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ABSTRACT

Blacks and Mexican Americans do not attain the levels of education enjoyed by whites. There are two possible explanations for these differences: (1) either the process of educational attainment varies across ethnic groups, or (2) if the process is invariant, whites start with social advantages not shared by blacks and Mexican Americans, and thus demonstrate higher educational outcomes. Using data from the 1972 National Longitudinal Study (NLS), and controlling for known differences in measurement errors across groups, statistically significant interactions in the attainment of blacks, Mexican Americans, and whites were found. These differences in structural parameters, however, were not large enough to account for the different educational outcomes across ethnic groups. Differences in outcome appear to be primarily a function of differential levels of social background. Thus, equality of educational opportunity will not equalize outcome when groups are not equal to begin with. (Author/APM)

EDUCATIONAL ATTAINMENT AMONG WHITES, BLACKS, AND MEXICAN-AMERICANS *

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EDUCATIONAL ATTAINMENT AMONG WHITES, BLACKS, AND MEXICAN-AMERICANS

ABSTRACT

The attainment of education is one of our society's most highly placed values, yet equality of educational outcomes remains an allusive goal. Blacks and Mexican-Americans do not attain the levels of education enjoyed by whites. There are fundamentally two possible explanations for why these differences exist: either the process of educational attainment varies across ethnic groups, or if the process is invariant, whites start with social advantages not shared by blacks and Mexican-Americans. Using data from the 1972 NLS, and the LISREL program to control for known differences in measurement errors across groups, we found statistically significant interactions across groups, but the differences in structural parameters were not large enough to account for the different outcomes. Differences in educational outcomes across ethnic groups are mostly a function of differential levels of social background.

EDUCATIONAL ATTAINMENT AMONG WHITES, BLACKS, AND MEXICAN-AMERICANS

The attainment of education is one of our society's most highly placed values. Not only is it a value unto itself, it plays an instrumental role in occupational and economic achievement. It is also the area to which the greatest commitment to equality of opportunity has been made. Despite this commitment, however, equality of educational outcomes remains an allusive goal. In 1979, the U.S. Bureau of the Census (1980) estimated that the median school years completed by whites age 25 years and over was 12.5. In comparison, blacks attained 11.9 years, and people of Spanish origin only 10.3. These differences result from differential access to education, and there are two ways to explain such differences. On the one hand, the social process by which people in this country translate their social and biological characteristics into years of schooling may be different from one group to another. For example, whites with fathers of higher socioeconomic status may possess an inflated opportunity for success over their white peers of lower status, while blacks may possess an equal (low) opportunity for success regardless of their status origins. On the other hand, the social process of educational attainment may be identically the same for the three groups, but because the groups start from different socioeconomic levels, the outcomes remain different. For example, the probability of completing college for offspring of professionally employed fathers may be the same for both whites and Mexican-Americans, but if there are many more whites proportionally in the professions than Mexican-Americans, there will be fewer Mexican-Americans

who complete college. The process is the same, but the outcomes are different. In its elemental form, that is the main issue addressed in this paper. Is the process of educational attainment the same for whites, blacks, and Mexican-Americans, or does it differ?

This paper is not the first to ask whether the process of socio-economic achievement differs among ethnic groups. The seminal work of Blau and Duncan (1967) gave birth to a number of analyses of interracial (e.g., Duncan, 1969; Jencks, et al., 1972) and interethnic (e.g., Duncan and Duncan, 1968; Featherman and Hauser, 1978) examinations of differences in status attainment. While a great deal of attention has been paid to occupational achievement, differences in the process of educational attainment have also received their share of attention. For example, based on Sewell and Hauser's (1975) social-psychological model of status attainment, both Portes and Wilson (1976) and Kerckhoff and Campbell (1977) concluded that the educational attainment process is different for blacks and whites. Both studies yielded evidence that social background variables are more important determinants of educational attainment among whites than among blacks. Other analyses have addressed the process of educational attainment among Mexican-Americans. Cantu (1975) studied a small sample of Mexican-Americans in Mercedes, Texas, and Featherman and Hauser (1978, p. 466) concluded that there were greater opportunities for blacks in the U.S. than for Mexican-Americans, because status origins were more of a handicap for Mexican-Americans than for blacks.

While all of the previous studies would lead to the same conclusion, namely that the educational attainment process differs among whites, blacks, and Mexican-Americans, they have all been based on the implicit assumption

that there were no measurement errors among the variables included in the analyses. Ignoring measurement errors, however, can result in systematic bias in parameter estimates, and when levels of measurement error differ between groups, interethnic comparisons of parameter estimates may exaggerate or understate interethnic differences. Mason, et al. (1976) demonstrated that measurement errors exist in socioeconomic variables, and outlined how these errors could affect parameter estimates in socioeconomic models of achievement. For example, random reporting errors in measured status origins will reduce the estimated effects of status origins on achievement. Random errors in measured status origins will also increase estimated effects of achievement variables on each other, because their mutual dependence on status origin will be underestimated. Following Mason, et al.'s (1976) analysis, Bielby, Hauser and Featherman (1977) found differences in reliability of measurement between blacks and whites, and concluded that measurement errors produced biased underestimates of the effects of background variables of nearly 20 percent for whites, but even greater underestimates for blacks. Thus, ignoring measurement error exaggerates racial differences in parameter estimates of educational returns to social origins. In addition, Wolfle (1979) showed that Hispanics also report social data with inherent measurement errors, and ignoring them would lead to estimates even more biased than among either whites or blacks.

Thus, while previous studies would lead one to believe there were interethnic differences in the educational process, the suspicion remains that these estimated differences may be an artifact of differential measurement errors. This study therefore reexamines interethnic differences

in the educational attainment process between whites, blacks, and Mexican-Americans using Jöreskog and Sörbom's (1978) general method for the analysis of covariance structures, a statistical procedure which controls for differential measurement errors in the estimation of structural parameters.

THE MODEL

The basic model of educational attainment used in this analysis considers education to be a function of father's occupational status, father's education, mother's education, number of siblings, sex, ability, academic preparation, and college plans. The model is shown diagrammatically in Figure 1; the theoretical variables of interest are shown with the ellipses. The five variables inside ovals on the left-hand side of Figure 1 are latent, exogenous variables; "latent" because they are not directly measured; "exogenous" because their causes, whatever they may be, are unanalyzed in this particular model. The latent ability variable is considered to depend on father's occupation and education, mother's education, number of siblings, and respondent's sex, plus a residual disturbance term, which represents all of the variation in ability not explained by the five independent variables. The disturbance term is assumed to be statistically independent of the five exogenous variables, and is also assumed to be independent of the disturbance terms attached to academic preparation, college plans, and educational attainment. The three parental status variables are expected to have positive effects on ability. While these relationships may be causally spurious due to the omission from this model of measures of parental ability, they are in any

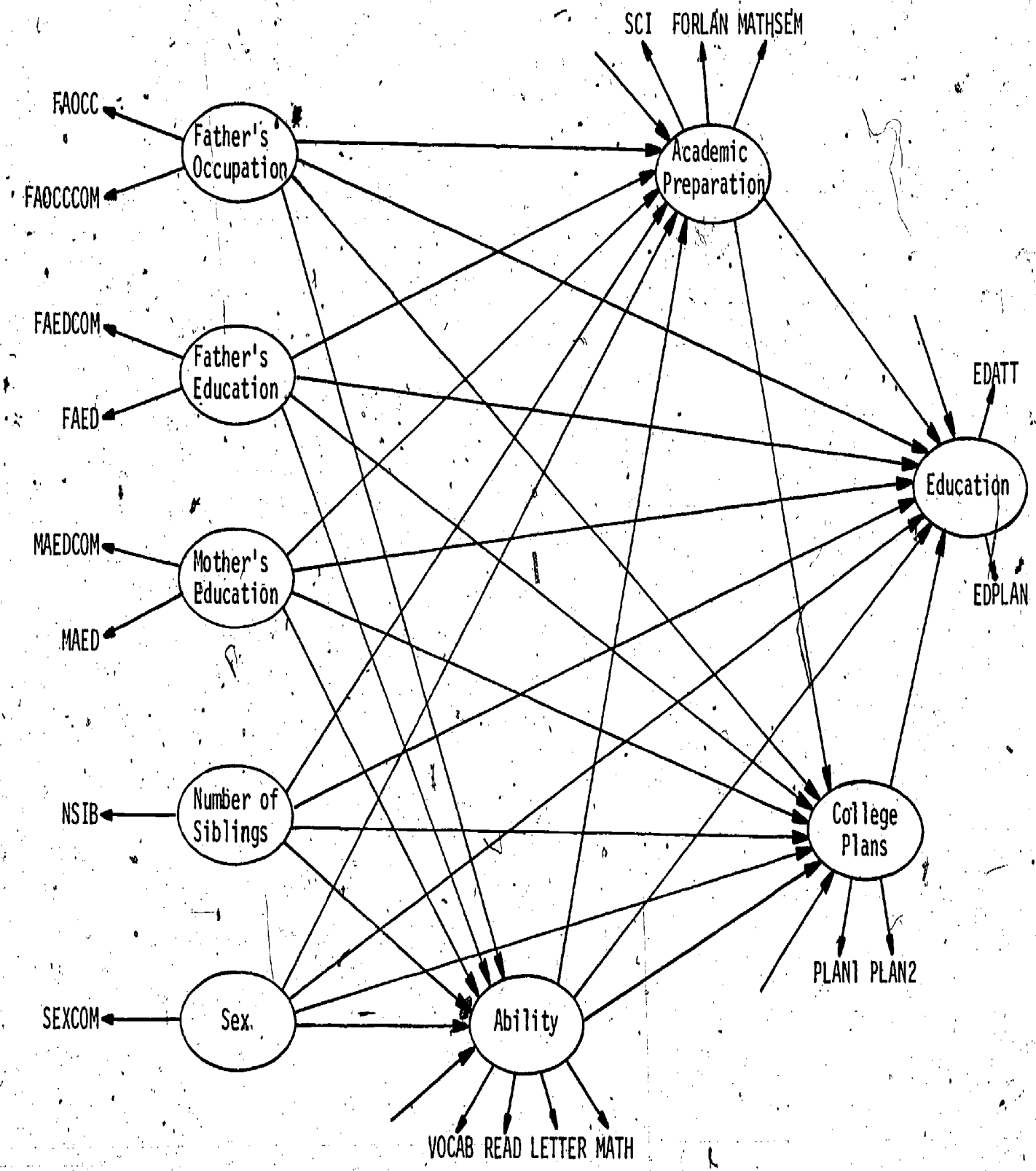

Figure 1. Structural Equation and Measurement Models of Educational Attainment Among 1972 Hjgh School Graduates

event expected to be positive (see Scarr and Weinberg, 1978). The effect from the number of siblings is expected to be negative. The effect of sex on ability is expected to be nearly zero; Wolfle (1980), for example, has summarized some of the literature on the relationship between sex and ability, and found most tests of ability were specifically constructed to produce a zero association between the test score and sex of the examinee.

Previous studies (Heyns, 1974; Alexander and McDill, 1976) considered curriculum placement to be the major mechanism by which secondary schools function to separate students into tracks that ultimately differentiate their adult roles. Yet as Alexander and McDill (1976) point out, it is not curriculum differentiation per se which differentiates students, but rather what happens to them in one curriculum track or another. In particular, students in college preparatory tracks complete a greater number of courses in academic subject matter, and thus develop the prerequisite skills and credentials necessary for college matriculation. In this analysis, therefore, the usual practice of measuring college preparatory curricular membership (a one-zero dummy variable) has been eschewed in favor of the number of semesters of academic courses completed in high school.

The academic preparation variable is considered to be causally determined by the five latent, exogenous variables, plus ability. Students with parents of higher socioeconomic status are expected to acquire more academic courses. A negative association is expected between sex and academic preparation, which implies women take fewer academic courses than do men. Both Heyns (1974) and Alexander and McDill (1976) report negative effects of the number of siblings on college preparatory curricular placement, and the same effect on academic preparation is expected here. Finally, higher

ability students are expected to complete more academic courses than students of lower ability.

College plans measure the respondents' intentions to continue their educations beyond high school. The variable may be considered to be dichotomously coded, so that a value of unity indicates the respondent expressed plans to attend college, and a value of zero indicates no plans to attend college. Positive effects from academic preparation, ability, and the three measures of parental socioeconomic status are expected, but no a priori predictions are made about the effects of sex and the number of siblings. Despite the fact that (until recently) more men than women actually attend college, no a priori guesses about the relationship of sex on college plans are offered, particularly when considered net of other causal influences.

Finally, education is considered to be dependent upon all of the preceding latent variables. Positive effects are expected from the three measures of parental status. Like previous studies, students with more siblings are expected to acquire less education, and women are expected to acquire less education than men, but this effect is not expected to be very large in absolute value. Ability, academic preparation, and college plans should all have positive effects on educational attainment; furthermore, college plans will probably have an effect larger in magnitude than the other variables due to its immediate effect of continuing educational attainment beyond high school.

METHOD

Sample

Data for this study are drawn from the National Longitudinal Study (NLS) of the High School of 1972 (see Levinsohn, et al., 1978). The NLS was designed to provide data on the development of educational, vocational, and personal aspects of the lives of adolescents as they make the transition from high school to the adult world. The total sample consists of 22,652 students selected from 1318 schools. The respondents were initially surveyed in the spring of 1972, their senior high school year. Subsequent follow-up surveys were conducted in the fall of 1973, 1974, and 1976. Logistical problems with the initial data collection effort prevented the inclusion of base-year information on nearly 6000 students; as a result, some important base-year responses are missing and the following analysis is based on the 16,683 remaining students. The sample was further restricted to subjects whose racial-ethnic identification was either white, black, or Mexican-American. In the latter case it was decided not to aggregate groups of Cubans, Puerto Ricans, and others of Spanish descent into a single group of Hispanics due to the diverse nature of their backgrounds and cultural heritage. Rather, only those who identified themselves as Mexican-American were included in these analyses. As with most other analyses of the process of socioeconomic achievement, pairwise present correlations were used to estimate the parameters of the model; the average number of whites in the analysis was 11,743; the average number of blacks was 1810; and the average number of Mexican-Americans was 493.

Bear in mind that the NLS sample is representative of high school seniors in 1972; it is therefore not necessarily representative of all youths of equivalent age. It is probable that a number of youths dropped out of high school prior to their senior year. It is also likely that the dropouts are systematically different than those who remained in high school. Moreover, fewer blacks and Mexican-Americans complete high school than whites. In 1977, for example, among people age 25 - 29, 86.8 percent of whites completed high school, but only 74.5 percent of blacks, and 58.1 percent of Hispanics (National Center for Education Statistics, 1979, p. 224). As a result, these data should not be assumed to be representative of all whites, blacks, and Mexican-Americans. Rather, they are representative of high school seniors, and the analysis of educational attainment reported in this paper relates to years of schooling attained after high school.

Variables

Father's Occupation. The first manifest measure of father's occupation (FAOCC; V0368) was a question administered in the base-year survey, which asked the respondents to indicate the kind of work done by their fathers. (The alphanumeric names refer to their labels in Levinsohn, et al., 1978.) The categories matched, more or less, the census major occupation groups. For this analysis the variable was recoded to the average Duncan (1961) socioeconomic index (SEI) score for the category, as revised to match the 1970 census occupation classification (Hauser and Featherman, 1977). The second variable (FAOCCCOM; V2468) was a composite of the individual's response to base-year and first-year follow-up questionnaire items indicating the father's occupation. This

variable was coded with the revised Duncan SEI score for detailed occupation groups. The exact construction of the composite variable is given in Levinsohn, et al. (1978, pp. 76-79).

Father's Education. Father's education was measured with two manifest variables. The first of these (FAEDCOM) was the NLS educational composite, V1627. The second (FAED) was the education question in the first follow-up, V1009. FAED was coded to match the categories used in FAEDCOM. Having done that, the category codes were recoded to years of educational attainment using midpoint interval estimates taken from the U.S. Bureau of the Census (1973). These two recoded variables now represent years of father's educational attainment in which gross categories have been coded to the midpoint value obtained from the distribution of educational attainment among males 25 years of age or over in 1970.

Mother's Education. Mother's education was measured in a similar fashion. First MAED (V1010) was recoded to the same values used in MAEDCOM (V1628). These were then recoded to years of educational attainment, using the midpoint value obtained from the U.S. Bureau of the Census's (1973) distribution of educational attainment among females 25 years of age or over in 1970.

Siblings. The number of siblings has but a single indicator (NSIB), computed as the sum of variables V1460, V1461, V1462, and V1463. These questionnaire items asked the respondent to indicate the number of older brothers, younger brothers, older sisters, and younger sisters, respectively. In handling missing data due to item nonresponse for these variables, a nonzero response to any of the four items accompanied by nonresponse to one or more of the other items was assumed to represent a zero response to the

nonresponse items. For example, individuals who indicated they had one older brother, but did not answer the other three questions, were assumed to have one sibling. If none of the four items was answered, the number of siblings was coded as missing data.

Sex. The respondent's sex was measured by the composite variable, SEXCOM (V1626). The variable was coded 0 if male, 1 if female. As a result, positive regression estimates emanating from this variable indicate a greater value of the dependent variable for females.

Ability. The latent variable of ability was measured by four manifest indicators of achievement: VOCAB (V0614), a scaled vocabulary score; READ (V0618), a scaled reading score; LETTER (V0619), a scaled letter-group score; and MATH (V0620), a scaled mathematics score. These variables were not recoded. They represent scores on standardized tests administered to the respondents during the spring of 1972, their senior year of high school.

Academic Preparation. The latent variable of academic preparation was measured by three manifest indicators. These were the number of semesters of science taken between July 1, 1969, and graduation (SCI; V0046), the semesters of foreign languages (FORLAN; V0053), and the semesters of math (MATHSEM; V0074). Gilmartin, et al. (1976) have shown that young men who plan scientific careers completed more math and foreign language courses in high school than young men who did not plan such careers. Moreover, young women planning such careers took more foreign language courses than would be predicted from their abilities. It is expected that completion of such courses will not only predict successful attainment of plans to enter scientific careers, but will also

predict the attainment of additional years of schooling. The manifest variables were not recoded.

College Plans. Two manifest variables were used to index college plans. The first of these was an NLS routing question (PLAN1; V0385), and was recoded unity for those people who planned to enter either a four-year college or university, or a two-year academic junior college, either full time or part time while working, apprenticing, or homemaking. Otherwise, the variable was coded zero. The second manifest measure of college plans (PLAN2; V0386) was based on responses to a question about what the respondents planned to do during the year after high school "if there were no obstacles." The variable was coded unity for those who said their plans were to attend either a two-year academic junior college or a four-year college or university. Otherwise, the variable was coded zero.

Educational Attainment. Educational attainment was measured with two manifest indicators. The first of these (EDATT) was V1854, actual educational attainment measured four years after high school graduation, and the second (EDPLAN) was V1855, planned educational attainment measured at the same time. The latent variable, educational attainment, is therefore a construction not only of actual years of education completed by the third follow-up survey, but also includes a component that measures plans for additional education. Both manifest measures were recoded to reflect years of schooling completed or expected to be completed. Following Featherman and Carter (1976) two years of attendance in a vocational, trade or business school were coded equal to one year of attendance in an academic school.

SPECIFICATION OF THE LISREL MODEL

Obtaining estimates for the model shown in Figure 1 was accomplished by using LISREL (linear structural relationships by the method of maximum likelihood), a computer program developed by Jöreskog and Sörbom (1978). The LISREL model assumes a causal structure among a set of unmeasured, latent variables, some designated as exogenous and others as endogenous. These unmeasured variables are also related to a set of observed variables such that the latent variables appear as causes of the observed variables. The LISREL model, therefore, consists of two parts: the measurement model and the structural equation model.

These two parts of the model have been described above in the vernacular. In LISREL teminology, two random vectors, $\underset{\sim}{\eta}' = (\eta_1, \eta_2, \eta_3, \eta_4)$, and $\underset{\sim}{\xi}' = (\xi_1, \xi_2, \xi_3, \xi_4, \xi_5)$, represent the latent endogenous and exogenous variables, respectively, in which η_1 = ability, η_2 = academic preparation, η_3 = college plans, and η_4 = educational attainment; furthermore, ξ_1 = father's occupation, ξ_2 = father's education, ξ_3 = mother's education, ξ_4 = number of siblings, and ξ_5 = sex.

The model specifies a fully recursive causal structure among the latent variables, such that:

$$\underset{\sim}{B}\,\underset{\sim}{\eta} = \underset{\sim}{\Gamma}\,\underset{\sim}{\xi} + \underset{\sim}{\zeta}$$

where $\underset{\sim}{B}$ (4×4), and $\underset{\sim}{\Gamma}$ (4×5) are matrices of structural coefficients in which $\underset{\sim}{\Gamma}$ is a full matrix relating the exogenous vector to each of the endogenous latent variables, and $\underset{\sim}{B}$ is a matrix relating each endogenous variable to those that follow it in the causal scheme. $\underset{\sim}{\zeta}' = (\zeta_1, \zeta_2, \zeta_3, \zeta_4)$ is a vector of randomly distributed residuals uncorrelated with each other and with $\underset{\sim}{\xi}$.

The vectors $\underset{\sim}{\eta}$ and $\underset{\sim}{\xi}$ are not observed, but $\underset{\sim}{y}' = (y_1, y_2, y_3, y_4, y_5, y_6, y_7, y_8, y_9, y_{10}, y_{11})$ and $\underset{\sim}{x}' = (x_1, x_2, x_3, x_4, x_5, x_6, x_7, x_8)$ are observed, such that:

$$\underset{\sim}{y} = \underset{\sim}{\Lambda}_y \underset{\sim}{\eta} + \underset{\sim}{\varepsilon}$$

and

$$\underset{\sim}{x} = \underset{\sim}{\Lambda}_x \underset{\sim}{\xi} + \underset{\sim}{\delta},$$

where $\underset{\sim}{\varepsilon}$ and $\underset{\sim}{\delta}$ are vectors of errors of measurement in $\underset{\sim}{y}$ and $\underset{\sim}{x}$, respectively. These errors of measurement represent both specific and random components of variation (see Alwin and Jackson, 1979). They are assumed to be uncorrelated with $\underset{\sim}{\eta}$, $\underset{\sim}{\xi}$, and $\underset{\sim}{\zeta}$, but may be correlated among themselves. The matrices $\underset{\sim}{\Lambda}_y$ (11 x 4) and $\underset{\sim}{\Lambda}_x$ (8 x 5) are regression matrices of $\underset{\sim}{y}$ on $\underset{\sim}{\eta}$ and of $\underset{\sim}{x}$ on $\underset{\sim}{\xi}$, respectively.

Let $\underset{\sim}{\Phi}$ (5 x 5) be the covariance matrix of $\underset{\sim}{\xi}$; let $\underset{\sim}{\Psi}$ be a diagonal matrix of variances for the disturbance vector, $\underset{\sim}{\zeta}$; and let $\underset{\sim}{\Theta}_\varepsilon$ and $\underset{\sim}{\Theta}_\delta$ be the covariance matrices of $\underset{\sim}{\varepsilon}$ and $\underset{\sim}{\delta}$, respectively. In application, some of the elements of the four regression matrices, and the four covariance matrices, are fixed and equal to preassigned values (often zero or unity). Other elements are free parameters to be estimated by the method of maximum likelihood. This estimation procedure requires that the estimates be maximized with respect to some known distribution, which in the case of LISREL is assumed to be the multivariate normal.

For the sake of simplicity in the following notations, consider X_1 = FAOCC, X_2 = FAOCCCOM, . . . , X_8 = SEXCOM, and Y_1 = VOCAB, Y_2 = READ, . . . , Y_{11} = EDPLAN. Furthermore, let $x_1 = (X_1 - \overline{X}_1), \ldots,$ $y_{11} = (Y_{11} - \overline{Y}_{11})$, so that all the variables are expressed as deviations

from their respective means. This transformation has no effect upon the regression slopes, but does serve to eliminate constant terms from the equations.

With this notation, the structural portion of the model is a fully recursive model among the latent variables, represented by the following structural equations:

$$\eta_1 = \gamma_{11}\xi_1 + \gamma_{12}\xi_2 + \gamma_{13}\xi_3 + \gamma_{14}\xi_4 + \gamma_{15}\xi_5 + \zeta_1$$

$$\eta_2 = \gamma_{21}\xi_1 + \gamma_{22}\xi_2 + \gamma_{23}\xi_3 + \gamma_{24}\xi_4 + \gamma_{25}\xi_5 + \beta_{21}\eta_1 + \zeta_2$$

$$\eta_3 = \gamma_{31}\xi_1 + \gamma_{32}\xi_2 + \gamma_{33}\xi_3 + \gamma_{34}\xi_4 + \gamma_{35}\xi_5 + \beta_{31}\eta_1 + \beta_{32}\eta_2 + \zeta_3$$

$$\eta_4 = \gamma_{41}\xi_1 + \gamma_{42}\xi_2 + \gamma_{43}\xi_3 + \gamma_{44}\xi_4 + \gamma_{45}\xi_5 + \beta_{41}\eta_1 + \beta_{42}\eta_2 + \beta_{43}\eta_3 + \zeta_4$$

Metrics for the latent variables have been established by fixing some elements in the Λ_y and Λ_x matrices to unity. In addition, ξ_4 and ξ_5 have been set exactly equal to their respective manifest indicators, implying that these variables were measured without error. As a result of these specifications, the metric of ability is measured in terms of MATH, the metric of academic preparation in terms of MATHSEM, the metric of college plans in terms of PLAN1 as recoded, and the metric of education in terms of EDATT as recoded. Among the endogenous variables, the metric of father's occupation is measured in terms of FAOCCCOM, father's education in terms of FAEDCOM, mother's education in terms of MAEDCOM, and, of course, siblings in terms of NSIB and sex in terms of SEXCOM.

In algebraic form, the measurement portion of the model is:

$$x_1 = \lambda_{11}\xi_1 + \delta_1$$

$$x_2 = \xi_1 + \delta_2$$

$$x_3 = \xi_2 + \delta_3$$

$$x_4 = \lambda_{42}\xi_2 + \delta_4$$

$$x_5 = \xi_3 + \delta_5$$

$$x_6 = \lambda_{63}\xi_3 + \delta_6$$

$$x_7 = \xi_4$$

$$x_8 = \xi_5$$

$$y_1 = \lambda_{11}\eta_1 + \varepsilon_1$$

$$y_2 = \lambda_{21}\eta_1 + \varepsilon_2$$

$$y_3 = \lambda_{31}\eta_1 + \varepsilon_3$$

$$y_4 = \eta_1 + \varepsilon_4$$

$$y_5 = \lambda_{52}\eta_2 + \varepsilon_5$$

$$y_6 = \lambda_{62}\eta_2 + \varepsilon_6$$

$$y_7 = \eta_2 + \varepsilon_7$$

$$y_8 = \eta_3 + \varepsilon_8$$

$$y_9 = \lambda_{93}\eta_3 + \varepsilon_9$$

$$y_{10} = \eta_4 + \varepsilon_{10}$$

$$y_{11} = \lambda_{11\,4}\eta_4 + \varepsilon_{11}$$

EDUCATIONAL ATTAINMENT AMONG WHITES

Assuming the joint distribution of the 19 variables in the model of educational attainment is multivariate normal, maximum likelihood estimates of parameters of the 23 structural and measurement model equations were obtained using Jöreskog and Sörbom's (1978) LISREL program. (In particular, the LISREL program used to estimate the models employed the corrected procedures in the computation of the gamma standardized solution matrix, and in the t-values for multiple group comparisons.) The estimates were computed from pairwise present correlations for white 1972 high school seniors. The correlations, means, and standard deviations among the 19 variables are shown in Table 1.

When the structural and measurement models were estimated for these data, a chi-square goodness-of-fit statistic was calculated, and is shown in the first row of Table 2. This value of 5975.59 suggests at first glance that the fit of the model is not acceptable. It is well-known, however, that, "in large samples virtually any model tends to be rejected as inadequate" (Bentler and Bonett, 1980). Thus, it is unlikely that any other theoretical model will fit short of saturation, but an alternative model, which is merely a specialized version of the original model, can be constructed; having estimated the two competing models, a chi-square difference test can be used to evaluate the statistical significance of the parameters that differentiate between the two competing models.

The question becomes which parameters of the model shall be changed. First of all, the structural equation model is already fully recursive, and to change it would defy the logic of the temporal and theoretical relationships among these variables. Second, the factor

Table 1. Correlations, Means, and Standard Deviations among Variables in a Model of Educational Attainment: White 1972 High School Graduates (N = 11,743)

	VOCAB	READ	LETTER	MATH	SCI	FORLAN	MATHSEM	PLAN 1	PLAN 2	EDATT	EDPLAN	FAOCC	FAOCCCOM	FAEDCOH	FAED	MAEDCOM	MAED	NSIB	SEXCOM
VOCAB	---																		
READ	.642	---																	
LETTER	.412	.517	---																
MATH	.525	.615	.615	---															
SCI	.285	.295	.244	.417	---														
FORLAN	.394	.353	.306	.372	.333	----													
MATHSEM	.269	.290	.268	.499	.552	.415	---												
PLAN 1	.367	.379	.315	.443	.332	.365	.351	---											
PLAN 2	.308	.314	.265	.368	.290	.296	.318	.645	---										
EDATT	.407	.399	.332	.483	.352	.400	.377	.630	.525	---									
EDPLAN	.425	.422	.318	.468	.346	.382	.366	.581	.481	.774	---								
FAOCC	.219	.186	.154	.218	.131	.209	.169	.280	.198	.288	.283	---							
FAOCCCOM	.210	.193	.165	.220	.132	.225	.174	.263	.190	.272	.289	.637	---						
FAEDCOM	.271	.249	.202	.256	.167	.251	.198	.320	.225	.355	.352	.579	.542	---					
FAED	.277	.258	.204	.258	.174	.253	.196	.326	.230	.355	.358	.574	.553	.900	---				
MAEDCOM	.253	.230	.185	.245	.162	.218	.181	.285	.196	.322	.315	.366	.349	.501	.486	---			
MAED	.264	.235	.195	.254	.164	.217	.181	.298	.213	.334	.326	.378	.357	.491	.518	.871	---		
NSIB	-.135	-.091	-.026	-.069	-.063	-.112	-.088	-.132	-.110	-.152	-.134	-.070	-.067	-.069	-.063	-.090	-.093	---	
SEXCOM	.034	.032	.139	-.123	-.165	.112	-.196	-.029	-.018	-.029	-.081	-.035	-.019	-.017	-.028	-.047	-.047	.020	---
Mean	52.34	52.14	52.20	52.21	3.65	2.77	3.95	.480	.377	13.47	14.94	42.61	43.74	12.03	12.02	11.88	11.86	2.93	.49
S.D.	9.63	9.28	8.76	9.34	1.91	2.27	2.00	.500	.485	1.56	2.50	22.31	22.81	3.39	3.36	2.72	2.63	1.99	.50

Table 2. Goodness of Fit for the Different Models of Educational Attainment of White 1972 High School Graduates (Pairwise Deletion N = 11753)

Model	χ^2	d.f.	Prob.	$\Delta\chi^2$	d.f.	Prob.
Uncorrelated errors	5975.59	121	0.0			
$\delta_{5,3}$ free	5595.16	120	0.0	380.43	1	0.0
$\delta_{5,3}$, $\varepsilon_{10,8}$ free	5584.22	119	0.0	10.94	1	.001

analytic structure could be altered by allowing, for example, a causal effect of the latent factor, mother's education, to influence one or more of the manifest variables measuring father's education. This suggestion is rejected also on the grounds of the logic of the measurement model. Third, the initial assumption that the measurement errors among the manifest variables were uncorrelated could be relaxed. In this case, there is no objection to altering the model, because the initial assumption of zero covariance is not only restrictive, but is also contrary to the findings of Bielby, Hauser and Featherman (1977) and Wolfle (1979).

Accordingly, the off-diagonal elements of the two error covariance matrices were examined to see which elements were most likely to be different from zero. Since the objective was to find which one of the a priori assumptions of zero covariance was least probable, the zero restriction was relaxed for that element of Θ_{ε} or Θ_{δ} which gave the largest decrease in the chi-square goodness-of-fit statistic. Following Sörbom (1975), an inspection of the LISREL-produced table of first-order derivatives (not shown here) suggested that $\Theta_{\delta_{53}}$ was probably not zero. This is the covariance between the errors of FAEDCOM and MAEDCOM, the composite education variables for father and mother, respectively.

Allowing the covariance between δ_3 and δ_5 to be free (i.e., a parameter to be estimated within the model), the model was reestimated and a new chi-square goodness-of-fit statistic calculated. As can be seen in the second row of Table 2, the difference in fits between the model assuming uncorrelated errors and a new model assuming $\Theta_{\delta_{53}}$ to be nonzero was 380.43. This is distributed as chi-square with one degree of freedom, and is, of course, highly significant. Apparently, in the

construction of the two composite education indices, a systematic component of error was introduced into the two measures. The correlation between these two error terms is .26.

The model with $\Theta_{\delta_{53}}$ free had a chi-square value equal to 5595.16 with 120 degrees of freedom. This is not a very good fit either; a new inspection of the derivative tables suggested that $\Theta_{\varepsilon_{10,8}}$ may be nonzero. A new model was therefore estimated allowing this parameter to be free, which resulted in a chi-square value of 5584.22 with 119 degrees of freedom. The resulting improvement in fit is marginal, even if statistically significant; because these variables are substantively unrelated, and the improvement in fit was not large in relative terms, the final model adopted for whites was the one in which δ_3 and δ_5 are correlated, but all other error covariances are specified to be zero.

Measurement Model: Whites

Having found the best-fitting, most logically plausible model for whites, attention is turned to the measurement properties of these variables among white respondents. The evidence indicates that reporting errors for whites are most probably random. Only one nonzero correlation among error terms was found, and that was between two NCES-constructed composite measures of parents' education. We consider this finding significant -- the correlation was not found between respondent's reports, but rather between two constructed indicators. Thus, the nonzero correlation is probably due to some specific covariance introduced into the composites, rather than to nonrandom errors in the original reports. The general finding agrees with the results published in Bielby, Hauser and

Featherman (1977), who concluded reporting errors were random for nonblack men.

Reporting errors may be random, but they are nonetheless substantial. Parameter estimates for the final measurement model for whites appear in columns 3 - 5 of Table 3. Column 6 shows the estimated reliability coefficients (the squared true score - observed score correlations estimated from the measurement model). These coefficients are striking in several ways. First, they are considerably lower than those previously reported for nonblack adults (Bielby, Hauser and Featherman, 1977) and for white twelfth-grade youths (Mason, et al., 1976). For example, Bielby, et al. (1977, p. 1258) found reliability coefficients for father's occupation of .85 and .89, while Mason, et al. (1976, p. 466) reported a coefficient of son's report of .91 . In comparison, the reliability coefficients for NLS indicators of father's occupation were only .67 and .60 . These differences cannot be explained here. They may be due to differences in the wording of the questionnaire items (see Featherman and Hauser, 1978; Kerckhoff, 1974), to differences in data collection procedures, to differential coding errors, to differential errors introduced during keypunching, or even to errors of reporting. In any event, reporting errors in the NLS data seem to be considerably more severe than among other data sets which address the achievement process.

Second, it is also striking that the reliability coefficients for the background variables are greater in value for the original questions than for the NCES-constructed composites. Apparently, the composites contain sources of error (possibly due to additional coding, or to keypunching errors) that are not contained in the original questions.

Table 3. Measurement Model Parameter Estimates for White 1972 High School Graduates (Pairwise N=11753)

Variable		Observed Variance σ_i^2	Error Variance σ_e^2	True Variance σ_τ^2	Relative Slope λ_{ij}	Reliability Coefficient $(\sigma_\tau^2/\sigma_i^2)\lambda_{ij}^2$
True	Observed					
Ability	VOCAB	92.80	47.20	59.30	.887	.50
	READ	86.10	33.83		.939	.61
	LETTER	76.78	41.85		.768	.46
	MATH	87.30	28.00		1.0*	.68
Academic Preparation	SCI	3.65	1.91	2.36	.856	.47
	FORLAN	5.14	3.50		.836	.32
	MATHSEM	4.01	1.64		1.0*	.59
College Plans	PLAN 1	.250	.056	.193	1.0*	.77
	PLAN 2	.235	.109		.808	.54
Education	EDATT	2.44	.443	1.99	1.0*	.82
	EDPLAN	6.26	1.620		1.525	.74
Father's Occupation	FAOCC	497.89	162.06	314.21	1.034	.67
	FAOCCCOM	520.39	206.18		1.0*	.60
Father's Education	FAEDCOM	11.47	1.36	10.13	1.0*	.88
	FAED	11.29	.91		1.012	.92
Mother's Education	MAEDCOM	7.41	1.33	6.12	1.0*	.83
	MAED	6.90	.55		1.018	.92

* Fixed value

Third, unlike Bielby, Hauser and Featherman (1977) who found that social background variables were reported with nearly equal reliabilities, it was found here that parent's education variables were measured with considerably greater reliability than father's occupation.

Among the schooling process variables, relatively low coefficients of reliability were also found. For the four measures of ability, the reading and math tests were more reliable than the vocabulary and letter-group tests. In measuring academic preparation, semesters of math and science were considerably more reliable indicators than semesters of foreign languages. In measuring college plans, a straightforward question (PLAN1) about college plans seemed to be a somewhat more reliable indicator than a question (PLAN2) which asked respondents to consider possible obstacles to their plans.

Finally, two manifest indicators were used to measure educational attainment. One question (EDATT) asked respondents in October 1976 -- four years after high school graduation -- to indicate their actual educational attainment. The second question (EDPLAN), presented at the same time, asked respondents to indicate their educational plans. The two variables were included for the purpose of capturing both actual attainments for those who had already terminated their education careers, and additional planned attainments for those who were still in the educational process. Of these two variables, actual education was slightly more reliable than educational plans. The reliability coefficient of .82 for educational attainment is compariable in size to reliability coefficients from the 1973 OCG. Bielby, Hauser and Featherman (1977, p. 1258) report a reliability coefficient of .89 for their initial survey, and .96 for an

OCG reinterview (mostly by telephone with the actual respondent). However, an OCG mailout-mailback questionnaire produced a reliability coefficient for education of .70. Thus, the NLS reliability coefficient of .82 is greater than the comparable coefficient derived from the OCG mailed questionnaire, but less than those obtained from the OCG personal interviews.

Structural Model: Whites

In this section the results of the structural equation portion of the model for whites are presented. The structural coefficients are shown in Table 4. Later these results will be compared to those obtained for blacks and Mexican-Americans; here the concern is only with the white portion of the NLS sample. As hypothesized above, father's occupation and both parent's levels of education positively influence respondent's level of ability. The number of siblings, as hypothesized, has a negative influence on ability. Knowing one's sex does not increase one's capacity to predict one's level of ability; even with a very large number of cases, this parameter estimate is statistically insignificant.

As discussed above, positive effects of parental status measures on academic preparation were expected, and a negative effect from number of siblings; women were expected to complete fewer academic courses than men, and higher ability students were expected to complete more academic courses. In several instances, these assumptions were shown to be wrong. While father's occupational status had a positive influence on the number of academic courses completed, the effects from both father's and mother's education were statistically indistinguishable from zero. Women students

Table 4. Maximum-Likelihood Estimates of Parameters of the Educational Attainment Process for White 1972 High School Graduates (N=11743)

Dependent Variable	Independent Variables: Father's Occup.	Father's Educ.	Mother's Educ.	Number Siblings	Sex	Ability	Academic Preparation	College Plans	R^2
Standardized Coefficients									
Ability	.108*	.157*	.188*	-.075*	.012				
Acad. Prep.	.073*	.024	.020	-.050*	-.172*	.608*			
College Plans	.105*	.054*	.071*	-.068*	.035*	.292*	.305*		
Education	.017	.072*	.063*	-.040*	-.011	.164*	.124*	.565*	
Regression Coefficients[a]									
Ability	.047 (.008)	.381 (.042)	.572 (.038)	-.262 (.037)	.185 (.145)				.16
Acad. Prep.	.006 (.002)	.011 (.008)	.012 (.007)	-.034 (.007)	-.510 (.028)	.119 (.002)			.45
College Plans	.003 (.000)	.008 (.002)	.013 (.002)	-.014 (.002)	.031 (.008)	.017 (.001)	.090 (.004)		.44
Education	.001 (.001)	.032 (.006)	.035 (.005)	-.025 (.005)	-.029 (.020)	.030 (.002)	.115 (.012)	1.786 (.042)	.68

[a]Standard errors in parentheses.

*Indicates absolute size of coefficient equals or exceeds 2.57 times its standard error.

completed about one-half of a semester fewer academic courses than their male counterparts. The higher ability students completed more academic classes -- this was the strongest of the six predictors.

Among the causal forces that influence the development of plans to attend college, Table 4 shows that ability and academic preparation are the strongest predictors. Among the background variables, father's occupation and both parent's education variables have positive and significant effects. Respondents with more siblings are less likely to express college plans. Finally, women tend to be more likely than men to express college plans, even after controlling for the other independent variables in the structural equation; this effect is small, but nonetheless statistically significant.

Finally, the estimated parameters of the structural equation of educational attainment were examined. Respondents who had fathers with higher occupational status were expected to themselves acquire more years of schooling. In the event, the coefficient was not statistically significant; net of other variables in the equation, father's occupation did not directly influence educational attainment. Both parent's educational attainment variables influence the acquisition of education with nearly equal effects. Respondents with more siblings acquire less schooling, although the effect is small once the influence of other variables is controlled. Despite the fact that women are slightly more likely than men to express college plans, in the event men and women appear to acquire nearly equal amounts of schooling, net of the influence of other variables in the model. Examining the influence of the endogenous variables, one may see that the most important predictor of years of

education among high school seniors is the variable which measures plans to attend college. This may seem trivially obvious, but it is certainly not trivial in its impact. Those students who expressed plans to attend college, net of ability, academic preparation, and other background variables, will acquire 1.8 years more schooling than their peers without college plans. Completing academic courses in high school also influences educational attainment, an effect nearly equal to that of the respondent's ability.

EDUCATIONAL ATTAINMENT AMONG BLACKS

This section presents results of the measurement and structural models of educational attainment among black NLS respondents. Once again, it is assumed that the 19 manifest variables have a joint multivariate normal distribution, and LISREL estimates of the model's parameters have been acquired using pairwise present correlations. These correlations, means, and standard deviations are shown in Table 5. Comparison of the means for blacks with the corresponding means for whites reveals that blacks have lower levels of socioeconomic background; that is, they have fathers with less occupational status and fewer years of education, and their mothers have less education. They also come from larger families than do whites. The average scores for blacks on the four ability tests were ten points below the average score for whites. Moreover, blacks completed fewer semesters of academic courses, and not as many blacks as whites expressed plans to attend college. Nonetheless, in 1976 blacks were planning to complete on the average 15.1 years of education compared

Table 5. Correlations, Means, and Standard Deviations among Variables in a Model of Educational Attainment: Black 1972 High School Graduates (N = 1810)

	VOCAB	READ	LETTER	MATH	SCI	FORLAN	MATHSEM	PLAN 1	PLAN 2	EDATT	EDPLAN	FAOCC	FAOCCCOM	FAEDCOM	FAED	MAEDCOM	MAED	NSIB	SEXCOM
VOCAB	---																		
READ	.613	---																	
LETTER	.418	.534	---																
MATH	.489	.585	.587	---															
SCI	.163	.213	.193	.277	---														
FORLAN	.396	.420	.366	.438	.416	---													
MATHSEM	.234	.257	.232	.369	.491	.477	---												
PLAN 1	.297	.333	.305	.372	.233	.378	.284	---											
PLAN 2	.277	.324	.275	.357	.219	.356	.257	.644	---										
EDATT	.307	.369	.293	.345	.237	.375	.287	.506	.498	---									
EDPLAN	.345	.411	.311	.394	.226	.379	.243	.468	.476	.666	---								
FAOCC	.218	.133	.127	.123	.035	.181	.093	.133	.121	.114	.116	---							
FAOCCCOM	.180	.162	.168	.152	.072	.242	.113	.217	.178	.192	.189	.486	---						
FAEDCOM	.208	.187	.125	.159	.047	.224	.129	.206	.192	.178	.225	.458	.419	---					
FAED	.220	.195	.158	.177	.075	.228	.147	.232	.213	.209	.255	.431	.450	.845	---				
MAEDCOM	.246	.213	.173	.198	.102	.265	.135	.199	.172	.169	.207	.335	.321	.480	.483	---			
MAED	.251	.223	.182	.213	.106	.290	.148	.238	.191	.193	.215	.359	.341	.480	.567	.834	---		
NSIB	-.209	-.173	-.122	-.161	-.050	-.261	-.134	-.193	-.167	-.159	-.162	-.168	-.175	-.255	-.263	-.304	-.317	---	
SEXCOM	.033	.041	.105	-.095	-.070	.080	-.078	.003	-.005	.288	-.014	-.045	-.005	-.025	-.022	-.045	-.024	.077	---
Mean	43.11	42.92	42.67	42.09	3.25	2.05	3.67	.418	.368	13.17	15.13	27.27	28.52	9.94	9.95	10.59	10.62	4.58	.58
S.D.	7.87	8.74	10.25	8.18	1.63	2.01	1.79	.493	.482	1.42	2.47	19.26	20.54	2.97	3.01	2.88	2.95	2.92	.49

to the average for whites of 14.9, even though they had actually attained 13.2 years compared to the average for whites of 13.5 years.

The following has been mentioned before, but it is worth repeating: the NLS sample is representative of high school seniors in 1972; it is therefore not necessarily representative of the corresponding age group of all youths. Among whites, for example, 49 percent of the NLS sample is male. This figure compared favorably with the 1970 census figures for the corresponding age cohort; in that year 50.4 percent of the white age cohort was male. In comparison, the 1970 census indicates that 49.3 percent of the Negro age cohort was male, but in the NLS sample only 42 percent of the black 1972 high school senior class was male. The discrepancy is due, in all likelihood, to more black males dropping out of school prior to graduation than either whites or black females. As a result, these data mask to some unknown extent the true differences between the races.

When the structural and measurement models were estimated for the black subsample, as shown in Table 6 a chi-square goodness-of-fit statistic was obtained equal to 851.52, indicating a poor fit between the estimated and the observed covariance matrices. As before, the model was reestimated after relaxing the assumption of uncorrelated errors of measurement. The first-order derivatives suggested that the correlation between ε_5 and ε_7 would most likely improve the fit of the model; these were the errors in measurement for semesters of science and semesters of math. With $\Theta_{\varepsilon_{75}}$ free, the fit of the model improved significantly. Another examination of the first-order derivatives suggested that $\Theta_{\delta_{53}}$

Table 6. Goodness of Fit for the Different Models of Educational Attainment of Black 1972 High School Graduates (Pairwise N = 1810)

Model	χ^2	d.f.	Prob.	$\Delta\chi^2$	d.f.	Prob.
Uncorrelated errors	851.52	121	0.0			
$\varepsilon_{7,5}$ free	706.67	120	0.0	144.85	1	0.0
$\varepsilon_{7,5}$, $\delta_{5,3}$ free	619.74	119	0.0	86.93	1	0.0
$\varepsilon_{7,5}$, $\delta_{5,3}$, $\varepsilon_{11,8}$ free	614.12	118	0.0	5.62	1	.018

should also be a free parameter. This is the correlation between the measurement errors of the parent's composite education variables. This became the final model; relaxing a further assumption about correlated error terms yielded a model only marginally better in its fit to the observed covariance matrix.

Measurement Model: Blacks

When Bielby, Hauser and Featherman (1977) examined the character of reporting errors among status variables, they found, "that reports of social background and achievement variables by nonblack males are subject to strictly random errors, while reports by black males appear subject to significant nonrandom error" (1977, p. 1242). While the present analysis appears to confirm their conclusions about the nature of measurement errors among whites, these data do not completely support their conclusions about blacks. As discussed above, the best fitting model for blacks suggested nonzero covariances between errors of measurement for two pairs of variables. One of these was between FAEDCOM and MAEDCOM, a result which matches the finding from the analysis of the white subsample, but these were NCES-constructed variables, and their correlation may result from some systematic error introduced in their construction, not necessarily in the initial reports of parental education. Another nonzero error covariance was found to exist between the errors of semesters of science and math. But these data were not supplied by the respondents; they were reported by the respondents' schools. Thus, the possitive correlation between ε_5 and ε_7, which equaled .32, indicates that the schools attended by blacks reported semesters of math and science with more consistency than

would have occurred if the reporting errors had been random. In other words, while evidence was found of some nonrandom errors of measurement among blacks, the variables involved were not self-reported. We would therefore not want to conclude, as did Bielby, Hauser and Featherman (1977), that reports among blacks were subject to significant nonrandom error.

There is plenty of evidence, however, to indicate the existence of random reporting errors among blacks. Table 7 gives the measurement model parameter estimates for blacks; column 6 of the table shows the estimated reliability coefficients. When the results for whites were previously compared to earlier studies, lower values of reliabilities were found in the NLS data. Among blacks, however, the reliability coefficients seem to be more in line with values previously published. For example, Bielby, Hauser and Featherman (1977, p. 1262) report reliability coefficients for black males in the neighborhood of .56 for father's occupation, and .89 for father's education. In comparison, the present model yields estimates of .50 and .49 for the two measures of father's occupation, and .78 and .94 for the two measures of father's education.

As seen above for whites, the reliabilities are greater for the original parental education questions than for the constructed composite measures. In addition, blacks report their parent's educational attainment with greater accuracy than they report their father's occupation.

Among the endogenous variables, the four ability measures seem to be as reliably measured for blacks as for whites. College plans are measured with equal reliability among blacks, and only modestly less than among whites. And as seen for whites, there seems to be some discrepancy among the reliability coefficients for academic preparation. Among whites,

Table 7. Measurement Model Parameter Estimates for Black 1972 High School Graduates (Pairwise N = 1810)

Variable		Observed Variance σ_i^2	Error Variance σ_e^2	True Variance σ_τ^2	Relative Slope λ_{ij}	Reliability Coefficient $(\sigma_\tau^2/\sigma_i^2)\lambda_{ij}^2$
True	Observed					
Ability	VOCAB	61.93	32.22	39.64	.866	.48
	READ	76.38	27.98		1.105	.63
	LETTER	105.01	56.56		1.105	.46
	MATH	66.92	27.28		1.0*	.59
Academic Preparation	SCI	2.64	2.04	.967	.791	.23
	FORLAN	4.06	.57		1.900	.86
	MATHSEM	3.19	2.22		1.0*	.30
College Plans	PLAN 1	.243	.083	.160	1.0*	.66
	PLAN 2	.232	.085		.958	.63
Education	EDATT	2.02	.64	1.38	1.0*	.68
	EDPLAN	6.11	2.07		1.711	.66
Father's Occupation	FAOCC	370.47	183.97	204.90	.954	.50
	FAOCCCOM	421.68	216.77		1.0*	.49
Father's Education	FAEDCOM	8.83	1.97	6.90	1.0*	.78
	FAED	9.05	.57		1.109	.94
Mother's Education	MAEDCOM	8.29	2.17	6.23	1.0*	.75
	MAED	8.69	.35		1.157	.96

*Fixed value

science and math courses were measured with greater reliability than courses in foreign languages. Among blacks, however, foreign language courses were measured with considerably more reliability than the other two subjects.

The measurement of educational attainment among blacks was not as reliable as among whites, although like whites actual attainment was slightly more reliable as an indicator of educational attainment than was planned education. Comparison of these reliabilities to those reported by Bielby, Hauser and Featherman (1977, p. 1262) revealed that blacks in the NLS sample report their schooling with considerably less reliability than blacks in the OCG study, even when the OCG question was included in a mailed questionnaire.

In sum, the suspicion that whites and blacks would exhibit differential levels of measurement error seems justified by the evidence. Particularly among the background variables, blacks seem to report status levels of their parents with greater error than do whites. Among the endogenous variables, there appears to be little difference in reporting errors, except for the mix of variables which measure academic preparation, and educational attainment.

Structural Model: Blacks

This section reports the results of the structural equation portion of the model for blacks. The structural coefficients are shown in Table 8. While the primary interest here is with the results for the black subsample, some comparisons will be made to whites. These comparisons will be informal here; later a formal statistical test will be developed to compare the relative sizes of the regression slopes across subsamples.

Table 8. Maximum-Likelihood Estimates of Parameters of the Educational Attainment Process for Black 1972 High School Graduates (N = 1810)

Dependent Variable	Independent Variables: Father's Occup.	Father's Educ.	Mother's Educ.	Number Siblings	Sex	Ability	Academic Preparation	College Plans	R^2
				Standardized Coefficients					
Ability	.183*	.006	.134*	-.078*	.032				
Acad. Prep.	.122*	-.038	.045	-.053*	.048*	.429*			
College Plans	.012	.119*	.007	-.030	-.016	.372*	.368*		
Education	-.044	.092*	-.020	.009	-.001	.235*	.121*	.525*	
				Regression Coefficients[a]					
Ability	.079 (.021)	.015 (.102)	.409 (.085)	-.272 (.067)	.487 (.314)				.13
Acad. Prep.	.010 (.003)	-.018 (.014)	.027 (.011)	-.037 (.008)	.142 (.042)	.084 (.006)			.43
College Plans	.000 (.001)	.017 (.006)	.001 (.005)	-.006 (.003)	-.014 (.019)	.021 (.002)	.105 (.015)		.36
Education	-.004 (.003)	.041 (.016)	-.011 (.013)	.006 (.009)	-.003 (.049)	.043 (.006)	.112 (.039)	1.659 (.102)	.61

[a]Standard errors in parentheses.

*Indicates absolute size of coefficient equals or exceeds 2.57 times its standard error.

Among blacks, the important predictors of ability are father's occupation, mother's education, and number of siblings. The first two have positive effects; the latter has a negative effect. These were predicted. Unlike whites, the influence of father's education among blacks has no influence on ability. And just as predicted, and already found to be true for whites, one's sex has no influence on ability for blacks. In other words, there seems to be little that differentiates blacks and whites in the structural coefficients for the determination of ability.

Neither does there seem to be much difference between these groups in the determination of the semesters completed of academic courses. Both blacks and whites with fathers of higher occupational status take more academic courses, but neither father's nor mother's education has any direct effect. Having more siblings decreases the number of academic courses completed in high school, but higher ability students, whether black or white, complete more academic courses. The only noticeable difference between these groups is in the effect of one's sex on academic preparation. Among whites, as already seen, women completed about one-half of a semester fewer academic courses than did men. Among blacks, women completed more academic courses. The metric coefficient suggests that black women complete about one-seventh of a semester more academic courses than black men, after controlling for the influence of all other variables in the equation. Nevertheless, to substantively interpret these differences between blacks and whites may be presumptous, because the differences may be an artifact produced by the measurement portion of the model. As previously seen among whites, math and science courses are measured with

greater reliability than courses in foreign languages. Among blacks, however, foreign language courses are measured with greater reliability than either science or math courses. As a result, the latent variable measuring academic preparation is weighted in favor of math and science for whites, but in favor of foreign language courses for blacks. It is also true that that women, regardless of race, complete more foreign language courses than men, but men take more science and math courses than women. Thus, the latent variable measuring academic preparation favors those courses taken by males among the whites, but favors courses taken by females among the blacks. In one sense, then, these differences are an artifact of how reliably academic courses are reported for blacks and whites, and should probably not be interpreted as representing substantive differences between the races.

Differences in the determination of college plans between blacks and whites do not appear to be important. For whites all the predictor variables were statistically significant. This is not the case for blacks, for whom only father's education, ability and academic preparation were significant. But when one compares their metric coefficients, one sees that they are not very different in size.

Examining the coefficients for educational attainment, once again the similarities between blacks and whites are more striking than their differences. For both groups, father's occupation has no significant direct effect on education, but father's education does. While the effects among whites for mother's education and number of siblings were significant and in the direction hypothesized, the effects of these variables among blacks were insignificant. For neither group is sex a

significant predictor of educational attainment. Among the endogenous variables, also, the effects across groups are nearly identical. Blacks who expressed plans to attend college completed on the average another 1.7 years of schooling; this compares to 1.8 years for whites. Moreover, for blacks, as for whites, the relative importance of ability exceeds that of academic preparation.

In sum, the similarities of effects between blacks and whites are more striking than the differences. The coefficients do, of course, vary between the groups, but the direction of effects are often identical, and the relative magnitudes are often close in value. There was one coefficient which exhibited a radical change between the groups. This was the net influence of one's sex on the number of academic courses completed in high school. Among whites, males took more academic courses, among blacks, females took more academic courses. As discussed above, however, male-female differences between blacks and whites are confounded with effects of differential reporting errors of manifest measures of academic preparation, and it should be with reluctance that anyone interpret this particular result substantively.

EDUCATIONAL ATTAINMENT AMONG MEXICAN-AMERICANS

In this section the results of the measurement and structural models of educational attainment among Mexican-American NLS respondents are presented. As before, it is assumed that the 19 manifest variables have a joint multivariate normal distribution, and LISREL estimates of the parameters have been acquired using pairwise present correlations. These correlations, means, and standard deviations for Mexican-Americans are shown in Table 9.

Table 9. Correlations, Means, and Standard Deviations among Variables in a Model of Educational Attainment: Mexican-American 1972 High School Graduates (N=493)

	VOCAB	READ	LETTER	MATH	SCI	FORLAN	MATHSEM	PLAN 1	PLAN 2	EDATT	EDPLAN	FAOCC	FAOCCCOH	FAEDCOM	FAED	MAEDCOM	MAED	NSIB	SEXCOM
VOCAB	---																		
READ	.600	---																	
LETTER	.338	.511	---																
MATH	.379	.624	.566	---															
SCI	.145	.203	.140	.290	---														
FORLAN	.155	.267	.228	.270	.176	---													
MATHSEM	.064	.217	.175	.402	.526	.292	---												
PLAN 1	.271	.346	.314	.334	.181	.170	.225	---											
PLAN 2	.154	.290	.219	.333	.238	.186	.243	.615	---										
EDATT	.254	.379	.247	.400	.155	.150	.259	.494	.381	---									
EDPLAN	.256	.407	.221	.422	.262	.211	.333	.512	.463	.678	---								
FAOCC	.096	.195	.109	.185	.062	-.092	.113	.075	.007	.094	.136	---							
FAOCCCOH	.080	.113	.111	.122	-.017	-.103	.061	.060	.072	.022	.051	.634	---						
FAEDCOM	.110	.109	.007	.081	.035	.020	.129	-.013	.083	.105	.130	.508	.482	---					
FAED	.134	.166	.031	.120	.015	.050	.089	-.071	.009	.056	.101	.488	.464	.801	---				
MAEDCOM	.106	.115	-.040	.032	.017	.083	.074	.024	.066	.084	.105	.311	.280	.506	.457	---			
MAED	.123	.120	-.005	.094	-.015	.105	.076	-.010	.022	.102	.084	.352	.303	.436	.481	.802	---		
NSIB	-.259	-.154	-.070	-.146	-.082	-.008	-.038	-.058	-.054	-.067	-.105	-.143	-.192	-.114	-.116	-.156	-.159	---	
SEXCOM	.065	-.036	.125	-.120	-.158	.075	-.184	-.088	-.084	-.021	-.079	.072	-.006	-.049	-.060	.038	.065	.061	---
Mean	44.19	43.83	45.13	44.42	2.92	2.32	3.32	.367	.315	12.97	14.39	28.84	28.47	9.35	9.48	9.11	9.18	4.51	.49
S. D.	7.62	9.34	10.12	9.12	1.56	1.95	1.77	.482	.465	1.32	2.25	20.85	20.55	2.80	2.82	2.32	2.35	2.48	.50

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Comparison of these means to those of whites suggests that Mexican-Americans have lower levels of socioeconomic background, and have ability scores about ten points less than those of whites. Mexican-American youths completed fewer semesters of academic courses than whites, and fewer Mexican-Americans expressed plans to attend college. These means are in close correspondence to those of blacks, who also exhibit lower levels of socioeconomic background and lower test scores. In 1976, when asked about their planned levels of educational attainment, Mexican-Americans responded on the average that they planned to complete 14.4 years of schooling; the corresponding figure for whites was 14.9, and for blacks was 15.1; at the same time, Mexican-Americans had actually attained 13.0 years of schooling, while blacks had attained 13.2 years, and whites had attained 13.5 years.

When the structural and measurement models were estimated for the Mexican-American subsample, as shown in Table 10 a chi-square goodness-of-fit statistic was obtained equal to 440.60 with 121 degrees of freedom. Examination of the first-order derivatives suggested, as for whites, that the error terms for FAEDCOM and MAEDCOM may be correlated. A new model with this parameter free resulted in a chi-square value of 411.91 with 120 degrees of freedom; this was a significant improvement in fit. A new examination of the first-order derivatives suggested ε_{10} and ε_8 may be correlated. The resulting improvement of fit is marginal, even if statistically significant; because these variables are substantively unrelated, and the improvement in fit was not large in relative terms, the final model adopted for whites was the one in which δ_3 and δ_5 are correlated, but all other error covariances are specified to be zero. This happens to be the identical model as adopted for whites, and suggests that neither

Table 10. Goodness of Fit for the Different Models of Educational Attainment of Mexican-American 1972 High School Graduates (Pairwise N=493)

Model	χ^2	d.f.	Prob.	$\Delta\chi^2$	d.f.	Prob.
Uncorrelated errors	440.60	121	0.0			
$\delta_{5,3}$ free	411.91	120	0.0	28.69	1	0.0
$\delta_{5,3}$, $\epsilon_{10,8}$ free	406.43	119	0.0	5.48	1	.019

whites nor Mexican-Americans report these variables with any substantial systematically correlated errors.

Measurement Model: Mexican-Americans

Reporting errors for Mexican-Americans are concluded to be basically random; the only two variables with substantial covariation between their errors are NCES-constructed variables, and the systematic component of covariance could have been introduced in their construction. While the errors of measurement may be random, they are nonetheless substantial. The coefficients are shown in column 6 of Table 11. Compared to the estimated reliabilities among whites, the Mexican-American respondents report their father's occupation as reliably, or as unreliably, as among whites, but both whites and Mexican-Americans exhibit reliability coefficients greater than among blacks; reliabilities of parental education were moderately less for Mexican-Americans than for either whites or blacks. For all three groups the reading and math subtests of ability were more reliable indicators than the vocabulary and letter-group subtests. In measuring academic preparation among Mexican-Americans, math and science courses are more reliable than foreign language courses. Between the two manifest measures of education, planned education is apparently somewhat more reliable as an indicator of educational attainment for Mexican-Americans than is actual education; this is in contrast to the results for both blacks and whites, for whom actual education was the more reliable indicator.

Some of these differences in reliabilities for the three groups are noteworthy. For example, both whites and Mexican-Americans report their father's occupations, college plans and education more accurately than do

Table 11. Measurement Model Parameter Estimates for Mexican-American 1972 High School Graduates (Pairwise N=493)

Variable		Observed Variance σ_i^2	Error Variance σ_e^2	True Variance σ_τ^2	Relative Slope λ_{ij}	Reliability Coefficient $(\sigma_\tau^2/\sigma_i^2)\lambda_{ij}^2$
True	Observed					
Ability	VOCAD	58.01	36.33	49.23	.664	.37
	READ	87.15	25.90		1.115	.70
	LETTERS	102.13	61.05		.913	.40
	MATH	83.00	33.78		1.0*	.59
Academic Preparation	SCT	2.43	1.47	2.20	.662	.40
	FORLAN	3.79	3.29		.478	.13
	MATHSEM	3.13	.93		1.0*	.70
College Plans	PLAN 1	.232	.058	.174	1.0*	.75
	PLAN 2	.216	.106		.792	.51
Education	EDATT	1.73	.66	1.07	1.0*	.62
	EDPLAN	5.06	1.14		1.918	.78
Father's Occupation	FAOCC	433.54	124.75	260.84	1.088	.71
	FAOCCCOM	421.39	160.55		1.0*	.62
Father's Education	FAEDCOM	7.85	1.67	6.16	1.0*	.78
	FAED	7.92	1.37		1.032	.83
Mother's Education	MAEDCOM	5.39	1.29	4.14	1.0*	.77
	MAED	5.49	.71		1.074	.87

* Fixed value

blacks, but the reports of parental education are in general more accurate for whites and blacks than for Mexican-Americans. Because the level of measurement error varies among the three groups, one should expect ordinary least-squares regression estimates to vary among the three groups as a function of differential measurement errors. It was our initial suspicion that there existed differences in measurement error that led us to eschew the usual regression approach in favor of LISREL. The suspicion has now proved to be a well-founded one. Whites, blacks and Mexican-Americans do not report educational process variables with equal reliabilities; as a result, past comparisons of the differences in the process of educational attainment, which were based on uncorrected regression coefficients, have probably resulted in exaggerated claims about ethnic-group differences in how individuals come to acquire different amounts of schooling.

Structural Model: Mexican-Americans

In this section the results of the structural equation portion of the model for Mexican-Americans are presented. The structural coefficients are shown in Table 12. Here these coefficients will be informally compared to those of whites, while postponing for the moment a formal test of statistical differences.

Among Mexican-Americans, the important predictors of ability are father's occupation and the number of siblings, but only the latter is statistically significant at the .01 level. Unlike blacks and whites, Mexican-American women scored lower on the ability factor than did men, but the net difference was not statistically significant. The detrimental influence of more siblings was greater among Mexican-Americans than among

Table 12. Maximum-Likelihood Estimates of Parameters of the Educational Attainment Process for Mexican-American 1972 High School Graduates (N=493)

Dependent Variable	Independent Variables: Father's Occup.	Father's Educ.	Mother's Educ.	Number Siblings	Sex	Ability	Academic Preparation	College Plans	R^2
Standardized Coefficients									
Ability	.183	.007	.007	-.130*	-.023				
Acad. Prep.	-.024	.051	.052	.036	-.197*	.443*			
College Plans	.112	-.254	.057	.031	-.079	.450*	.167*		
Education	-.083	.145	.016	-.013	.031	.166*	.114*	.437*	
Regression Coefficients[a]									
Ability	.080 (.037)	.016 (.252)	.023 (.221)	-.455 (.142)	-.351 (.693)				.07
Acad. Prep.	-.002 (.008)	.024 (.053)	.031 (.047)	.025 (.030)	-.582 (.147)	.087 (.012)			.22
College Plans	.003 (.002)	-.036 (.014)	.010 (.012)	.006 (.008)	-.069 (.040)	.026 (.004)	.049 (.018)		.30
Education	-.007 (.004)	.064 (.030)	.009 (.026)	-.008 (.017)	.085 (.084)	.030 (.009)	.106 (.038)	1.381 (.171)	.58

[a]Standard errors in parentheses.

*Indicates absolute size of coefficient equals or exceeds 2.57 times its standard error.

either whites or blacks, but the influence of parental education and father's occupation among Mexican-Americans were statistically indistinguishable from zero.

In the determination of academic preparation, among Mexican-Americans none of the social background variables were statistically significant, but ability and sex were. Increments to ability produced additional academic preparation of the same order of magnitude as among both whites and blacks. The Mexican-American men, like the whites, completed more academic courses than did women. Note here, as in the equation predicting ability, that only about one-half as much variance is explained for Mexican-Americans than for whites.

In developing plans to attend college, the Mexican-Americans, as do whites and blacks, depend primarily on ability and academic preparation. None of the social background variables are statistically significant. In this instance, all three groups seem to be nearly the same.

Examining the coefficients for educational attainment, the similarities among the three groups are more striking than their differences. Among the background variables, there are some differences, but it cannot be said that they consistently favor or disfavor any group. Thus, for none of the groups is the effect of father's occupation statistically significant; for Mexican-Americans the influence of father's education is greater than that for blacks and nearly twice that exhibited among whites, but the influence of mother's education among both Mexican-Americans and blacks is negligible, whereas among whites there exists a positive effect of mother's education on educational attainment. Additional brothers or sisters is less a detriment to Mexican-Americans

and blacks than to whites. For none of the groups are either men or women advantaged in terms of educational attainment net of the influence of other variables in the model. For all three groups, the most important predictors of education were the endogenous variables, ability, academic preparation, and college plans. The effect of ability was slightly larger among blacks than among the other two groups, while the effect of academic preparation was marginally greater for whites than for Mexican-Americans. Those whites and blacks who expressed plans to attend college actually attained about 1.7 additional years of schooling on the average, but Mexican-Americans attained only 1.4 years.

In sum, the similarities of effects among Mexican-Americans, blacks, and whites are more impressive than the differences. The coefficients do, of course, vary among the three groups, but the direction of statistically significant effects are most often identical, and the relative magnitudes are often close in value. The question that remains unanswered is whether the differences that do exist between the two groups reflect substantive differences, or whether they are differences that might have been expected to occur by chance. It is to this question that attention will now be directed.

COMPARISON OF STRUCTURAL EFFECTS ACROSS GROUPS

Having obtained estimates of the parameters of the model of educational attainment for whites, blacks, and Mexican-Americans, it can now be asked whether the differences in estimated parameters result from random sampling fluctuations, or whether the differences result from real differences in the process of educational attainment among the three

groups. To effect this analysis, the model is estimated for all three groups, specifying for each group the most appropriate measurement model. While the same structural model is specified for all three groups, the estimates of the structural parameters are allowed to vary as they will across the three groups. Then a new model is constructed, specifying that the parameter estimates in the gamma matrix (effects from exogenous to endogenous variables) and beta matrix (effects from endogenous to subsequent endogenous variables) are invariant across the three groups. If the goodness-of-fit statistics between these two models do not vary significantly it may be concluded that specifying invariant structural effects across the groups does not seriously erode the ability to fit the model to the data. If, however, the two chi-square values are significantly different, then it would be concluded there are differences among the groups large enough to seriously erode the model's ability to reproduce the observed covariance matrix. This approach, in fact, tests for statistical interactions among the structural coefficients across the three groups. Rejecting the hypothesis of invariant gamma and beta coefficients is therefore equivalent to concluding that the structural coefficients of the process of educational attainment varies among whites, blacks, and Mexican-Americans.

When all three groups were considered together, and the gamma and beta matrices were specified to be invariant across the groups (that is, no differences in the process of educational attainment), a chi-square goodness-of-fit statistic was obtained equal to 7009.19 with 412 degrees of freedom. This reflects, of course, a poor fit, but the question of interest is whether the fit is any less worse than a model that does not

constrain the gamma and beta coefficients to be invariant across groups. When such a model was estimated, the chi-square coefficient was 6655.50 with 360 degrees of freedom. The difference between these values is equal to 353.69, which is also distributed as chi-square with 52 degrees of freedom. At the .01 level of probability, it may be concluded that the coefficients in the model of educational attainment differ to a degree not attributable to sampling error.

DISCUSSION

It may be concluded that the process of educational attainment varies among whites, blacks, and Mexican-Americans to a statistically significant degree. But now the common distinction should be drawn between differences that are statistically significant and differences that are substantively important. It is well known that practically any difference, no matter how small, may be statistically significant if the sample size is large enough. In the NLS sample, the sample size is very large indeed. There were nearly 12,000 whites in the sample analyzed here. As a result, one may be very confident in the stability of parameter estimates for whites, and may extend this confidence to the comparison of these coefficients to corresponding coefficients for blacks and Mexican-Americans. It is another question, however, as to whether these differences are substantively important. As noted throughout this paper, the similarities among the structural coefficients across the three groups have been more striking than have been the differences. Thus, our overall impression is that the process of educational attainment may differ among the three groups, but not to a degree worthy of much notice.

Yet it should be made clear to whom these conclusions pertain. The population to which these results are generalizable consists of high school seniors, students who were still in school in the spring of 1972. When one examines the social mechanisms by which high school seniors convert their human capital into additional years of schooling, it does not seem to matter very much whether a person is white, black, or Mexican-American. Those high school seniors who accumulated additional years of postsecondary schooling were primarily those who possessed higher ability scores, who accumulated academic courses in high school, and who had developed plans to attend a college of university.

These results have led to conclusions clearly unlike those of previous analyses of ethnic group differences in the process of educational achievement. Portes and Wilson (1976) concluded that parental status and ability were relatively more important effects of educational attainment among whites than among blacks. Similarly, Kerckhoff & Campbell (1977) found that social background, while important for whites, had practically no explanatory power for blacks. And when Featherman and Hauser (1978) examined ethnic-group differences in educational attainment, they found that the effects of status origins were less important for Mexican-Americans than for blacks. Thus, the conventional wisdom is that net of one's ability and performance in school, whites of higher status backgrounds will acquire more years of schooling than lower status whites. In contrast, previous research suggests that social background for blacks and Mexican-Americans plays only a minor role in explaining their educational differentiation.

These are not, however, the conclusions reached here. It was at first suspected, and later confirmed during the progress of this study, that these groups reported status and educational process variables with differing degrees of measurement error. Because measurement error will bias uncorrected regression estimates, and because different amounts of measurement error among groups will unpredictably exaggerate or understate true differences in structural estimates, this study compared the educational processes for whites, blacks, and Mexican-Americans with reference to structural estimates corrected for measurement error. Examination of these coefficients has led us to conclude that net of ability, academic preparation, and college plans, social background plays about the same role for blacks and Mexican-Americans as for whites. Even for the effects of ability, academic preparation, and college plans on education, such differences as exist are relatively minor.

One should not perhaps leap to broad conclusions from these findings, but they are at least suggestive. First, they suggest that a lot of what we have come to believe about interethnic differences in the process of educational attainment has been based on the mistaken assumption that the size of measurement errors were negligibly small, or at least invariant across groups. But such assumptions are unwarranted, and these results suggest that there is a lot less to be made of differences among ethnic groups than has been previously suggested. Second, these results may indicate the emergence of a heretofore nonexistent class structure within the black and Mexican-American ethnic groups. This is the thesis of Wilson (1978), who believes that for blacks socioeconomic background is emerging as a more important determinant of achievement than race per se.

In any event, a word of caution is in order. This analysis has been based on a cohort of high school seniors who in 1976 had not all finished their educational careers. In measuring years of schooling, both educational attainment and educational expectations in 1976 were used as manifest indicators. Other studies have used educational attainment alone as dependent on social background and educational process variables. While we believe the latent education variable used in this analysis measures the amount of schooling this cohort will eventually attain more accurately than a single manifest measure of educational attainment would have by itself, readers should be warned that comparison of these results to those of previous studies are confounded to some small degree by changes in how educational attainment was defined. These definitional differences are not responsible for the major changes in substantive conclusions reached in this study. Separate analyses not reported here, which were based on educational attainment alone as a single manifest variable, did not yield substantive conclusions any different from those reported here.

It is one thing to say that the process of educational attainment varies but little among the three groups. It is another to say that the outcomes will be the same. Because blacks and Mexican-Americans have mean ability scores nearly ten points below those of whites, take fewer academic courses in high school, have fewer members who plan to attend college, have parents of lower socioeconomic status, and more siblings, thus will the 1972 high school cohort of blacks and Mexican-Americans attain less schooling than their white peers. Equality of educational opportunity will not equalize outcomes when the groups are not equal to begin with. Before blacks and Mexican-Americans can be expected to match

whites in terms of educational attainment, either the groups must be equalized in terms of socioeconomic and school-processing variables, or the process by which these variables translate to years of schooling must be unequalized. Strangely enough, this means decreasing the influence of these variables upon schooling for blacks and Mexican-Americans; as long as blacks and Mexican-Americans have mean values on the independent variables less than those of whites, structural effects equal to those of whites mean less schooling for blacks and Mexican-Americans.

Until the completion of this analysis, there was no comprehensive comparison of the process of educational attainment between majority whites and minority blacks and Mexican-Americans; particularly when one focuses on a recent cohort of high school graduates, and measures differences in structural coefficients net of differences in measurement error. When we did so, the process of educational attainment was found to be essentially invariant among whites, blacks, and Mexican-Americans. The recent entry of this cohort into the labor market has precluded extension of the analysis to the examination of the processes of achievement of occupations and earnings. From what we know about differences in the processes for whites and blacks, however, it would be premature to make any conclusions about differences among whites, blacks, and Mexican-Americans in the achievement of occupations and earnings on the basis of these results.

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