

**THE SURVEY OF INCOME AND
PROGRAM PARTICIPATION**

**THE EFFECTS OF BEING UNINSURED
ON HEALTH CARE SERVICE USE:
ESTIMATES FROM THE SIPP**

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TABLE OF CONTENTS

	Page
I. INTRODUCTION.....	1
II. A MODEL OF VOLUNTARY JOB MOVES.....	3
III. A MODEL OF INVOLUNTARY JOB MOVES.....	5
IV. WAGE IMPACTS OF STAYING AND MOVING.....	7
V. ECONOMETRIC MODEL.....	8
VI. DATA SOURCES.....	10
VII. WAGE EQUATION ESTIMATES.....	16
VIII. WAGE GAINS AND LOSSES FROM MOVING.....	20
IX. DETERMINANTS OF JOB CHANGING BEHAVIOR.....	23
X. CONCLUSIONS.....	28
REFERENCES.....	31
ENDNOTES.....	34

I. INTRODUCTION

Job mobility has traditionally served as a primary adjustment process for workers and employers. The economic conditions affecting job mobility decisions range from discretionary choice factors such as a desire for better pay, working conditions, and opportunities for advancement to exogenous factors beyond the control of workers--for example, plant relocations and shutdowns, recessionary or structural declines in the demand for labor, and technological change. This distinction between mobility emulating from discretionary factors and mobility resulting from exogenous changes suggests that in analyzing job changing behavior, it is useful to separate workers into three groups: voluntary movers, involuntary movers, and stayers. Analysis of determinants of the job changing behavior of members of the first two groups is important in gaining a better understanding of worker mobility and labor market flexibility. For example, worker response to wage differences between current jobs and potential alternative jobs helps determine whether such differences serve as an efficient labor supply adjustment mechanism. A related concern is the role of specific and general human capital on job moves and wages.

Using a unified conceptual framework, this paper outlines a sequential two-stage probability model for job moves. This model is estimated on a nationally representative sample of white adult male workers from the 1984 panel of the Survey of Income Program Participation. The model's first stage, which is described in Section II, focuses on voluntary job moves and the second, which is discussed in Section III, pertains to involuntary job moves. Stage I is treated as supply-driven; workers are viewed as selecting themselves as either voluntary movers or erstwhile stayers. Stage II is modelled as demand-driven; employers are viewed as selecting some of the erstwhile stayers to lay-off or fire. Hence, some erstwhile stayers become involuntary movers. The distinction between voluntary separations as a function of alternative wages in the external labor market and involuntary separations based on internal conditions within the firm is also found in employment contract and wage bargaining models (Hall and Lazear, 1984).

The probability of making a job change at either Stage I or Stage II is modelled in terms of wage differentials, a set of human capital determinants, and industry and

occupation variables. The model focuses on the role of wages, rather than on non-pecuniary, job satisfaction factors (Gottschalk and Maloney, 1984; Akerloff, Yellen and Rose, 1988). We particularly emphasize the importance of the gap--either positive or negative--between workers' wages on their present job and their alternative or opportunity wages--that is, the best wage they could receive on an alternative job. If this gap does, in fact, influence decisions on whether or not to change jobs, wage levels subsequent to these decisions will, in turn, be affected. This topic is examined in Section IV.

Then, after discussing certain estimation issues in Section IV and describing the data we use in Section VI, we begin the empirical work in this paper in Section VII by presenting wage equations that are estimated separately for voluntary movers, involuntary movers, and stayers. Parameter estimates from these wage outcome relations are then used in Section VIII to predict the wages movers would have received had they stayed and stayers would have received had they moved. Comparisons of these predicted wages to each group's observed wages yield measures of the economic gains and losses associated with job changes. Finally, in Section IX, we use probit estimates of the probability of making either voluntary or involuntary job moves to examine whether wage differentials, human capital, and job characteristics, influence decisions on whether or not to change jobs. Brief conclusions are presented in Section X.

Thus, the emphasis in this paper is on how wage differentials influence both voluntary and involuntary job moves, and on the economic benefits and costs associated with such moves. Compared to other recent contributions to the job mobility literature, the paper is unique in its separate, but entirely symmetrical, treatment of voluntary and involuntary job changes. More typically, researchers have focused on one type of job move to the exclusion of the other. Nevertheless, our research is consistent with earlier empirical work on job mobility. For example, our estimates of the wage losses associated with involuntary job moves are similar to other estimates for laid off workers (D'Amico and Golon, 1986; Podgursky and Swain, 1987). And our cross section estimates of post-move wage rates are adjusted for differences in job tenure, thereby capturing some of the wage dynamics emphasized in studies based on longitudinal data bases (NLS data for young men, Antel, 1986; PSID data, Ruhm, 1987).

II. A MODEL OF VOLUNTARY JOB MOVES

In the standard human capital model of job mobility, wage differentials between jobs are the primary determinant of voluntary job moves.¹ Workers evaluating possible job moves are viewed as comparing the potential discounted stream of wages on the alternative job, net of costs associated with changing jobs, to the discounted stream of the earning loss from discarded specific human capital on the current job. If the former is greater than the latter, workers will move to the alternative job. In this model, wages are endogenous to the job move decision. Workers who voluntarily move are likely to enjoy higher future wages than if they had stayed, but workers who stay are likely to receive higher wages than if they had moved. Thus, workers non-randomly self-select themselves as either voluntary movers or erstwhile stayers on the basis of future potential wages on their current and alternative jobs.

To formalize, an i^{th} worker's decision to voluntarily change jobs at decision point t^* may be viewed in terms of a comparison between the expected present value of wage alternatives defined by $V_i(W_{ai}^*) \geq V_i(W_{si}^*) + C_i^*$, where $V_i(W_{ai}^*)$ is the discounted expected payoff of his alternative wage stream, $V_i(W_{si}^*)$ is the discounted expected payoff of his current wage stream, and C_i^* is expected search and moving costs. The cost of job changing is determined by both job search costs and the psychological disutility associated with the change (Moretensen, 1986). This disutility varies among workers, depending on individual adaptability to change and to uncertainty associated with a new boss and work environment and, perhaps, less stable earnings (Ruhm, 1987). Voluntary job changers presumably perceive their expected earnings on their alternative job, net of the cost of moving, as exceeding that on their current job. Whether this occurs is determined by a random draw from a wage offer distribution, $F(W)$, representing the wage offers made by all firms in the market (Burdett, 1978). For a worker to actually receive a wage offer above his current wage (W_{si}), a situation given by $1-F(W_{si})$, requires him to match the human capital requirements of a job paying such a wage.

Firm-specific human capital, generally thought to be positively associated with longer tenure with a particular employer, affects mobility through wage growth (Hasimoto, 1981). Worker differences in job-specific skills and the corresponding

variance in wages generate different propensities to change jobs (Antel, 1986; Bartel and Borjas, 1981; Mincer and Jovanovic, 1981). Thus, estimates of human capital parameters show steeply declining rates of job change with working age and even more steeply declining rates with respect to length of current job tenure (Mincer and Jovanovic, 1981). Even after adjusting for heterogeneity across worker samples and within worker groups, steeper declines are found for workers with high rates of job specific skill accumulation than for workers who acquire little specificity in their human capital. (Mincer and Jovanovic, 1981)

The i^{th} worker's decision rule at a decision point t^* is given by

$$(1) [1-F(W_s)]E\{V_i(W_{ai} | W_{ai} > W_{si})\} + F(W_s)V_i(W_{si}) - C_i > V_i(W_{si})$$

or

$$(1a) [1-F(W_s)]E\{V_i(W_{ai} | W_{ai} > W_{si})\} - C_i > 1-F(W_s)V_i(W_{si}).$$

In this formulation, if the worker is to voluntarily change jobs, the expected payoff from switching jobs minus the cost of search and moving must be greater than the discounted expected wage stream from staying in the old job. The equations imply that since the wage offer distribution varies across labor markets, the net expected payoff from changing jobs depends upon the industries and occupations in which the worker searches. They also imply that the net expected payoff from changing jobs is affected by various exogenous factors--for example, age, job tenure, and education--since such factors influence both the wage offer distribution and current wages.

The decision rule in equation (1a) can be written as linear combinations of exogenous variables R_{1i} and R_{2i} for the probability of the expected payoffs on the left and right hand sides of the inequality. Thus, the worker's decision rule can be restated as

$$(2) R_{1i}\beta_1 + e'_{1i} > R_{2i}\beta_2 + e'_{2i}$$

which provides the basis for the following unobserved index of utility:

$$(2a) \quad M_{ii}^* = (R_{1i}\beta_1 - R_{2i}\beta_2) + (e'_{1i} - e'_{2i}) = (R_{1i}\beta_1 - R_{2i}\beta_2) + (\epsilon_{1i})$$

where if $M_{ii}^* > 0$, the worker would potentially receive a net benefit from moving and, hence, has an incentive to change jobs.

This formulation allows for the possibility that workers voluntarily change jobs through a self-selection process that is systematically affected by unobserved individual attributes. The effect of wage differentials and pecuniary and non-pecuniary moving costs, as well as their interaction with such exogenous factors as age, job tenure, and education, are combined in the index function. The outcome of the index function, denoted by M_I , classifies individual workers into the voluntary mover and erstwhile stayer categories, respectively, according to $M_I = 1$ if $M_{ii}^* > 0$ and $M_I = 0$ otherwise. In the following section, we derive an analogous set of conditions for involuntary job moves, but one that is premised on employer decision making, rather than worker decision making, and hence, is represented by a second index function.

III. A MODEL OF INVOLUNTARY JOB MOVES²

In the absence of a firm going out of business or union imposed constraints, the probability of involuntary job moves is largely determined by employers' perceptions of individual workers and their productivity. Employers may be usefully viewed as ranking workers in a job separation queue on the basis of differences between their wages and marginal revenue products. Thus, workers with low productivity relative to the wage they are receiving go to the front of the job separation queue (Weiss, 1980).³ Such workers may well have been hired during periods of excess labor demand and, relative to their current market value, subsequently became "overpaid" as labor demand conditions changed.

Consequently, the wage they can potentially receive on their best alternative job is likely to be considerably below their current wage.

Suppose a worker's current wage is above his best alternative wage. In modelling employer termination decisions, we assume that the larger this difference is, the

more likely it is that his current wage exceeds his marginal revenue product and, therefore, the higher his risk of termination. To see this, let us examine an employer of a group of workers who have identical job responsibilities and who are also similar in terms of age, job tenure, education, and other such characteristics; but who nonetheless differ in terms of productivity -- perhaps, because of innate differences in ability or motivation. The employer will wish to retain the more able of these workers and, thus, will keep their wage rates relatively close to or even above the wages they could potentially receive on alternative jobs. But unless the employer is able to pay its low productivity workers substantially less than its more able workers--and, as a practical matter, employers are usually constrained in the degree to which they can adjust wages to reflect individual merit not related to such "objective" factors as job tenure and education--terminating low productivity workers may offer the employer a labor cost savings. Thus, holding age, job tenure, and other workers' characteristics constant, the sign and magnitude of differences between the workers' current and alternative wages can be viewed as leading to a ranking in a queue for involuntary separations. Large, positive magnitudes should occur for those at the front of this queue⁴ and zero or even negative values for workers at the back of the queue.⁵

Using our earlier notation, we therefore hypothesize that an employer will lay off workers for whom the expected wage streams are given by

$$(3) V_i(W_{si}) > F(W_s)E\{V_i(W_{ai} | W_{ai} < W_{si})\} + 1-F(W_s)V_i(W_{si})$$

or

$$(3a) F(W_s)V_i(W_{si}) > F(W_s)E\{V_i(W_{ai} | W_{ai} < W_{si})\}$$

Analogous to our formulation for voluntary moves, the two sides of this inequality are determined by linear combinations of exogenous variables that can be used to defined an unobserved index function:

$$(4) M_{III}^* = (S_{1i}\alpha_1 - S_{2i}\alpha_2) + (E_{1i} - E_{2i}) = (S_{1i}\alpha_1 - S_{2i}\alpha_2) + (\epsilon_{IIIi})$$

As before, this formulation allows the selection decision to depend upon a set of exogenous factors, S --for example the age, job tenure, and education of individual i and the cost to the employer of hiring and training a replacement worker for him--as well the gap between the current and the alternative wage. If $M_{III}^* > 0$, then the labor cost saving to an employer of terminating a workers would be positive, corresponding to an observed outcome of $M_{II} = 1$. Negative values of this index imply a potential revenue loss resulting from termination. Thus, for example, if a negative wage differential for a worker who is relatively productive combines with other exogenous variables to result in $M_{III}^* < 0$, this corresponds to an observed outcome of $M_{II} = 0$, signifying employer selection of the worker as a stayer. Although this selection process puts the burden of choice on the employer, certain worker characteristics and behavior undoubtedly aid the employer in this process.

IV. WAGE IMPACTS OF STAYING AND MOVING

Equation (1a), which is based on standard human capital theory, implies that voluntary movers are likely, on average, to enjoy wage gains as a result of moving and that stayers receive higher wages than they would have had they moved. In contrast to workers who voluntarily move, involuntary movers are erstwhile stayers who are forced to change jobs. Thus, involuntary movers may make larger investments in job-specific human capital at their old job than voluntary movers since their anticipated tenure is greater. Moreover, involuntarily movers may engage in little on-the-job search since such activity is contrary to their commitment to the employer.⁶ In contrast to voluntary movers then, the potential greater loss of job specific human capital, sub-optimal job search, the likelihood (as implied by equation (3a)) that they were overpaid by their previous employer, and the stigma associated with being fired or laid off all tend to impose wage losses on involuntary movers.

Thus, to summarize, we hypothesize that in terms of wages: a) voluntary movers will be better off moving than staying, b) stayers will be better off staying than moving, and c) involuntary movers are erstwhile stayers who would have been better off not moving since they were previously receiving higher wages than those available to them elsewhere. These alternative outcomes are estimated using a set of three wage equations that are interdependent with the selection process

summarized by equations (1)-(4):

$$(5) \ln W_{1i} = a'_1 X_{ki} + e_{1i} \quad \text{iif } M_I = 1$$

$$(6) \ln W_{2i} = a'_2 X_{ki} + e_{2i} \quad \text{iif } M_I = 0 \text{ and } M_{II} = 1$$

$$(7) \ln W_{3i} = a'_3 X_{ki} + e_{3i} \quad \text{iif } M_I = 0 \text{ and } M_{II} = 0$$

where W_{1i} and W_{2i} are the current, post-move wage received by voluntary and involuntary movers, respectively; W_{3i} is the current wage received by stayers; a'_1 , a'_2 , and a'_3 denote unknown parameter vectors; e_{1i} , e_{2i} , and e_{3i} denote normal random error terms with zero means and finite variances; $k=1, 2, 3$ denotes the three mover-stayer subgroups; and X_{ki} is a vector of job and personal characteristics including length of job tenure on current job, years of work experience prior to current job, training and education, time unemployed, health status, marital status, industry, and occupation.

V. ECONOMETRIC MODEL

A major issue in estimating equations (5)-(7) is that they are potentially subject to serious selectivity bias if, as seems likely, workers were assigned to the three mover-stayer groups partially on the basis of non-observable variables that were also related to wage outcomes. This possibility can be formally expressed in the form of the following regression function for an m subsample ($m=1,2,3$):

$$(8) E(\ln W_{mi} | X_{ki}, \epsilon_{ji}) = a'_m X_{ki} + E(e_{mi} | X_{ki}, \epsilon_{ji})$$

for $j = I, II$ and $k = 1, 2, 3$

If $E(e_{ki} | X_{ki}, \epsilon_{ji}) \neq 0$, then OLS estimates of $\ln W_{ki}$ on the vector X_{ki} are biased by the interdependence between the assignment of workers to the three mover-stayer categories and the respective wage outcomes.

To outline our approach for treating this potential bias, we begin by deriving a reduced form probit index function that combines the wage-related and exogenous

variables in equation (2) or (4) into a single vector:

$$(9) M_{ji}^* = g_{ji}'Z_{ji}^* + v_{ji}^* \quad \text{for } j = I, II$$

where the error term v_{ji}^* is normally distributed with zero mean and variance σ_2^* and Z_{ji}^* is a combined vector of exogenous variables that affect the worker's decision rule ($j=I$) and the employer's selection rule ($j=II$). The g vector of coefficients measure the net effect of general and specific human capital and job characteristics on the observable outcome of the latent variable index.

Given separate, univariate estimates and a multivariate normal distribution of error terms for the two probits-- $j=I,II$ --specified in (9), the conditional expected values for the error terms in (8) are written as:

$$(10) E(e_{1i} | M_{Ii}^* > 0) = \sigma_{e1vI} [\phi(g_I'Z_{Ii}^*)/\Phi(g_I'Z_{Ii}^*)]$$

$$(11) E(e_{2i} | M_{Ii}^* < 0, M_{IIi}^* > 0) = \sigma_{e2vI}[\phi(g_I'Z_{Ii}^*)/1-\Phi(g_I'Z_{Ii}^*)] \\ + \sigma_{e2vII}[\phi(g_{II}'Z_{IIi}^*)/\Phi(g_{II}'Z_{IIi}^*)]$$

$$(12) E(e_{3i} | M_{Ii}^* < 0, M_{IIi}^* < 0) = \sigma_{e3vI}[\phi(g_I'Z_{Ii}^*)/1-\Phi(g_I'Z_{Ii}^*)] \\ + \sigma_{e3vII}[\phi(g_{II}'Z_{IIi}^*)/1-\Phi(g_{II}'Z_{IIi}^*)]$$

where $g_j'Z_{ji}^*$ is a standardized normal variable and the Mills ratios reflect the truncated normal distributions of the error terms over the set of wage outcome equations. The assignment of multiple Mills ratios follows a scheme used in other studies of multiple outcomes (Fishe, Trost, and Lurie, 1981; and Tunali, 1986). The coefficients, σ , are covariance terms on ϕ and Φ , the standard normal density and distribution functions. Their signs indicate the direction of distributional shifts of the error terms from wage equations in which the hypothesized truncation effect is present (Maddala, 1983; Greene, 1990).

These conditional expectations, which are based on a multivariate extension of the Mills ratio, lead to augmented wage equations with selection bias variables that are

now written as

$$(13) \ln W_{1i} = a'_1 X_{ki} + \sigma_{e1vI} [\phi(g_I^* Z_{Ii}^*) / \Phi(g_I^* Z_{Ii}^*)] + \mu_{1i}$$

$$(14) \ln W_{2i} = a'_2 X_{ki} - \sigma_{e2vI} [\phi(g_I^* Z_{Ii}^*) / 1 - \Phi(g_I^* Z_{Ii}^*)] \\ + \sigma_{e2vII} [\phi(g_{II}^* Z_{IIi}^*) / \Phi(g_{II}^* Z_{IIi}^*)] + \mu_{2i}$$

$$(15) \ln W_{3i} = a'_3 X_{ki} - \sigma_{e3vI} [\phi(g_I^* Z_{Ii}^*) / 1 - \Phi(g_I^* Z_{Ii}^*)] \\ - \sigma_{e3vII} [\phi(g_{II}^* Z_{IIi}^*) / 1 - \Phi(g_{II}^* Z_{IIi}^*)] + \mu_{3i}$$

The conditional means for the wage disturbance terms are contingent on the selectivity biases in the probit assignment equations. Computing these means for use in equations (13)-(15) requires calculating the estimated values of the index function from the probits as standard normal density and distribution functions, ϕ and Φ , following procedures outlined in Heckman (1979).

This multi-stage estimation procedure addresses several analytical and statistical issues. First, unobserved worker characteristics affecting job changing are included in the latent variable index specified in the probits. Second, wage rates are treated as interdependent with the selection process, thereby departing from the traditional role of wage differentials as exogenously given thermostatic inducements to change jobs. Third, once the wage relations are adjusted for selectivity bias, they can be used to predict the wages movers would have received had they stayed and stayers would have received had they moved. These predicted wages can, in turn, be compared to observed wages to estimate wage gains and losses to workers resulting from the move-stay decisions observed in the data set. As will be seen, these estimates provide evidence on the importance of wage differentials in guiding the reallocation of labor resources in labor markets.

VI. DATA SOURCES

The data used in estimating the model described above come from the Survey of Income and Program Participation (SIPP), a longitudinal survey of households

conducted over a 34-month period beginning in October 1983. The survey was designed to collect detailed data on income, especially on earnings and transfer receipts, and labor force status. Information on training, education and schooling, job characteristics, and work history was also collected at different points throughout the survey (Nelson, McMillen, and Kasprzyk, 1985). Much of the data we use come from a special module to the third wave of the survey, which contains information on the most recent of any changes in employers made during the ten years prior to the third wave (Ryscavage, 1986, and Ryscavage and Feldman-Harkins, 1987).

SIPP is based on a nationally representative, area probability sample of 21,000 households comprising 53,726 persons. The detailed information on job changing, education, and work experience, which was collected from May to August 1984, was for persons sixteen years or older. For estimation purposes, we used a relatively homogenous subsample of white male wage earners, 35 to 55 year old, who were employed in industries other than agriculture and construction. By focusing on prime age male workers, we limit job moving due to school-to-work and work-to-retirement-to-work job transitions. The race restriction was imposed to minimize the influence of discrimination. The industry restrictions were intended to eliminate workers employed in economic sectors where frequent job changes are virtually institutionalized and, hence, not contingent upon the sort of decision-making analyzed here. We also excluded a small number of cases for which there was non-reporting of both hourly wages and monthly earnings and, consequently, a wage measure could not be constructed.

This resulted in an estimation sample of 3,097 individuals who were employed and reported earnings at the time of the third wave SIPP survey in Summer 1984. About 27 percent of this sample changed employers voluntarily and about 8 percent changed employers involuntarily at least once during the ten years prior to the third wave of the survey. Over 65 percent of the voluntary moves were attributed to dissatisfaction with the wages, working conditions, or location of the previous job, with the remaining moves due to family problems and other reasons.⁷ About 90 percent of the involuntary moves were due to layoffs, with the remaining involuntary moves attributable to firings.

Table 1 presents means of the personal characteristics of the stayer and the two mover subsamples as of 1984. As shown by the table, average educational achievement was above the high school level for all three worker groups, although lowest for involuntary movers. Age and total years of work experience was a bit higher for stayers than for either group of movers. The much shorter tenure with the employer of record in 1984 for job movers than job stayers reflects the fairly recent year in which these persons changed employers--which, on average, was 1979--and the fact that some job movers experienced lengthy periods of unemployment upon leaving their former employers. Prior work experience, defined as total work experience minus years of tenure on the current job, was substantially higher for movers than for stayers, reflecting their more recent employment with another employer. A much larger proportion of stayers and involuntary movers than voluntary movers worked in blue collar jobs and in manufacturing, although substantial numbers of stayers and involuntary movers were also employed in white collar and service jobs. In addition, stayers were more likely to be union members and less likely to work part-time than either group of movers.

TABLE 1: Sample Means in 1984

	Stayers	Voluntary Movers	Involuntary Movers
Hourly Wages	\$13.13	\$12.13	\$ 9.42
Years of Tenure			
On Current Job	16.7	3.6	3.1
Total Years of			
Work Experience	25.5	23.4	24.4
Years of Education	13.2	14.0	12.5
Age	44.1	42.1	42.7
Occupation (%)			
Blue Collar	43.9	27.9	53.4
Sales/Office	14.6	21.3	17.7
Service	6.1	7.5	8.0
Professional/			
Technical	35.2	43.1	20.7
Changed Occupations	0.0	61.2	66.6
Industry (%)			
Manufacturing	40.2	30.0	42.2
Transportation	15.5	9.8	8.7
Trade	12.9	18.0	21.7
Service	19.1	31.2	21.5
Public Administration	12.9	11.0	5.9
Changed Industries	0.0	65.1	70.3
Other Characteristics (%)			
Union Member	35.7	14.7	22.3
Married	87.3	83.2	85.1
Veteran Status	55.3	54.2	56.9
Part-time	1.6	3.4	4.7
Sample Size	1995	851	250

For purposes of this study, several variables including length of tenure, length of work experience, and age had to be computed at t^* , the year at which the move-stay decision was made. The value of t^* is available in SIPP for workers who did actually change jobs during the ten years over which SIPP permits job moves to be observed, but the dates of job change decision points are obviously not observed for stayers. Thus, so that stayers' work histories could be appropriately compared to those of movers, each stayer was randomly assigned a synthetic decision point year between 1974 and 1984. This assignment was made by using the distribution of decision years for voluntary movers as the numerical seed values for a tabled probability mass function. Random draws of decision years from this mass function were generated and then used to assign synthetic decision years randomly to stayers. Thus, to permit a reasonable comparison of movers and stayers, the distribution of assigned synthetic years for stayers parallels the observed frequency distribution of decision years for voluntary movers over the ten year interval. This resulted in a mean decision year of 1979 for both voluntary movers and stayers.

The mean variable values shown in Table 2 are measured at t^* , the observed decision year for movers and the synthetic decision year for stayers. Thus, for example, job tenure for movers is computed from the year of job start for the most recently held previous job to the year this job was terminated; for stayers, this measure is computed from the year of job start to the synthetic year of decision and pertains to the job still held in 1984. Work experience for members of all three groups is computed as the difference between the first year in which a worker reported employment for six months or longer--that is, the year in which the worker became a permanent member of the labor force--and the decision year. Age is calculated as the difference between the year of birth and the decision year.

TABLE 2: Sample Means at Year of Decisions, t*

Variable	Sample Means		
	Stayers	Voluntary Movers	Involuntary Movers
t* -decision year	1979	1979	1980
TEN* -tenure on job	11.0	6.4	7.6
WORKEXP* -years since started work for more than six months	21.6	19.5	20.6
EDUC* -years of education	13.2	14.1	12.5
AGE* -age at decision year	40.8	38.8	39.4
Occupation & Industry (%)			
Blue Collar*	43.9	31.1	58.1
Sales/Office*	14.6	21.9	15.3
Service Jobs*	6.1	8.4	5.2
Professional/Technical*	35.2	38.5	21.1
Manufacturing*	40.2	23.7	47.8
Sample size	1995	851	251

As can be seen in Table 2, stayers had much longer job tenure spells than movers at t*, even after the adjustment for the decision year and even though total work experience and age are similar for the three subgroups. This suggests that either because of self-selection or employer-selection or both, stayers were considerably less mobile than movers, even prior to the job changes reported for the latter in the SIPP data. As was also true in 1984, at the year of decision, blue collar and manufacturing jobs were considerably more heavily represented among members of the involuntary job mover and the job stayer groups than among members of the voluntary job changer group, while white collar and non-manufacturing jobs were more frequently held among members of the voluntary job mover subsample.

VII. WAGE EQUATION ESTIMATES

Estimates of wage equations for each of the three subsamples appear in Table 3. These equations are estimated on the log of the hourly wage rate in 1984. Thus, the coefficient estimates indicate the percentage effect on wages of one unit changes in the independent variables. As outlined in Section V, the wage regressions in Table 3 include Mills ratios, which were constructed by using parameter estimates from reduced-form probits, to control for selection bias due to the non-random assignment of workers to the three subgroups.⁸ The wage equation estimates contain a number of important implications that organize around three topical areas: a) tenure, work experience, and job match; b) structural, institutional, and demographic determinants of wages; and c) selectivity biases. We begin by discussing the first of these topics.

TABLE 3: OLS Estimates of Wage Equations

Variables	Stayers	Voluntary Movers	Involuntary Movers
Constant	6.212 (.093)**	6.29 (.142)**	6.83 (.314)**
TEN: years on current job	.008 (.001)**	.031 (.006)**	.022 (.010)*
EDUC: Years of Schooling	.052 (.005)**	.050 (.006)**	.032 (.011)**
PRIOR EXP: Years of work experience elsewhere	.004 (.0018)**	.002 (.002)	.001 (.003)
TRAIN=1 if previous job related training received	.023 (.019)	-.019 (.033)	.070 (.053)
SOUTH=1 if lives in south	-.024 (.021)	-.043 (.038)	-.022 (.065)
VET=1 if veteran	.025 (.018)	.033 (.032)	-.043 (.052)
METRO=1 if live in metro area	.114 (.020)**	.111 (.038)**	.111 (.058)*
UNION=1 if union member on current job	-.006 (.020)	-.016 (.047)	.188 (.063)**

UNEMP=1 if were unemployed 6 months or more in past	-.064 (.027)**	-.113 (.037)**	-.095 (.055)*
Transportation=1	-.003 (.041)	.028 (.057)	-.155 (.095)
Trade=1	-.197 (.043)**	-.147 (.049)**	-.166 (.071)**
Services=1	-.095 (.038)**	-.150 (.040)**	-.283 (.077)**
Public Admin=1	-.094 (.040)**	-.215 (.052)**	-.056 (.112)
Sales/Office Jobs=1	-.103 (.031)**	-.126 (.046)**	-.312 (.089)**
Blue Collar Jobs=1	-.178 (.038)**	-.245 (.049)**	-.323 (.080)**
Non-Professional Service Jobs	-.347 (.043)**	-.547 (.067)**	-.475 (.111)**
MOVEIND=1 if job move between industries		-.019 (.038)	-.058 (.061)
OCCMOVE=1 if job move between occupations		-.093 (.037)**	-.189 (.058)**
MOSS=1 if employed at current job since entering labor force	.024 (.031)		
Part-Time=1 if weekly hours < 30	-.203 (.089)*	-.009 (.098)	-.123 (.147)
Health=1 if health affects ability to work	-.017 (.033)	-.099 (.061)	-.074 (.088)
Marriage=1 if married	.072 (.026)**	.109 (.043)**	.108 (.073)
Mills Ratio	.021	-.037	.123
Probit I	(.051)	(.024)	(.072)*
Mills Ratio	.024		-.039
Probit II	(.051)		(.063)
F value	37.97	18.85	6.83
R ²	.29	.32	.35
sample size	1995	851	251

Standard errors are shown in parentheses with ** denoting significance at the 1 percent level and * denoting significance at the five percent level.

The effects of job-specific human capital on earnings are represented by the estimated coefficients on years of tenure at current job. As can be seen, this coefficient is positive and significant for all three groups, but is largest for voluntary movers, smaller for involuntary movers, and smallest for stayers. Stayers, with average tenure levels four to five times greater than that of either type of mover, have probably mined much of their potential job-specific skill accumulation. Consequently, they have moved sufficiently far along their concave earnings-tenure function that they receive relatively low returns on additional years of tenure.⁹ Job movers, who are at a relatively early stage of job-specific skill accumulation, realize considerably higher wage payoffs for each additional year of tenure. This is especially true for voluntary job changers, persons who presumably made a rational decision to change employers.

Table 3 provides some evidence of the effective transfer of prior work experience and skill accumulation. For stayers, there is a positive and significant coefficient on the variable measuring work experience prior to current job, a coefficient that implies that an additional year of work experience on a prior job increases wages on the current job by 0.4 percent. In contrast to stayers, for both groups of movers, the coefficients on work experience prior to the current job are small and insignificant, although they are positive. Thus, the estimated effect on wages of years of prior work experience is almost the mirror image of the effect on wages of years of tenure with the current employer. One explanation for this is that movers had about 20 years of work experience on their prior jobs, on average, while stayers had less than 10 years. Consequently, the return to an additional year of prior work was greater for stayers than for movers.

Formal schooling, measured by years of education, dominates the human capital determinants of hourly wages. The positive and statistically significant coefficients for stayers and voluntary movers imply that, at the mean, each year of additional schooling increases hourly wages by about five percent. The smaller, but still positive and statistically significant, education coefficient for laid off and fired workers attests to the portability of formal educational credentials within the labor market, even in the face of economic adversity.

Training investments outside of formal schooling have a positive, but statistically insignificant impact on the wages of stayers and involuntary movers. The very small negative and insignificant coefficient on the training dummy for voluntary movers is suggestive of a lack of transferability of prior training investments to new jobs.

Wages in the four industry sectors listed in Table 3 are all compared to wages in manufacturing, the omitted sector. It appears, not surprisingly, that workers in manufacturing receive considerably higher wages than workers in either the trade or services sectors. However, blue collar jobs pay lower wages than professional and technical jobs, the omitted occupational group. Professional and technical workers also fare better than sales workers, office workers, and general service workers.

More importantly from the perspective of this paper, movers who crossed occupational lines fared considerably worse than movers who did not, suggesting the cost of abandoning specific human capital. Indeed, the coefficient estimates imply that involuntary movers who also changed occupations received wages that were 19 percent lower than involuntary movers who continued to work within the same broad occupational category and that voluntary movers who changed occupations received wages that were 9 percent lower than voluntary movers who did not. In addition, there is some hint in Table 3 that movers who crossed industry lines also fared worst than other movers, although these estimated relations are not statistically significant.

The negative coefficient on UNEMP, a dummy variable that equals one if a worker has had at least one jobless spell of over six months since entering the labor force, suggests that lengthy spells of unemployment depress wage rates, especially for movers. Possible explanations for this finding include the erosion of human capital while unemployed, stigma effects, and a declining reservation wage while unemployed.

The remaining coefficient estimates reported in Table 3 are generally consistent with expectations, although they are not always statistically significant. For example, men employed in metropolitan labor markets have wages that are about

11 percent higher than their non-urban counterparts, and married workers have wages that are 7 to 11 percent above those of single workers. Table 3 also suggests that health limitations and part-time employment (less than 30 hours per week) may reduce hourly earnings. Union membership has a large positive and highly significant effect on the wage rates of involuntary movers, but appears to have virtually no effect on the wages of stayers and voluntary movers, a result that has no obvious explanation, except, perhaps, the general weakness of the labor movement during the mid-1980s.

Only one of the coefficients on the Mills ratios in the wage equations reported in Table 3 is statistically significant: the coefficient on the first probit Mills ratio in the wage equation for involuntary movers.¹⁰ The large positive sign on this coefficient implies a leftward shift in the wage distribution of involuntary movers relative to that for the entire sample, suggesting that adverse selection bias reduces the expected value of wage rates for involuntary movers.

VIII. WAGE GAINS AND LOSSES FROM MOVING

Information on the returns to voluntary job moves and losses from involuntary moves can be developed from the wage equation estimates that appear in Table 3. Using these estimates, we predict hourly wage values for each observation had outcomes other than those actually observed occurred. These predictions are developed for stayers had they voluntarily or involuntarily moved, for voluntary movers had they stayed, and for involuntary movers had they been able to stay or had they voluntarily moved rather than being forced to move through layoffs or firings. These predicted wage rates are then compared to those actually observed for each subgroup.

The computation of mean hourly wages for voluntary job movers (denoted by vm) had they stayed is based on

$$(16) E(\log W_{3vm} | v_{Ivm}^* > -g_I' Z_{Ivm}^*) = a_3' X_{vm} - \sigma_{e3vII} [\phi(g_I' Z_{Ivm}^*) / \Phi(g_I' Z_{Ivm}^*)].$$

A similar calculation for stayers (denoted by s) had they changed jobs voluntarily is based on

$$(17) E(\log W_{1s} | v_{1s} < -g_I' Z_{1s}^*) = a_1' X_s + \sigma_{e1vI} [\phi(g_I' Z_{1s}^*) / (1 - \Phi(g_I' Z_{1s}^*))].$$

The predicted wage for stayers had they been forced to move involuntarily is given by

$$(18) E(\log W_{2s} | v_{1s} < -g_I' Z_{1s}^*, v_{II s} < -g_{II}' Z_{II s}^*) = a_2' X_s \\ + \sigma_{e2vI} [\phi(g_I' Z_{1s}^*) / 1 - \Phi(g_I' Z_{1s}^*)] + \sigma_{e2vII} [\phi(g_{II}' Z_{II s}^*) / 1 - \Phi(g_{II}' Z_{II s}^*)].$$

Predicted wages for involuntary movers (denoted fm) are computed using similar algorithms, with appropriate substitution of Mills ratio terms. The predicted wage for involuntary movers had they stayed is calculated by

$$(19) E(\log W_{3fm} | v_{Ifm} < -g_I' Z_{Ifm}^*, v_{II fm} > -g_{II}' Z_{II fm}^*) = a_3' X_{fm} \\ + \sigma_{e3vI} [\phi(g_I' Z_{Ifm}^*) / 1 - \Phi(g_I' Z_{Ifm}^*)] \\ - \sigma_{e3vII} [\phi(g_{II}' Z_{II fm}^*) / 1 - \Phi(g_{II}' Z_{II fm}^*)].$$

Finally, wages for involuntary movers had they moved voluntarily is predicted by

$$(20) E(\log W_{3fm} | v_{Ifm} < -g_I' Z_{Ifm}^*) = a_1' X_{fm} \\ + \sigma_{e3vI} [\phi(g_I' Z_{Ifm}^*) / 1 - \Phi(g_I' Z_{Ifm}^*)].$$

These algorithms account for selection bias by incorporating the Mills ratio coefficients from the wage regressions for stayers and the two groups of movers. To use equations (16) and (19), it was necessary to rescale the tenure and prior work experience measures for voluntary and involuntary movers to what the values of these variables would have been had these persons not changed jobs. Similarly, in using equations (17) and (18), it was necessary to rescale these variables to what they would have been had stayers changed jobs in t^* , their synthetic year of decision.

Comparisons of the predicted wages just described with actual observed wages are shown for each subgroup in Table 4. These results show patterns that are revealing and plausible. Stayers appear to have made a wise decision for themselves. They would lose from leaving their current job, especially by moving involuntarily. The mean wage gain from voluntary moves is positive and, hence consistent with the theory of wage differentials. However, the magnitude of the differential is small, perhaps too small to offset costs associated with moving. One possible explanation for the small size of the differential is that many job moves are motivated by reasons other than opportunities for wage improvements (see Akerloff, Rose and Yellen, 1988; Bartel, 1982). A second reason may be poor labor market information; many voluntary job leavers may be overly optimistic about their opportunities elsewhere. A third possible explanation is that our "voluntary" subsample almost surely contains persons who left their jobs in anticipation of being laid off or in lieu of being fired.

TABLE 4: Mean Wage Outcomes for Alternative Job Mobility Decisions*

	Mean Observed Wage	Predicted Mean Wage "IF" Stayed	Predicted Mean Wage "IF" moved Voluntarily	Predicted Mean Wage "IF" Moved Involuntarily
Stayers	\$13.13 (6.57)	n.a.	\$12.85 (2.98)	\$10.72 (2.93)
Voluntary Movers	12.13 (7.64)	\$11.69 (2.81)	n.a.	n.a.
Involuntary Movers	9.42 (4.62)	10.74 (2.59)	10.01 (2.84)	n.a.

*Standard deviations in parentheses
n.a.: not applicable

Table 4 suggests that involuntary movers suffer substantial wage losses. Apparently, however, these losses could have been partially mitigated by voluntary moves in anticipation of being terminated. The wage rate involuntary movers could have received had they been able to stay as compared to their observed wage after being forced to change employers is a strong argument for a model of job move decision making, such as ours, that treats involuntary movers as erstwhile stayers.

IX. DETERMINANTS OF JOB CHANGING BEHAVIOR

In Section II, we pointed out that standard human capital theory implies that the probability that a worker will change jobs voluntarily is positively related to the difference between the worker's potential wage on his best alternative job and his wage on his current job. And in Section III, we hypothesized that the probability of an involuntary job move is positively related to the difference between a worker's current wage and his potential wage on his best alternative job. In this section, we use probit estimates of the effects of wage differences and other factors on the probability of voluntary and involuntary job moves to test these two hypotheses, and, in addition, examine how such exogenous factors as education, job tenure, and age influence job changing decisions.

These probit estimates appear in Table 5. Model I, which is for voluntary moves, is based on the full sample. Thus, voluntary movers are compared to both stayers and involuntary movers. Model II, which pertains to involuntary moves, is based on comparisons of involuntary movers with stayers.¹¹ As is evident, many of the variables used in the wage equations reported in Table 3 are not also part of the probit specification. The reason for this is that the former are based on data for 1984, while the latter are based on data for the year of decision, t^* . SIPP provides much more complete information for 1984 than for the year of decision.

TABLE 5: Probit Estimates

	<u>Model I</u>	<u>Model II</u>
CONSTANT	1.147 (.271)**	.647 (.388)
WAGE DIFFERENCE	.0167 (.004)**	.0215 (.007)**
EDUC*	.034 (.010)**	-.018 (.015)
TEN*/WORKX*	-1.47 (.079)**	-1.53 (.128)**
AGE*	-.035 (.004)**	-.027 (.006)**
<u>Industry and Occupation Variables</u>		
Manufacturing*	-.349 (.059)**	.143 (.081)
Blue Collar*	-.088 (.075)	.383 (.109)**
Sales/Office*	.164 (.077)*	.267 (.125)*
Services*	.110 (.111)	.067 (.183)
Average Wage Difference	-\$0.12	\$0.34
log likelihood ratio	557.5	200.5
pseudo R ²	.15	.13
sample size	3097	2246

Standard errors in parentheses with ** denoting significance at the 1 percent level and * denoting significance at the 5 percent level.

To construct a wage differential variable for voluntary movers, we subtracted the wage each worker would receive from staying from the wage the worker would receive from voluntarily moving. For example, this variable was constructed for stayers by subtracting their observed wage from the wage they would have received had they voluntarily moved. The value of this latter, hypothetical wage was predicted on the basis of equation (17). For voluntary movers, the wage differential variable was obtained by subtracting their predicted wage for staying, which was based on equation (16), from their observed wage. And for involuntary movers, it was constructed by using equation (19) to predict their wage had they stayed and then subtracting this value from the predicted value of the wage they would have received had they voluntarily moved, which was based on equation (20).

The small negative mean value of this variable, which is shown in Table 5, implies that voluntary changes in jobs would result in wage losses for many members of the sample. The positive and statistically significant sign on the coefficient estimate of the wage differential variable in the Probit I regression equation suggests that the probability of voluntary job moves increase, the larger the potential wage returns from changing jobs. The strength of this coefficient is determined by using an algorithm based on the derivative of the expected value of $y=1$ and the standard normal probability density. Using this procedure, we find that a \$1 increase in wages, *ceteris paribus*, increases the probability of a voluntary move by one-third of one percentage point.

Our examination of the relation between wage differentials and involuntary mobility is similar to our test of whether wage differentials influence voluntary mobility. We first constructed a wage differential variable by subtracting each worker's wage on their best alternative job--that is, the wage for voluntarily moving--from the wage they would receive if they stayed. Thus for stayers, the variable is constructed by using equation (17) to predict the wage they would have received had they voluntarily changed jobs and then subtracting the resulting value from their observed wage. And for involuntary movers, we subtracted the predicted value of the wage they would have received had they voluntarily moved (a prediction based on equation 20) from the predicted value of their wage had they been able to stay (a prediction based on equation 19).

The positive mean value of the variable resulting from these procedures, which is reported in Table 5, implies that most members of the estimation sample used to estimate the Probit II regression, a sample that is dominated by stayers, would have been better off not changing jobs. The sign on the coefficient estimate of this variable in the Probit II regression is positive and statistically significant, supporting the employer selection hypothesis upon which the second probit index function is based. Increasing the wage differential used by employers to rank order their workers in a termination queue by \$1, increases the probability of an involuntary job move by about one-half of one percentage point, an effect that is somewhat larger than that reported above for voluntary move probabilities.

The findings just discussed suggest that wage differentials do play the role we hypothesized on the job moves of prime age white males. On the one hand, as the standard human capital model implies, differences between these workers' best alternative wage and the wage on their current job are positively related to voluntary job changes. On the other hand, differences between their current wage and their best alternative wage, which we hypothesize indexes the extent to which they are overpaid, appear positively related to involuntary job movements. Neither of the estimated relationships, especially the former, were particularly large in magnitude, however, suggesting that even substantial wage differentials do not engender large labor mobility responses.

The remaining independent variables used in the probits reported in Table 5 are intended to capture the human capital characteristics of workers in our sample, as well as the characteristics of the jobs they held at the year of decision, t^* . Educational attainment, as a portable credential with transferable skills, should facilitate voluntary moves to better jobs, while, perhaps, reducing the possibility of being discharged. Age, by reducing the time horizon for receiving wage gains, should be negatively related to voluntary job changes and, if employers are sympathetic to older workers or take account of seniority in their decisions of whom to terminate, may also be negatively related to involuntary job changes.

We use the ratio of tenure to total work experience ($TEN^*/WORKX^*$) in the probit mover-stayer maximum likelihood equations to proxy employee commitment to their jobs and employer commitment to workers and to capture heterogeneity

among workers with respect to mobility propensity and heterogeneity among jobs with respect to layoff propensity. The length of tenure-spells with a particular employer, as a proportion of total time in the work force, is an index of commitment and job specific capital investments by both employee and employer. Workers whose tenure with a particular employer is a relatively high proportion of their total time in the labor force have made a greater investment commitment to that employer than workers with relatively low proportions. Moreover, such workers may be innately less mobility-prone than workers with relatively low proportions. High ratios also represent relatively substantial commitments by employers to particular workers. In addition, high ratios are only feasible for workers who hold jobs where firings and permanent layoffs are infrequent. Consequently, we anticipate that the tenure-work experience ratio will be negatively related to the probability of both voluntary and involuntary job changes.

The propensity to change jobs should also vary among industries and occupations. For example, manufacturing jobs offer relatively high wages, but have dwindled in number in recent years. Thus, one might expect fewer voluntary moves emulating from manufacturing jobs, but a greater number of involuntary moves.

The estimates in Table 5 are generally consistent with our expectations concerning the effects of job specific human capital investments, education, and age and demonstrate their importance as determinants of both voluntary and involuntary job moves. For example, the negative and highly significant coefficients on the tenure-work experience ratio in both models show the lower likelihood of changing jobs, the greater the proportion of time since joining the labor force a worker has spent on a given job. The negative coefficient on age, which is highly significant in both the voluntary mover and involuntary mover equations, indicates that older, more experienced workers are less likely to move, either voluntarily or as a result of being forced to do so by their employer.¹²

The marginal effects of age and tenure can be calculated on the basis of $\delta E[M]/\delta Z^* = \phi(g'Z^*)g$, where the derivative of the expected value for $y=1$ is defined on the predicted index based on ϕ , the standard normal density. Taking the coefficient for age from Table 5 times the pdf for the predicted index, we find that an increase in age from 40, approximately the mean age at t^* , to 50 lowers the

probability of a voluntary move by about seven percentage points and the probability of an involuntary move by about five percentage points. Increasing the tenure ratio by 25 percent lowers the probability of either a voluntary or an involuntary move by about three percentage points. When compared to job movement among members of our sample--27 percent of the sample used in Model I moved voluntarily and 11 percent of the sample used in Model II moved involuntarily--these results suggest that labor mobility is quite sensitive to both age and job tenure.

While the direction and size of the effects of tenure and age on voluntary job moves is similar to that on involuntary moves, the remaining variables differ dramatically in their effects on the two types of moves. Formal schooling, for example, appears to increase the probability of a voluntary move, but decrease the likelihood of an involuntary job change, although in the latter case the coefficient is statistically insignificant.

In the probit equations, the mobility of manufacturing workers is compared to that of workers in all other industries, while the mobility of blue collar, sales, office, and service workers is compared to that of professional and technical workers, the omitted occupational group. It appears that blue collar workers face a considerably greater risk of having to make an involuntary job move than do professional and technical workers, while workers employed in manufacturing may face a greater risk of termination than workers in other industries. These results are indicative of the displaced worker phenomena in American labor markets throughout the past two decades. Indeed, manufacturing workers are apparently much more reluctant to move voluntarily than workers in other industries and, although the relation is statistically insignificant, there is some evidence in Table 5 that blue collar workers are more reluctant to change jobs voluntarily than workers in other occupations.

X. CONCLUSIONS

In this section, we first summarize our conclusions for voluntary job moves. We then briefly examine our findings for involuntary job moves.

Our findings for voluntary job moves imply that although some workers do move towards better paying jobs, relatively few face strong monetary incentives to make voluntary job changes. Actual wage gains made by those in our sample who did move were small, particularly for those who changed occupations, while those who preferred to remain with their current employers--both stayers and involuntary movers--potentially face wage losses from job moves.

However, it also appears that even those workers in our sample who could potentially enjoy large wage gains from job moves were not very responsive to these incentives. Why? There are a number of possible reasons including the lack of information on the part of workers concerning potential alternative wages, uncertainties, search costs, and psychological costs associated with job changes. Furthermore, our empirical estimates imply that age and length of tenure with an employer relative to time since first entering the labor force both have a retarding effect on labor mobility. This finding has especially important implications for our sample since the men in it were already 25 to 45 years old at the beginning of the ten year period over which our data on job moves pertain, and many had been employed by the same firm for all or most of their working lives. Moreover, non-pecuniary factors and negative signals from current employers may play a more important role in many voluntary job change decisions than potential monetary rewards.

The involuntary movers in our sample, especially those forced to change occupations, apparently suffered large reductions in wage rates as a result of being terminated by their former employers. And it appears that at best these job changers could only partially have ameliorated this effect by leaving in anticipation of being laid off or fired. Thus, it seems rational for at least some workers to respond sluggishly to signals that their current job may end, hanging on to the job as long as possible. Moreover, the involuntary movers in our sample were without work for 31 weeks, on average, after leaving their previous job, while the voluntary movers were without work for only 15 weeks, on average. However, while the costs to individuals who are forced to change jobs are substantial--costs that could, perhaps, be mitigated somewhat by retraining--our empirical evidence suggests that the risk of termination is positively related to the extent to which workers are overpaid. Thus, involuntary job movements would appear to have the attractive

allocative property of falling most heavily on the poorest matches between individuals and jobs.

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ENDNOTES

1. In this section, we treat involuntary job changers identically to stayers on the grounds that in contrast to voluntary movers, both stayers and involuntary movers do not wish to change jobs. In the following section, we contrast involuntary movers with stayers.
2. In this analysis, we focus on involuntary movers who are re-employed, ignoring those who are not. Determinants of the probability of re-employment and wage effects for dislocated workers have been estimated by Podursky and Swain (1987), Swain and Podursky (1990), Addison and Portugal (1989), and Howland and Peterson (1988), using Dislocated Worker Survey data from the January 1984 supplement to the Current Population Survey.
3. Implicit contract theory, however, has suggested circumstances under which workers receive wages in excess of their marginal revenue product, yet are retained by their employers. As Hutchens (1989) has pointed out, however, such contracts are likely to apply to no more than a relatively small subset of workers.
4. Instead of a large positive gap between a worker's current and alternative wage indicating that he is overpaid, such a gap could, in principle, result from the worker's acquisition of large amounts of firm-specific human capital. If so, the worker's employer may, in contrast to our argument in the text, have strong incentives to retain him. Another possibility is that employers fail to make strong attempts to retain workers with negative gaps between their current and alternative wage because such workers are likely ultimately to become dissatisfied and leave voluntarily. Consequently, the employer has little to lose from laying off such workers. If either of these circumstances are sufficiently common, a negative (rather than positive) relation could exist between involuntary mobility and the gap between current and alternative wages. Later, however, we attempt to estimate this relation and find evidence that it is, in fact, significantly positive.
5. Although we have developed our model in terms of an individual employer selecting some of its workers as involuntary movers and some as stayers, it should be apparent that a similar framework can be applied to plant closings and firm failures. For example, if as a result of union pressure and downward shifts in demand, wages at a particular plant far exceed the marginal revenue products of workers at the plant, and wages are rigid downward, the plant's entire workforce may be at risk of involuntary separation.
6. In this analysis, we treat involuntary job changes as completely unanticipated although, in reality, many workers have either direct signals or some subjective probability concerning their longevity with their current employer. Thus, some workers who appear in our data as "voluntary" movers are, in actuality, involuntary movers who simply beat their employers to the punch. Indeed, those categorized as "voluntary" movers were about as likely to have timed their move during a recessionary period as those categorized as "involuntary" movers. Our data cover three recessionary periods, 1974-75, 1980, and 1981-1982. During these years, around 27 percent of all the "voluntary" moves in our sample were made, as were around 27 percent of all the "involuntary" moves.

7. The self-reported nature of the data introduces some bias in these responses since, as previously suggested, some voluntary moves may be disguised layoffs and firings. These problems are common to all studies using self-reported data.
8. In Table 5, we present estimates for probit equations that are identical to those used to construct the Mills ratios, except that only those in Table 5 include wage differential variables constructed from the estimated wage equations. The coefficient estimates in these two sets of probits are very similar in sign and magnitude. Table 5 is discussed in Section IX.
9. When squared terms for tenure and net work experience were added to the variables reported in Table 3, the coefficient estimates on these terms confirmed the concavity of the earnings function. These coefficients, however, were not statistically significant, possibly because of the relatively narrow age range of the sample used in this study.
10. The possibility of heteroskedasticity associated with OLS estimates of selection bias is relevant when selection bias is significantly different from zero. We tested for homoscedasticity with the Breusch-Pagan test (Greene, 1990), which is based on a Lagrange multiplier test. This test implies acceptance of homoscedasticity of the error term in the wage equation for involuntary movers.
11. We test for the more general case where $Cov = [\epsilon_{II} \ \epsilon_{III}] \neq \rho = 0$ with a joint outcome based on the distribution parameter $\Xi = (\epsilon_{ki}, \epsilon_{II}, \epsilon_{III})$. Our test leads to acceptance of separate probit estimates. The correlation between the error terms from the two probits ($\rho = -.05$) is not significantly different from zero, thereby supporting the argument for a separate, sequential probit estimating framework.
12. When total work experience was used in place of age in the regressions appearing in Table 5, the level of significance and the magnitude of its coefficient was similar to that of the coefficient on age.