseload. In some States, depending on support laws, a woman's children from a previous marriage are eligible for AFDC even though the family is now headed by a stepfather.

17 It is known that income reported to the CPS falls substantially short of aggregate benchmark estimates derived from administrative statistics. Hence, another source of upward bias in the estimates of AFDC eligibles.

(3) Finally, payments may be limited by setting a fixed maximum (M) which generally varies by family size.

In 1970, 7 States employed two of the reduction methods and 34 others employed one—the most commonly used method being the fixed maximum.[18] Both the fixed maximum and ratable reduction have the effect of lowering overall payments, albeit in rather different ways. The ratable reduction reduces payments to all families by a fixed percent, but the fixed maximum affects only those with high benefits and generally the least amount of other income. The reduced standard, on the other hand, not only lowers payments but also limits eligibility.

The procedure described thus far may be summarized by the general algebraic formula:

$$P = r^d\ (pS - Y^u - Y^e)$$

Subject to the constraint that if $P > M$ set $P = M$.

Where:

P=Amount of the AFDC payment to a family of a given size.

S=State's full standard of need for a family of a given size.

p=Percentage by which the standard of need (S) is reduced for the purpose of calculating benefits. For most States p has a value of 1.

r^d=Percentage by which the budget deficit is reduced for the purpose of calculating benefits. Again, for most States, r^d has a value of 1.

Y^u=All unearned income received by the assistance unit such as social security, alimony, child support, unemployment insurance, etc.

Y^e=All earned income of the assistance unit such as wages, salaries, self-employed income, etc. (Ignoring for the moment the earnings disregards legislated over the 1960's.)

M=Fixed maximum on payments which generally varies by family size.

The above formula—without the work incentive provisions legislated in the 1960's—offers little incentive for welfare recipients to go to work since payments are reduced dollar for dollar for additional earnings. In fact, if the costs of working such as transportation, child care, clothing, et cetera, are taken into account, the financial "advantage" of working may actually be negative. Such was the case with AFDC until the 1961 amendments to the Social Security Act. The emphasis on "rehabilitation" of welfare recipients that began to emerge from Congress in the late 1950's was reflected in these amendments. Open-ended matching funds were authorized for States providing social and rehabilitation services to present, former, and potential welfare recipients; and by 1962 States were required to deduct work expenses (including child care) from earnings before determining eligibility and calculating payments. The latter, at least, prevented recipients from being worse off by going to work but maintained the dollar for dollar reduction in benefits for earnings above the amount of the deductions.

[18] See app. A, sec. XIV, and adaptation I.

After 4 years of "rehabilitation" services, the AFDC caseload had increased by 45 percent and the rate of growth was accelerating. Thus, in 1967, Congress initiated the stronger Work Incentive or WIN program (which required all employable recipients to sign up for training or employment services) and required all States, by July 1969, to disregard the first $30 and one-third of the remainder of monthly earnings in calculating AFDC benefits. The "thirty and a third" provision is designed to provide financial work incentive to the welfare mother,[19] by allowing her to earn $30 each month with no reduction in benefits. Furthermore, each dollar of earnings above $30 now reduces benefits by only 67 cents. Previously, a mother earning $90 a month would have her family's benefits reduced by the full $90. Now, she "keeps" $50 because benefits are reduced by only $40.

Since the "thirty and a third" provision was intended to encourage current recipients to work and not to expand eligibility (as reviewed earlier, the "30 and ⅓" has the potential effect of extending benefits to families with higher incomes), the procedures for determining initial eligibility for AFDC were revised. Now, before a payment is calculated, the potential recipient family must pass a "full standard" test of eligibility. The "full standard" test is a pass-fail type test which simply compares current resources less work expenses to the State's "full standard of need." The term "full standard" is used to indicate that, for this test, States may not apply reduced standards or any other reduction method—but must use the full value of the standard of need.

The changes regarding economic eligibility which were legislated over the 1960's may now be summarized algebraically. From 1962 to July 1969, economic eligibility and payments were determined as follows:

$$P=r^d(pS-Y^u-\max\,[0,\ Y^e-WRE-CCE]) \qquad (1)$$

Subject to the constraint that if $P>M$ set $P=M$.

Where:

WRE=monthly work-related expenses as determined by the States.

CCE=monthly child care expenses also determined by the States.

Beginning in July 1969, each potential recipient family must first qualify on the basis of the full standard test:

If $S \geq Y^u-\max\,(0,\ Y^e-WRE-CCE)$, the unit passes the full standard test. (2)

Then a payment is calculated:

$$P=r^d(pS-Y^u-\max\,[0,\ .67(Y^e-30)-WRE-CCE]) \qquad (3)$$

Subject to the constraint that if $P>M$ set $P=M$.

Where the 0.67 and 30 represent the disregard of the first $30 of earnings and one-third of the remainder.

[19] HEW's initial interpretation of the work incentive provision limits its application to female-headed families who are recipients.

The latter formulation, to date, has not been changed and hence describes the current program. Although the "30 and 1/3" provision theoretically improves the financial work incentive inherent in the payment calculation, the procedure is not without drawbacks—stemming mainly from the "full standard" test of eligibility. In general, the full standard test is a more stringent economic screen than the payments formula because the disregards on earnings are not applicable.[20] Consequently, a potential recipient family with earnings may qualify for benefits on the basis of the payments formula but fail to qualify on the basis of the "full standard" test. Only after a family is on the rolls may earnings rise above the point of "full standard" eligibility because of the disregards. States supposedly recalculate payments periodically to eliminate units whose income has increased beyond the point of eligibility (in terms of the payment formula) and to adjust payments for lesser changes in income, but the "full standard" test is never reapplied to families on the rolls. Such a set of procedures yields the inequitable effect of barring benefits to categorically eligible families with incomes equal to or even less than that of families already on the rolls.

To estimate AFDC eligibles, a variation of formula (1) is applied to the CPS file for 1967 and variations of formulas (2) and (3) for 1970. The major problem in applying these computational formulas to the CPS file is the difference in income accounting periods used by the CPS and the AFDC program. The CPS measures annual income but welfare eligibility is based on monthly income. A family earning $10,000 for 11 months of the year is legally eligible for AFDC benefits during the 12th month, assuming no assets or other sources of income. Simple annualization of the computational formulas to accord with the CPS accounting period would miss such families and presumably underestimate the eligible pool. In short, families who appear to be ineligible for benefits on the basis of annual income may actually be eligible for benefits for some shorter period during the year because the receipt of their income is spread unevenly over the 12-month period.

Thus two computational procedures were devised—a simple annual accounting period procedure and a more complicated part-year procedure. The annual method simply annualizes the appropriate program variables and parameters such as standards, disregards, work related and child care expenses, etc., to accord with the annual income variables of the CPS. The part-year procedure, on the other hand, disaggregates the CPS annual data for each categorically eligible family into two periods on the basis of the weeks worked and weeks not worked of the head. Separate payments are then computed for each period assuming the heads earnings were received during the weeks worked period—annual payments being the sum of the positive payments for each period. Both of these procedures are described in greater detail in appendix A.

Validity of the Methodology

Validity, in this case, is primarily a function of the accuracy of the decision rules and computational formulas in approximating the

[20] This may not hold for units with sizable amounts of unearned income, besides public assistance, that live in States where a reduced standard applies; i.e., p has a value less than 1.

operation of the AFDC program; and the accuracy of the CPS in representing the population in question, i.e., AFDC eligibles. Although insuring the latter is not within the scope of this study, a word of caution is in order. The sampling procedures of the CPS are designed to produce estimates which are descriptive of the national population. Therefore, estimates which are descriptive of a limited subsample—such as AFDC eligibles—should be approached with circumspection. Although the aggregate estimates are within the design of the CPS sample, small sample size becomes a problem very quickly if the subsample population is disaggregated in too great detail. The Census Bureau does not insure the validity of results with respect to the detailed characteristics of limited subsamples. While elaborate procedures are utilized for correcting the characteristics of the national sample to agree with controls, understandably, this cannot be done for every conceivable subsample which may be drawn from the CPS.

Aside from the fundamental problems of the data base, the estimates will be biased to the extent the decision rules and computational formulas do not accurately reflect the real world operation of the AFDC program. Given the complexity of the program and the lack of information on the details of operation, it is difficult to assess, a priori, the biases. One obvious drawback of the approach is the inability to evaluate the effect of administrative discretion which may have been an important factor affecting AFDC caseloads over the 1960's.

However, confidence in the estimates is increased if they agree with other independent measures of AFDC eligibles. Since no other measures of eligibles are available, the best that can be done is to test for the consistency of the estimates of eligibles with the measures of the participating population that are available. The populations are considered "consistent" if the eligible population includes a population that "looks like" the participating caseload on a number of important characteristics such as average payments, categorical types of families, geographic location, and so forth. Detailed tables and comparisons on such characteristics are presented in appendix B. Only the conclusions of that analysis will be presented here.

The validation analysis suggests that the best estimates of AFDC eligibles are those based on the part-year accounting period procedures. It is these estimates that are used in the subsequent analysis of participation.[21] However, the various validation comparisons also suggest that the model may be overestimating average payments and consequently may be overestimating eligibles as well. Unfortunately, this conclusion, at least for the present, must be classified as tentative. The validation analysis of the payments estimates is based on several key

[21] The 1970 estimates of eligibles are slightly downward biased because of the full-standard test. The tables of appendix B show the total effect on eligibility of the full standard test, but the number of families on the rolls above the point of full standard eligibility cannot be identified given the current set of computational procedures. Thus, the 1970 estimates of eligibles, which are based on the full standard test as well as the payments formula, do not include such families. However, the number of such families cannot be large. In January of 1971 only 8.3 percent of welfare mothers worked full time while 5.0 percent worked part time. The average earnings for welfare mothers in that month were $221.25—less the average monthly work expense of $65.30—leaves $155.95 of countable earnings (in terms of the full standard test) which is below even the lowest State standard for a family of four of $176 in the State of Arkansas.

assumptions and estimates which themselves are subject to error, such as the nature of underreporting of public assistance and the estimates of the annual AFDC caseload. Thus, at this point, no further technical work has been undertaken to improve the estimates because no certain measure of what is better exists. However, the tentative conclusion should be kept in mind and used as a caveat to temper any analysis based on the estimates of eligibles.

III. Results

Poor female-headed families with children, generally considered an approximate measure of potential eligibles for AFDC, numbered about 1.7 million in 1970. In contrast, by the end of that same year close to 1.9 million female-headed families were receiving AFDC benefits and another estimated one-half million received benefits at some time during the year. These substantial discrepancies indicate that even for the female-headed portion of AFDC, poor female-headed families are an inadequate measure of potential eligibles or, being less generous, that the problems of administering AFDC have truly reached crisis proportions. The figures of table 1 suggest the former to be the case, not the latter. The estimates of eligibles generated by the methodology of section II show that, indeed, there are more families eligible for AFDC than are now participating. For both 1967 and 1970 the aggregate figures on monthly and annual caseloads are well within the counts of eligibles. Unlike the approximate comparison above, the 1970 estimate of eligible female-headed families of 2.7 million more than accounts for the annual caseload of 2.4 million. Perhaps more interesting than the statistical reconciliation of caseloads and eligibles is the apparent nonpoverty status of a substantial proportion of the participating caseload. The figures suggest that about 30 percent of the female-headed families receiving AFDC benefits in 1970 were not poor (assuming that all of the 1.7 million poor female-headed families with children are both eligible and on the rolls). While this observation serves to illustrate the point made earlier—that the current rules of eligibility for AFDC include the potential of extending benefits to families that are not poor according to official Government definitions of poverty—it does not necessarily indicate that AFDC recipients are "doing well" on welfare. Poverty lines are arbitrary standards set primarily for the purpose of analyzing the changes in and composition of the population at the low end of the income scale. The point is that poor female-headed families with children are not an adequate measure of eligibility for AFDC.

The reconciliation of eligibles and caseloads, may be extended to more detailed comparisons with the same general conclusion as above. The figures in table 2 show that in general the estimates of eligibles account for the participating caseload on the characteristics of race, place of residence, region and age of family head. However, for 1970 there are several instances—namely, central cities, nonwhites, female heads under 25, and the western census division—where the numbers of eligibles can account for the monthly caseload only, and fall short of the larger annualized figures. The interpretation of such discrepancies should be approached with caution and the proverbial caveats of technical research.

TABLE 1.—*1967 and 1970 estimates of eligibles and caseloads by categorical type*

	Female headed	Unemployed fathers	Other	Total
1967:				
Eligibles	2, 183	282	781	3, 246
Monthly caseload	962	66	227	1, 254
Annual caseload	1, 385	95	327	1, 808
1970:				
Eligibles	2, 714	523	788	4, 025
Monthly caseload	1, 922	153	382	2, 457
Annual caseload	2, 460	196	489	3, 144

NOTE.—The monthly caseload estimates were taken from the previously cited 1967 and 1971 AFDC studies which give detailed characteristics on the November–December caseload for 1967 and the January caseload for 1971. The caseload figures in this table exclude stepfather cases. See app. B, pp. 71–74 for a detailed description of the derivation of the annual caseload estimates. Also, in app. B a more detailed breakdown of the "Other" category is provided for both eligibles and caseloads.

TABLE 2.—*1967 and 1970 estimates of eligibles and caseloads by place of residence, region, race, and age of head*

[Numbers in thousands]

	1967			1970		
	Eligibles	Monthly caseload [1]	Annual caseload [1]	Eligibles	Monthly caseload [1]	Annual caseload [1]
Total	3, 246	1, 278	1, 840	4, 025	2, 524	3, 231
Place of residence:						
Metropolitan:						
Center city	1, 243	684	985	1, 555	1, 428	1, 828
Noncenter city	709	218	314	1, 070	449	575
Nonmetropolitan	1, 234	372	536	1, 400	640	820
Region:						
Northeast	798	369	531	1, 038	694	847
North Central	752	258	372	943	500	663
South	1, 177	340	490	1, 206	607	791
West	519	274	395	838	605	891
Race:						
White	2, 046	676	973	2, 658	1, 220	1, 502
Latin American	NA	185	266	NA	357	514
Other white	NA	491	707	NA	863	1, 048
Nonwhite	1, 200	594	855	1, 367	1, 140	1, 459
Unknown	NA	8	12	NA	164	210
Age of head: [2]						
0 to 19	72	45	65	93	138	177
20 to 24	319	146	210	411	388	497
25 to 29	324	168	242	471	359	460
30 to 34	325	158	228	437	308	394
35 to 39	287	144	207	346	244	312
40 to 44	273	112	161	305	185	237
45 to 54	369	124	179	440	202	259
55 to 64	131	26	37	143	37	47
65 to 99	82	0	0	69	2	3

[1] Detailed estimates do not add to totals for certain characteristics because of unknown categories and derivation adjustments.

[2] These figures refer to female-headed families only. The caseload estimates were derived by assuming that the age distribution of female-headed AFDC units (see table 1) is the same as the age distribution of all AFDC mothers (mothers in the home for 1967).

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Although the discrepancies probably do suggest that participation in AFDC for these groups has reached relatively high levels, the results cannot be interpreted as statistical evidence of welfare cheating—a recent issue of intense public concern. The methodology underlying these discrepancies is simply too imprecise at the detailed level to warrant such conclusions. Aside from the already mentioned problems of the CPS,[22] one of the more obvious defects in this case involves the annual AFDC caseload estimates. Since no official statistics are collected on annual caseloads, the annual totals are derived from HEW official statistics on monthly caseloads and case openings. The detailed figures are then calculated by assuming that the annual caseload is distributed exactly like the monthly caseload. No a priori reason exists for making such an assumption except that it is the only alternative in the absence of the actual facts. As such, the derived estimates of the annual caseload are somewhat questionable for the detailed characteristics and are better suited for analysis at the aggregate level. The detailed estimates, the caveats withstanding, are presented here because they are helpful in assessing overall patterns of participation in AFDC. The various estimates viewed as a whole for the patterns they present are less likely to misrepresent reality than any one of the estimates viewed in isolation.

For the purpose of analyzing participation, participation rates have been constructed by expressing the annual caseload as a percent of the eligible population. The resulting rates for 1967 and 1970 are shown in table 3 along with percentage change figures for caseloads and eligibles. Although the number of eligibles increased by 24 percent, the rise was modest compared to that of the caseload which doubled. Consequently, the overall rate of participation increased from 56 percent in 1967 to 78 percent in 1970. Similarly, the rate for female-headed families increased from 63 percent to 90 percent. Indeed, virtually every geographic and demographic group shown in table 3 exhibits a substantial increase in participation over the 1967 to 1970 period. The one major exception is the unemployed father category, where the growth in eligibles kept apace of the growth in the caseload, a result consistent with the generally higher level of unemployment in the economy in 1970 than in 1967.[23]

[22] In addition to the already mentioned problems of the CPS, the more general decennial census problem of undercounting the nonwhite population may be affecting the estimates presented here. Since CPS sample estimates are adjusted to control totals derived from the decennial census, the problem of nonwhite undercounts also affects the CPS.

[23] See the *Economic Report of the President*, transmitted to the Congress, February 1971, U.S. Government Printing Office, Washington, D.C., 1971, pp. 36–39.

TABLE 3.—*AFDC growth, 1961–70*

[Figures based on counts of eligible units and/or recipient units]

	1967 participation rate	1970 participation rate	Monthly[1] caseload percentage change, 1961–67	Monthly caseload percentage change, 1967–70	Eligibles percentage change, 1967–70
Total	56	78	45	97	24
Unit type:					
Female-headed	63	91	72	100	24
Unemployed fathers	34	37	43	132	85
Other	42	62	−19	68	01
Place of residence:					
Metropolitan, total	65	92	70	108	30
Central city	79	118	NA	109	25
Noncentral city	41	54	NA	106	39
Nonmetropolitan	43	59	0	72	13
Region:					
Northeast	67	82	90	88	30
North Central	49	70	30	94	25
South	42	66	16	96	02
West	76	106	96	121	61
Race:					
White	48	59	57	80	30
Latin American	NA	NA	NA	93	NA
Other white	NA	NA	NA	76	NA
Nonwhite	71	107	50	2	14
Unknown	NA	NA	−84	1,950	NA
Age of head:[2]					
0 to 19	90	190	115	207	29
20 to 24	66	121	85	166	29
25 to 29	75	98	58	114	45
30 to 34	70	90	33	95	34
35 to 39	72	90	31	69	21
40 to 44	59	78	37	65	12
45 to 54	49	59	30	63	19
55 to 64	28	33	42	42	09
65 to 99	0	0.4	([3])	([3])	0.16

[1] 1961 figures on monthly caseloads do not include estimates for the State of Massachusetts.
[2] See footnote 2 of table 2.
[3] Base too small to compute.

The widespread increase in participation, however, appears to have had little effect on relative patterns. For both 1967 and 1970, participation rates are highest for the same demographic and geographic groups. Female-headed families—especially those headed by younger females, nonwhites, central city residents, and families located in the Northeast and West—have maintained a higher probability of being on AFDC than their counterparts. This pattern is not unexpected and in general accords with the pattern of caseload growth over the decade. Albeit all geographic and demographic groups demonstrated high rates of caseload growth between 1967 and 1970, these groups, nonwhites excepted,[34] exhibited high rates of growth in the earlier part of the decade as well.

[34] The AFDC administrative statistics on race are very difficult to interpret given the volatile changes in the unknown category, which accounted for about 10 percent of the growth in caseloads between 1967 and 1970, and the incomplete data on the Latin American minority group.

That these prolonged, high rates of growth involve substantial absolute numbers is evidenced by the change in the distribution of the caseload over the decade. In 1961, 58 percent of the caseload was located in metropolitan areas, 38 percent was located in the Northeast and West, and 63 percent was female headed. By 1971, the comparable statistics were 74 percent in metropolitan areas, 50 percent in the Northeast and West, and 78 percent female headed. The circumscribed nature of the 1961–67 caseload increase is most dramatic for the place of residence characteristic. All of the new growth of the period can be accounted for by the growth within metropolitan areas.

The high absolute levels of participation for central cities (118%), female heads under 25 (134%) non-whites (106%), and the West (106%) in 1970 suggest that for these groups saturation of the eligible pool is being approached or has been attained. Such a conclusion, however, should be treated as highly tentative, given the aforementioned problems of deriving the detailed estimates. The estimates for all female heads, on the other hand, are more reliable since they make up a large proportion of the total estimate. Hence, they allow some slightly less tentative speculation regarding future caseload growth. The 90 percent level of participation for female heads in 1970 is a fairly reliable indication that a high proportion of eligible female-headed families are now receiving AFDC benefits and that any substantial future growth from this group will have to come from the creation of new eligibles.

To a limited extent, new eligibles appear to have played a roll in caseload growth between 1967 and 1970. In 1967, 2.2 million female-headed families were eligible for AFDC, while approximately 1.4 million received benefits some time during the year—leaving close to 800,000 non-participants. But between 1967 and 1970 the number of female-headed recipient families increased by about one million, suggesting that at least 200,000 of the total 531,000 "new eligibles" (see table 1) had to join the welfare rolls between 1967 and 1970. Although the magnitude of creation indicated here is not overwhelming, the issue is one of crucial policy importance.

That the structure of the AFDC program provides incentives for intact families to split and for single females to have illegitimate children is well documented.[25] Although no definitive evidence exists to show that this is in fact happening, it is generally agreed that a program which provides such incentives is undesirable. On the other hand, new eligibles resulting from program changes that simply expand eligibility up the income scale or from mere population growth can hardly be viewed as disturbing. The figures in table 4 suggest that the expansion of income eligibility was the dominant phenomenon over the 1967–70 period.

[25] For brief discussion, see The President's Commission on Income Maintenance Programs, *Background Papers*, U.S. Government Printing Office, 1970, pp. 276–277. For more in depth analysis, see Elizabeth Durbin, *Welfare Income and Employment: An Economic Analysis of Family Choice*, Praeger, New York, 1969.

TABLE 4.—*Female-headed families eligible for AFDC, 1967 and 1970*

Family money income excluding public assistance	Female-headed families categorically eligible for AFDC[1] 1967 (thousands)	1970 (thousands)	Percent change, 1967–70	Percent of total change[3]	Female-headed families economically eligible for AFDC[2] 1967 (thousands)	1970 (thousands)	Percent change, 1967–70	Percent of total change[3]
0	530	735	39	35	530	735	39	39
0 to $999	524	568	08	08	524	568	08	08
$1,000 to $1,999	534	472	−12	−11	470	465	01	−01
$2,000 to $2,999	415	384	−07	−05	268	324	21	11
$3,000 to $3,999	467	380	−19	−15	189	243	29	10
$4,000 to $4,999	370	377	02	01	88	168	91	15
$5,000 to $5,999	275	379	38	18	49	86	76	07
$6,000 to $6,999	229	276	21	08	28	67		07
$7,000 to $7,999	132	217	64	15	17	29		02
$8,000 to $8,999	70	135	93	11	9	10		00
$9,000 to $9,999	33	93	182	10	2	10	98	02
$10,000 to $11,999	44	103	134	10	7	5		00
$12,000 to $14,999	23	77	235	09	0	3		01
$15,000+	18	54	200	06	0	1		00
Total	3,665	4,249	16	100	2,183	2,714	24	100

[1] All female-headed families with children regardless of income are considered categorically eligible for AFDC. The definition of family in this table is equivalent to the AFDC definition of assistance unit. (See page 6 on categorical eligibility).

[2] Female-headed families with children that meet the various State income tests of eligibility are considered economically eligible for AFDC.

[3] Absolute change 1967–70 in each income class is computed as a percent of total absolute change 1967–70.

While the number of categorically eligible female-headed families below $6,000 increased by 180,000 between 1967 and 1970, those under $6,000 that were economically eligible for AFDC increased by 470,000—accounting for 89 percent of the total increase in economically eligible female-headed families.[26] In fact, the largest percentage and absolute increases in categorical eligibility—i.e., female-headed families with minor children regardless of income—occurred at the upper end of the income scale where the incentives of AFDC are least likely to have an effect. Categorical eligibles increased by 74 percent above $6,000 but by less than 6 percent below $6,000. Even in absolute terms, the higher income group accounted for 70 percent of the overall increase in categorical eligibility. These figures suggest that the dominant phenomenon accounting for the growth of AFDC eligibles between 1967 and 1970 was not the creation of female-headed families but rather expansion of income eligibility. At the same time the figures do not prove the absence of a behavioral link between AFDC and the creation of female-headed families, as the aggregate data may conceal significantly different statistical patterns for subgroups of the total population.

The aggregate data does, however, reveal some interesting changes in the income distribution below $6,000. The $1,000 through $4,000 income classes exhibit absolute declines in the number of families, while the 0 and $5,000 to $6,000 categories increased by 38 and 39 percent respectively. One possible interpretation of these data is that a substantial number of female-headed families with low incomes either improved their economic status and moved up the income scale or stopped working and went on welfare, since it seems reasonable to assume that most of those with no other income are receiving AFDC. Again the magnitude of the numbers involved is not large—205,000 new eligibles in the zero income class—but the issue is one of significant policy concern. The implied trend is that female heads of families are choosing total dependence on AFDC rather than combining work and welfare as was the intent of the work incentive legislation of the 1967 amendments to the Social Security Act. Such a tentative hypothesis is consistent with the observed trend in welfare caseloads. Although caseloads doubled between 1967 and 1970 there was no significant change in the proportion of welfare mothers who worked. By the end of 1970, 13.9 percent of welfare mothers were employed either full or part-time, up only 0.2 percentage points from the end of 1967.[27] It should be emphasized that this interpretation of the data is offered here, not as a conclusion, but as a hypothesis for future investigation. Without more detailed analysis it is impossible to describe with any precision the dynamics of the change being observed in the data.

[26] Note that the measure of income being used here excludes public assistance income.

[27] The 1967 and 1971 AFDC studies show that in January of 1971, 13.9 percent of welfare mothers were employed either full or part-time up only 0.2 percentage points from the end of 1967.

Appendix A

A COMPUTATIONAL MODEL FOR ESTIMATING AFDC ELIGIBLES

The Basic Model for Data Year 1970

SECTION I. SPECIFICATION OF DATA BASE

The basic model is specifically designed for the 1971 Current Population Survey tape files for data year 1970. Adaptions of the basic model to other data years are specified in the Adaptions section of the specifications. All variable codes refer to the CPSEO Codebook prepared by Lou Koenig of The Urban Institute, available at The Urban Institute as working paper 718.2.

SECTION II. GENERAL SEQUENCE OF THE CORE MODEL

To be fully understood, this section should be reread after reading all other sections of the specifications.

1. First check each CPS primary, secondary, and sub-family for the presence of minor children as defined in Section III.2.

2. If minor children are present, then determine if the family or persons within the family meet the criteria of one of the six different types of categorical units defined in Section IV. Determination of categorically eligible units should be made according to the sequence outlined in Section IV.

3. Next compute an annual payment for all types of categorical units except "other related" child units (children living with relatives other than parents. See Section IV.5.) as directed by Section V.1. The payments calculations of Section V are based on an annual aggregation of all variables and program parameters. Estimates for work related and child care expenses may be included or excluded from these calculations for any one computer run. Once the payments have been computed apply the full standard test of Section V.2 to all units that received a positive payment. The application of the payments formula and the full standard test will enable the identification of the following populations:

(*i*) Categorically eligible units that receive a positive payment.

(*ii*) Categorically eligible units that receive a positive payment and are not screened out by the full standard test.

(*iii*) Categorically eligible units that receive a positive payment and are screened out by the full standard test.

(*iv*) Categorically eligible units that do not receive a positive payment.

4. When the above procedures have been completed, the alternative method of calculating payments and the full standard test described in Section VI will then be applied to the same set of categorical units. The payments calculations of Section VI disaggregate the annual variables and program parameters into two parts corresponding to the intra-year variation in the flow of earnings of the different types of categorical units.

Because of the dual payments and full standard test of Section VI, some units may lose one payment but maintain the other. For these units total annual payments must be recomputed after the full standard test to reflect the partial loss.

Again, payments are calculated first and then the full standard tests are applied to all units that receive a positive payment. These procedures will identify the following populations:

(*i*) Categorically eligible units that receive a positive payment.

(*ii*) Categorically eligible units that receive a positive payment and are not screened out by the full standard test.

(*iii*) Categorically eligible units that receive a positive payment but are screened out by the full standard tests.

(*iv*) Categorically eligible units that do not receive a positive payment.

5. Next impute payments to all categorically eligible units that are comprised exclusively of "other related" children (defined IV.5.) as directed by Section VII.1. Then apply the economic screen to the families with whom they reside as described in Section VII.2. These procedures will enable the identification of the following populations:

(*i*) All "other related" child units.

(*ii*) All "other related" child units living with families that have limited resources as determined by the economic screen.

SECTION III. DEFINITIONS

1. CPS primary families, secondary families, and sub-families will all be analyzed as separate family units.

2. The first requirement of categorical eligibility is that a family contain at least one minor child. The precise definition of a minor child varies by state mainly on the basis of age and school attendance. A simplification of these variations puts all states in one of the two following categories:

2.1 Minor children may be defined as own or other related (with respect to the family head) never married children under 18. In terms of CPSEO variables this definition is as follows:

For a primary family:

(i) Own children of the primary family head under 18 never married, not in a sub-family (23.13=04);

(ii) Other relative of primary family head under 18, never married, not in a sub-family (23.13=16).

For a secondary family:

(iii) Own children of the secondary family head under 18, never married (23.13=27);

(iv) Other relatives of secondary family head under 18, never married (23.13=28).

For a sub-family:

(v) Own never married children under 18 (23.13=15, 13). This definition will be applied to the following states (11.14): Georgia (58); Mississippi (64); Missouri (43); Alaska (94); Delaware (51); Florida (59); Wisconsin (35); Connecticut (16); Minnesota (41); South Dakota (45)*.

Source: U.S. Dept. of Health, Education and Welfare, Social Rehabilitation Service, Assistance Payments Administration, *Characteristics of State Public Assistance Plans: General Provisions—Eligibility, Assistance, Administration*, 1970 Edition, Washington, D.C.

2.2 Alternatively, minor children may be own or other related (with respect to the family head) never married children under 18 and 18–20 if attending school. In terms of CPSEO variables this definition is as follows:

For a primary family minor children are:

(i) Own children of the primary family head under 18, never married, not in a sub-family (23.13=04).

(ii) Own children of the primary family head 18 and over, never married, not in a sub-family (23.13=09), if age is 18–20 (23.31=18, 19, 20) and attending school. Attending school will be determined by the following method: worked part-year, part or full time because of school (24.34=4) or did not work last year because of school (24.35=3).

(iii) Other relative of primary family head under 18, never married, not in a sub-family (23.13=16).

(iv) Other relative of primary family head 18 and over, never married, not in a sub-family (23.13=21), if age is 18–20 (23.31=18, 19, 20) and attending school (23.34=4 or 24.35=3).

For a secondary family minor children are:

(v) Own children of the secondary family head under 18, never married (23.13=27).

(vi) Other relatives of secondary family head under 18, never married (23.13=28).

(vii) Other relatives of secondary family head 18 and over never married (23.13=30) if age is 18–20 (23.31=18, 19, 20) and attending school (24.34=4 or 24.35=3).

For a sub-family:

(viii) Own children of the sub-family head, never married, and under 18 (23.13=15, 13).

*State codes for these States have been assigned because they are not identified in the CPS file. (See section XIII.)

This definition applies to the following States (11.14):

Arizona (86)
Hawaii (95)
Idaho (82)*
Kansas (47)
Maine (11)*
Maryland (52)
Massachusetts (14)
Michigan (34)
Nebraska (46)*
New York (21)
Ohio (31)
Oregon (92)
Pennsylvania (23)
Rhode Island (15)
South Carolina (57)
Tennessee (62)
Texas (74)
Kentucky (61)
Nevada (88)*
Oklahoma (73)
Arkansas (71)
Utah (87)
W. Virginia (55)
Wyoming (83)*
D.C. (53)
New Jersey (22)
California (93)
Alabama (63)
Illinois (33)
Montana (81)*
Washington (91)
Vermont (13)*
North Dakota (44)*
Louisiana (72)
New Hampshire (12)*
New Mexico (85)
Virginia (54)
Colorado (84)
Indiana (32)
North Carolina (56)
Iowa (42)

Source: Same as for subsection 2.1 above.

3. UP States in data year 1970 were as follows: Var. 11.14 = Cal. (93) Colo. (84); Del (51); D.C. (53); Hawaii (95); Ill. (33); Kan. (47); Maine (11)*; Md. (52); Mass. (14); Mich. (43); Mo. (43); Nebr. (46)*; N.J. (22); N.Y. (21); Ohio (31); Okla. (73); Oreg. (92); Pa. (23); R.I. (15); Utah (87); Vt. (13)*; Wash. (91); W. Va. (55).

Source: Same as for sub-section 2.1 above.

4. The nine census divisions are:

(i) *New England:* Maine (11)*, New Hampshire (12)*, Vermont (13)*, Mass. (14), R.I. (15), Conn. (16).

(ii) *Middle Atlantic:* N.Y. (21), Penn. (22), N.J. (23).

(iii) *South Atlantic:* Del. (51), D.C. (53), W. Va. (55), Va. (54), N.C. (56), S.C. (57), Ga. (58), Fla. (59), Md. (52).

(iv) *East South Central:* Ky. (61), Tenn. (62), Miss. (64), Ala. (63).

(v) *West South Central:* Okla. (73), Ark. (71), Texas (74), La. (72).

(vi) *East North Central:* Mich. (34), Ohio (31), Ind. (32), Ill. (33), Wis. (35).

(vii) *West North Central:* Minn. (41), Iowa (42), Mo. (43), Kans. (47), Nebr. (46)*, S. Dak. (45)*, N. Dak. (44)*.

(viii) *Mountain:* Mont. (81)*, Wyo. (83)*, Idaho (82)*, Nev. (88)*, Utah (87), Colo. (84), Ariz. (86), N. Mex. (85).

(ix) *Pacific:* Wash. (91), Oreg. (92), Calif. (93), Alaska (94)*, Hawaii (95).

5. Adults: Adults are all persons not classified as minor children including all unrelated individuals.

6. Race: White (23.32 = 1.4); Non-white (23.32 = 2, 3, 5, 6).

7. Place of Residence: SMSA, center city (12, 21 = 1), SMSA, not in center city (12.21 = 2); Not in SMSA (12.21 = 3).

8. Census Region: Northeast (11.13 = 1), North Central (11.13 = 2), South (11.13 = 3), West (11.13 = 4).

SECTION IV. CATEGORICAL ELIGIBILITY

1. Female headed units: If a family contains minor children the head is female, then the head and all own and other related minor children constitute a categorically eligible AFDC unit. All other persons in the family are excluded. Call these units AFDC–FH.

*State codes for these States have been assigned because they are not identified in the CPS file. (See section XIII.)

2. Unemployed father units: If a family contains minor children, lives in a state with an AFDC–UP program, and the head is a male eligible for the AFDC–UP program then the head, spouse, and all own and other related minor children constitute a categorically eligible AFDC unit. All other persons in the family are excluded. Call these units AFDC–UP.

2.1 Fathers will be considered eligible for the AFDC–UP program if they worked part-year (24.33–4, 3) and the reason for part-year work was looking (24.34=1). They must also have worked less than 40 weeks last year (24.31–2, 3, 4).

3. Incapacitated father units: If a family contains minor children, and the head is a male who fits the definition of an incapacitated father then the head, spouse, and all own and other related minor children constitute a categorically eligible AFDC unit. All other persons in the family are excluded.

3.1 If the father works zero weeks (24.31=1) because of illness (24.35=1) he will be considered incapacitated. Call these units AFDC–IF_1.

3.2 If the father worked part-year (24.33=3, 4) because of illness (24.34=2) and weeks worked is less than 40 (24.31=2, 3, 4) he is considered incapacitated. Call these units AFDC–IF_2.

4. Male headed single parent units: If the family contains minor children and the head is a male with no spouse present (23.37=1, 3, 5, 6, 7) then the head and all own or related minor children constitute a categorically eligible AFDC unit. All other persons in the family are excluded. Call these units AFDC–SP.

5. Other related child units: If a family contains at least one other related minor child and is not categorically eligible under rules 1–4, then the other related minor children, by themselves, constitute a categorically eligible AFDC unit. Call these units AFDC–RC.

SECTION V. ANNUAL PAYMENTS FORMULA AND FULL STANDARD TEST

1. Payments Formula:

$$P = r^d\,(pS - Y^u - max\,[O, r^e\,(Y^e - Y^D) - u_1\,WRE - u_2\,CCE])$$

Subject to the constraint that if $P > M$ set $P = M$, and if $P < O$ set $P = O$.

1.1 Where the variables P, Y^u, Y^e, WRE, CCE, u_1, u_2 are defined as follows:

P is the annual AFDC payment computed on the basis of an annual accounting period.

Y^u is the sum of annual unearned income less public assistance (p.a.) over all members of the AFDC filing unit. For each person annual unearned income less p.a. is the sum of CPSEO variables 25.14, 25.15, 25.17, and 25.18.

Y^e is the sum of earned income of the head and spouse (if present). For each person earned income is the sum of CPSEO variables 25.11, 25.12, and 25.13.

WRE and CCE represent estimates of work related expenses and child care expenses for weeks worked. WRE and CCE are estimated only if Y^e is positive. The estimating procedures for WRE and CCE may vary by type of categorical unit and by the method of calculating payments. Precise rules for estimating WRE and CCE are outlined in Sections IX and X.

u_1 and u_2 are parameters which may be set equal to 0 or 1 so that the model has the option of including or excluding estimates for WRE and CCE.

1.2 The program parameters r^d, p, r^e, Y^D, M, and S are defined as follows:

r^d is the budget deficit reduction rate

p is the state standard reduction rate

S is the state full standard which varies by family size

r^e is the tax rate on earnings, which has a value of .67

Y^D is the earnings disregard which has an annual value of 360

M is the state maximum payment which varies by family size

The values of the program parameters r^d, p, S and M for 1970 are given in Section XIV, Tables I, II, and III. Rule for estimating S by family size are outlined in Section XII.

2. Full Standard Test

All categorical units that receive a positive payment according to the formula above will then be subjected to the full standard test. Application of the full standard test is in effect a more stringent economic screen than the payments formula. The population of categorical eligibles that receives a positive payment and passes the full standard test should be smaller than the population that receives positive payments. Apply the full standard test as follows: If $S \geq Y^u + \max\,(O, Y^e - u_1\,WRE - u_2\,CCE)$ for a given unit, the unit passes the test. WRE and CCE may be included or excluded but the treatment must be consistent with the payments formula.

SECTION VI. PART-YEAR PAYMENTS FORMULAS AND FULL STANDARD TESTS

The part-year payments formulation involves calculations for different parts of the year depending on the work experience of the head or spouse of the categorical unit. Two different methods of computing payments will be employed depending on the type of categorical unit. For *AFDC–FH*, *AFDC–SP* and *AFDC–IF*$_1$ separate payments calculations and full standard tests will be applied for weeks worked (W) and weeks not worked (NW) of the principal earner—defined as the head for *AFDC–FH* and *AFDC–SP* and the wife for *AFDC–IF*$_1$. This method is described in Section 1 below and shall be termed the "weeks worked calculations."

For *AFDC–UP* and *AFDC–IF*$_2$ payments calculations and means test will be made only for that portion of the year that the unit are categorically eligible. Categorical eligibility (see Section III) for these units is determined by weeks not worked of the principal earner—defined as the head in both cases. This method is described in Section 2 below and shall be termed "weeks eligible calculations."

1. The "weeks worked calculation" involves separate payments for weeks worked (W) and weeks not worked (NW) as determined in Section VII. The total annual AFDC payment is represented by the following formula: $P^* = P_1 + P_2$

Where: P^* is the annual *AFDC* payment computed on the basis of a part-year accounting period.

P_1 is the payment for weeks worked.

P_2 is the payment for weeks *not* worked.

1.1 Weeks Worked Payment Formula

Weeks worked (W) refers to the principle earner=head for *AFDC–FH* and *AFDC–SP* units, but spouse for *AFDC–IF*$_1$ units. In the latter case, if spouse is absent set $W=0$ and hence $NW=52$, suppressing the weeks worked calculations.

$P_1 = r^d\,(apS\text{-}aY^{u*} - \max\,[0,\ r^e\,(Y^e - aY^D) - u_1\,WRE - u_2\,CCE])$

Subject to the constraints that if—

$P_1 > aM$ set $P_1 - aM$ and if $P_1 < 0$ set $P_1 = 0$.

Where:

$a = W/52$

y^{u*} is unearned income as defined above except all unemployment insurance (UI) of principal earner—as determined in Section XI—is excluded.

y^e is earnings of the principal earner.

WRE, if included, should be estimated for the W of the family earner.

CCE, if included, should be estimated for *AFCD–FH* and *AFDC–SP* units only as outlined in Section X.

1.2 Weeks Not-Worked Payments Formula: $P_2 = r^d\ (1-a)\ [pS - Y^{u*}]\text{-}UI)$

Subject to the constraints that if $P_2 > (1-a)M$ set $P_2 = (1-a)M$ and if $P_2 < 0$ set $P_2 = 0$.

Where: UI is total amount of unemployment insurance of family earner.

1.3 P^* before full standard test.

Compute P^* on basis of estimates of P_1 and P_2 obtained from sec's 1.1 and 1.2. above.

1.4 Weeks Worked Full Standard Test Unit passes the full standard test if:

$aS \geq aY^{u*} + \max(0, Y^e - u_1\ WRE - u_2\ CCE)$

1.5 Weeks Not-Worked Full Standard Test

The unit passes if:

$(1-a)S \geq (1-a)\,Y^{u*} + UI$

1.6 P* after the full standard test

Because of the dual payments and full stanadrd tests of this section, the full standard tests may screen out one payment but not the other. To reflect this possible reduction, P* must be recomputed after the application of the full standard tests. In this recomputation P_1 and/or P_2 should be set equal to zero if the unit fails the corresponding full standard test. $P^{*\prime} = P_1 + P_2$

Where: $P^{*\prime}$ is the annual AFDC payment computed on the basis of a part year accounting period, after the full standard tests.

P_1 and/or P_2 is set equal to zero if the unit fails the corresponding full standard test.

2. The "weeks eligible calculation" involves a payments calculation for AFDC–UP and AFDC–IF$_2$ units for the weeks not worked (NWH) of the principal earner/head. Where NW_H corresponds to the number of weeks these units are categorically eligible. Because of the "worked, less than 40 weeks" requirement for categorical eligibility (see Section III) NW_H must be >12.

2.1 Payments Formula:

$P^* = r^d(bpS - bY^{u*} - UI - \max\ [0, r^e(cY^e{}_w - bY^D) - u_1\ WRE])$ Subject to the constraint that if $P^* > bM$ set $P^* = bM$ and if $P^* < 0$ set $P^* = 0$.

Where: P^* is the annual AFDC payment based on part-year computations.
b is the fraction of the year that the principal earner did not work i.e. fraction of year unit is eligible. $b = \frac{(NW_H)}{52}$

Y^{u*} is the units annual unearned income excluding UI of principal earner/head.

UI is the total annual UI of principal earner; c is the proportion of the wife's annual earnings that will be counted as earned during the weeks eligible.

$c = 1$ if $W_w < NW_H$, but

$c = NW_H/W_w$ if $W_w > NW_H$.

where W_w is wife's weeks worked.

Y_w^e is the total annual earnings of the wife WRE, if included should be estimated for W_w or NW_H whichever is less.

CCE are not allowed for this types of unit.

2.2 Full Standard Test

Unit passes the full standard test if $bS \geq by^{u*} + UI + \max\ (0, cy_w^e - u_1\ WRE)$.

SECTION VII. ECONOMIC TREATMENT OF RELATED CHILD UNITS (AFDC-RC)

The economic treatment of AFDC-RC units is unique from that of other units—the payments formulas and full standard tests outlined in the previous sections will not be applied to AFDC-RC units.

Rather:

1. Impute to all AFDC-RC units an average payment which varies by census division and number of children. See Section XIV, Table IV for a schedule of average payments for 1970.

2. To screen out AFDC-RC payments to units that reside with higher income families the following economic screen will be applied to the relative families—counting the other related children and their unearned income as part of the family and family income respectively.

2.1 Apply the annual payments formula of Section V to the family (family in terms of CPS definitions) with whom the AFDC-RC unit resides. Count Y^u of all family members but Y^e of only adults i.e. exclude Y^e of any minor children. Estimate WRE for the weeks worked (W) of the adult who had the greatest number of weeks worked. No CCE are allowed. The number of AFDC-RC units living with families who receive positive payments according to these calculations provides a rough measure of the proportion of all AFDC-RC units that are in financial need.

SECTION VIII. DETERMINATION OF WEEKS WORKED (W) AND WEEKS NOT WORKED (NW)

1. W will be determined by midpoints of the CPSEO variable (24.31) for weeks worked last year, according to the following schedule:

Value of 24.31	Weeks worked	Midpoint
1	0	0
2	1 to 13	7.0
3	14 to 26	20.0
4	27 to 39	33.0
5	40 to 47	43.5
6	48 to 49	48.5
7	50 to 52	[1] 52.0
8	Full year	52.0

[1] Arbitrarily set midpoint equal to 52 to prevent part-year calculations for persons who worked 50 to 52 weeks.

2. NW is simply 52−W.

SECTION IX. IMPUTATION OF WORK RELATED EXPENSES (WRE)

1. Estimating procedures for annual payments calculations of Section V. If Y^e is positive impute WRE as follows:

$$WRE = \frac{(MWRE_i X12)W}{52}$$

Where: $MWRE_i$ is the average monthly WRE for census division i as reported by the 1971 AFDC Survey. See Section XIV, Table V for a schedule of MWRE by census division.

W is the weeks worked of the head or spouse whichever is greater. (See Section VII for method of determining numerical value of W.)

2. Estimating procedures for "weeks worked calculations" of Section VI.1. If Y^e is positive, impute WRE as above except W refers to weeks worked of head for AFDC–FH and AFDC–SP and to weeks worked of spouse for AFDC–IF_1.

3. Estimating procedures for "weeks eligible calculations" of Section VI.2. If Y^e of *spouse* is positive, impute WRE as above except W is weeks worked of spouse or NW_H whichever is less.

SECTION X. IMPUTATION OF CHILD CARE EXPENSES (CCE)

Impute CCE if:

(*i*) Unit is AFDC–FH or AFDC–SP (i.e. no other types of units will be allowed CCE).

(*ii*) Unit is a primary family (i.e. sub-families and secondary families will not be allowed CCE).

(*iii*) Unit has children under 12.

(*iv*) There is no adult present in the household (census definition) who worked zero weeks last year (24.31=1).

For both the annual (Sec. IV) and part-year (Sec. V) payments and full standard tests:

$$CCE = \frac{(MCCE_i X12)}{52} W$$

Where: $MCCE_i$ is average monthly child care expenses for census division i. See Section XIV, Table V for values of MCCE by census division as reported by the 1971 AFDC Survey.

W is weeks worked of head.

SECTION XI. ALLOCATION OF UNEARNED INCOME SOURCE D TO UNEMPLOYMENT INSURANCE

1. If 25.17 (Unearned Income Source D: unemployment insurance, workmen's compensation, government employee's pension, and veteran's payments) equals 0, then the dollar amount of unemployment insurance equals 0 also.

2. If 25.17 equals a positive amount and 25.19 (variable which designates detailed sources of unearned income) equals 7 (unemployment insurance) and not 4 (government employee's pensions), 5 (workmen's compensation), or 6 (veteran's payments), then the dollar amount of unemployment insurance equals the full positive amount of 25.17.

3. If 25.17 equals a positive amount and 25.19 equals 7 and 4, 5 or 6 or any combination of 7 and the latter three, then the dollar amount of unemployment insurance will be determined as follows:

(*i*) If 25.19 = 7 and 4, then *17%* of $ amount of 25.17 is designated UI.
(*ii*) If 25.19 = 7 and 5, then *43%* o. $ amount of 25.17 is designated UI.
(*iii*) If 25.19 = 7 and 6, then *51%* of $ amount of 25.17 is designated UI.
(*iv*) If 25.19 = 7, 4 and 5, then *14%* of $ amount of 25.17 is designated UI.
(*v*) If 25.19 = 7, 4 and 6, then *15%* of $ amount of 25.17 is designated UI.
(*vi*) If 25.19 = 7, 5 and 6, then *30%* of $ amount of 25.17 is designated UI.
(*vii*) If 25.19= 7, 4, 5, and 6, then *12%* of $ amount of 25.17 is designated UI.

The above percentages were determined by taking the ratio of the annual mean unemployment insurance payment to the sum of the annual mean payments of the various combinations of types of Source D unearned income. For example, if the codes of variable 25.19 indicate the receipt of unemployment insurance (7), government pensions (4) and workmen's compensation (5), then the annual mean payments of unemployment insurance, government pensions, and workmen's compensation were summed and the annual mean unemployment insurance payment is computed as a percentage of that sum. The resulting percentage is then applied to the positive dollar amount of variable 25.17 to determine the proportion to be allocated to unemployment insurance. The mean payments in 1970 for each type of Source D income were:

Unemployment Insurance = $2,926
Workmen's Compensation = 600
Government Pensions = 575
Veterans Payments = 804

Source: U.S. Dept. of Health, Education, and Welfare, Social Security Bulletin Annual Statistical Supplement. 1969." U.S. Government Printing Office, Washington, D.C.

SECTION XII. ESTIMATION OF STATE STANDARDS BY ASSISTANCE UNIT SIZE

Given published data for monthly values of S by states for assistance units for size 2 and 4, annual values of S by states for assistance units of size 3, 5, 6, 7, 8, 9, and 10 will be estimated as follows:

$$S_{ik}=(S^m_{i4}+(k-4)\Delta_i)\cdot 12$$

where:

S_{ik} is annual S for state i and family size k.
S^m_{i4} is monthly S for state i and family size 4.
k is family sizes 3 and 5–10.
Δ_i is an estimate of state i's increase in monthly S for each additional person in the assistance unit.

Estimate Δ_{i1} as follows:

$$\Delta_i=\frac{S^m_{i4}-S^m_{i2}}{2}$$

where S^m_{i2} is monthly S for state i and family size 2.

Values of S^m_2, S^m_4, and Δ by states are given in Section XIV, Table II.

SECTION XIII. ALLOCATION OF POPULATION TO UNIDENTIFIED STATES (STATES WITH CODE 99)

Because the CPS does not specify state codes for the residents of eleven states, the following procedures were devised to randomly allocate families to specific states on the basis of population. Population percentages reflect the distribution of population in 1970, as reported by the *Statistical Abstract of the United States 1971*, Table 12, p. 14.

A. Maine, Vermont, New Hampshire.

If a family lives in the Northeast region (11.13 = 1) and the state code is 99 (Maine, New Hampshire, and Vermont) randomly assign:

1. 46% of such families to Maine and assign the state code 11.
2. 34% of such families to New Hampshire and assign the state code 12.
3. 20% of such families to Vermont and assign the state code 13.

B. South Dakota, North Dakota, and Nebraska.

If a family lives in the North Central region (11.13 = 2) and the state code is 99 (South Dakota, North Dakota, and Nebraska) randomly assign:

1. 22% of such families to North Dakota and assign the state code 44.
2. 24% of such families to South Dakota and assign the state code 45.
3. 54% of such families to Nebraska and assign the state code 46.

C. Montana, Idaho, Wyoming, Nevada, and Alaska.

If a family lives in the West region (11.13 = 4) and the state code is 99 (Montana, Idaho, Wyoming, Nevada, Alaska) randomly assign:

1. 28% of such families to Montana and assign the state code 81.
2. 28% of such families to Idaho and assign the state code 82.
3. 13% of such families to Wyoming and assign the state code 83.
4. 19% of such families to Nevada and assign the state code 88.
5. 12% of such families to Alaska and assign the state code 94.

TABLE I.—*Values of r^d and p by State as of July 1970 (r^d and p do not vary by family size)*

State	r^d	p	State	r^d	p
Alabama	0.35	1.0	Montana	1.0	1.0
Alaska	1.0	1.0	Nebraska	1.0	1.0
Arizona	.65	1.0	Nevada	1.0	1.0
Arkansas	1.0	1.0	New Hampshire	1.0	1.0
California	1.0	1.0	New Jersey	1.0	1.0
Colorado	1.0	1.0	New Mexico	.9	1.0
Connecticut	1.0	1.0	New York	1.0	1.0
Delaware	1.0	1.0	North Carolina	.86	1.0
District of Columbia	1.0	.75	North Dakota	1.0	1.0
Florida	.60	1.0	Ohio	1.0	.78
Georgia	1.0	1.0	Oklahoma	.85	1.0
Hawaii	1.0	1.0	Oregon	1.0	.80
Idaho	1.0	1.0	Pennsylvania	1.0	1.0
Illinois	1.0	1.0	Rhode Island	1.0	1.0
Indiana	1.0	1.0	South Carolina	.52	1.0
Iowa	.812	1.0	South Dakota	1.0	1.0
Kansas	1.0	1.0	Tennessee	1.0	1.0
Kentucky	.87	1.0	Texas	1.0	.75
Louisiana	.51	1.0	Utah	1.0	1.0
Maine	1.0	1.0	Vermont	1.0	.895
Maryland	1.0	1.0	Virginia	1.0	.90
Massachusetts	1.0	1.0	Washington	1.0	1.0
Michigan	1.0	1.0	West Virginia	1.0	.52
Minnesota	1.0	1.0	Wisconsin	1.0	.829
Mississippi	.3	1.0	Wyoming	1.0	1.0
Missouri	1.0	1.0			

Source: Tables 2 and 3 of "State Maximums and Other Methods of Limiting Money Payments," NCSS report D-3 (10/70). Published by the U.S. Department of Health, Education, and Welfare, National Center for Social Statistics.

TABLE II.—Values of $S^m{}_2$, $S^m{}_4$, and Δ by States for 1970

State	$S^m{}_2$	$S^m{}_4$	Δ	State	$S^m{}_2$	$S^m{}_4$	Δ
Alabama	148	230	41.0	Montana	143	250	53.5
Alaska	300	400	50.0	Nebraska	235	330	47.5
Arizona	164	256	46.0	Nevada	220	317	48.5
Arkansas	122	176	27.0	New Hampshire	221	294	36.5
California	289	432	75.5	New Jersey	234	347	56.5
Colorado	153	235	41.0	New Mexico	135	203	34.0
Connecticut	200	330	65.0	New York	219	336	58.5
Delaware	181	287	53.0	North Carolina	147	184	18.5
District of Columbia	181	280	49.5	North Dakota	192	284	46.0
Florida	143	223	40.0	Ohio	173	258	42.5
Georgia	148	208	30.0	Oklahoma	141	218	38.5
Hawaii	189	263	37.0	Oregon	180	281	50.5
Idaho	206	272	33.0	Pennsylvania	218	313	47.5
Illinois	228	282	27.0	Rhode Island	202	263	30.5
Indiana	203	322	58.0	South Carolina	123	198	37.5
Iowa	186	300	57.0	South Dakota	220	300	40.0
Kansas	166	267	50.5	Tennessee	142	217	37.5
Kentucky	156	264	54.0	Texas	150	239	44.5
Louisiana	130	213	41.5	Utah	182	217	44.5
Maine	205	349	72.0	Vermont	241	327	43.0
Maryland	183	302	59.5	Virginia	196	279	41.5
Massachusetts	210	314	52.0	Washington	228	303	37.5
Michigan	181	263	41.0	West Virginia	186	265	39.5
Minnesota	302	299	48.5	Wisconsin	189	255	33.0
Mississippi	168	232	32.0	Wyoming	177	277	50.0
Missouri	230	325	47.5				

Source: Table 3 of "OAA and AFDC: "Standards for Basic Needs for Specified Types of Assistance Groups"; July 1970; NCSS Report D-2 (7/70), U.S. Dept. of Health, Education, and Welfare.

TABLE III.—*State monthly[1] maximums (M) by number of persons in assistance unit for 1970*

State	1	2	3	4	5	6	7	8	9	10
Alabama	50	80	110	140	170	170	170	170	170	170
Alaska	125	175	225	275	325	375	425	475	525	525
Arkansas	81	91	101	111	121	131	141	146	146	146
California	148	162	211	253	290	320	345	363	376	382
Colorado	153	194	235	278	321	352	383	411	445	476
Delaware	125	137	149	161	171	181	191	200	209	218
Georgia	71	102	133	164	161	164	164	164	164	164
Indiana	100	125	130	175	200	225	250	275	300	325
Kentucky	320	320	320	320	320	320	320	320	320	320
Maine	76	109	142	175	208	241	274	307	340	373
Mississippi	30	48	60	72	84	96	108	108	108	108
Missouri	80	105	130	155	180	205	230	255	280	305
Nebraska	110	140	170	200	230	260	290	320	350	380
New Mexico	200	200	200	200	200	200	200	200	200	200
Oklahoma	141	180	218	249	283	309	332	353	353	353
Tennessee	97	113	129	145	161	161	161	161	161	161
Utah	128	165	189	214	241	274	299	315	331	347
Virginia	305	305	305	305	305	305	305	305	305	305
West Virginia	182	182	182	182	183	182	182	182	182	182
Wyoming	193	227	227	244	244	244	261	261	261	261

[1] Multiply by 12 to obtain annual values.

TABLE IV.—*Average monthly payments for AFDC–RC units for 1970*

Census division	All AFDC–RC units	AFDC-RC units by the number of children in unit					
		1	2	3	4	5	6 or more
New England	113.94	[1] 72.46	[1] 116.03	[1] 152.24	[1] 179.94	[1] 220.65	[1] 213.11
Middle Atlantic	139.15	96.16	154.82	[1] 185.97	[1] 219.81	[1] 269.54	[1] 260.33
East North Central	98.29	55.64	93.08	[1] 131.30	[1] 155.26	[1] 190.39	[1] 183.89
West North Central	89.13	55.01	76.13	111.08	[1] 140.79	[1] 172.65	[1] 166.75
South Atlantic	70.47	43.16	74.08	104.27	[1] 111.32	[1] 136.50	[1] 131.84
East South Central	52.31	37.18	50.06	[1] 69.91	[1] 82.63	[1] 101.33	[1] 97.87
West South Central	54.16	36.65	[1] 55.17	[1] 72.38	[1] 85.55	[1] 104.91	[1] 101.33
Mountain	90.60	59.79	92.29	[1] 121.09	[1] 143.12	[1] 175.50	[1] 169.50
Pacific	124.85	81.12	137.38	198.69	[1] 197.22	[1] 241.84	[1] 233.58
Total	90.85	47.79	92.54	121.42	143.51	175.98	169.97

[1] These data were estimated by making the ratio of the average payment for a unit of a given size and census division to the average payment for all units of the same census division equal to the analogous ratio for the whole United States.

Source: Adapted from table 82 of "Findings of the 1971 AFDC Study, Pt. II. Financial Circumstances," U.S. Department of Health, Education, and Welfare, NCSS, Jan. 12, 1972.

TABLE V.—*Average monthly WRE and average monthly CCE by Census Division for 1970*

Census Division	MWRE	MCEE
New England	51.78	83.16
Middle Atlantic	82.76	74.40
East north-central	53.17	68.34
West north-central	64.25	68.09
South Atlantic	34.36	49.78
East south-central	29.18	43.93
West south-central	31.66	[1] 47.97
Mountain	42.42	----------
Pacific	101.85	72.40

[1] Do not impute any CCE to States in the mountain census division.

Source: Table 72 in "Findings of the 1971 AFDC Survey, pt. II, "Financial Circumstances." U.S. Department of Health, Education, and Welfare, NSCC, Jan. 12, 1972.

ADAPTION OF THE BASIC MODEL TO PROGRAM YEAR 1967

The specifications below outline section by section the changes which must be made to the Basic Model (specified for program year 1970) to estimate AFDC eligibles for 1967. These changes are incorporated into the model as options so the revised model has the capability to estimate eligibles for 1967 and 1970.

SECTION I. SPECIFICATION OF DATA BASE

1968 CPS tape for income data year 1967.

SECTION II. GENERAL SEQUENCE OF THE CORE MODEL

Sequence of the model is unchanged for 1967 except eliminate the full standard tests from both the annual and part year calculations. The full standard test was not in effect for program year 1967.

SECTION III. DEFINITIONS

Definitions are unchanged with the following exceptions:

1. New classification of states by the two alternative definitions of minor children.

Definition 2.1 from Section III of the Basic Model: Alaska (94); Arizona (86); Delaware (51); Florida (59); Georgia (58); Idaho (82)*; Mississippi (64); Missouri (43); Nevada (88)*; South Dakota (45)*.

Definition 2.2 from Section III of the Basic Model: Alabama (63); Arkansas (71); California (93); Colorado (84); Connecticut (16); District of Columbia (53); Hawaii (95); Iowa (42); Kansas (47); Kentucky (61); Louisiana (72); Maine (11)*; Maryland (52); Massachusetts (14); Michigan (34); Minnesota (41); Nebraska (46); New Hampshire (12)*; New Jersey (22); New Mexico (85); New York (21); Ohio (31); Oklahoma (73); Oregon (92); Pennsylvania (23); South Carolina (57); Tennessee (62); Texas (74); Utah (87); Vermont (13)*; Virginia (54); Illinois (33); Indiana (32); Montana (81); North Carolina (56); North Dakota (44)*; Rhode Island (15); Washington (91); West Virginia (55); Wisconsin (35); Wyoming (83)*.

Source: U.S. Dept. of Health, Education, and Welfare, Social and Rehabilitation Service, Assistance Payments Administration, *Characteristics of State Public Assistance Plans: General Provisions—Eligibility, Assistance, Administration*, 1967 Edition, Washington, D.C.

2. New list of UP states: California (93); Colorado (84); Connecticut (16); Delaware (51); Hawaii (95); Illinois (33); Kansas (47); Pennsylvania (23); Rhode Island (15); Utah (87); Washington (91); West Virginia (55); Wisconsin (35); Maryland (52); Massachusetts (14); Michigan (34); Nebraska (46)*; New York (21); Ohio (31); Oklahoma (73); Oregon (92).

Source: Same as for sub-section 1 above.

SECTION IV. CATEGORICAL ELIGIBILITY

No change.

SECTION V. ANNUAL PAYMENTS FORMULA AND FULL STANDARD TEST

Omit the full standard test in simulating program year 1967. The basic structure of the payments formula remains the same. Definition of variables P, Y^u, Y^e, WRE, CCE, u_1 and u_2 are the same.

The general definitions of program parameters r^d, p, r^e, Y^D, M and S are the same but all parameters have new values for 1967. For 1967 r^e, tax rate on earnings, has a value of 1 and Y^D, the income disregard, has a value of 0. 1967 values for r^d, p, S, and M are given in Section XIV, Tables I, II, and III.

SECTION VI. PART YEAR PAYMENTS FORMULA AND FULL STANDARD TESTS

For the "weeks worked calculations" follow the specifications as written for 1970 with the following exceptions:

1. Omit the full standard tests (paragraphs 1.4 and 1.5) and recalculation of P^* after the full standard test (paragraph 1.6).
2. Use 1967 values for parameters r^e, Y^D, r^d, p, S, and M.

For the "weeks eligible calculations" follow specifications as written for 1970 with the following changes for 1967:

1. Omit the full standard test (paragraph 2.2).
2. Use 1967 values for parameters r^e, Y^D, r^d, p, S, and M.

SECTION VII. ECONOMIC TREATMENT OF RELATED CHILD UNITS (AFDC-RC)

Economic procedures for AFDC-RC units are unchanged with the following exceptions:

1. Substitute 1967 schedule of average payments for RC units for imputing payments. See Section XIV, Table IV.
2. In applying the economic screen to the caretaker families use parameter values for 1967.

SECTION VIII. DETERMINATION OF WEEKS WORKED (W) AND WEEKS NOT WORKED (NW)

No change.

SECTION IX. IMPUTATION OF WORK RELATED EXPENSES

General procedures are unchanged. Substitute 1967 input data for values of $MWRE_i$ (average monthly WRE for census division i). See Section XIV, Table V.

SECTION X. IMPUTATION OF CHILD-CARE EXPENSES

General procedures are unchanged. Substitute new 1967 input data for values of $MCCE_i$ (average monthly CCE for census division i). See Section XIV, Table V.

SECTION XI. ALLOCATION OF UNEARNED INCOME SOURCE D TO UNEMPLOYMENT INSURANCE

The general procedures are the same but new allocation percentages are substituted for 1967. (*i*) 19%; (*ii*), 52%; (*iii*), 40%; (*iv*), 16%; (*v*), 15%; (*vi*), 27%; and (*vii*), 13%.

Mean payments for each type of source D income in 1967 were: Unemployment Insurance, 2,362; Workmen's Compensation, 562; Government Pensions, 638; Veterans Payments, 526.

Source: U.S. Dept. of Health, Education, and Welfare, "Social Security Bulletin Annual Statistical Supplement, 1969." U.S. Government Printing Office, Washington, D.C.

SECTION XII. ESTIMATION OF STATE STANDARDS BY ASSISTANCE UNIT SIZE

Procedures for estimating S by family size are the same—using 1967 input data—see Section XIV, Table II.

SECTION XIII. ALLOCATION OF POPULATION TO UNIDENTIFIED STATES (STATES WITH CODE 99)

Procedures are the same substituting new population percentages for 1967 as shown below.

State:	Percent	State:	Percent
Maine	47	Montana	29
New Hampshire	33	Idaho	28
Vermont	20	Wyoming	13
South Dakota	24	Nevada	19
North Dakota	23	Alaska	11
Nebraska	53		

Source: *Statistical Abstract of the United States*, Table 12, p. 14.

SECTION XIV. DATA TABLES FOR 1967

TABLE I.—*Values of r^d and p by State for 1967 (r^d and p do not vary by family size)*

State	r^d	p	State	r^d	p
Alabama	0.50	1.0	Montana	1.0	1.0
Alaska	1.0	1.0	Nebraska	1.0	1.0
Arizona	1.0	1.0	Nevada	1.0	1.0
Arkansas	1.0	1.0	New Hampshire	1.0	1.0
California	1.0	1.0	New Jersey	1.0	1.0
Colorado	1.0	.75	New Mexico	.95	1.0
Connecticut	1.0	1.0	New York	1.0	1.0
Delaware	1.0	1.0	North Carolina	1.0	1.0
District of Columbia	1.0	1.0	North Dakota	1.0	1.0
Florida	1.0	1.0	Ohio	1.0	1.0
Georgia	1.0	1.0	Oklahoma	1.0	1.0
Hawaii	1.0	1.0	Oregon	1.0	1.0
Idaho	1.0	1.0	Pennsylvania	1.0	1.0
Illinois	1.0	1.0	Rhode Island	1.0	1.0
Indiana	1.0	1.0	South Carolina	1.0	1.0
Iowa	.95	1.0	South Dakota	1.0	.90
Kansas	1.0	1.0	Tennessee	1.0	1.0
Kentucky	.87	1.0	Texas	1.0	1.0
Louisiana	1.0	1.0	Utah	1.0	1.0
Maine	1.0	1.0	Vermont	1.0	1.0
Maryland	1.0	1.0	Virginia	1.0	.90
Massachusetts	1.0	1.0	Washington	1.0	1.0
Michigan	1.0	1.0	West Virginia	1.0	.65
Minnesota	1.0	1.0	Wisconsin	1.0	1.0
Mississippi	.27	1.0	Wyoming	1.0	1.0
Missouri	1.0	1.0			

Source: Table 5 of "Money Payments to Recipients of Public Assistance, October 1967," U.S. Department of Health, Education, and Welfare, National Center for Social Statistics, NCSS report D-4 (October 1967).

TABLE II.—*Values of S^m_4 and Δ by States for 1967*

State	S^m_4	Δ[1]	State	S^m_4	Δ[1]
Alabama	177.00	29.5	Montana	219.00	47.0
Alaska	255.47	52.0	Nebraska	276.50	47.5
Arizona	232.00	47.5	Nevada	262.25	52.0
Arkansas	174.00	25.0	New Hampshire	204.00	26.0
California	220.20	57.5	New Jersey	280.00	54.5
Colorado	216.00	41.0	New Mexico	193.00	32.5
Connecticut	257.00	64.5	New York	262.15	54.0
Delaware	236.00	43.5	North Carolina	147.75	13.5
District of Columbia	182.00	15.5	North Dakota	251.00	39.5
Florida	196.00	35.0	Ohio	232.00	29.0
Georgia	187.60	28.5	Oklahoma	163.00	27.5
Hawaii	219.75	31.0	Oregon	203.25	41.5
Idaho	211.60	28.0	Pennsylvania	197.40	36.0
Illinois	181.12	40.0	Rhode Island	225.00	36.0
Indiana	271.40	52.0	South Carolina	155.80	32.5
Iowa	192.00	48.0	South Dakota	248.30	38.0
Kansas	234.00	41.5	Tennessee	198.00	34.5
Kentucky	190.00	44.0	Texas	163.95	41.0
Louisiana	161.75	39.0	Utah	185.00	23.5
Maine	254.00	43.5	Vermont	209.50	40.0
Maryland	171.50	51.25	Virginia	195.00	27.0
Massachusetts	250.00	48.0	Washington	209.35	33.0
Michigan	223.00	36.5	West Virginia	222.60	40.0
Minnesota	215.00	45.0	Wisconsin	218.15	35.0
Mississippi	194.09	32.0	Wyoming	240.30	57.5
Missouri	225.46	47.5			

[1] Δ For 1967 was computed from State standards data for 1968—the 1st year standards data was available for unit's of size 2 and 4.

Note: 1967 standards for assistance units of size 4 are not precisely comparable to those for 1970. In 1967 the standards refer to a family consisting of a mother, father, and 2 children while the standards for 1970 refer to a family consisting of a mother and 3 children.

Source: U.S. Department of Health, Education, and Welfare, Social and Rehabilitation Service, "Old [illegible]sistance and Aid to Families With Dependent Children: Tables on Percent of Basic Needs Met for [illegible]d Types of Cases, January 1967," August 1967.

TABLE III.—*State monthly[1] maximums (M) by number of persons in assistance unit for 1967*

State	2	3	4	5	6	7	8	9	10
Alabama	40	65	90	115	140	140	140	140	140
Alaska	50	80	110	140	170	200	230	260	290
Arizona	80	107	134	161	188	215	220	220	220
Arkansas	65	75	85	95	105	115	125	125	125
California	148	178	208	238	268	298	328	358	386
Delaware	125	137	149	161	171	181	191	201	211
Florida	32	55	78	85	85	85	85	85	85
Georgia	67	96	125	154	154	154	154	154	154
Indiana	50	75	100	125	150	175	200	225	250
Kentucky	260	260	260	260	260	260	260	260	260
Louisiana	80	99	116	133	145	163	163	163	163
Maine	80	110	137	164	191	218	245	250	250
Maryland	250	250	250	250	250	250	250	250	250
Mississippi	25	40	50	60	70	80	90	90	90
Missouri	66	90	114	138	162	186	210	234	258
Nebraska	110	140	170	200	230	260	290	320	350
Nevada	61	92	123	154	185	216	247	278	309
New Mexico	190	190	190	190	190	190	190	190	190
Oklahoma	120	150	175	200	220	239	255	277	277
South Carolina	45	66	87	108	125	125	125	125	125
Tennessee	90	105	135	135	135	135	135	135	135
Texas	72	93	114	135	135	135	135	135	135
Utah	138	163	185	205	226	246	260	274	288
Virginia	215	215	215	215	215	215	215	215	215
Washington	325	325	325	325	325	325	325	325	325
West Virginia	165	165	165	165	165	165	165	165	165
Wyoming	170	200	200	215	215	215	230	230	230

[1] Maximums must be multiplied by 12 to obtain annual values.

TABLE IV.—*Average monthly payments for AFDC–RC units for 1967*

Census division	Average monthly payments, by number of children per unit						
	1	2	3	4	5	6+	Total
New England	$55.00	$81.40	$106.68	[1] $143.90	[1] $165.97	[1] $195.98	$88.28
Middle Atlantic	69.76	115.31	173.75	[1] 185.82	[1] 214.32	[1] 253.08	114.00
East North Central	51.56	88.00	117.45	[1] 146.13	[1] 168.54	[1] 199.02	89.65
West North Central	46.20	78.18	102.85	139.93	[1] 154.76	[1] 182.75	82.32
South Atlantic	38.14	67.48	94.61	106.45	134.49	113.62	65.93
East South Central	30.98	41.40	56.25	[1] 69.73	[1] 80.43	[1] 94.97	42.78
West South Central	39.73	64.55	82.58	[1] 102.82	[1] 118.59	[1] 140.04	63.08
Mountain	76.60	38.01	72.16	100.23	[1] 142.13	[1] 167.83	75.60
Pacific	92.44	57.20	102.66	140.49	186.79	[1] 205.22	92.44
Total	45.56	77.13	105.86	123.00	141.84	167.65	75.53

[1] These data were estimated by making the ratio of the average payment for a unit of a given size and census division to the average payment for all units of the same census division equal to the analogous ratio for the whole United States.

Source: U.S. Department of Health, Education, and Welfare, "Findings of the 1967 AFDC Study," table 128, August 1970.

TABLE V.—*Average monthly WRE and average monthly CCE by census division for 1967*

Census division	MWRE	MCCE
New England	54.21	61.02
Middle Atlantic	62.93	66.34
East North Central	40.15	53.06
West North Central	44.56	52.92
South Atlantic	23.62	39.59
East South Central	15.91	28.49
West South Central	27.29	41.37
Mountain	35.54	50.79
Pacific	63.10	67.85

Source: U.S. Department of Health, Education, and Welfare, Social and Rehabilitation Service, National Center for Social Statistics, "Findings of the 1970 AFDC Study," table 72, January 1972, and "Findings of the 1967 AFDC Study," table 118, August 1970.

Appendix B

VALIDATION OF THE COMPUTATIONAL MODEL FOR ESTIMATING AFDC ELIGIBLES

Validation of the computational model, described in the text and Appendix A, involves (1) an internal consistency check of the various estimates of AFDC eligibles, and (2) a comparison of these estimates with outside controls. Since no control estimates of eligibles exist, the best that can be done is to test the estimates for consistency with estimates of the participating caseload. The estimates of eligibles should at least include a population that "looks like" the participating caseload.

INTERNAL CONSISTENCY

Tables I and II show the estimates of eligible families and average payments by categorical type for 1967 and 1970. The various estimates show the effect of an annual versus a part-year accounting period, and for 1970, the effect of the full standard test. As expected, the part-year accounting period expands eligibility and raises total payments, while the full standard test restricts eligibility and lowers total payments. However, the effect of both is proportionally greater on the estimates of eligibles than on total payments. This is due to the fact that the full standard test "screens out" and the part-year accounting period "screens in" families with higher incomes and lower payments than the overall average. The average payments data in Table II illustrate that this is occurring.

Table I.—*1967 and 1970 estimates of eligible units by categorical type*

[In thousands]

	FH	UP	IF_1	IF_2	SP	RC[1]	Total
1967 estimates:							
Annual	1,907	71	191	115	111	111	2,509
Part-year	2,183	282	205	308	157	111	3,246
1970 estimates:							
Annual—without FST[2]	2,996	330	260	178	153	109	4,025
Passed FST[2]	2,467	192	246	119	83	109	3,215
Failed FST[2]	529	138	14	59	71	0	810
Part-year—without FST[2]	3,103	581	262	345	182	109	4,582
Passed FST[2]	2,714	523	257	296	126	109	4,025
Failed FST[2]	389	57	5	49	56	0	557

[1] Although the economic screens for RC units differ from the annual and part-year procedures, the estimates of poor RC units are included with estimates for both methods for the purpose of determining totals.
[2] FST—Full standard test.

Table II.—*1967 and 1970 estimates of payments to eligible units by categorical type*

	Annual average payments per unit							Total payments (in millions)
	FH	UP	IF_1	IF_2	SP	RC	Total	
1967 estimates:								
Annual	1,471	1,422	1,467	1,261	1,222	774	1,418	3,556
Part year	1,376	1,016	1,416	951	1,011	774	1,269	4,118
1970 estimates:								
Annual—without FST	2,005	1,977	2,132	1,741	1,205	890	1,941	7,812
Passed FST[1]	2,279	2,789	2,216	2,272	1,853	890	2,246	7,223
Failed FST[1]	726	844	630	668	576	0	727	589
Part year—without FST[1]	1,938	1,149	2,110	1,107	1,160	890	1,729	7,924
Passed FST[1]	2,079	1,241	2,127	1,236	1,368	890	1,857	7,475
Failed FST[1]	742	323	929	329	544	0	644	3,588

[1] FST—full standard test.

Analysis at a more detailed level indicates that the economic leniency of a part-year accounting period and the economic restrictiveness of the full standard test apply almost exclusively to families in which the head has work experience and earnings. A review of the computational formulas of Sections V and VI of Technical Appendix A shows that the part-year computations reduce to an annual computation for units in which the head has no work experience or earnings. For most such units the annual payments formula reduces to the full standard test formula.[1] Thus, most of the variation in the estimates shown in Tables I and II represents units in which the head works at some time during the year.

The estimates for UP and IF_2 units indicate that the computational formulas are operating as expected with respect to work experience. The disproportionate affect of the part-year computation and the full standard test on the UP and IF_2 estimates follows from the fact that the heads of such units, by definition have work experience and earnings. The same appears to apply for the SP units which, being male headed, are likely to be characterized by working heads.

DESCRIPTION OF CONTROLS

Two independent measures of welfare caseloads are available for validation—HEW administrative statistics and CPS figures on families that report the receipt of public assistance to the CPS. In both cases, the statistics represent measures of actual participating caseloads rather than measures of potential eligibles. Thus, both are imperfect validation controls, and, as previously mentioned, the best one can do is to test for consistency.

However, even as measures of actual caseloads both controls fall short of ideal. The CPS lumps reporters of all types of public assistance into one category so that recipients of AFDC are not readily distinguishable. Administrative statistics on the other hand refer to monthly rather than annual caseloads. Thus, assumptions must be made and manipulations applied to obtain controls which are comparable to the model estimates of eligibles.

To identify AFDC reporters in the CPS file, the categorical rules of the computational model are simply applied to all families with minor children (as defined in Section III of Technical Appendix A) reporting receipt of public assistance. Families that do not meet the criteria of any of the six categorical screens are assigned to a residual category termed General Assistance (GA).

To estimate annual caseloads from HEW administrative statistics, openings during the year are added to the monthly caseload at the beginning of the year, the logic being that the caseload at the beginning of the year plus all openings during the year yields an estimate of all units that participated at some time during the year. This procedure yields the following annual caseloads for 1967 and 1970, respectively:

December 1966 caseload [1]	1, 127, 000
Openings in 1967 [2]	703, 000
Annual caseload for 1967	1, 830, 000
December 1969 caseload [3]	1, 875, 000
Openings in 1970 [4]	1, 395, 948
Annual caseload for 1970	3, 270, 948

[1] U.S. Dept. of Health, Education, and Welfare; Social and Rehabilitation Service, National Center for Social Statistics, Report A-6 (6/7), "Program Facts." 1970.

[2] Op. Cit., NCSS Report A-5 (6/67 and 12/67), "Reasons for Opening and Closing Public Assistance Cases."

[3] Op. Cit. NCSS Report A-2 (12/69). *Public Assistance Statistics.*

[4] Op. Cit. NCSS Quarterly Report A-9, *Applications, Cases Approved, and Cases Discontinued for Public Assistance.* In cases where states did not report openings in a particular quarter, the number of openings reported in the nearest reporting quarter was imputed.

One flaw in this procedure is the potential of double counting families with more than one opening during the year. In a research note appearing in the September–October 1970 issue of *Welfare in Review,* Bradley Schiller estimated that the rate of intra-year recidivism was about 1.8 percent of the annual average

[1] This does not hold in states where a reduced standard or ratable reduction apply; i.e., r^4 and p have values of less than one.

caseload.[2] With annual average caseloads in 1967 and 1970 of 1.200 [3] million and 2.208 [4] million respectively, the adjustments for intra-year recidivism are as follows:

1967 annual unadjusted caseload	1, 830, 000
Recidivism adjustment	21, 600
1967 annual adjusted caseload	1, 808, 400
1970 annual unadjusted caseload	3, 270, 948
Recidivism adjustment	33, 744
1970 annual adjusted caseload	3, 237, 204

NOTE.—Where necessary these figures are further adjusted for comparability with cell detail of particular tables.

These procedures yield an aggregate estimate of the annual caseload but provide no distributional detail regarding the characteristics of the caseload. Because no known source of this type of information exists, the assumption is made that the annualized caseload looks like the monthly caseload described in the 1967 and 1971 AFDC studies. Thus, the annual estimates by geographic and demographic characteristics are obtained by multiplying the monthly estimates (from the AFDC Studies) by the ratio of the annual caseload to the end of year monthly caseload of the AFDC Studies. This procedure involves the somewhat dubious assumption that the annual rate of turnover is the same for all geographic and demographic groups.[5] Turnover, however, is most likely a function of economic alternatives to AFDC which in turn are a function of geographic and demographic characteristics. Hence, the detailed annual caseload estimates should be approached with this potential bias in mind.

The only exception to this assumption is the distribution of the 1970 caseload by the Census region characteristic. Since 1970 data on monthly caseloads and openings are published on a state by state basis, the procedures described above to obtain the aggregate annual estimates may be applied at the regional level to obtain estimates of annual AFDC caseloads for each Census region.

[2] His estimate is based on data for female headed units in 1968. This is not as up to date and as precise as one would like but is the only known information available. Schiller's estimate assumes that no families are on public assistance for more than two noncontiguous episodes during a year. He notes that this assumption holds up well on an annual basis.

[3] Calculated from Table 3 of NCSS Report A-2 (12/70).

[4] From Schiller's Article.

[5] Turnover here refers to the difference between the number of families that are receiving benefits at the end of the year and the number that received benefits at sometime during the year.

TABLE III.—*1967 and 1970 estimates of eligible and participating AFDC units by categorical type*

[In thousands][1]

	FH	UP	IF[1]	IF[2]	SP	RC	GA	Total
67 estimates:								
Annual estimates of eligibles	1,907 (76)	71 (3)	194 (8)	115 (5)	111 (4)	111 (4)	NA	2,509 (100)
Part-year estimates of eligibles	2,183 (67)	282 (9)	205 (6)	308 (10)	57 (15)	111 (3)	NA	3,216 (100)
Monthly administrative caseload [2]	962 (77)	66 (05)	153 (12)		24 (02)	50 (04)	NA	1,254 (100)
Annual administrative caseload	1,385 (77)	95 (05)	220 (12)		35 (02)	72 (04)	NA	1,808 (100)
AFDC reporters and GA	819 (57)	38 (3)	119 (8)	59 (4)	31 (2)	47 (3)	315 (22)	1,431 (100)
AFDC reporters	819 (73)	38 (3)	119 (11)	59 (5)	31 (3)	47 (4)	NA	1,116 (100)
Computed AFDC reporters [3]	774 (77)	37 (4)	97 (10)	51 (5)	21 (2)	18 (2)	NA	1,003 (100)
70 estimates:								
Annual estimates of eligibles W/O FST	2,996 (74)	330 (8)	260 (6)	178 (4)	153 (4)	109 (3)	NA	4,025 (100)
Annual estimates of eligibles W/FST	2,467 (77)	192 (6)	246 (8)	119 (4)	83 (3)	109 (3)	NA	3,215 (100)
Part-year estimates of eligibles W/O FST	3,103 (68)	581 (13)	262 (6)	345 (8)	152 (4)	109 (2)	NA	4,582 (100)
Part-year estimates of eligibles W/FST	2,714 (67)	523 (13)	257 (6)	296 (7)	126 (3)	109 (3)	NA	4,025 (100)
Monthly administrative caseload [2]	1,922 (78)	153 (6)	256 (10)		23 (1)	114 (5)	NA	2,457 (100)
Annual administrative caseload [2]	2,441 (78)	188 (6)	313 (10)		31 (1)	156 (5)	NA	3,129 (100)
AFDC reporters and GA	1,370 (60)	127 (06)	168 (07)	89 (04)	28 (01)	88 (04)	397 (18)	2,268 (100)
AFDC reporters	1,370 (73)	127 (07)	168 (09)	89 (05)	28 (01)	88 (05)	NA	1,871 (100)
Computed AFDC reporters [3]	1,343 (76)	125 (07)	159 (09)	84 (05)	22 (01)	38 (02)	NA	1,771 (100)

Numbers in parentheses represent percentage distributions by categorical type.
[1] The administrative statistics, taken from the 1967 AFDC study and the 1971 AFDC study, were adjusted for comparability with the categorical definitions of the model.
1970 adjustments.—Some 66,000 stepfather cases, which could not be simulated from the CPS data base, were subtracted from the total caseload of 2,524,000 to get the 2,457,000 shown in the table.
The estimate of RC units was obtained by subtracting the 66,000 stepfather cases from the study estimate of 251,000 no adult units, leaving 185,000 units. This number was further reduced by 71,000 incapacitated mother units, which NCSS says are also no adult units. Thus, the estimate of 114,000 RC units shown in the table.

The estimates for UP, SP, and IF units can be taken directly from the Study with no adjustments—no attempt was made to distinguish between IF_1 and IF_2 units. Given the Study estimates for RC, UP, and SP, units, FH units are computed as a residual.
Administrative estimates have not been adjusted to exclude recipients residing in Puerto Rico and the Virgin Islands (about 60,000 in 1970). The population of these areas is not included in the CPS.
1967 adjustments.—The adjustments for 1967 are the same as for 1970 substituting figures from the 1967 AFDC study.
[3] CPS reporters of AFDC whose eligibility is verified by the computations of the model.

TABLE IV.—*1967 estimates of eligible and participating AFDC units*

[In thousands] [1]

	Annual estimates of eligibles	Part year estimates of eligibles	Monthly administrative caseload [2]	Annualized administrative caseload [2]	AFDC reporters
Total	2,508 (100)	3,246 (100)	1,278 (100)	1,840 (100)	1,116 (100)
Region:					
Northeastern	598 (24)	798 (25)	369 (30)	531 (30)	278 (25)
North Central	513 (22)	752 (23)	258 (21)	372 (21)	269 (24)
Southern	986 (29)	1,177 (36)	340 (27)	490 (27)	382 (34)
Western	381 (15)	519 (16)	274 (22)	395 (22)	187 (17)
Place of residence:					
SMSA-CC	964 (38)	1,213 (38)	684 (54)	985 (51)	501 (45)
SMSA-NCC	527 (21)	769 (24)	218 (17)	314 (17)	200 (18)
Not SMSA	1,018 (41)	1,231 (38)	372 (29)	536 (29)	415 (37)
Unit size:					
1	74 (3)	74 (2)	53 (04)	76 (04)	31 (03)
2	650 (26)	812 (25)	262 (21)	377 (21)	199 (18)
3	588 (23)	789 (24)	269 (21)	387 (21)	236 (21)
4	416 (17)	597 (18)	223 (18)	321 (18)	225 (20)
6	294 (12)	379 (12)	171 (13)	246 (13)	115 (13)
6 plus	486 (19)	595 (18)	298 (23)	429 (23)	277 (25)
Level of income: [3]					
0	589 (23)	589 (18)	(4)	(4)	377 (34)
$1 to $999	669 (27)	670 (21)	(4)	(4)	283 (25)
$1,000 to $1,999	622 (25)	638 (20)	(4)	(4)	184 (16)
$2,000 to $2,999	322 (13)	406 (13)	(4)	(4)	83 (07)
$3,000 to $3,999	184 (07)	291 (09)	(4)	(4)	68 (06)
$4,000 to $4,999	66 (03)	203 (06)	(4)	(4)	46 (04)
$5,000 to $5,999	29 (01)	141 (04)	(4)	(4)	22 (02)
$6,000 to $7,999	26 (01)	170 (05)	(4)	(4)	35 (03)
$8,000 to $9,999	2 (00)	58 (02)	(4)	(4)	11 (01)
$10,000 plus	0 (00)	77 (02)	(4)	(4)	7 (01)

[1] Numbers in parentheses represent the percentage distribution of units by region, place of residence, unit ıze, and level of income.

[2] Detailed estimates do not add to totals for certain characteristics because of unknown categories and de·ivation adjustments.

[3] Level of income refers to all income except public assistance received by persons included in the assistance unit.

[4] Not available.

TABLE V.—*1970 estimates of eligible and participating AFDC units*

[In thousands] [1]

	Annual estimates of eligibles without FST	Annual estimates of eligibles with FST	Part-year estimates of eligibles without FST	Part-year estimates of eligibles with FST	Monthly Administrative caseload [2]	Annualized Administrative caseload [2]	AFDC reporters
Total	4,025 (100)	3,215 (100)	4,582 (100)	4,025 (100)	2,521 (100)	3,231 (100)	1,871 (100)
Region:							
Northeast	1,052 (26)	825 (26)	1,192 (26)	1,038 (26)	691 (28)	817 (27)	536 (29)
North Central	875 (22)	679 (21)	1,079 (24)	943 (23)	500 (20)	663 (21)	413 (22)
Southern	1,250 (31)	1,052 (33)	1,313 (29)	1,206 (30)	667 (27)	791 (25)	536 (29)
Western	818 (21)	659 (20)	968 (21)	838 (21)	605 (26)	891 (28)	386 (21)
Place of residence:							
SMSA—CC	1,631 (41)	1,314 (41)	1,784 (39)	1,555 (39)	1,428 (57)	1,828 (57)	894 (48)
SMSA—NCC	1,009 (25)	758 (24)	1,245 (27)	1,070 (27)	449 (18)	575 (18)	429 (23)
Not SMSA	1,385 (34)	1,143 (35)	1,553 (34)	1,400 (35)	610 (25)	820 (25)	548 (29)
Unit size:							
1	67 (02)	67 (02)	67 (01)	67 (02)	109 (04)	140 (04)	44 (02)
2	1,113 (28)	833 (26)	1,215 (27)	1,010 (25)	621 (25)	795 (25)	377 (20)
3	1,077 (27)	817 (25)	1,255 (27)	1,070 (27)	592 (23)	758 (23)	481 (26)
4	705 (18)	588 (18)	836 (18)	768 (19)	449 (18)	575 (18)	372 (20)
5	453 (11)	369 (11)	528 (12)	471 (12)	301 (12)	385 (12)	230 (12)
6 plus	610 (15)	541 (17)	681 (15)	639 (16)	452 (18)	579 (18)	368 (20)
Level of income: [3]							
0	821 (20)	821 (26)	821 (18)	821 (20)	NA	NA	613 (33)
$1 to $999	673 (17)	673 (21)	673 (15)	673 (17)	NA	NA	413 (22)
$1,000 to $1,999	624 (16)	622 (19)	624 (14)	623 (15)	NA	NA	294 (16)
$2,000 to $2,999	473 (12)	432 (13)	478 (10)	465 (12)	NA	NA	148 (08)
$3,000 to $3,999	447 (11)	317 (10)	452 (10)	391 (10)	NA	NA	115 (06)
$4,000 to $4,999	344 (09)	181 (06)	378 (08)	278 (07)	NA	NA	79 (04)
$5,000 to $5,999	288 (07)	88 (03)	344 (08)	210 (05)	NA	NA	63 (03)
$6,000 to $7,999	273 (07)	72 (02)	417 (09)	268 (07)	NA	NA	76 (04)
$8,000 to $9,999	69 (02)	8 (00)	212 (05)	162 (04)	NA	NA	35 (02)
$10,000 plus	13 (00)	1 (00)	183 (04)	131 (03)	NA	NA	35 (02)

[1, 2, 3] See footnotes of table IV.

VALIDATION OF ESTIMATES OF FAMILIES

Tables III, IV and V show the estimates of eligible families and the control estimates of participating families tabulated by various demographic, geographic, and economic characteristics. The figures presented here are mostly a repetition of those in the text, the major addition being the estimates of families reporting the receipt of AFDC to the CPS—henceforth referred to as Reporters. The general conclusion here, as in the text, is that the estimates of eligibles are consistent with the controls in that more families are eligible than are receiving payments. For virtually all of the characteristics shown, at least one of the estimates of eligibles is large enough to account for CPS Reporters and for either the monthly or annual administrative estimates of caseloads.

One minor exception to the general conclusion is the 1970 estimate of related child (RC) units. Since little *a priori* information was available for determining the economic rules for RC units, a stringent economic screen on caretaker families was chosen to avoid an unrealistically high estimate. The results suggest that the screen is overly stringent and a future refinement might be to devise a more lenient screen. The effect on the total estimate, however, should be negligible. The low 1970 estimates of eligibles for family size one in Table V reflect the problems of estimating RC units.

The figures on Reporters in Table III allow several observations regarding the validity of the computational model. The first row of figures shows the results of screening all families with minor children reporting receipt of public assistance to the CPS by the categorical rules of the model. The number of CPS reporters falls substantially short of administrative estimates suggesting that a major problem on the CPS is nonreporting of public assistance by a number of recipient units. But, of the AFDC recipients reporting to the CPS, the model verifies eligibility for 90 percent in 1967 and 95 percent in 1970.[6] Although the residual general assistance category (GA) appears high (22 and 18 percent of all reporting families), the figures are not inconsistent with monthly GA caseloads which reached 827,000 by the end of 1967 and 500,000 by the end of 1970. Also, the high income families picked up by the part-year calculations (Tables IV and V) do not seem unrealistic given the distribution of reporters by level of income.

VALIDATION OF ESTIMATES—PAYMENTS

Tables VI, VII, and VIII show model and control estimates of average payments broken down by region, place of residence, unit size, and categorical type. The annual administrative average payment for both years was derived by dividing total expenditures on AFDC money payments by the annualized caseload estimate. The monthly administrative payments, on the other hand, represent average monthly payments [7] inflated by a factor of twelve. The monthly figures are presented here to show the "annual rate" of monthly payments and to point out that this differs from a more appropriately estimated annual average payment. The annualized monthly estimates do not capture the downward effect of part-year participation on the overall average. Mere growth in the caseload implies that a substantial amount of part-year participation took place in both 1967 and 1970.

TABLE VI.—*1967 and 1970 estimated payments by categorical type*

	Average payments per unit							Total payments (millions)
	FH	UP	IF_1	IF_2	SP	RC	Total	
1967 estimates:								
Annual	$1,471	$1,422	$1,407	$1,261	$1,222	$774	$1,418	$3,556
Part year	1,376	1,016	1,410	951	1,011	774	1,269	4,118
Monthly administrative [1]	1,913	2,837	1,753		1,379	906	1,873	NA
Annual administrative [2]	NA	NA	NA		NA	NA	1,244	2,250
AFDC reporters	1,641	950	1,548	1,003	1,024	800	1,520	1,697
Computed AFDC reporters [3]	2,024	1,400	1,692	1,148	1,118	865	1,890	1,896
1970 estimates:								
Annual W/O FST	2,005	1,977	2,132	1,741	1,265	890	1,941	7,812
Annual W/FST	2,279	2,789	2,216	2,272	1,853	890	2,246	7,223
Part year W/FST	1,938	1,149	2,110	1,107	1,160	890	1,729	7,924
Part year W/FST	2,079	1,241	2,127	1,236	1,368	890	1,857	7,475
Monthly administrative [1]	2,205	2,852	2,138		1,661	1,086	2,144	NA
Annual administrative [2]	NA	NA	NA		NA	NA	1,517	4,851
AFDC reporters	1,975	1,214	2,232	1,272	1,342	1,388	1,875	3,509
Computed AFDC reporters [3]	2,587	1,655	2,018	1,535	2,012	1,091	2,434	4,310

[1] Monthly administrative figures in this table represent monthly payments reported to the 1967 and 1970 AFDC survey multiplied by 12. Thus the figures for this row represent the annual rate of monthly payments for the various types of categorical units in the survey month. True annual average payments would be lower because part year participation would be taken into account, assuming benefits levels do not change dramatically over the year. For FH, SP, and RC units the payments from the surveys were for units not precisely comparable to those of the model. For FH units, payments for all units with 1 adult were used; for SP units payments data for an "other status" residual category were used; and for RC units payments data for units with no adults were used.

[2] The 1967 annual administrative total is the total spent on AFDC money payments in calendar year 1967 as reported in NCSS report A-4(69), "Trend Report." The 1967 average total figure was derived by dividing 2,250,000,000 by the annual estimate of participants of 1,800,000.

Similarly, the 1970 annual administrative total is the total spent on AFDC money payments in calendar year 1970 as reported in NCSS report A-2 (December 1970). The average total figure was obtained by dividing 4,851,000,000 by the annual estimate of participants of 3,200.

[3] Computed amounts refer to the part year method of calculating payments—without the FST for 1970.

[6] This represents application of the part-year payments formula (no full standard test in 1970) to CPS reporters excluding the general assistance category. See AFDC Computed Reporters in Table III.

[7] From the *1967 and 1971 AFDC Studies.*

TABLE VII.—*1967 estimated annual average payments per unit*

	Annual	Part-year	Monthly administrative	Annual administrative	AFDC reports
Total	$1,418	$1,269	$1,873	$1,244	$1,520
Region:					
Northeast	1,870	1,654	2,416	NA	1,907
North Central	1,580	1,361	1,989	NA	1,519
Southern	1,001	906	1,204	NA	1,083
Western	1,554	1,365	2,077	NA	1,812
Place of residence:					
SMSA-CC	1,678	1,487	NA	NA	1,803
SMSA-NCC	1,420	1,243	NA	NA	1,550
Not SMSA	1,169	1,064	NA	NA	1,116
Unit size:[1]					
1	599	599	547	NA	659
2	851	786	1,323	NA	1,047
3	1,214	1,064	1,639	NA	1,259
4	1,543	1,316	1,950	NA	1,523
5	1,661	1,516	2,218	NA	1,664
6 plus	2,285	2,081	2,676	NA	2,113

[1] The payments for unit sizes 2 to 6 plus were derived from table 129 of pt. II of the *1967 AFDC Study*. Hence they represent November/December 1967 payments multiplied by 12 for units with 1 adult. The payment for unit size 1 was derived from table 128 for units with no adult recipients.

TABLE VIII.—*1970 estimated annual average payments per unit*

	Annual W/O FST	Annual W/FST	Part year W/O FST	Part year W/FST	Monthly administrative	Annual administrative	AFDC reporters
Total	$1,941	$2,216	$1,729	$1,857	$2,144	$1,517	$1,875
Region:							
Northeast	2,558	3,020	2,294	2,487	3,018	NA	2,535
North Central	1,933	2,205	1,667	1,801	2,204	NA	1,867
Southern	1,268	1,433	1,206	1,286	1,294	NA	1,233
Western	2,179	2,520	1,829	1,960	2,185	NA	1,861
Place of residence:							
SMSA-CC	2,220	2,566	2,031	2,202	2,326	NA	2,206
SMSA—NCC	1,980	2,388	1,677	1,810		NA	1,870
Not SMSA	1,577	1,785	1,421	1,510	1,611	NA	1,340
Unit size:[1]							
1	661	661	661	661	693	NA	1,109
2	1,396	1,661	1,301	1,413	1,567	NA	1,497
3	1,674	1,980	1,492	1,617	1,968	NA	1,505
4	2,142	2,414	1,871	1,953	2,362	NA	1,839
5	2,315	2,636	1,995	2,113	2,711	NA	2,188
6 plus	3,033	3,296	2,654	2,783	3,230	NA	2,594

[1] The payments for unit sizes 2-6 plus were derived from table 93 of pt. II of the 1972 AFDC Survey. Hence they represent January 1971 payments multiplied by 12 for units with 1 adult. The payment for unit size 1 was derived from table 82 for units with no adult recipients.

Reporters presumably present a better picture of annual average payments since the CPS asks how much public assistance was received over the last year. However, the figures in Table VI indicate a problem of underreporting of AFDC benefits in the CPS. The total reported amount of AFDC falls over a billion dollars short of the annual administrative figure of 4.9 billion dollars while the reported average payment of 1877 is greater than the annual administrative average of 1517. The apparent contradiction is explained by the fact that the major problem with the CPS estimates is non-reporting of small benefits. Several studies on the reporting of public assistance income by Rockwell Livingston suggest that over and under-reporting of public assistance by units that do report tend to cancel out, and that the shortcomings of the total reported amounts are due to units receiving small benefits that do not report at all.[9] Even with this shortcoming, CPS reported average payments along with the annual administrative average payment represent the best control estimates for analyzing the model estimates.

On the basis of *a priori* logic the part-year accounting period estimates of the model are the most appealing since they represent a closer approximation of the actual monthly accounting period of the AFDC program. For 1970, at least, the part-year estimates more closely approximate those for Reporters and the annual caseload. This is especially true for the UP and IF_2 units which are most affected by the part-year calculations because of their work experience. (See Table VI.)

However, one would expect average payments of the eligible population to be lower than those for both Reporters and the annual caseload since the eligible population includes non-participants, who are most likely to be families eligible for smaller than average benefits. The high average payment for the eligible population (1857) in 1970 relative to that for the annual caseload (1517) suggests that the model is over-estimating average payments. A conclusion corroborated by the computed payments for reporters that are higher than reported amounts.

For 1967, both the annual and part-year model estimates are fairly close to reported amounts. Given the conclusion that reported amounts are biased upward suggests, however, that the part-year estimates are still a better choice since on the whole they are lower than the annual estimates. Also, the over-all part-year average payment is very close to the annual administrative estimate. This still suggests an upward bias in the model estimates of average payments as they are expected to be lower than caseload estimates. Computed payments for Reporters, again, corroborate this conclusion.

Given the recognized inadequacies of the data base and the methodology, there are several reasons why payments may be overestimated. Others may exist, but are not recognized.

Payments may be overestimated because other components of income—wages, salaries, transfer payments excluding public assistance, etc.—are under-reported to the CPS. Dr. Nelson McClung formerly of the Urban Institute has developed procedures to correct for such under-reporting. When the procedures are applied to the CPS tape, the effect on the AFDC estimates can be determined by re-estimating AFDC eligibility from a corrected CPS tape.

Also, most state standards of need include an amount for rent equal to the maximum the state will allow. If a unit's rent is lower than this maximum, the standard of need is lowered accordingly for the purpose of determining eligibility and benefits. Thus, to the extent that the rent allowance is less than the maximum for a large proportion of the caseload, the model over computes payments by implicitly assuming that all units receive the maximum.

Finally, the difference between control and model estimates of average payments may reflect a time lag in the application and receipt of benefits on the part of eligible units. That is, an eligible unit may defer the decision to apply for welfare for a month or two in the hope that their economic situation will improve. Thus, on an annual basis a unit's benefits are less than the maximum amount they could have received had they applied for benefits immediately.

The tentative conclusion to be drawn from this analysis is that eligibles may also be biased, upward. To the extent the bias in payments is due to underreporting or improper rent allowances the estimates of eligibles will be biased upwards along with payments. However, to the extent the payments bias is due to a time-lag in the receipt of benefits only payments and not eligibles are biased upwards.

Unfortunately, this conclusion, at least for the present, must be classified as tentative. The validation analysis of the payments estimates is based on several key assumptions and estimates which themselves are subject to error, such as the nature of underreporting of public assistance and the estimates of the annual AFDC caseload. Thus at this point no further technical work has been undertaken to improve the estimates because no certain measure of what is better exists. However, the tentative conclusion should be kept in mind and used as a caveat to temper any analysis based on the estimates of eligibles.

[9] See Livingston, Rockwell, "Evaluation of the Reporting of Public Assistance Income in the Special Census of Dane County, Wisconsin: May 15, 1968", *Research and Statistics to Meet Today's and Tommorrow's Challenges*, Proceedings of the Ninth Workshop on Public Welfare Research and Statistics, New Orleans, La., August, 1969; Livingston, Rockwell, "Interpreting Census Statistics on Public Assistance," presented at the American Statistical Association's Annual Meeting, August 23, 1968, Pittsburgh, Pennsylvania; and Livingston, Rockwell; "Evaluating the Reporting of Public Assistance Income in the 1966 Survey of Economic Opportunity", *The Influence of Research and Statistical Reporting on Policy Making*, Proceedings of the Tenth Workshop on Public Welfare Research and Statistics, Washington, D.C., August 1970.

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