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ABSTRACT

Using data from the National Longitudinal Survey of Youth's continuous work history files, this paper examines how individual and market characteristics influence the unemployment rates of Hispanic youth. The results show that family income, marital status, post-school vocational education, age, and local unemployment rates significantly influence unemployment, especially among women. Hispanic youth joblessness rates are found to be quite high, between 30 percent and 40 percent, due primarily to relatively long spells of nonwork after a job loss. Sex differences in labor turnover results are also found, primarily due to the fact that female nonwork duration is nearly 50 percent longer than that of young Hispanic males. (CMG)

Labor Market Turnover and Joblessness for Hispanic American Youth

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Labor Market Turnover and Joblessness for Hispanic American Youth

This paper estimates the determinants of joblessness of Hispanic American youth with the use of national panel data and an analysis which focuses on the rates of entering and leaving work and nonwork. A main issue is to estimate how individual characteristics and labor market characteristics affect the labor turnover rates of Hispanic youths. For instance, education and skill training for individuals are at the heart of several federal policies to reduce joblessness by improving labor supply, whereas tighter aggregate labor market conditions are associated with efforts to reduce youth joblessness by focusing on labor demand--i.e., maintaining strong aggregate demand for workers primarily by monetary and fiscal policies. Low family income and age have also been used as factors in "targeting" federal employment funds; and recent research has stressed the heterogeneity within the Hispanic community. Higher local unemployment rates are found to reduce male rates of job finding rather sharply.

Consideration of employment policy issues in a turnover model context is in keeping with several other recent studies of youth labor markets. Leighton and Mincer (1979), Heckman and Borjas (1980), Flinn and Heckman (1981), and Stephenson (1982) each used a turnover analysis approach to examine the determinants of high youth rates of joblessness and short periods of job tenure. Each of these studies extends to youth labor markets the basic premise that understanding the relatively high rates of youth joblessness begins with examining the determinants of the rates of entering and leaving spells of work and nonwork. This general agreement, which can be traced to the work of Hall (1972), and, more recently, Clark and Summers (1979), is referred to as the turnover hypothesis of unemployment. This basic premise underlies the current work as well.

Another similarity in the Heckman and Borjas, Flinn and Heckman, and Stephenson studies is the use of maximum likelihood methods to study labor market spell duration determinants. This approach is especially appropriate in that individual labor market duration data are frequently censored and thus cannot be properly studied with standard regression techniques. The advantage of this approach is apparent in several recent empirical papers dealing with unemployment duration--Burdett et al. (1981), Lancaster (1979), Lancaster and Nickell (1980), and Tuma and Robins (1980). This paper also uses a maximum likelihood approach to estimate parameters in several models of determinants of exit rates from work and nonwork, using continuous time, individual data. The data are from the New Youth Cohort, a national panel of nearly 13,000 youths aged 14 to 21, collected by NORC in 1979 and 1980. One-seventh of the youths are Hispanics; they are the subject of this study.

We will first consider several theoretical issues, and then present two different empirical models: a constant hazard rate model, and a model which allows for time dependence. The data are then described. The next section considers empirical results for each model. The final section summarizes implications of the research for Hispanic youth labor policy.

THEORETICAL ISSUES

The purpose of this section is to provide a theoretical framework for the empirical analysis. We consider job finding and job leaving in a stationary world, and briefly discuss nonstationary implications. The discussion focuses on the single individual and assumes a two-state environment in which the individual either works or searches for work.

Job Finding

Simple job search models have been offered as a foundation for the recent empirical studies of unemployment duration by Bjorklund and Holmlund (1981), Flinn and Heckman (1981), and Lancaster (1979). We begin with a similar model.

Assume that an income-maximizing individual, who is not working, searches for work and receives job offers which are sorted into acceptable and nonacceptable offers. Job offers arrive as a random process which we assume to be described by a Poisson process with parameter h, or $h(t)$, $t > 0$. Let $h(t)dt$ be the probability of a job offer in a short interval, $(t,t+dt)$, and let $F(w)$ be a known distribution of wage offers. We assume that accepted jobs last forever and that job offers cannot be hoarded, i.e., a once-refused job offer cannot be later accepted, and workers live essentially forever. The key behavioral decision by the searcher is the determination of a reservation wage w^* at time t, because a choice sequence of $w^*(t)$ leads to a sequence of transition probabilities which may be interpreted as job-finding probabilities.[1] The transition probability, μ, in a short interval $(t,t+dt)$, equals the product of two components, $h(t)$, the job offer probability in that interval, and $[1-F(w^*(t))]$, the acceptance probability, or

(1) $$\mu = [1-F(w^*(t))]\ h(t)dt.$$

A function Θ, called the hazard, or failure, rate, is the limiting value of μ as $dt \to 0$. This limiting value provides a linkage between individual search policy and observed spells of unemployment durations.

Let $G(t)$ be the probability of job finding by an unemployed person at any time before t. Thus, $1-G(t)$, often called the survivor function, is the probability that a person who began an unemployment spell at a time t remains unemployed until time t+dt. We express the relationship between μ and Θ as follows:

$$(2) \qquad \Theta(t) = \lim_{dt \to 0} \mu(t,t+dt)/dt$$

$$\Theta t = \lim_{dt \to 0} \Pr(\text{at job at } t+dt \mid \text{unemployed at } t)/dt$$

Equation (2) can be expressed in terms of the survivor function, $1-G(t)$, and $g(t)$, an associated density function,

$$(3) \qquad \Theta(t) = g\ (t)dt/1-G(t),$$

and, on integration,

$$(4) \qquad 1-G\ (t) = \exp\ [-\int_s^t \Theta(u)du].$$

Equation (4) is the fundamental relation connecting a search policy with unemployment duration; more specifically, equation (4) relates the sequence of job-finding probabilities associated with choice of $w^*(t)$ to the distribution of unemployment duration.[2] If transition rates are constant over time, a product of the stationary search model (Flinn and Heckman, 1981, p. 7), then

(5) $1-G_1(t) = \exp[-\Theta u]$

where u, $u = t-s$, is the duration time in the state.

Furthermore, as is well known, the assumed exponential distribution of search times (u) means that the expected duration of nonwork (D) can be written as the reciprocal of the hazard rate, or

(6) $D = 1/\Theta$

or

$$D = 1/\lim_{dt \to 0}[1-F(w^*(t))] \cdot h(t)dt.$$

An optimal search policy, if one assumes an infinite time horizon and a discount rate, r, involves solving for a reservation wage in the familiar expression (see Lippman and McCall, 1976).

(7) $$c + w^* = (\lambda/r) \int_{w^*}^{\infty} (w-w^*)f(w)dw,$$

where c is the instantaneous (and constant) search cost and f(w) is the known distribution of wage offers. If $w^* < w$, search stops and the offer is accepted. Equation (7) suggests that the searcher should select that w^* which will equate expected marginal costs and marginal revenue from continued search. This is a stationary search process even though w^* may change as other values in (7) change.

A decline in w^* can arise via a leftward shift in the wage offer distribution, an increase in the cost of search (c), a decline in the rate of arrival of job offers (h), or an increase in the discount rate (r). Associated with these effects, as Flinn and Heckman have noted (p. 7), are hazard rate changes. The hazard rate $\Theta(t)$ will increase with a rise in search costs, an increase in the discount rate, or a leftward shift in the distribution of wage offers, which means that each of these three effects, other things equal, would reduce expected nonwork duration, D.

An increase in h, the rate of job offers, ceteris paribus, would produce two effects of different sign: (1) an increase in D via an increase in w^*, and (2) a decrease in D via an increase in Θ, the instantaneous transition rate. Which effect of an increase in h dominates D cannot be determined a priori. Feinberg (1977), however, notes that the second effect dominates in normal and rectangular wage offer distributions.

These theoretical issues can be linked to the main analytical point, training vs. aggregate demand policy as strategies to enhance job finding for Hispanic youth. We expect a reduction in the local unemployment rate to increase the job offer arrival rate. This effect on expected nonwork durations is, however, ambiguous, for reasons just stated. More training would also increase the rate of job offers, but we also expect more training to operate as a rightward shift in the job offer distribution. If w^* did not increase enough to offset this distribution shift, then we would expect that the net effect of greater training would be to reduce nonwork duration. Which effect, training or greater labor demand, would have the greater impact on reducing D is an empirical issue.

Job Leaving

In the job finding discussion, we built on recent developments in job search models used to examine unemployment arising from turnover. To model the rate of leaving a job is more complicated. On the one hand, one might consider a search model of an employed worker similar to the job finding model in that a currently employed individual would be assumed to compare the best rewards from alternative time uses (w*) vs. keeping a current wage, w. Yet such a model has an extra complication; one has to consider the potential actions of both the worker and the current employer in terms of changing the effective rules regarding the quantity or quality of work as well as wage adjustments (Okun, 1981, Chap. II). To develop such a general model is beyond the scope of the present paper. On the other hand, more formal presentations such as that of Flinn and Heckman (1981, pp. 27-30) or Burdett et al.—papers which utilize dynamic programming methods to derive instantaneous utility-maximization rules for leaving a job—are somewhat disappointing in terms of predictive content. That is, according to Flinn and Heckman (p. 30), if one continues to assume time stationary value functions, then the harzard function associated with job leaving is independent of time spent at the job! This seems a stiff price to pay in order to achieve a tractable model, but to drop the stationarity assumption sharply undermines one's ability to derive testable propositions.

A reasonable alternative is to estimate the rate of job leaving in an empirical model which is based loosely on economic theory and to test for the presence or not of time dependence, among other determinants. That is, based on past research, e.g., Burdett et al.(1981), we expect that greater wage rates will be associated with a reduced rate of job leaving. Hispanic youth with relatively more work experience, education, and skill training, variables which may be closely associated with a relatively greater market wage, will therefore be expected to have a lower rate of job leaving than other persons. Similarly, we expect that job separations will be affected by several aspects of the labor market. First, the rate of job leaving should be affected by the overall tightness of the labor market; yet the nature of the effect is unclear, a priori. In an economic downturn, layoffs increase, but voluntary quits presumably will decrease. Similarly, in geographic areas where market and nonmarket alternatives are relatively numerous, such as a large urban area, we would expect job separations to exceed that for persons from rural areas. Finally, we expect that a number of demographic characteristics and past work efforts may affect the rate of job leaving. For instance, if the individual worker's earnings are relatively important to a family, as might be the case in a low-income family, then we would expect job leaving rates to be relatively low. Being married, greater fluency in English, older youth, and a more stable past work history, all may reduce the rate of job leaving.

As per the effect of job tenure on the rate of job leaving, the frequent observation is that persons with relatively more time on the job will have a reduced rate of job leaving—e.g., Leighton and Mincer (1979). Jovanovich (1979), however, presents a theoretical model of worker and firm sorting in which the separation probability at first rises early in the tenure period and then begins to decline with more and more time on the job. This time-dependence effect is tested below.

EMPIRICAL MODELS OF LABOR TURNOVER

The Basic Model

In this section we present the basic stochastic model[3] used to study the determinants of early post-school labor mobility. We assume, following Heckman and Borjas (1980), Robins, Tuma, and Yaeger (1980), and Tuma and Robins (1980), who have presented similar labor turnover models, that the individual is in one of two states at any time, employed or not employed.

We begin by describing an individual's work history in some total observation perid (0, T). Within this overall time period, one may consider an infinite number of smaller time periods and record the individual's employment state, employed or not employed, in each interval. A spell is a continuous period of time in a state. We consider persons in state $\underline{i}$ at time $\underline{t}$ and ask what is the probability that they are in state $\underline{j}$ at some later time $\underline{t + \Delta t}$. We assume stochastic movement over time from one state to another. Specifically, we assume a standard first-order, finite state, continuous-time Markov process generates the distribution of state outcomes over time. The probability that a worker who is in state i at time t then switches to state j at a later time, $t + \Delta t$, is the transition probability $p_{ij}(t, t + \Delta t)$. The transition rate, $\theta_{ij}(t)$, is thus defined as:

(8) $$\theta_{ij}(t) = \lim_{\Delta t \to 0} \text{pr (in state } j \text{ at } t + \Delta t \mid \text{in state } i \text{ at } t)/\Delta t$$

$$= \lim_{\Delta t \to 0} p_{ij}(t, t + \Delta t)/\Delta t,$$

where $i \neq j$. The rate of leaving one state $\theta_i(t)$ is the rate of entering the second state j. The denominator in equation (8), the probability of remaining in state i until time t, is really $1 - G_i(t)$, where $G_i(t)$ is the probability of leaving state i at any time t. The term $1 - G_j(t)$ is called a survivor function, when it gives the probability that a person in state i remains in that state between a start time s and time t. As noted in equation (5), if the transition rates are time independent, then the survivor function is expressed as:

$$(9) \quad G_i\,(t \mid s) = e^{-ur_i},$$

where $u = t - s$. That is, the probability that a nonworking youth remains jobless declines exponentially as the length of joblessness increases. Even though θ_i is assumed time independent, the probability of leaving a state varies over time. According to Tuma and Hanna (1979), this is one of the main advantages to modeling social processes by transition rates and not probabilities of change.

In this paper we assume that the same θ_{ij} exists only for persons of the same values of an observable, fixed, exogenous vector of X variables. We assume a log-linear relationship between θ_{ij} and X, or

$$(10) \quad \ln \theta_{ij} = X\beta_{ij}.$$

We then use the estimated β_{ij} to derive individual θ_{ij}. The log-linear transformation restricts the θ_{ij} to be positive.

Alternative Models

Two alternative model parameters are estimated in this paper for Hispanic American youth.[4] It is instructive to present and briefly describe each model.

Model	Description
(11) Model 1 $\theta_{jk}(t) = e^{(\beta_{jk}\underline{X})}$	This is the time-independence (time-invariant) model just presented as the Basic Model. Transition rates, θ_{jk}, are postulated to be log-linear functions of the observed variable vector, $\underline{X}$.
(12) Model 2 $\theta_{jk}(t) = e^{(\beta_{jk}\underline{X}+\gamma t)}$	This is a time-dependent model which postulates that transition rates decline exponentially over time until some asymptote is reached. I assume a zero asymptote and that the e_{jk} are the same in each period, but time in a spell does alter exit rates.

Estimation. We estimate the β_{ij} by a maximum likelihood method and data on observed spell length. Let Y_i be the observed duration of the ith spell. A spell ends when a state change takes place within the reference period or at the end of the sample reference period, in either case $Y_i=1$; otherwise let $Y_i=0$. In this two-state case, if we assume time-independent transition rates and independence of observed spells, then the likelihood function for leaving the nonwork state j is:

$$(13) \quad L_j = \prod_{i=1}^{n} f_j(u_i \mid \beta_{ij}, \underline{X_i})^{Y_i} \cdot (1-F_j[u_i \mid \beta_{ij}, \underline{X_i}])^{1-Y_i}$$

where n is the observed number of spells in state j. Maximizing with respect to β_{ij} gives maximum likelihood estimates of β_{ij}. With these β_{ij} we can predict individual specific transition rates. In turn, these transition rates can be used to derive various estimates of Hispanic American youth labor mobility, such as the expected work duration, the expected nonwork duration, and the steady-state employment probability.[5]

DATA

The primary data sources of this study are the first two waves of the National Longitudinal Survey of Youth (NLS-Youth) which were collected in 1979 and 1980 by the National Opinion Research Center (NORC) in cooperation with the Center for Human Resources Research at Ohio State University. These data are particularly suited to the research goals stated above. First, the overall sample size, 12,693 youths aged 14 to 21 years, includes 1,924 Hispanics. This relatively large sample size permits disaggregation by sex and the application of criteria which are consistent with employment policy analysis. A second advantage is that the sample is national in scope. A third advantage is that the survey design accounts for all time between January 1, 1978, and the spring 1980 interview—that is, all work and nonwork spells are accounted for in this period. These detailed data have been processed for this study into specific periods of three work-history categories: (1) working, (2) not working owing to layoff, and (3) not working for other reasons.[6] A final advantage is the availability of person-specific environmental variables, such as SMSA and county employment rates, industrial characteristics, and labor demand measures, from the City-County Databook. These data were matched with the NLS-Youth data.

Sample means of the study group of Hispanic youth used here are shown in Table 1. These are individual sample means although the unit of analysis is a spell of work or nonwork and one individual may have more than one spell.

The main data screens used were age and enrollment in school. Persons selected became 16 years old on or before the spring 1979 inter-

Table 1

Explanatory Variables: Sample Means and Standard Deviations

Variables and Definitions	Mean (SD)	
	Men	Women
If 1979 interview was conducted in Spanish	.15 (.36)	.09 (.29)
If believe problem with getting a good job is due to POOR ENGLISH (interview conducted in Spanish)	.26 (.64)	.26 (.57)
Percentage Spanish in county, 1979	1.98 (1.90)	1.97 (1.84)
If MARRIED, spouse present, 1979 (incl. common law marriage)	.39 (.17)	.31 (.19)
Local UNEMPLOYMENT RATE, 1979	5.33 (1.50)	5.17 (1.48)
Percentage population change (1970-75) in county, 1979	60.31 (92.62)	68.91 (93.81)
EDUCATION COMPLETED		
If 0-9 years	.49 (.50)	.35 (.48)
If 10 or 11 years	.27 (.45)	.19 (.40)
If 12 or more years	.24 (.42)	.45 (.50)
AGE in years	21.16 (1.30)	21.41 (1.19)
If not U.S. resident at age 14	.27 (.45)	.24 (.43)
If VOCATIONAL EDUCATION received between Jan. 1, 1978 and spring 1980	.05 (.22)	.09 (.29)

(table continues)

Table 1 (cont.)

Explanatory Variables: Sample Means and Standard Deviations

Variables and Definitions	Mean (SD)	
	Men	Women
ETHNIC ORIGIN		
If Mexican or Mexican American	.69 (.47)	.70 (.46)
If Puerto Rican	.13 (.34)	.10 (.31)
If other Hispanic	.18 (.36)	.20 (.21)
INCOME, net family, 1978	$9,817 (9,663)	$10,188 (9,663)
If work-limiting HEALTH problems, 1979	.07 (.25)	.05 (.22)
If ever STOPPED, BOOKED or CONVICTED of CRIME, 1979	.42 (.49)	.18 (.38)
If ever SUSPENDED from school	.23 (.43)	.14 (.34)
λ	.8193 (1.073)	.5176 (.749)
Number in sample	115	96

view and did not attend college or high school after January 1, 1978. The sample may be thus described as Hispanic youth aged 15 to 21 years on January 1, 1978, more than one-half of whom had left school prior to high school graduation.[7] In fact, roughly 30% had at most 9 years of formal schooling and a large proportion of young men and women had either been suspended from school and/or had a possible criminal record.

EMPIRICAL RESULTS

This section is organized into four parts. We first provide a brief rationale for the empirical specifications and then present the transition rate results. Next, because the transition rate coefficients may not be readily interpretable, we present several derivations with employment policy implications; these results were calculated with the Model 1 transition rate results. We then present results from Model 2, the time-dependence form of the transition model.

Specification

Two empirical models, job finding and job leaving, were estimated with two forms of the transition rate model, the time-invariant and time-dependent specifications, shown as equations (11) and (12), respectively. The same set of observed variables in the vector X were used in each model. The choice of X variables was guided by concern for economic and demographic issues. The X vector was fixed. That is, in general, we did not include X vector terms whose values changed over particular employment spells. Admittedly, however, some terms, such as marital status, were first measured in the New Youth Cohort only in the spring 1979 interview; consequently they may involve a change since January 1, 1978, the start of the employment history reference period.

In the theoretical discussion of job finding, the search costs, rate of job offer arrival, and discount rate were linked to the rate of job finding. Direct measures of job search costs and discount rates are not available in the data. We expect, however, that several aspects of psychic costs of job search may be captured in a set of survey questions regarding perceived problems in obtaining (and holding) a good job. These problems may include language problems and not having lived in the United States very long. Thus, we include whether or not the 1979 interview was in Spanish and if the youth lived outside the United States at age 14. As for the discount rate, we expect that youth who have been suspended from school or have had an adverse encounter with police—e.g., those who have been stopped, booked, or convicted—to attach relatively greater weight to immediate gratification of needs. This, in turn, may be an indicator of a greater personal rate of time preference. One might thus expect such persons to have shorter nonwork durations. Yet job search also involved employers' choices and early school leaving, or a police encounter may lead to fewer job offers by employers (and/or an early dismissal if hired). The net effect on <u>job finding</u> of these proxy measures of discount rate level is thus unclear.

A greater rate of positive job offer arrivals, h, is also measured by proxy terms, including a lower local unemployment rate, higher individual educational level, relatively greater age, and the absence of a work-limiting health problem. We expect each term to be associated with a faster rate of job finding.

The final specification also included a number of demographic and environmental terms which may alter the individual's relative taste for work, the individual's ability to allocate time for market work, or the level of market wage rates available to the individual. These factors, which include marital status, family income level (net of the respondent), ethnic origin, educational level, and post-school vocational educational training, may also affect the rate of job finding and job leaving.

Results. Separate transition-rate estimates for Model 1 for the Hispanic male and female youths are shown in Tables 2 and 3. Each of the models were highly significant statistically as measured by a chi-square ratio. The coefficients indicate changes in the logarithm of the transition rate. As such, it may be more convenient to interpret some coefficients in percentage terms. For example, in Table 2, col. 3, we note that Puerto Rican men had a statistically significant and lower rate of job finding than other Hispanic young men. The antilog of the -0.79 coefficient implies that young Hispanic men who listed ethnic origin as Puerto Rican had job-finding rates which were 55% lower than those of otherwise similar Hispanic men in other locations. This particular result is important in terms of one of the main goals of this paper; namely, to examine ethnic differences within the Hispanic American group. Ethnic group differences, however, were not found for Hispanic young women.

Age of the youths in January 1978 varied from 15 to 21 years. As frequently observed in other youth labor studies, age has an important and statistically significant effect on female Hispanic youth labor turn-

Table 2

Determinants of Rates of Job Findings, by Sex

	Young Women		Young Men	
	Model 1	Model 2	Model 1	Model 2
Constant	-21.64** (10.51)	-19.51** (10.54)	-2.32 (8.34)	-4.13 (8.58)
AGE	.76 (.52)	.63 (.52)	-.05 (.40)	.02 (.41)
INCOME	.04** (.02)	.05** (.02)	-.01 (.01)	-.01 (.01)
If EDUCATION, 0-9 yrs	.98 (.91)	.68 (.91)	-.58 (.91)	-.48 (.94)
If EDUCATION, 10 or 11 yrs	.13 (.31)	.10 (.31)	-.09 (.57)	-.06 (.58)
If EDUCATION, 13-18 yrs	-2.17* (1.26)	-2.17* (1.26)	-1.17 (.83)	-1.11 (.83)
If not in U.S. at age 14	.23 (.29)	.37 (.29)	-.15 (.65)	.00 (.66)
If Spanish interview, 1979	.25 (.41)	.25 (.41)	-.32 (.30)	-.34 (.30)
If Mexican American, Chicano, or Mexican	.25 (.29)	.38 (.29)	-.13 (.26)	-.04 (.26)
If Puerto Rican	-.06 (.43)	-.02 (.43)	-.79*** (.30)	-.68** (.31)
If other Hispanic	—	—	—	—
If work-limiting HEALTH	-.48 (.67)	-.55 (.67)	.74** (.28)	.77** (.28)
If MARRIED, 1979	.16** (.06)	.17*** (.05)	-.04 (.06)	-.05 (.06)
If ever SUSPENDED	.30 (.33)	.35 (.34)	-.013 (.219)	.18 (.20)

(table continues)

Table 2 (cont.)

Determinants of Rates of Job Findings, by Sex

	Young Women		Young Men	
	Model 1	Model 2	Model 1	Model 2
If ever STOPPED, BOOKED, or CONVICTED	.33 (.34)	.46 (.36)	-.24 (.20)	-.02 (.20)
If VOCATIONAL EDUCATION received, 1979	-.83** (.45)	-.99** (.46)	-1.27** (.60)	-1.13** (.60)
Local UNEMPLOYMENT RATE, 1979	-.007 (.082)	-.03 (.08)	-.16** (.06)	-.16** (.06)
λ	-.98 (.76)	-.85 (.76)	.10 (.54)	.01 (.55)
Time dependence, γ	— —	.0013*** (.0004)	— —	.0009*** (.0003)
Log likelihood x (-2)	30.31**	38.03***	39.90***	46.57***
Number of spells	105	105	163	163

Note: Standard errors are in parentheses. Data base in 1979 National Longitudinal Survey of Youth. See Table 1 for means of variables.

*Statistically significant at the 10% level.
**Statistically significant at the 5% level.
***Statistically significant at the 1% level.

Table 3

Determinants of Rates of Job Leaving, by Sex

	Young Women		Young Men	
	Model 1	Model 2	Model 1	Model 2
Constant	18.45* (11.34)	15.59 (11.80)	-11.47* (7.25)	-13.6 (7.35)
AGE	-1.19** (.56)	-1.08* (.58)	.22 (.35)	.28 (.36)
INCOME	-.03** (.02)	-.03** (.02)	-.0004 (.0104)	.005 (.010)
If EDUCATION, 0-9 yrs	-.94 (.97)	-.80 (1.00)	1.02 (.86)	1.07 (.87)
If EDUCATION, 10 or 11 yrs	.50 (.33)	.41 (.34)	.74 (.55)	.77 (.54)
If EDUCATION, 13-18 yrs	2.11 (1.34)	1.99 (1.37)	-1.34 (1.10)	-1.29 (1.10)
If not in U.S. at age 14	.02 (.28)	.03 (.29)	-.49 (.60)	-.40 (.60)
If Spanish interview, 1979	-.26 (.39)	-.25 (.40)	-.03 (.29)	.08 (.30)
If Mexican American, Chicano, or Mexican	-.17 (.33)	-.17 (.32)	-.08 (.25)	.015 (.25)
If Puerto Rican	.21 (.45)	.20 (.44)	.11 (.32)	.19 (.32)
If other Hispanic	—	—	—	—
If work-limiting HEALTH	-.66 (.78)	-.72 (.77)	.42 (.29)	.34 (.29)
If MARRIED, 1979	-.04 (.06)	-.02 (.06)	.12** (.05)	.14 (.06)
If ever SUSPENDED	.09 (.31)	.07 (.31)	.16 (.21)	.26 (.21)

(table continues)

Table 3 (cont.)

Determinants of Rates of Job Findings, by Sex

	Young Women		Young Men	
	Model 1	Model 2	Model 1	Model 2
If ever STOPPED, BOOKED, or CONVICTED	.84** (.35)	.84** (.35)	-.20 (.18)	-.26 (.21)
If VOCATIONAL EDUCATION received, 1979	-.15 (.39)	-.21 (.39)	-.98** (.73)	-1.13** (.60)
Local UNEMPLOYMENT RATE, 1979	-.06 (.09)	-.06 (.09)	.12* (.07)	-.16** (.06)
λ	1.54* (.81)	1.34 (.85)	-.35 (.50)	.009 (.55)
Time dependence, γ	— —	.0014*** (.0005)	— —	.0020*** (.0004)
Log likelihood x (-2)	32.51**	41.67***	39.43***	69.35***
Number of spells	99	99	153	153

Note: Standard errors are in parentheses. Data base in 1979 National Longitudinal Survey of Youth. See Table 1 for means of variables.

*Statistically significant at the 10% level.
**Statistically significant at the 5% level.
***Statistically significant at the 1% level.

over rates. Older female youth found jobs more quickly and left jobs more slowly than younger persons.[8]

Family income also has a positive and statistical effect on the rate of job finding of Hispanic young women and a negative and statistical effect on their rate of job leaving. To the extent such women consider nonmarket activities like child care or home production as normal goods, such a result is somewhat unexpected. That is, a young woman whose family has a relatively greater income may have less need to work in the market and can "afford" to do other things. Yet these young women were selected for inclusion in the sample only if they were not attending school. As noted in Table A.1, selected women had lower average income than school attenders. Also, the sample mean family income for women was only about $10,000 in 1977, a figure which is well below the U.S. average for white or black families (and 11% below that for other Hispanic families). Thus the positive association of family income and job finding rate should be interpreted cautiously because of sample selection criteria and possible nonlinear income effects.

Education and training, two important employment policy alternatives of the federal government, are measured here for effects on labor turnover. Education is measured with a set of dummy variables: the reference group has 12 years of education. The only significant relative educational difference is for women. Those with over 12 years of school find jobs more slowly and leave jobs faster than women with 12 years of schooling. Just why this result emerges is not altogether clear, but it is consistent with women with some higher education being relatively more willing to job-shop. Further research is needed on this point, however.

As for training effects on job turnover, work experience prior to Jaunary 1, 1978, was tried in earlier versions of the model but was omitted here due to some measurement problems. Work experience or age have been used by economists as proxy measures for on-the-job training: here we use age. Training is also measured by a dummy variable equal to 1 if the youth was in a post-school vocational or technical training program. Such training negatively and significantly reduces the rate of job finding for both young men and young women relative to other persons who did not receive this training. This result may be due to such persons being more selective, such persons being less desirable from the employer's viewpoint, or some combination of supply and demand considerations. More research is needed to disentangle these effects.[9]

Unemployment rate at the local level was measured as the 1978 county unemployment rate, a term which is a proxy for the overall "tightness" of the job market. The intention is that this term will reflect differences in labor demand level or differences in job offer flow between locations, but obviously, to the extent that supply-related factors also add to unemployment rates, the measure is not exact. Results here are statistically significant only for Hispanic men: the job finding rate is slowed if the unemployment rate is greater. Having found a job, however, means that the rate of leaving the job is positively associated with unemployment in Model 1 and negatively in Model 2. The latter effect is probably due to an interaction between unemployment rate level and the time-varying parameter which changes over duration in a state. The possible interaction, which is not modeled in this paper, means that the negative sign on unemployment should not be interpreted alone and that greater unemployment rates may still lead to faster rates of job leaving, e.g., more layoffs than quits.

English proficiency and sociocultural adjustments of a recent immigrant are also likely to affect individual job search behavior and potential employment tenure. We assume that persons who answered the 1977 NORC interview in Spanish and were living outside the United States at age 14 had such problems. Results obtained here were not statistically significant for either factor. Perhaps NORC needs to develop a better measure of the extent to which English proficiency is a problem.

Several minor results, especially those which were relatively large and statistically significant, should be listed. Marriage, defined here to include living with a nonrelated adult of the opposite sex, is associated with a faster rate of job finding for women. Another result is that a work-limiting health problem is associated with an increased job finding rate for men. As for problems with police, it appears that having been stopped, booked, or convicted has a large and significant effect on female rates of job leaving. Specifically, young women who had an adverse police encounter left jobs 130% faster than other women. Whether such women are the first to be asked to leave by employers or whether they quit more readily cannot be determined here. We can only note that a police encounter will increase female chances of being jobless.

Processed Results

One of the advantages to estimating transition rates is that one may use the rate estimates to predict various outcome measures. In this section, we present the expected duration of work, the expected duration of nonwork, and the long-run (or steady state) probability of joblessness.[10] The predictions are calculated as follows:

The expected duration in state k = $1/\theta_{kj}$, and

The steady state probability of being in state k = $\dfrac{\theta_{jk}}{\theta_{jk} + \theta_{kj}}$.

These outcomes measures are computed here by predicting case-specific θ_{jk} from the β weights in Tables 2 and 3 and case-specific X values. The average duration of work was 56 weeks for women and 53 weeks for men. Nonwork durations were 33 weeks for women and 22 weeks for men. These nonwork duration differences by sex were the main reason for the steady state joblessness rate differences, 40% for women vs. 32% for men. The high rates of joblessness do vary by economic and demographic factors.

Six different criteria were used to sort the data: ethnic group, local unemployment rate, age, education, English proficiency, and family income. Results shown in Table 4 by different groups thus represent not only the differential β weights associated with the criterion variable in question, but also reflect case-specific values of the X terms.

Subgroup differences among the specific Hispanic subgroups listed in Table 4 were in general not statistically significant at conventional levels in the transition-rate estimates. The Table 4 entry differences for Hispanic subgroups are therefore somewhat tentative. Still, we can note certain differences. Other Hispanic groups, including Cubans, Spanish, and others, do relatively much better in terms of male joblessness rates than Mexican groups or Puerto Ricans. In turn, Mexican groups stay longer at jobs and find jobs faster than Puerto Ricans.

Table 4

Processed Results from Transition Rate Estimates, by Sex and Hispanic Group

	Young Women				Young Men			
	N	Expected Work Duration (Weeks)	Expected Nonwork Duration (Weeks)	Steady-State Nonwork Probability	N	Expected Work Duration (Weeks)	Expected Nonwork Duration (Weeks)	Steady-State Nonwork Probability
Ethnic Group								
All	281	56.03	33.41	.40	426	52.81	21.99	.32
		(38.42)	(23.84)	(.20)		(34.73)	(11.97)	(.14)
Mexican American, Chicano, or Mexican	195	54.60	32.11	.39	286	52.94	22.21	.32
		(37.89)	(24.30)	(.21)		(35.70)	(12.36)	(.14)
Puerto Rican	29	47.57	35.53	.39	55	38.03	27.98	.43
		(36.81)	(21.76)	(.18)		(16.05)	(8.57)	(.11)
Other Hispanic	57	65.22	36.80	.39	85	61.96	17.36	.23
		(39.99)	(23.19)	(.20)		(37.30)	(10.71)	(.09)
Local Unemployment Rate								
If 0 to 5.9%	184	58.87	32.72	.40	261	49.12	17.79	.29
		(42.23)	(18.81)	(.20)		(33.97)	(9.66)	(.12)
If 6% or more	97	50.64	34.72	.40	165	58.66	28.63	.36
		(29.32)	(31.30)	(.20)		(35.27)	(12.30)	(.15)
Age								
If ≤ 18 years	12	16.83	35.94	.67	22	36.47	29.71	.45
		(4.74)	(13.89)	(.05)		(11.49)	(8.20)	(.12)
If 19+ years	269	57.78	33.30	.39	404	53.71	21.57	.31
		(38.33)	(24.20)	(.20)		(35.37)	(12.01)	(.14)

(table continues)

Table 4 (cont.)

Processed Results from Transition Rate Estimates, by Sex and Hispanic Group

	Young Women				Young Men			
	N	Expected Work Duration (Weeks)	Expected Nonwork Duration (Weeks)	Steady-State Nonwork Probability	N	Expected Work Duration (Weeks)	Expected Nonwork Duration (Weeks)	Steady-State Nonwork Probability
n								
ears	111	32.18 (10.07)	36.12 (16.31)	.52 (.17)	214	47.05 (23.92)	25.93 (8.99)	.38 (.13)
11 years	65	47.24 (27.37)	31.58 (19.56)	.42 (.19)	136	49.33 (41.41)	16.21 (8.92)	.29 (.13)
ears	105	86.69 (42.44)	31.69 (31.62)	.27 (.16)	76	75.28 (38.78)	21.22 (18.28)	.21 (.09)
Problem								
sh interview	24	57.29 (39.4[illegible])	33.17 (24.34)	.40 (.20)	58	68.94 (38.15)	28.89 (15.97)	.31 (.09)
anish interview	257	42.57 (21.16)	35.91 (17.73)	.46 (.18)	368	50.27 (33.54)	20.90 (10.85)	.32 (.15)
ncome								
e ≤ $10,000	157	42.58 (20.89)	41.48 (26.58)	.49 (.17)	304	48.57 (31.13)	21.38 (10.67)	.32 (.13)
e > $10,000	124	73.06 (47.79)	23.19 (14.46)	.29 (.18)	122	63.39 (40.70)	23.50 (14.66)	.30 (.16)

Local unemployment rate was measured here as a continuous variable in the analysis but split into a dummy variable to develop the Table 4 entries. A greater unemployment rate is associated with a much greater long-run joblessness rate for men, 36% vs. 29%, a result which is primarily due to an increased length of an expected nonwork spell. No direct effect of a local unemployment rate change was found on the joblessness rate of women. Still, the component parts, work and nonwork durations, did change.

Age of the youth is a proxy for a number of employment-related factors. Some employers may prefer older youth or be prevented by state laws or insurance clauses from hiring youth aged 16 or 17 years. Also, older youths may simply be more willing to stay longer at a job, especially if they have car payments, family obligations, and other financial needs. Age was a highly significant determinant of rates of entering and leaving jobs. Results in Table 4 show these effects dramatically. A three-year age difference (three years is the difference in the average age in the above 18 year group, 20 years, and the below 18 year old group, 17 years) is associated with a threefold increase in the expected duration of a work spell for women and a similar but less sharp change for men. For women, the expected duration of work increases from 17 weeks to 58 weeks between ages 17 and 20 years. The length of time not working also appears to fall in this period. As a result of both factors, shortened nonwork spells and lengthened work spells, the steady-state joblessness rates fall sharply.[11]

Three other results presented in Table 4 concern educational attainment, English proficiency, and family income. We focus here on education and family income, two potential target criteria for employment policies. We do not discuss the English proficiency results because we feel that they were poorly measured.

For both sexes, the long-run joblessness rate for high school graduates and youth with college is about one-half that for youths with at most 9 years of formal education. Greater family income is also associated with a lower joblessness rate, especially for women. The policy implications are that Hispanic youth from low-income families should be aided in some manner, be it training or job-finding assistance or some other scheme. Also, Hispanic youths who have left school prior to secondary school completion should be encouraged to return to school so as to enhance their subsequent employment chances.

Time Dependence. The results presented so far have been for Model 1, which assumes that transition rates do not vary over time. Yet there are several reasons why such an assumption may not be appropriate. For instance, a change in economic conditions during a spell of work (or nonwork) may cause a change during the spell in the rate of job finding (or job leaving). Also, a decline in the reservation wage over the duration of time not working may increase the rate of job finding. If such effects are the only source of time variation, then the time-invariant model has biased constant terms, but the bias in other coefficients is usually slight.[12] We therefore show here the effect of a time-varying parameter only on the constant rate.

The time-varying parameter estimates shown in Tables 2 and 3 are highly significant statistically for young men and young women. For both sexes and both work and nonwork categories, exit rates increase over time in the state. For youth in a nonwork state, such a result is consistent with several aspects of job search theory, including a declining reservation wage rate and an increasing spatial distance in job search efforts. As for employed workers, sorting by firms or employees during early tenure could account for this time dependence. Firms need to decide if they wish to keep the worker, while the young worker needs to decide if the job matches his or her career goals. Similar ideas were mentioned earlier by Jovanovich (1979) as to why the rate of job leaving for employer persons need not be monotonically declining, but may increase early in the tenure period. For a sample of mainly teenaged youth, it is not really surprising that positive time dependence is obtained.

CONCLUSION

In this study we have considered the determinants of the rates of entering and leaving work for a national sample of young Hispanic men and women. Data studied were continuous work histories for individuals in the period from January 1978 to spring 1980. Youth studied here, aged 15 to 21 years at the start of the period, did not attend school in this two-year period and were unlikely to return to school. Roughly 70% did not have a high school diploma and 43% had at most 9 years of education. Also, 26% of the youth lived abroad at age 14, and 35% were married. To adjust for special sample selection criteria, we estimated and included Heckman's lambda, which is presented in Appendix A.

We have examined one aspect of Hispanic youth employment problems: the association of high joblessness rates with high labor turnover rates.

Three aspects of the study are important. First, relatively little research has been directed at Hispanic youth employment. This study adds to that literature by describing Hispanic youth labor turnover behavior and by relating a number of economic and demographic issues to this behavior. Family income, marital status, and post-school vocational education, for example, were found to have serious and statistically significant effects on turnover rates, especially for women. Age and local unemployment rate levels also were associated with differential rates of labor turnover. Prior studies have also found these factors, plus family income and others, to be important determinants of labor market behavior.

Second, several policy alternatives were implicitly considered to see how they might affect Hispanic youth rates of entering and leaving employment--e.g., labor demand variation (as measured by local unemployment rate) and education and training provision. While above-average local unemployment rates were associated with lower rates of job finding for men, but not for women, no clear picture emerges as to whether or not this policy or that is better. Instead, one is left with a set of policy-relevant observations:

- Hispanic youth joblessness rates are quite high, between 30 and 40%, and these rates are due primarily to relatively long spells of nonwork after a job loss.
- Age, education, and family income level all sharply affect Hispanic youth employment behavior and thus call for "targeting" employment policies according to these criteria.
- Sex differences in labor turnover results also were found, primarily due to the fact that female nonwork duration was nearly 50% longer than that of young Hispanic men. Employment policy targeting by sex for Hispanic youth may therefore also be appropriate.

- English-language training may be needed for Hispanic youth, but results obtained here do not support such a policy. Data better suited to measure this effect may suggest that such training is appropriate.

A third and final comment concerns the method of analysis. Most of the results presented were for a time-invariant model which assumed an exponential distribution of "wait" times at work or nonwork. A time-varying transition rate model was also presented in which exit rates were found to increase during time at work or not at work. Yet the earlier results obtained with the constant rate model were affected only slightly in that the main change was in the constant term and not, for example, the relative education effects on job finding. More research is needed to understand more fully the nature of this time dependence.

APPENDIX A. AN ADJUSTMENT FOR POTENTIAL SELECTION BIAS

The main focus of this paper is the early post-school labor market behavior of Hispanic youth. To create an analysis file from the original longitudinal data file, only youth who had left regular school on or before January 1, 1978, were included. The risk is that systematic subgroup differences in the characteristics associated with school-attenders vs. school-leavers may blur one's ability to obtain an unbiased estimate of the relationship between a youth's particular characteristic and rate of job finding (or job leaving). The problem cannot be overcome merely by adding more and more right-hand-side variables, since unobserved subgroup differences may also lead to this bias.

James Heckman (1979) refined a statistical method which enables consistent parameter estimates to be obtained in the case in which one, first, has a binary choice, include/not include, and second, has an ordinary least-squares regression for the outcome variable. In the present paper, the situation is somewhat different. Heckman assumed a bivariate normal distribution of the error terms in the binary choice and the outcome variable models. In this paper, we estimate g, Heckman's selection bias adjustment factor, by maximum-likelihood probit methods. This much is exactly as Heckman developed it. The difference arises in the second step, in that the outcome variable(s) estimated here is the instantaneous rate of finding or leaving a job, an assumed continuous-time Markov process which we also estimate by maximum likelihood methods. The statistical properties of Heckman's approach in the context of such a turnover analysis have yet to be developed. See Stephenson (1982) for a related application. Intuition suggests that less bias will be present with λ included than it it were omitted.

Table A.1 presents sample means for the selected and nonselected subgroups. As noted, the youth here were older, from lower-income families, and had less formal education than youths continuing in school or college. In addition, from the other differences listed it appears that early school leavers may have sharp social, economic, and cultural differences from the nonselected youth. Early school leaving appears to be associated with having lived outside the United States at age 14 and other potential English-language problems, which may in turn be related to early post-school and labor market success.

Table A.2 shows maximum likelihood estimates computed by Heckman's lambda-probit routine. The specification is intended to reflect tastes for schooling and budget constraints. Several points should be noted. First, each model is highly significant as indicated by a chi-square statistic (which is, -2 times the difference between the log likelihood ratio of the estimated model from the likelihood based only on the intercept). Second, for both young men and young women, age and, to some extent, education, are the dominant variables determining continued enrollment in regular school or not. In addition, for young Hispanic men, not having been in the United States at age 14 is associated with a lower rate of school retention.

These probit coefficients in Table A.2 were used to predict the probability of being in school for all youth, $F(Z)$, and a λ for each youth was computed as $\frac{f(Z)}{1-F(Z)}$, where $f(Z)$ is the density function evaluated at the estimated probability. This λ was then used an an instrument in the exit rate empirical estimations.

Table A.1

Sample Means of Selected and Nonselected Hispanic Youth Aged 16-21 Years in 1979[a]

	Selected	Not Selected
Age	21.33 (1.25)	19.23 (1.59)
Family income, 1978 dollars (000)	9.986 (9.642)	11.092 (10.462)
If education, 0-9 years	.43 (.49)	.25 (.43)
If education, 10 or 11 years	.24 (.43)	.45 (.49)
If education, 13-18 years	.04 (.19)	.13 (.33)
If not in U.S. at age 14	.26 (.44)	.04 (.21)
If married	.40 (.49)	.06 (.25)
If interviewed in Spanish	.12 (.33)	.04 (.18)
If problems in getting a job due to English	.30 (.46)	.14 (.35)
Number in sample	211	433

[a]The main sample selection criterion was not to have attended school or college after January 1, 1978. The selected sample includes 115 men and 96 women.

Table A.2

Probit Coefficient Results for Sample Selection

	Men		Women	
	Probit Estimates	Mean	Probit Estimates	Mean
Constant	14.47*** (3.45)	1.00	22.69*** (3.99)	1.00
Age/10	6.64*** (1.61)	2.00	-10.68*** (1.88)	1.99
Family income/($000)	-.008 (.005)	10.75	-.005 (.009)	10.71
If education, 0-9 years	1.96 (4.42)	.33	-4.38 (4.86)	.29
If education, 10-11 years	10.08** (5.11)	.40	-5.00 (5.11)	.35
If education, [illegible]-18 years	1.50*** (.41)	.09	2.22*** (.43)	.11
If not in U.S. at age 14	-1.27*** (.40)	.12	.10 (.31)	.11
If education, 0-9 years*Age	-.18 (.22)	6.45	.13 (.24)	5.45
If education, 10-11 years*Age	-.52** (.25)	7.82	.23 (.25)	6.77
χ^2 with 8 d.f.	238.77***		200.03***	
Number in sample	321		323	

** and *** indicate statistical significance at 1% level and 5% levels, respectively.

NOTES

[1]If $w^* < W$, the market wage offer, the job is accepted and search stops.

[2]Lancaster (1979, pp. 940-941).

[3]The Basic Model description closely follows that in Stephenson (1982).

[4]This section is similar to that in Tuma (1979).

[5]See Tuma and Robins (1980) concerning the mathematical derivations of these outcome measures.

[6]In the empirical work, I tried to examine three, not two, states. This choice is technically feasible and exploits the available data more fully.

[7]Because of potential selection bias due to having screened out youth still in school, an adjustment factor was created using a routine developed by Heckman (1979). The auxiliary equations used for that calculation are presented in the Appendix.

[8]Inclusion of this age term is also important as a way of mitigating estimation problems resulting from not controlling for initial conditions.

[9]These education and training effects are described here as person-specific. In fact, the unit of analysis was spells of work and nonwork. To the extent that education and the number of spells are related, these results may be over- or understated.

[10]Details regarding the mathematical derivations of these expressions are in Tuma and Robins (1980).

[11]Of course, some of these processed age results may be due to the effect of other factors such as education or marriage. For example, if older youths are more likely to have graduated from high school and youths with this amount of education leave jobs more slowly, then an age-specific subsample work-exit prediction really reflects not only differences in subsample ages weighted by the age coefficient, but subsample differences in education attainment weighted by the work-exit rate coefficient for education. To decompose these components is beyond the scope of this paper.

[12]Robins, Tuma, and Yaeger (1980, p. 564). This relatively slight change in rate coefficients between Model 1 and Model 2 is found here, with the exception of the unemployment rate effect in the male results for job leaving.

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