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MIDDLE-AGE JOB MOBILITY: ITS DETERMINANTS AND CONSEQUENCES

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MIDDLE-AGE JOB MOBILITY: ITS DETERMINANTS AND CONSEQUENCES

Ann P. Bartel and George J. Borjas*

It is well known that job mobility is an important characteristic of the working life histories of individuals. Labor Department data indicate that on the average a young man at age twenty can expect to have 6.6 job changes during the next forty years of his working life. While the expected number of job changes declines over the life cycle, it is interesting to note that men aged 45-54 still expect to change jobs an additional 1.4 times prior to retirement.[1] Due to lack of microdata, early studies of job mobility were unable to analyze individual mobility patterns; rather, these studies examined the determinants of inter-industry job separation.[2]

The recent availability of longitudinal data on earnings and job histories has allowed researchers to significantly expand the study of job mobility. For example, Parnes and Nestel (1975) studied the

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[1] See U.S. Department of Labor (1964), Manpower Report No. 10.

[2] For example, see Stoikov and Raimon (1968), Burton and Parker (1969), and Pencavel (1972).

determinants and consequences of job changing using the National Longitudinal Survey of Mature Men (NLS). Their most important findings were that the probability of quitting a job was systematically related to job tenure, job satisfaction, the existence of private pension plans and the individual's current wage. Using the same data set, Borjas (1975) analyzed the effects of differential lifetime mobility patterns on the current earnings of mature men by estimating a "segmented" earnings function--that is, relating the individual's earnings to his job history. The study suggested the existence of a strong positive relationship between human capital investment and job duration. Finally, Bartel (1975), using the Coleman-Rossi Retrospective Life History Study, was able to decompose post-school earnings growth into gains occurring on the job and gains due to job mobility. It was shown that while inter-firm mobility had a positive effect on earnings growth through the wage gain obtained <u>across</u> jobs, the more mobile individuals had significantly lower wage growth per time period <u>within</u> the job. The use of longitudinal (life-cycle) data, therefore, has provided economists with new insights into an important characteristic of labor markets, namely job mobility.

Our paper uses the wealth of information available in the NLS to expand on previous work in several ways. First, we investigate whether there is a meaningful distinction among types of job separations. Traditional analysis has categorized job separations as either employee-initiated (quits) or employer-initiated (layoffs). We question whether this dichotomy is correct. For example, a person who quits his job for personal (exogenous) reasons, such as health, has a different motivation than a quitter in search of a better job. This argument would suggest the need for a more detailed breakdown of quits. On the other hand, it has recently been argued that it is irrelevant which party to the "contract" initiates the separation.[3] A job can be viewed as a match (or marriage) between employer and employee. Since the separation is solely determined by the existence of an improper match, it is unnecessary to know which party initiated the separation in order to know the factors determining the separation. This line of reasoning would, of course, lead to the conclusion that the quit-layoff breakdown is meaningless and that separation rates should be the focus of the analysis.

The National Longitudinal Survey data is especially useful for studying the relationship between wages and the probability of quitting. Most theoretical work on the determinants of job separation concludes that the probability of changing jobs is related to a reservation wage. The NLS data set allows us to test this relationship since it includes information on the individual's "hypothetical wage"--that is, the wage required to induce the individual to accept another job. Given this information, we are able to compare the effects of different measures of the individual's price of time (e.g. the current wage and the reservation wage) on the probability of quitting. In addition, we analyze the role of human capital variables, job related characteristics and family background in the determination of job mobility.

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[3]For a discussion of this hypothesis in terms of the marriage market, see Becker, Landes, and Michael (1976).

The analysis of the determinants of job separations in the cross-section naturally leads to an investigation of the relationship between previous separations and future separations. In particular, we consider whether such a relationship exists, and whether the nature of previous separations is a good predictor of the nature of future separations.

Finally, we analyze the effects of job mobility on earnings and on job satisfaction. We distinguish between the immediate gains to mobility and the future gains to mobility, and also consider whether the nature of the separation is an important determinant of the consequences of job mobility.

Part II of the paper presents a brief framework for the analysis of job mobility. It will review various theories that have been presented in the literature to explain quit-layoff phenomena. In Part III, we conduct an empirical analysis of the determinants of job separations and consider whether the distinction between quits and layoffs is indeed appropriate. Part IV analyzes the effects of job mobility on wage growth both in the short- and long-runs. A summary of the empirical results is presented in Part V.

I. A Framework for the Analysis of Job Mobility

Economic theory predicts that an individual will attempt to sell his services in the market which offers him the highest return. This simple concept was first applied by Sjaastad (1962) to the analysis of labor mobility in his study of internal migration in the United States.[4] The

[4] See Polachek and Horvath (1976) for a more recent analysis of geographic mobility using individual, instead of aggregate, data.

individual is assumed to calculate his discounted net return from migrating at time t, and base his decision to move on whether the net return is positive. In the case of job mobility, the individual will engage in a similar calculation of the discounted net returns from leaving his current job. Hence for the i^{th} worker, the probability of quitting in his i^{th} year of job tenure, q_{ij}, will be positively related to the gains from job mobility. That is:

$$q_{ij} = f(\overline{Y}_{ij} - Y_{ij}) \tag{1}$$

where: $\overline{Y}_{ij}$ is the present value of the expected real income stream if the individual leaves his current job in job year j, Y_{ij} is the present value of the expected real income stream in the current job calculated at job year j.[5] Equation (1) suggests the following implications: (1) the higher the value of the current job, the less likely the individual is to quit his job; (2) the better the alternatives available to the individual relative to the current job, the more likely a quit will take place; (3) the longer the time remaining until retirement, the greater the gain from quitting since the returns to mobility can be collected over a longer period of time.

Of course, not all job separations need be initiated by the employee. At job year j, the employer will compare his estimate of the worker's marginal revenue product, MP_{ij}, to the wage he is currently paying the

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[5]Of course, equation (1) implicitly nets out the costs of mobility which may vary across individuals.

worker, w_{ij}. The employer will then lay off those workers whose marginal revenue products fall below the wage. Thus the probability of laying off the i^{th} worker after j years of job tenure, ℓ_{ij}, can be expressed as:[6]

$$\ell_{ij} = g(MP_{ij} - w_{ij}) \tag{2}$$

Clearly the higher the labor costs, the higher the probability of a layoff since ℓ_{ij} is negatively related to the difference between marginal revenue product and labor costs.[7]

We now turn to an analysis of the process by which the differentials in equations (1) and (2) lead to job separations.

A. The Matching Hypothesis

One approach to the study of a job mobility is to view turnover as the result of an imperfect match between employer and employee.[8] According to this hypothesis, the worker and the firm learn about each other

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[6]We can interpret MP_{ij} as the stream of marginal products received by the firm during the worker's tenure. Similarly, w_{ij} can be interpreted as the discounted stream of all labor costs (e.g. wages, fringe benefits, etc.) Note that the worker's perception of the value of the job, Y_{ij}, includes w_{ij} and any other job consumption benefits obtained by the worker in that job.

[7]Note that our analysis focuses on permanent layoffs. For a discussion of temporary layoffs, see Feldstein (1976).

[8]See Becker, Landes, and Michael (1976) for an application of the matching hypothesis to marital instability; Jovanovic (1976) develops a model of job matching in the labor market.

during the first few years of the worker's tenure. If they determine that they have been imperfectly matched with one another (i.e. the worker is not suitable for the job), a separation will occur. Thus this approach predicts a negative relationship between job tenure and separations. In the matching model developed by Jovanovic (1976), the employer adjusts wages to the worker's productivity as he learns about the latter over time. If the worker's productivity (and hence the wage) falls below the level which is attainable in other firms, the worker will quit. Since those employees that remain at the firm will be those whose productivity (and hence wages) are high, there will be a positive correlation between wage levels and job tenure. Thus the matching hypothesis predicts a negative relationship between wages and separations, not holding job tenure constant. Once tenure is held constant, we would expect this negative relationship to be weaker, since for a given tenure level there will be a smaller variation in productivity across individuals within the firm.

Note that since the wage is assumed to be adjusted to equal the worker's productivity, layoffs will never occur in this model. The fact that a separation might be officially recorded as a "layoff" is basically a matter of semantics and might in fact be due to the reluctance to cut wages below the going entrance wage or to eligibility constraints in unemployment compensation programs. Therefore an important conclusion of the matching hypothesis is that there is no useful distinction between a quit and a layoff. Finally, it is important to note that because learning about the job is likely to take a relatively short period of time, the matching hypothesis is most relevant for understanding turnover in the early years of job tenure.

B. The Specific Training Hypothesis

Alternative models of job separation have focused on the concept of fixed costs of employment.[9] It is argued that when a firm hires a worker it incurs certain fixed costs in the form of hiring and training costs. The immediate implication of the existence of these fixed costs is that workers with a higher degree of "fixity" are less likely to be laid off during periods of slack demand since the employer has an incentive to recoup his investment. This model, of course, can be generalized such that fixed costs are borne by both workers and employers (e.g. specific training), and therefore workers with a higher degree of fixity are less likely to quit.[10]

To illustrate the effects of the existence of specific training on labor turnover, consider:

$$q_i = \alpha_0 - \alpha_1 S_w \tag{3}$$

$$\ell_i = \beta_0 - \beta_1 S_f \tag{4}$$

where

S_w = worker financed specific training

S_f = firm financed specific training

Equations (3) and (4) state that the probability of quitting (layoff) is

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[9]See Oi (1962), Rosen (1968) and Becker (1975).

[10]In Parsons' (1972) model of specific training, implications are derived for both quit and layoff behavior.

a negative function of worker (firm) financed specific training. We can express S_w as a function of the wage and skills (e.g. education, E) by noting that an individual's wage can be defined as:

$$W = \gamma_0 + \gamma_1 E + \gamma_2 S_w \tag{5}$$

Equation (5) implies:

$$S_w = \frac{W}{\gamma 2} - \frac{\gamma_1}{\gamma_2} E - \frac{\gamma_0}{\gamma_2} \tag{6}$$

Substituting (6) into (3) yields:

$$q_i = (\alpha_0 - \frac{\alpha_1 \gamma_0}{\gamma_2}) - \frac{\alpha_1}{\gamma_2} W + \frac{\alpha_1 \gamma_1}{\gamma_2} E \tag{7}$$

Therefore the specific training hypothesis predicts that the probability of quitting is inversely related to the wage rate and positively related to education (i.e., skills). It can be seen that if S_w and S_F are positively correlated, the wage and education coefficients will be qualitatively similar in the quit and layoff equations. However if S_w and S_F are negatively correlated, then wages will have a positive effect while education will have a negative effect on layoffs. One would expect to observe a positive correlation between S_w and S_F, however, since substantial specific training investment is likely to take place only after the matching process has been completed.

C. The Job Search Hypothesis

Another approach to the analysis of job mobility is suggested by the existence of imperfect information in the labor market.[11] It has been shown that imperfect information on the part of employers and employees creates a non-degenerate distribution of wage offers for given skills.[12] This wage dispersion is likely to affect the labor turnover decisions of both individuals and firms.

We can easily derive the implications of the existence of wage dispersion on the probability of quitting by considering equation (1). Clearly the higher the current wage relative to alternatives the lower the probability of quitting. In the case of layoffs, the effect of the wage is not as clear: The worker may be receiving a higher relative wage in this firm than elsewhere since his productivity in this particular firm may also be relatively higher. The effect is nil if wages are adjusted to productivity, but positive if discrepancies between wages and productivity are larger at higher wage levels and are not eliminated.

D. Compensating Differentials

It has been argued that the relationship between wages and separation rates is another example of the theory of compensating wage differentials.[13] Workers who are employed in industries that have high layoff

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[11]For basic models of job search see McCall (1970), Mortensen (1970) and the survey article by Lippman and McCall (1976).

[12]For a derivation of this distribution see Mortensen (1974).

[13]For example, see Hall (1970). Of course, this argument dates back to Adam Smith who specifically mentioned job stability as a determinant of wage differentials across types of jobs.

rates will demand a wage premium to compensate them for the job instability. Thus we would expect to find a positive correlation between wages and the probability of a layoff.[14]

E. Summary

We have discussed four hypotheses that examine the relationships between wages and quit and layoff rates. As an aid to the reader we summarize these predictions in Table 1 below.

TABLE 1
Predicted Signs of the Wage Coefficient

Hypothesis	Quit Rate	Layoff Rate
Matching	-	-
Specific training	-	-
Job search	-	?
Compensating differentials	?	+

Note that these predictions have been made under the assumption that job tenure is not held constant in the equation. If job tenure were held

[14] Pencavel (1972) argues that each employer selects a particular wage-quit strategy such that the lower the quit rate an employer is willing to tolerate the higher the wage rate he must pay, thus resulting in a negative correlation between the probability of quitting and the wage rate. The problem with this prediction is that those industries which carry out this policy might be precisely those industries with high quit rates.

constant, we have shown earlier that the negative effects of the wage on the quit and layoff rates would be weaker or possibly vanish according to the matching hypothesis. For the other hypotheses, however, the predictions of signs are invariant with respect to job tenure.

II. The Determinants of Job Mobility

In this section we present an empirical analysis of the determinants of job mobility. The data set used is the National Longitudinal Survey of

Mature Men (aged 45-59 in 1966) which provides continuous information on work and earnings histories between 1966 and 1971.[15] To simplify the empirical analysis, we restrict our sample to consist of all white men who reported a wage rate in 1966, who reported key variables such as education and job mobility patterns, and who were still in the labor force in 1971. We also limit our sample to individuals who experienced no geographic mobility during the period in order to focus on job mobility within the local labor market. We avoid the problem of individuals being recalled from a layoff by defining job mobility as a change in employers. Thus recalls would not be measured as job separations. Moreover, by deleting migrants and retirees from the sample of job changers, we further underestimate the true separation rate. Any additional restrictions on the sample will be discussed as the results are presented.

Table 2 shows the mean probabilities of job separation for our sample. It can be seen that despite the age range of the sample we observe a high degree of job separation. In fact, within a five-year period, 22 percent of the sample changed employers. Note, however, that about half of these separations were employer-initiated--i.e. layoffs or firings. This leaves a remarkable amount of quits considering the average age of the sample. When we segment the five-year period into shorter time spans, 1966-69 and 1969-71, we see clear evidence of the effect of a downturn in the business cycle on the type of job separations that occur. For example, in the 1966-69 period, 57 percent of all job separations were quits, while in

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[15] See U.S. Department of Labor (1970), Manpower Research Monograph No. 15 for a complete discussion of the survey.

TABLE 2

Probabilities of Job Separation by Education

Variable	s ≤ 8	9 ≤ s ≤ 11	s = 12	s ≥ 13	All Men
1966-69 Separation rates	.212	.167	.165	.129	.175
1969-71 Separation rates	.115	.095	.100	.102	.105
1966-71 Separation rates	.236	.212	.222	.180	.217
Percent of 1966-69 separations that are quits	.514	.551	.618	.674	.570
Percent of 1969-71 separations that are quits	.356	.316	.490	.441	.393
Percent of 1966-71 separations that are quits	.530	.542	.586	.617	.558
Percent of 1966-69 quits that are job-related	.669	.598	.618	.621	.635
Percent of 1969-71 quits that are job-related	.683	1.000	.490	.533	.649
Percent of 1966-71 quits that are job-related	.640	.565	.592	.567	.599
Sample size	641	400	491	333	1,865

1969-71, the statistic was 39 percent. Although job separations decline with educational attainment for this sample, it is interesting to note that the proportion of job separations that are quits <u>increases</u> with education. This pattern, however, is not as clear during the economic downturn of 1969-71. In Table 2, we also find that about 60 percent of all quits are due to "job-related" reasons. An individual is assumed to have quit for job-related reasons if his quit is due to: (1) dissatisfaction with wages, hours, working conditions, type of work, (2) difficulty in getting along with employer and/or fellow employees, or (3) finding a better job. An individual leaves for personal reasons if his quit is due to (1) dislike of location or community, or (2) health or family reasons.

Our analysis will be conducted in two steps. Although in the previous section we argued that the matching process would only be relevant in explaining turnover in jobs of short duration, we will initially focus on mobility from all jobs using the traditional dichotomy of quits and layoffs, as well as a more refined definition of quits in terms of job-related and personal quits. We will also analyze the determinants of job separations without distinguishing between quits and layoffs. Finally, we differentiate between short and long jobs and discuss in detail the relevance of the matching process.

Our discussion in Part II showed that we expect to observe a negative correlation between the wage and the probability of quitting and an ambiguous correlation between the wage and the probability of a layoff, ceteris paribus. Recall, however, that the quit rate was affected by the real wage [see equation (1)]. This suggests the use of a measure of the price of time which captures the value of the job to the individual. The NLS provides us with such a measure in terms of the reservation wage--i.e., the wage that would induce the individual to leave his current job. In the case of layoffs, the firm makes its decision based on labor costs [see equation (2)] which are better measured by the actual wage. Moreover, there might exist personal, job and/or family characteristics which affect the differentials given in both equations (1) and (2) and hence are likely to affect the quit and layoff rates. These characteristics will be included in our empirical formulation of the quit and layoff equations.[16]

A. The Determinants of Quits

Table 3 presents the estimated quit regressions using the reservation wage as the wage variable. Since the dependent variable is defined as being equal to unity if the individual quit his job and zero otherwise,

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[16]The probability of quitting can be written as:

$$q = \int_{\hat{w}}^{\infty} f(w)\, dw$$

where $\hat{w}$ is the reservation wage and $f(w)$ is the wage offer distribution facing individuals of given skills. In principle, for a given wage offer distribution the quit rate would be exactly determined by the reservation wage. Since our measure of the reservation wage is correlated with $f(w)$ across individuals of varying skills, the variables measuring human capital also serve to standardize for the wage offer distribution.

TABLE 3*
Maximum-Likelihood Logit Regressions on the Probability of Quitting Between 1966 and 1969, Using the Reservation Wage

	3.1 All Quits		3.2 Quit for Job-Related Reason		3.3** Quit Because Found Better Job		3.4 Quit for Personal Reason	
	b	t	b	t	b	t	b	t
$\hat{w}$	-.0173	(-3.17)	-.0148	(-2.93)	-.0071	(-2.32)	-.0052	(-1.62)
NOTTAKE	-.0896	(-4.43)	-.0620	(-3.65)	-.0459	(-3.31)	-.0369	(-2.65)
STEADY	.0242	(.79)	-.0005	(-.02)	.0118	(.76)	.0261	(1.42)
ACCEPT	.0147	(.23)	-.0506	(-.73)	-	-	.0394	(1.20)
OTHER	-.0623	(-1.34)	-.0733	(-1.53)	-	-	-.0050	(-.20)
LIKE	-.0646	(-2.73)	-.0422	(-2.15)	.0133	(.70)	-.0296	(-1.92)
PENS	-.0565	(-3.36)	-.0470	(-3.28)	-.0265	(-2.66)	-.0155	(-1.39)
PUBLIC	-.0685	(-2.03)	-.0810	(-2.34)	-.0235	(-1.24)	-.0048	(-.25)
DEVP	.0006	(.54)	.0006	(.58)	.0005	(.74)	.00003	(.04)
DEVN	.0004	(.13)	.00005	(.02)	.0010	(.62)	.00005	(.03)
WKS	-.0012	(-.53)	-.0009	(-.51)	-.0041	(-1.83)	-.0005	(-.37)
SPELLS	.0803	(4.04)	.0514	(3.14)	.0381	(3.31)	.0381	(3.21)
EDUC	.0036	(1.30)	.0030	(1.29)	.0024	(1.54)	.0009	(.53)
REM	.0023	(1.40)	.0029	(1.49)	.0016	(1.65)	.0002	(.22)
HLTH	-.0156	(-.82)	-.0071	(-.44)	-.0009	(-.07)	-.0098	(-.79)
LIQ	.0001	(.19)	.0007	(-.67)	-.0001	(-.53)	.0003	(1.16)
OWN	-.0203	(-1.06)	-.0131	(-.81)	-.0030	(-.26)	-.0110	(-.90)
RES	-.0002	(-.59)	-.0002	(-.53)	.0001	(.40)	-.0001	(-.34)
MAR	-.0113	(-.37)	-.0043	(-.16)	.0043	(.21)	-.0081	(-.42)
WLFP	.0334	(1.48)	.0258	(1.38)	.0149	(1.22)	.0127	(.81)
WW	.0006	(.07)	.0040	(.60)	.0031	(-1.83)	-.0040	(-.59)
N	1724		1654		1588		1608	
X^2	118.33		96.207		52.843		42.523	

**The nonlinear constraints in the logit procedure resulted in the deletion of ACCEPT and OTHER.

*Key:

$\hat{w}$ = the reservation wage rate as of 1966

w_o = the actual wage rate as of 1966

NOTTAKE = 1 if individual would not accept a new job at any wage

STEADY = 1 if individual would accept a steady job

ACCEPT = 1 if individual would accept a job at an unknown wage

OTHER = 1 if individual gave any other response to the reservation wage question

LIKE = 1 if individual liked his job "very much" or "fairly well"

PENS = 1 if private pension plan existed at the firm

PUBLIC = 1 if individual was employed by the government

DEVP = difference between usual hours of work in the current job and mean hours of work if the difference is positive

DEVN = absolute value of this difference if it is negative

WKS = number of weeks unemployed in 1965-66

SPELLS = number of spells of unemployment in 1965-66

EDUC = years of schooling

REM = remaining years of work experience

HLTH = 1 if individual was in good health

LIQ = liquid assets in thousands of dollars

OWN = 1 if individual owned a house

RES = years living in the current residence as of 1966

MAR = 1 if individual was married

WLFP = 1 if individual's wife was employed

WW = wife's wage rate

the estimation technique utilized is maximum likelihood logit.[17] It is important to utilize this technique since ordinary least squares does not take account of the restriction that the probability of quitting should lie in the [0, 1] interval. The logit method of estimation assumes that the probability that the i^{th} individual quits his job is given by the logistic function:

$$q_i = 1/[1 + e^{-\beta x}] \qquad (8)$$

where x is a vector of independent variables, such as the wage, job characteristics and human capital variables.

The logit coefficient, β_j, shows the percentage change in the odds of quitting for a one unit change in x_j. The marginal effect of x_j on q_i is given by:

$$\frac{\partial q_i}{\partial x_j} = \beta_j \, q_i \, (1 - q_i) \qquad (9)$$

These marginal effects, evaluated at the mean, are the logit coefficients presented in the tables.

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[17] For a theoretical discussion of the problems encountered in estimating equations with dichotomous dependent variables see Nerlove and Press (1973).

The regressions in Table 3 examine the determinants of the probability of quitting between 1966 and 1969.[18] Some regressions on the probability of quitting in the five-year period, 1966-71, are presented below.

1. The Reservation Wage

Our measure of the reservation wage is based on the question:[19]

> Q: Suppose someone in this area offered you a job in the same line of work you are in now. How much would the new job have to pay for you to be willing to take it?

Individuals responded to this question by either giving a numerical wage or by answering that: (a) they would not accept a job at any wage; (b) they would accept a steady job at the same or less pay; (c) they would accept a job but did not know at what wage; and (d) any other response. About half of our sample responded with an actual reservation wage. For those individuals who gave one of the above reasons, we set the reservation wage equal to the actual wage but at the same time we standardize with a set of dummies indicating the actual response. The dummies we use to correspond to the above answers are (a) NOTTAKE, (b) STEADY, (c) ACCEPT, and (d) OTHER.[20]

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[18] Note that the sample sizes in Table 3 are different in each column. This is because for each regression we defined the relevant sample as those individuals who did not change jobs plus those who changed for the particular reason under analysis. We use this method in order to answer the question of what determines a particular type of separation versus staying on the job.

[19] NLS 1966 Questionnaire, Question 29a.

[20] In effect what we are doing is to use the best available information (i.e. the actual wage) for those individuals who did not report a reservation wage. The dummies capture the fact that the true reservation wage was unavailable for this group of individuals. See Dagenais (1973) for a discussion of the econometric problems encountered with missing information.

In regression 3.1 we estimate the equation for all quits; that is, the dependent variable is coded as unity if the individual quit his job in 1966-69, and zero if he did not change jobs at all. We find that the effect of the reservation wage is negative and significant.[21] Its magnitude indicates that a one dollar increase in the amount required to induce the individual to change jobs decreases the probability of quitting by about 16 percent in this sample.[22] It is also interesting to note the effects of the dummies indicating the individual's response to the reservation wage question. For example, those individuals who responded that they would not accept a job at any wage (NOTTAKE) are 83 percent less likely to quit a job than individuals who gave a numerical reservation wage. Thus the qualitative response to the reservation wage question in 1966 was as important as the quantitative response in indicating which individuals were more likely to quit in the next three years.[23]

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[21]There are two reasons for our using the 1966 wage rate even though we examine mobility during the subsequent three years: (1) it is important to have a base period in order to assign a wage to those individuals who did not change jobs; (2) in this age group the wage rate at any point in time should be a stable measure of the individual's stock of human capital.

[22]This number is calculated by dividing the coefficient on $\hat{w}$ by the mean probability of quitting which is .11.

[23]Note that since the effect of NOTTAKE is five times the effect of a one dollar increase in the wage, individuals who responded that they would not take a new job at any wage were, in effect, indicating they would require a five dollar wage increase to change jobs. Since the mean wage is under four dollars, this group of individuals requires more than a doubling of the wage in order to change jobs.

We can extend the empirical analysis by noting that the NLS provides detailed information on the reasons for quitting. We segment the sample into two major categories: quitting for job-related reasons and quitting for personal reasons. One could argue that the reservation wage should have a weaker effect in the case of personal (or exogenous) quits since when unexpected personal problems arise, the "value" of the job, as measured by the reservation wage, becomes a less critical factor in the individual's decision to quit. This is, in fact, what the results in Table 3 indicate. In regression 3.2, we find that a one dollar increase in the reservation wage significantly lowers the probability of a job-related quit by 21 percent. On the other hand, in regression 3.4, we see that a one-dollar increase in the reservation wage lowers the probability of an "exogenous" quit by only 12 percent.

We can isolate from the men who quit for job-related reasons a small group of individuals who quit because they found a better job--that is, men who were "pulled" from the current job by a better job offer. It would appear that the reservation wage should have a strong negative effect on the probability of finding a better job, since these individuals most closely resemble the typical decision maker in search models. The results in Table 3 confirm this expectation. From regression 3.3, we can calculate that a one-dollar increase in the reservation wage makes the individual 23 percent less likely to find a better job.

In order to make our results comparable with those from other data sets, and because the reservation wage is defined for only half our sample, we estimated the logit regressions using the actual wage as the relevant measure of the price of time. The estimated equations are shown in Table 4. The results are qualitatively similar to those obtained using the reservation wage: A one-dollar increase in the wage reduces the probability of a job-related quit by 26 percent, but reduces the probability of an exogenous quit by only 6 percent. The similarity between the two sets of coefficients should not be surprising. The basic difference between $\hat{w}$ and w_o is that the reservation wage incorporates the value of nonpecuniary aspects of the job. Since our vector of standardizing variables includes a measure of job satisfaction, we are in a sense holding constant these nonpecuniary differences; in fact, we may be "double-counting" variations in these differences thereby weakening the effect of $\hat{w}$ on the quit rate.

2. Job Characteristics

A vector of variables describing the characteristics of the individual's current (1966) job is included in the regressions in Tables 3 and 4. We find that individuals who liked their jobs (as measured by LIKE) were 60 percent less likely to quit in the next three years. It is important to note that the probability of being "pulled" from the job is not affected by the individual's level of job satisfaction; LIKE has an insignificant effect in equation 3.3. This result is intuitive since an individual may like his current job very much but if a better job offer is found he will accept it.

It has been argued that the existence of private pension systems inhibits job mobility.[24] While our results in equation 3.1 strongly support this hypothesis (that is, the probability of quitting is inversely related

[24] See Parnes and Nestel (1975) for empirical evidence of this hypothesis.

TABLE 4
Maximum Likelihood Logit Regressions on the Probability of Quitting Between 1966 and 1969 Using the Actual Wage

	4.1 All Quits		4.2 Quit for Job-Related Reason		4.3* Quit Because Found Better Job		4.4 Quit for Personal Reason	
	b	t	b	t	b	t	b	t
w_o	-.0160	(-2.33)	-.0183	(-2.74)	-.0128	(-2.70)	-.0025	(-.69)
NOTTAKE	-.0754	(-3.82)	-.0505	(-3.05)	-.0407	(-2.97)	-.0323	(-2.36)
STEADY	.0382	(1.26)	.0110	(.41)	.0172	(1.10)	.0305	(1.68)
ACCEPT	.0286	(.45)	-.0409	(-.59)	-	-	.0452	(1.38)
OTHER	-.0484	(-1.05)	-.0619	(-1.29)	-	-	-.0008	(-.03)
LIKE	-.0674	(-2.85)	-.0445	(-2.27)	.0142	(.74)	-.0310	(-2.00)
PENS	-.0569	(-3.35)	-.0447	(-3.09)	-.0229	(-2.27)	-.0163	(-1.45)
PUBLIC	-.0676	(-2.00)	-.0784	(-2.26)	-.0196	(-1.03)	-.0050	(-.26)
DEVP	.0007	(.61)	.0005	(.47)	.0003	(.39)	.0002	(.20)
DEVN	-.0001	(-.05)	-.0003	(-.13)	.0009	(.59)	-.0003	(-.19)
WKS	-.0008	(-.39)	-.0008	(-.45)	-.0040	(-1.87)	-.0003	(-.20)
SPELLS	.0774	(3.90)	.0499	(3.04)	.0369	(3.24)	.0358	(3.06)
EDUC	.0027	(1.00)	.0029	(1.28)	.0029	(1.83)	.0003	(.14)
REM	.0022	(1.30)	.0019	(1.40)	.0015	(1.53)	.0002	(.20)
HLTH	-.0175	(-.92)	-.0071	(-.44)	.0001	(.01)	-.0109	(-.87)
LIQ	.00004	(.97)	-.0007	(-.69)	-.0001	(-.53)	.0002	(.92)
OWN	-.0218	(-1.13)	-.0132	(-.82)	-.0009	(-.08)	-.0113	(-.88)
RES	-.0002	(-.54)	-.0002	(-.50)	.0001	(.39)	-.0001	(-.26)
MAR	-.0129	(-.42)	-.0026	(-.10)	.0074	(.35)	-.0072	(-.37)
WLFP	.0358	(1.59)	.0254	(1.37)	.0125	(1.02)	.0153	(.95)
WW	-.0006	(-.08)	.0035	(.55)	.0035	(.89)	-.0046	(-.69)
N	1724		1654		1588		1608	
x^2	112.37		94.33		55.29		40.01	

*The nonlinear constraints in the logit procedure resulted in the deletion of ACCEPT and OTHER.

to the existence of a private pension plan in the firm), it is interesting to consider the differential effects of pensions (PENS) on the probability of quitting for job-related versus personal reasons. From equation 3.2 we find that individuals who had a private pension plan were 67 percent less likely to quit for job-related reasons; yet in equation 3.4 the existence of a pension plan reduces the probability of exogenous quits by only 36 percent and is statistically insignificant.[25] It should be noted that the negative effect of PENS might be the result of a simultaneous relationship between the existence of private pension plans and the probability of quitting. If the availability of a private pension plan is dependent on job tenure and if future separation rates are correlated with job tenure, then the pension coefficient could reflect the influence of job tenure on the existence of a private pension plan. As will be seen below, once job tenure is introduced into the equation, the effect of PENS on the probability of quitting is, in fact, diminished.

We also have evidence that institutional factors have strong effects on job mobility. For example, we find that individuals in the public sector are 69 percent less likely to quit their jobs. This result suggests that either public employment inhibits job mobility or that individuals who prefer job stability choose public sector jobs. We can also analyze the effect of union membership on the probability of quitting. Since the NLS does not provide a measure of union

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[25] A more complete study of the effects of pension plans on separation rates would take into account the type of vesting provisions in the plan. Unfortunately, the NLS data do not provide this information.

membership until the 1969 survey, we can only analyze its effect on the probability of quitting between 1969 and 1971. We find that an individual who was a member of a union in 1969 was 75 percent less likely to quit his job during the next two years, not holding job tenure constant.[26,27]

Finally, we also included two variables to measure the extent of unemployment that the individual has undergone during the past year: WKS, the number of weeks unemployed, and SPELLS, the number of spells of unemployment. Generally, we find that WKS has an insignificant effect on the probability of quitting with one important exception: The probability of being pulled from the job by getting a better job offer is inversely related to WKS, holding SPELLS constant. This finding could be interpreted as evidence that people who have long periods of unemployment might be viewed as undesirable job applicants by firms, thus lowering their probability of being pulled from the job. In fact, an additional week of unemployment leads to a decrease of 13 percent in the probability of finding a better job offer. On the other hand, SPELLS has a strong positive effect on the probability of quitting for all groups. On the average, an additional spell of unemployment, holding WKS constant, roughly doubles the probability of quitting

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[26] For the sake of brevity, these regressions are not given in the tables.

[27] There is a possibility that the negative effect of union membership on the quit rate is due to a simultaneous relationship in that people who have little job stability would have no incentive to join unions.

the current job. This effect is probably due to the fact that SPELLS is a proxy for job separations that occurred within the past year prior to entering the current job. Thus it indicates that mobility is most likely at early stages of tenure.

This effect is even more strongly observed through the use of the variable TENURE (current job tenure as of 1966) in Table 5. This table contains ordinary least squares regressions using the reservation wage as the wage variable; the maximum likelihood logit program would not converge in the estimation of these regressions.[28] The effect of TENURE on the probability of quitting is strongly negative for all samples, although for the sake of brevity we only show the equation for all quits. For example, from equation 5.1 we obtain the fact that an additional year at the current job reduces the probability of quitting by 15 percent. This result can be explained through the use of the specific training hypothesis. That is, there is a positive correlation between the volume of specific training and job duration thus inhibiting individuals with longer job tenure, *ceteris paribus*, from quitting. As explained in Part II, an alternative hypothesis is that individuals and employers view the first few years of a job as a trial "match." If either the employer or employee find the match incompatible then job separation will occur. According to this hypothesis, once this initial trial period has elapsed, we would expect mobility not to be

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[28] To enable the reader to properly interpret the results in Table 5, Appendix Tables A-1 and A-2 present OLS regressions replicating Tables 3 and 4 in the text. Appendix Tables A-3 and A-4 present OLS regressions that include job tenure and use both the reservation wage and the actual wage for all samples.

TABLE 5
Effect of Job Tenure on the Probability of Quitting
Between 1966 and 1969*

	5.1 All Quits		Length of Current Job in Years	5.2 All Quits**	
	b	t		b	t
$\hat{w}$	-.0032	(-1.14)	0 - 2	.4116	(12.34)
NOTTAKE	-.0372	(-2.36)	3 - 5	.0802	(2.31)
STEADY	.0316	(.91)	6 - 8	.0501	(1.44)
ACCEPT	-.0059	(-.08)	9 - 11	.0462	(1.28)
OTHER	-.0568	(-1.60)	12 - 14	-.0212	(-.60)
LIKE	-.0862	(-3.24)	15 - 17	-.0094	(-.27)
PENS	-.0300	(-2.00)	18 - 20	.0048	(.13)
PUBLIC	-.0645	(-2.61)	21 - 23	-.0162	(-.46)
DEVP	.0002	(.19)	24 - 26	-.0038	(-.10)
DEVN	-.0010	(-.38)	27 - 29	-.0048	(-.12)
WKS	-.0026	(-1.05)	30 - 32	-.0018	(-.04)
SPELLS	.0929	(3.82)	33+	Left out group	
EDUC	.0028	(1.16)			
REM	-.0020	(-1.31)			
HLTH	-.0154	(-.87)			
LIQ	.0003	(.67)			
OWN	-.0158	(-.85)			
RES	.0003	(.77)			
MAR	.0036	(.12)			
WLFP	.0362	(1.86)			
WW	-.0052	(-.82)			
TENURE	-.0089	(-12.48)			
R^2	.15			.26	

*These coefficients are obtained from OLS regressions.

**This regression includes all the variables that are in regression 5.1 except TENURE.

affected by job tenure. This implication can be tested by breaking up the TENURE variable into a set of dummies. We also do this in Table 5 where we only show the coefficients of the job tenure dummies, although all the variables shown in equation 5.1 were included in the regression. The results are striking. The probability of quitting the job within the first three years of job tenure is 378 percent higher than that of quitting after 33 years of job tenure. This percentage effect drops dramatically to 74 percent in the second three years of job tenure. Thereafter, no significant relative effects of job tenure on the probability of quitting are observed, although weak positive effects exist until eleven years of job tenure. These results support the "matching" view of mobility since the effect of tenure is much stronger in the early years.

It is important to note that the introduction of job tenure into the quit regression reduces the significance of several variables, for example, the wage rate which becomes insignificant and pension plans which become less significant. In our discussion of the effects of private pension plans, we had indicated that the pension plan variable could be simultaneously related to the quit rate through job tenure. The results in Table 5 confirm this hypothesis. Similarly, once job tenure is held constant, the wage effect is diminished since wage levels and job tenure are strongly and positively correlated.[29] The question arises, however, as to whether a correctly specified quit

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[29] For a detailed discussion of the relationship between job tenure and wage levels, see Borjas (1975).

function should include job tenure as an exogenous variable. It can be argued that the same process which determines the probability of quitting in 1966-69 also determines job tenure as of 1966, since tenure in the current job is the result of the process determining mobility in earlier periods. To the extent that the variables determining job mobility in the 1966-69 period also determined tenure in the current job as of 1966, it is not surprising that the coefficients on the other variables are affected significantly. In fact, it is worth noting that the R^2 obtained in estimating a regression of the quit rate on job tenure is only slightly smaller (.12) than the explanatory power obtained by including job tenure in addition to the personal, human capital, and job-related variables.

3. Personal Characteristics

The regressions in Table 3 also include a set of variables describing the individual's background, finances, marital status and other characteristics. Overall, these variables have little effect on the probability of quitting. For example, while education has a positive effect on quits, it is always insignificant. Similarly, time remaining in the labor force (defined as the expected age of retirement minus current age) has a weak positive effect on the probability of quitting. This is consistent with an investment view of job mobility. That is, the longer the time remaining in the labor force the larger the payoff to any investment in mobility; thus the more likely the individual is to quit his job. Moreover, REM has no effect on the probability of quitting for personal reasons since for these individuals quitting is not an investment decision, but the result of exogenous factors.

Finally, a dummy indicating the wife's participation in the labor force (WLFP) has the strongest effect of all the personal characteristics. Its positive effect can be interpreted as evidence of an intra-family substitution effect. That is, individuals whose wives have a close attachment to the labor force are more likely to have a weaker attachment to their jobs. It is interesting to note that WLFP is weakest for individuals who quit their jobs for personal (exogenous) reasons. This is consistent with the hypothesis that quitting for personal reasons is a response to an exogenous shift in the individual's opportunity set and cannot be readily explained by systematic shifts in economic variables.

4. Quits Between 1966 and 1971

Table 6 presents the equations estimating the determinants of the probability of quitting between 1966 and 1971. This modifies the previous empirical analysis by extending the period under study to include a downturn in the business cycle. By comparing Tables 3 and 6, it can be seen that with one exception the results for 1966-71 are identical to those for 1966-69. The exception is the estimated effect of education. Recall that EDUC had a weak effect on the probability of quitting in 1966-69; in Table 6, however, EDUC has a positive and significant effect. We had shown in Part II that education would have a positive effect on the probability of quitting because of the existence of specific training.

TABLE 6

Maximum Likelihood Logit Regressions on the Probability of Quitting Between 1966 and 1971

	6.1 All Quits		6.2 Quit for Job-Related Reason		6.3* Quit Because Found Better Job		6.4 Quit for Personal Reason	
	b	t	b	t	b	t	b	t
$\hat{w}$	-.0169	(-3.04)	-.0137	(-2.69)	-.0040	(-1.39)	-.0060	(-1.77)
NOTTAKE	-.0760	(-3.73)	-.0447	(-2.61)	-.0259	(-2.14)	-.0420	(-2.88)
STEADY	.0002	(.04)	-.0094	(-.30)	.0091	(.50)	.0094	(.43)
ACCEPT	-.0342	(-.42)	-.0527	(-.68)	-	-	.0043	(.09)
OTHER	-.0495	(-1.07)	-.0835	(-1.56)	-	-	.0052	(.22)
LIKE	-.0373	(-1.39)	-.0333	(-1.50)	.0015	(.09)	-.0124	(-.70)
PENS	-.0718	(-4.05)	-.0574	(-3.76)	-.0238	(-2.36)	-.0234	(-1.99)
PUBLIC	-.0828	(-2.38)	-.1045	(-2.72)	-.0269	(-1.35)	-.0077	(-.41)
DEVP	.0008	(.61)	.0009	(.90)	.0009	(1.59)	-.0004	(-.42)
DEVN	.0006	(.20)	.0002	(.07)	-.0002	(-.11)	.0003	(.13)
WKS	-.0007	(-.03)	-.0001	(-.03)	-.0024	(-.89)	.0000	(.00)
SPELLS	.0545	(2.50)	.0314	(1.68)	.0130	(.71)	.0262	(2.05)
EDUC	.0068	(2.36)	.0039	(1.64)	.0023	(1.45)	.0038	(1.99)
REM	.0026	(1.48)	.0036	(2.34)	.0014	(1.34)	-.0007	(-.62)
HLTH	-.0076	(-.38)	.0004	(.02)	-.0001	(-.01)	-.0080	(-.62)
LIQ	.0001	(.22)	-.0001	(-.10)	-.0001	(-.31)	.0002	(.70)
OWN	.0130	(-.64)	-.0075	(-.44)	.0004	(.04)	-.0063	(-.47)
RES	-.0004	(-.84)	-.0004	(-1.15)	.00004	(.16)	.00002	(.06)
MAR	-.0080	(-.24)	-.0202	(-.74)	-.0103	(-.54)	.0008	(.04)
WLFP	.0411	(1.75)	.0233	(1.20)	.0119	(.94)	.0217	(1.37)
WW	.0006	(.08)	.0032	(.46)	.0012	(.27)	-.0025	(-.40)
N	1654		1585		1510		1530	
X^2	95.313		80.822		29.850		36.136	

*The nonlinear constraints in the logit procedure resulted in the deletion of ACCEPT and OTHER.

B. The Determinants of Layoffs

Our analysis of mobility has concentrated on separations initiated by the individual. In this section we focus on separations initiated by the firm. The NLS data provide two categories of firm-initiated separations: layoffs and discharges. Since most of these separations are layoffs, we ignore the distinction between the two categories. Table 7 presents the set of layoff regressions both for 1966-69 and 1966-71.

1. The Wage

In studying worker-initiated mobility, we argued that the relevant wage variable underlying the individual's decision to quit was the reservation wage. Clearly when we analyze firm-initiated separations the relevant wage variable should be the actual wage, since the actual wage is more positively correlated than the reservation wage with the firm's labor costs. Note that the actual wage has a positive (but insignificant) effect on the probability of being laid-off. This finding is consistent with a specific-training hypothesis only if there is a negative correlation between firm-financed and worker-financed specific training. A more likely explanation is provided by the hypothesis discussed in Part II that workers are compensated for working in jobs that have high layoff rates by receiving higher wages. We can test this hypothesis further by including a set of industry dummies in the regression in order to capture the inter-industry differences in layoff rates. Within an industry one would expect a weaker relationship between layoff rates and wages. In

TABLE 7
Maximum Likelihood Logit Regressions on the Probability of Being Laid Off

	7.1* 1966-69		7.2 1966-71	
	b	t	b	t
w_o	.0057	(1.38)	.0048	(1.05)
NOTTAKE	-.0427	(-2.45)	-.0343	(-1.74)
STEADY	.0497	(1.82)	.0747	(2.26)
ACCEPT	-.0163	(-.20)	.0746	(1.13)
OTHER	.0030	(.09)	.0328	(.84)
LIKE	-.0055	(-.21)	.0010	(.03)
PENS	-.0479	(-3.10)	-.0891	(-4.89)
PUBLIC	-	-	-.1680	(-3.41)
DEVP	.0023	(2.44)	.0024	(1.66)
DEVN	-.0017	(-.72)	.0010	(.39)
WKS	.0022	(1.49)	.0018	(.82)
SPELLS	.0710	(5.00)	.0683	(3.52)
EDUC	-.0076	(-3.10)	-.0020	(-.67)
REM	.0008	(.55)	.0027	(1.49)
HLTH	-.0021	(-.12)	.0621	(2.52)
LIQ	-.0037	(-2.21)	-.0032	(-2.17)
OWN	.0289	(1.43)	.0005	(.03)
RES	-.0001	(-.38)	.0001	(.32)
MAR	-.0222	(-.72)	-.0219	(-.63)
WLFP	.0014	(.08)	-.0094	(-.46)
WW	.0092	(2.48)	.0089	(1.79)
N	1679		1671	
X^2	125.53		113.53	

*The nonlinear constraints in the logit procedure resulted in the deletion of PUBLIC.

fact, the wage coefficient is reduced by about 50 percent once industry is held constant.[30]

2. Job Characteristics

While job satisfaction (LIKE) had a negative and significant effect on quits, it has no effect on the probability of being laid off. *A priori*, however, this result is not as obvious as it seems. That is, one could argue that individuals who like their jobs are better workers and are less likely to be laid off. Alternatively, if LIKE is positively correlated with fringe benefits and hence labor costs, the layoff rate would be positively related to LIKE. The observed coefficient then is the result of two opposing forces. It is also worthwhile to note the negative and significant effect of pension plans on layoff rates. This would be consistent with the hypothesis that individuals who have pension plans at their firms strive for better job performance in order to reduce the probability of being laid off.

The results in Table 7 show that unemployment in the last calendar year (as measured by WKS and SPELLS) is positively related to the probability of layoff. These variables, of course, could be proxies for some undesirable characteristics of the worker thus increasing the probability of involuntary turnover.

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[30]The only positive and significant industry coefficient was construction (the omitted industry was agriculture) which had a coefficient of .106 and a t-value of 2.26. It is interesting to note that the industry variables do not significantly affect the wage coefficient in the quit regression (for example, see Table 10).

Job tenure is introduced into the layoff equation in Table 8. Regression 8.1 shows that job tenure has a strong negative effect on the probability of layoff: An additional year of job tenure decreases the probability of layoff by 12 percent. Again this could be due to the existence of firm-financed specific training, although as can be seen in Table 8, the relative effect of job tenure becomes zero past the first three years of employment. This finding would tend to support the "matching" hypothesis discussed earlier. It is interesting to note the difference in the effects of the job tenure dummies between quits and layoffs. There is a much sharper decline in the effects of the tenure dummies in the layoff regression. This could be evidence of some degree of job seasonality. Due to the construction of the data set, termination of temporary jobs could not be distinguished from actual layoffs. Thus it could be that the strong effect of short job tenure on the probability of being laid off is merely evidence of a large proportion of "layoffs" that are actually seasonal jobs.[31]

We also found that union membership had a weak positive effect on the probability of being laid off between 1969 and 1971, not holding job tenure constant. This could occur because in unionized firms the employer would have less flexibility in reducing wages. Therefore, the only alternative open to him in the face of a reduction in product demand might be to cut employment.[32]

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[31]Appendix Table A-5 presents OLS regressions on the layoff rate in 1966-69 with and without job tenure.

[32]For the sake of brevity, these results are not presented in the tables.

TABLE 8
Effect of Job Tenure on the Probability of Being Laid Off Between 1966 and 1969

	8.1 Layoffs*		Length of Current Job in Years	8.2 Layoffs**	
	b	t		b	t
W_o	.0086	(1.99)	0 - 2	.3165	(10.70)
NOTTAKE	-.0168	(-.91)	3 - 5	-.0057	(-.18)
STEADY	.0437	(1.49)	6 - 8	-.0403	(-1.30)
ACCEPT	-.0291	(-.35)	9 - 11	.0027	(.08)
OTHER	-.0064	(-.17)	12 - 14	-.0533	(-1.73)
LIKE	-.0235	(-.85)	15 - 17	-.0197	(-.64)
PENS	-.0228	(-1.40)	18 - 20	.0018	(.05)
PUBLIC	-	-	21 - 23	-.0281	(-.91)
DEVP	.0014	(1.38)	24 - 26	-.0457	(-1.41)
DEVN	-.0028	(-1.16)	27 - 29	-.0398	(-1.12)
WKS	-.0002	(-.10)	30 - 32	-.0321	(-.95)
SPELLS	.0673	(4.17)	33+	Left out group	
EDUC	-.0074	(-2.83)			
REM	-.0008	(-.51)			
HLTH	.0007	(.04)			
LIQ	-.0031	(-1.77)			
OWN	.0353	(1.74)			
RES	.0001	(.23)			
MAR	-.0127	(-.41)			
WLFP	.0065	(.37)			
WW	.0064	(1.63)			
TENURE	-.0098	(-8.48)			
N	1679				
X^2	231.05				

*Maximum likelihood logit regression. The nonlinear constraints in the logit estimation procedure resulted in the deletion of PUBLIC.

**Ordinary least squares. This regression includes all the variables that are in regression 8.1 except TENURE.

3. Personal Characteristics

One striking difference between the quit and layoff regressions is the effect of education. As discussed in the previous section, education had a positive and insignificant effect on the probability of quitting. The results in Table 7 show that education is an important determinant of layoff rates: an additional year of schooling reduces the probability of being laid off by 9 percent. This finding is consistent with a specific training hypothesis only if there is a negative correlation between firm-financed and worker-financed specific human capital. A more likely explanation is that jobs requiring more human capital are less unstable. Note that the effect of education on layoff rates is weaker in 1966-71. It appears that during the 1969-71 downturn in the business cycle the more educated men were more likely to be laid-off than during the 1966-69 period, thus more closely resembling the less educated individuals.[33] Finally, note that the positive effect of the wife's wage is consistent with our earlier evidence of an intra-family substitution effect.

C. The Determinants of Separations

In this section we analyze the determinants of job separations without distinguishing between quits and layoffs. It will be recalled that the matching hypothesis discussed in Part II predicted that there is no

[33]It is important to note that the 1969-71 recession was not typical of other business cycle downturns. Usually a downturn in the cycle increases the differential between the layoff rates of the less educated and highly educated workers.

useful distinction between quits and layoffs in studying the determinants of job mobility. Table 9 presents the regressions estimating the determinants of the separation rate. An interesting conclusion to be drawn from Table 9 is that the explanatory power of the independent variables in the separation rate equation (whether measured in terms of X^2 or R^2) is higher than the explanatory power of these variables in the separate quit and layoff equations.

The coefficients in equation 9.1 are mostly weighted averages of the coefficients observed in the separate quit and layoff equations reported in regressions 4.1 and 7.1. Note, however, that although the matching hypothesis predicts a negative correlation between separation rates and the wage rate when job tenure is not held constant, the results in Table 9 do not support this prediction.

D. Separations from Short and Long Jobs

Contrary to the matching hypothesis, our analyses in Sections A and B indicate that it is important to distinguish between quits and layoffs in analyzing the determinants of job separations. Moreover, the results in Section C only partially support the matching view of job mobility. However, we argued in Part II that the matching hypothesis is most relevant during the early years of job tenure. Therefore, in order to more accurately test the matching hypothesis, we divide our sample into individuals whose job tenure as of 1966 is less than or equal to three years (short jobs) and individuals whose tenure is longer than three years (long jobs). It is interesting to note that the probability of a job separation for individuals in short jobs is 53 percent, and 56 percent of these separations are quits. For longer jobs the respective statistics are 7 percent

TABLE 9

Determinants of the Probability of Separating Between 1966 and 1969

Variable	9.1* b	9.1* t	9.2** b	9.2** t	9.3** b	9.3** t
w_o	-.0052	(-.79)	-.0035	(-.75)	.0039	(.87)
NOTTAKE	-.1006	(-4.39)	-.0823	(-4.50)	-.0460	(-2.64)
STEADY	.0628	(1.70)	.0759	(1.88)	.0582	(1.53)
ACCEPT	.0104	(.12)	.0163	(.19)	-.0301	(-.38)
OTHER	-.0454	(-.91)	-.0420	(-.99)	-.0532	(-1.33)
LIKE	-.0714	(-2.28)	-.0816	(-2.57)	-.0828	(-2.76)
PENS	-.1015	(-5.00)	-.0966	(-5.43)	-.0550	(-3.23)
PUBLIC	-.1854	(-3.74)	-.1292	(-4.21)	-.1241	(-4.28)
DEVP	.0028	(2.10)	.0030	(2.22)	.0014	(1.10)
DEVN	-.0002	(-.06)	-.0002	(-.08)	-.0020	(-.72)
WKS	.0015	(.63)	.0029	(1.08)	-.0002	(-.07)
SPELLS	.1270	(5.47)	.1655	(6.53)	.1408	(5.86)
EDUC	-.0026	(-.80)	-.0034	(-1.19)	-.0022	(-.82)
REM	.0027	(1.35)	.0025	(1.39)	-.0024	(-1.40)
HLTH	-.0158	(-.68)	-.0145	(-.68)	-.0085	(-.42)
LIQ	-.0018	(-1.40)	-.0005	(-.98)	-.0002	(-.37)
OWN	-.0002	(-.01)	-.0010	(-.04)	.0076	(.37)
RES	-.0004	(-.88)	-.0004	(-.93)	.0003	(.66)
MAR	-.0325	(-.86)	-.0252	(-.72)	-.0036	(-.11)
WLFP	.0309	(1.34)	.0290	(1.35)	.0283	(1.39)
WW	.0089	(1.52)	.0075	(1.23)	.0024	(.41)
TENURE					-.0119	(-14.97)
R^2			.11		.21	
χ^2	198.99					
N	1865		1865		1865	

*Maximum likelihood logit.

**Ordinary least squares.

and 59 percent. To test the relevance of the matching model, we estimated separation, quit, and layoff regressions separately for short and long jobs. The wage coefficients from these regressions are given in Table 10.

Once we take account of the compensating differential effect on the probability of a layoff by holding industry constant, we find that the actual wage has no effect on the probability of separating, quitting, or being laid off from a short job. In Part II, we showed that, according to the matching hypothesis, there was no distinction between quits and layoffs, and that the wage rate would not be related to either when analyzing separations from jobs of short tenure. The results in Table 10 seem to indicate that the matching process is useful for understanding the determinants of separations from short jobs.

For longer jobs, wages are negative and significant in determining the probability of separating or quitting, and have a weak, negative effect on the probability of a layoff. Once we realize that the layoff coefficient may have a positive bias if we are not adequately controlling for compensatory differentials, these findings are consistent with a specific training hypothesis that assumes a positive correlation between worker-financed and firm-financed specific training investment.

It is important to note that the results in Table 10 are not affected by the introduction of job tenure into the regressions. The wage coefficients from these regressions are shown in Appendix Table A-6. In the case of long jobs, job tenure is often insignificant and the wage coefficient does not change. Indeed, in these regressions the $\overline{R}^2$ is not

TABLE 10

Effects of the Wage Rate on the Probability of Separating from Short and Long Jobs*

	Separations			Quits			Layoffs		
	Short Jobs	Long Jobs	All Jobs	Short Jobs	Long Jobs	All Jobs	Short Jobs	Long Jobs	All Jobs
				A. Not Holding Industry Constant					
$\hat{w}$	.0060	-.0056	-.0058	-.0039	-.0049	-.0076	.0147	-.0011	.0010
	(.64)	(-2.05)	(-1.68)	(-.40)	(-2.22)	(-2.59)	(1.57)	(-.62)	(.40)
$\bar{R}^2$	.051	.034	.105	.032	.013	.060	.068	.048	.070
w_o	.0132	-.0059	-.0035	-.0014	-.0064	-.0079	.0231	.0001	.0040
	(1.03)	(-1.58)	(-.75)	(-.10)	(-2.11)	(-1.97)	(1.74)	(.03)	(1.13)
$\bar{R}^2$	.053	.033	.104	.032	.012	.058	.070	.048	.094
				B. Holding Industry Constant**					
$\hat{w}$	.0069	-.0058	-.0063	-.0022	-.0046	-.0074	.0148	-.0015	.0004
	(.74)	(-2.11)	(-1.87)	(-.22)	(-2.08)	(-2.54)	(1.63)	(-.84)	(.14)
$\bar{R}^2$	.073	.048	.136	.029	.019	.074	.151	.058	.133
w_o	.0115	-.0063	-.0057	.0008	-.0060	-.0078	.0168	-.0006	.0017
	(.89)	(-1.67)	(-1.21)	(.05)	(-1.97)	(-1.94)	(1.30)	(-.22)	(.48)
$\bar{R}^2$	.073	.047	.135	.029	.018	.073	.148	.058	.133
Sample size	434	1431	1865	332	1392	1724	305	1374	1679

*The coefficients presented in this table are OLS coefficients.

**We use a set of dummies at the one-digit industry level.

substantially increased when job tenure is held constant. Therefore, within the set of long jobs, there appears to be little need to standardize for job tenure. In the case of short jobs, however, tenure has a significant negative effect and increases $\bar{R}^2$ substantially although the wage coefficients are still insignificant in the quit equation. This result, therefore, indicates that the matching process may take place in a very short span of time within the job.

The reader might wonder why the results in our earlier tables point out the need to differentiate between quits and layoffs even though most mobility occurs early in the job and the matching hypothesis explains this turnover. The answer relates to the fact that in the pre-retirement years, most individuals are in long jobs, and a regression pooling the two groups (short and long jobs) would be heavily weighted towards an analysis of the determinants of mobility from long jobs.

E. Serial Correlation in Job Mobility

In our discussion of job tenure, we have already found that individuals who have short current job duration are much more likely to separate from the job either through quit or layoff. This would suggest that there is a strong serial correlation in job mobility. We can further explore this relationship by examining the effects of mobility in 1966-69 on the probability of job separation in 1969-71.

In Table 11, we present selected coefficients from regressions estimating the probability of quitting and being laid-off in 1969-71. The results are quite striking. The probability of quitting the 1969 job in the 1969-71 period is 227 percent higher if a quit occurred in 1966-69,

TABLE 11*

The Effects of Previous Mobility on Mobility in 1969-71

	Probability of Quitting		Probability of Being Laid Off	
	b	t	b	t
Quit in 1966-69	.1003	(5.89)	.1036	(5.10)
Laid off in 1966-69	.1259	(6.06)	.2102	(8.96)
Sample size	1747		1788	

*All of these coefficients were obtained from ordinary least squares regressions.

while it is 286 percent higher if a layoff occurred in 1966-69. On the other hand, the probability of being laid off in 1966-71 is 157 percent higher if a quit occurred in the previous three years, while it is 318 percent if a layoff occurred in the 1966-69 period. We can summarize these results by noting that the probability of quitting does not depend on the nature of the previous job separation, while the probability of layoff is more strongly related to a previous layoff than to a previous quit. This result, of course, could be due to the fact that terminations of seasonal jobs are coded as layoffs in the NLS. More generally, the results in Table 11 show that certain individuals are chronic movers. It is well known that turnover rates decline with age, yet we find that within this group of middle-aged men there exists a subset of individuals who are moving continuously. Moreover, it is interesting that the nature of the earlier separation does not strongly determine the nature of the subsequent separation once we account for some degree of seasonality in the layoff group.

In summary, the evidence from Tables 10 and 11 suggest the importance of the matching process in the first few years of job tenure once we abstract from the effect of compensating differentials in jobs that have high layoff rates. As predicted by the matching hypothesis, we observe an insignificant wage coefficient in the regressions explaining quits and layoffs from short jobs. Moreover, the importance of the matching process <u>within</u> the job is highlighted by the fact that the nature of the previous separation does not determine the nature of the separation from the current (short) job.

III. The Consequences of Job Mobility

The previous sections have focused on an analysis of the determinants of job mobility. We now extend our study by concentrating on the effects of job mobility on the earnings profile. Previous work on the NLS by Borjas (1975) and Parnes and Nestel (1975) has found that quitters had wage growth over the survey period at least as great as the non-changers. It is important to note, however, that these studies examined the wage growth of these different groups contemporaneously with the job change. Thus their results captured a mix of both the immediate and future gains in wages from job mobility. We extend their work in several respects. First, we distinguish between the immediate gains from job mobility and the future gains from job mobility. We compare individuals who quit their jobs in 1966-67 (and did not change jobs thereafter) with individuals who stayed with the same employer throughout the five-year period. The relevant comparisons are the immediate wage gains (i.e. 1966-67) and future wage growth (1967-71). This enables us to answer the question of whether the effect of mobility on the earnings profile is a parallel upward shift. The second contribution of our analysis is to distinguish between job-related quits and quits due to exogenous reasons as well as between quits and layoffs. Of course, the *a priori* expectation is that if a quit is to pay at all, it should pay for those individuals who left their job because of a better job offer. Finally, we will briefly analyze the effects of the different types of job separation on job satisfaction. In interpreting the effects of job mobility, however, one should bear in mind that individuals who leave their job might leave precisely because they have lower wage growth. Moreover another problem arises in that we compare the wage growth in the early years of job tenure for movers with the wage growth in the later years of job tenure for stayers. Thus comparing the subsequent wage growth of movers to that of the stayers might yield a biased estimate of the true effect of job mobility.

A. Immediate Wage Gains

Columns 12.1 and 12.2 in Table 12 present the regressions explaining percentage wage growth in the 1966-67 period for individuals who either did not change jobs at all in 1966-71 or who changed jobs in 1966-67 only. By including a set of dummies indicating the nature of job mobility we are able to measure the immediate gains from mobility since each dummy variable gives us the percentage difference in wages between changers and non-changers.[34] Columns 12.3 and 12.4 present similar regressions explaining 1966-69 percentage wage growth. In these latter regressions, the sample includes individuals who either did not change jobs at all in 1966-71 or who changed jobs in 1966-69 only. The dummy variables indicating job mobility therefore capture the gains to mobility that occurred in 1966-69. Note that this definition of immediate wage gains entails a longer period, and is therefore less exact than the analysis of 1966-67 wage growth.

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[34] Note that the sample sizes in this section have declined slightly. This is due to our restricting the sample to men who reported their wage in periods subsequent to 1966. Also note that our use of percentage wage growth (and not absolute growth) is suggested by the human capital model of wage determination. For an exposition of the model see Mincer (1974).

TABLE 12*

Regressions on Immediate Wage Gains from Job Mobility

	12.1 1966-67		12.2 1966-67		12.3 1966-69		12.4 1966-69	
	b	t	b	t	b	t	b	t
LAYOFF	.1044	(2.12)	.0901	(1.58)	.0158	(.41)	.0420	(.98)
PERSONAL	.0730	(1.04)	.0417	(.55)	.0290	(.54)	.0098	(.16)
PULL	.1180	(1.46)	.0927	(1.10)	.1847	(3.17)	.1797	(3.01)
PUSH	-.1076	(-1.15)	-.1537	(-1.56)	.0143	(.27)	.0076	(.14)
EDUC			-.0066	(-2.68)			.0023	(.86)
REM			.0040	(2.33)			.0036	(1.94)
MAR2			.1809	(2.50)			.0490	(.88)
MAR1			-.2316	(-3.09)			-.0834	(-1.43)
TENURE			.0003	(.36)			-.0013	(-1.45)
WKS			.0005	(.18)			-.0000	(-.04)
SPELLS			-.0227	(-1.00)			-.0489	(-1.68)
DOCC			.0471	(1.39)			.0068	(.38)
DDUNC			.0007	(.53)			.0029	(4.28)
Sample size	1289		1289		1383		1383	
R^2	.01		.03		.01		.03	

*Key to additional variables:

LAYOFF	=	1 if individual was laid off in the relevant time period
PERSONAL	=	1 if individual quit because of personal reasons
PULL	=	1 if individual quit because he found a better job
PUSH	=	1 if individual quit because he was dissatisfied with his job
MAR2	=	1 if individual was married at the end of the time period
MAR1	=	1 if individual was married at the beginning of the time period
DOCC	=	1 if individual changed occupations during the relevant time period
DDUNC	=	change in Duncan Occupational Index that occurred during the relevant time period

In Table 12, four dummy variables are used to capture the effect of job mobility on wage growth: LAYOFF, indicating whether or not the individual was laid off from the 1966 job; PERSONAL, indicating if the individual quit his 1966 job for exogenous reasons; PULL, if the quit took place because he found a better job; and PUSH, indicating that the quit took place because of dissatisfaction with his 1966 job. The excluded group, of course, are those individuals who did not change employers.

Generally, the dummy variables indicating type of quit are insignificant except for the effect of PULL on 1966-69 wage growth. We find that individuals who quit the 1966 job because they found a better job have percentage wage gains that are 18 percent higher in 1966-69 (or 6 percent higher per year) than those who stayed. In the 1966-67 period, however, the effect of PULL is 12 percent per year. Moreover, in this shorter period the effect of PUSH is negative and slightly significant. Note that these results point out the need for distinguishing between different types of quits. That is, in order to correctly estimate the effects of job mobility on wage growth, one needs to know the motivating force behind the individual's decision to quit. This finding is even more strongly observed when we utilize a more detailed breakdown of quits as in Table 13. One striking result is the difference in the effects of the several categories which composed PUSH in Table 12: (a) individuals who quit due to dissatisfaction with wages; (b) individuals who quit due to dissatisfaction with working conditions; and (c) quitting due to interpersonal relations. It is worth noting

TABLE 13

Regressions on Immediate Wage Gains from Job Mobility
Using Detailed Reason of Quit

	13.1 1966-67		13.2 1966-69	
Change Was Due to:	b	t	b	t
Layoff	.1044	(2.13)	.0158	(.41)
Health	.1264	(.91)	.0694	(.71)
Disliked location	.0071	(.03)	.0913	(.78)
Disliked wages	.1385	(.99)	.1232	(1.58)
Disliked work	-.2226	(-1.60)	-.0850	(-1.02)
Interpersonal relations	-.6319	(-2.27)	-.0539	(-.39)
Found better job	.1180	(1.46)	.1847	(3.18)
Other reasons	.0648	(.73)	.0256	(.32)
Family problems			-.3589	(-1.16)
Sample size	1289		1383	
R^2	.01		.01	

that both quits due to dissatisfaction with work and quits due to interpersonal problems in the job have strong negative effects on the immediate wage gain, while quits due to dissatisfaction with wages have a positive effect which becomes significant in the 1966-69 period. Thus if an individual disliked a nonpecuniary aspect of his job, he is willing to trade away some of his wages. As will be seen below, the trade is "fair" since these individuals gain significantly in terms of job satisfaction. These findings show that to examine the effects of quits on wage growth it is important to distinguish between the different types of quits; otherwise, the net impact of quitting will be a conglomeration of many diverse effects.

The effect of LAYOFF on the immediate wage growth is interesting. We find that individuals who were laid off in 1966-67 had a significant increase in their wage growth as compared to the non-movers. Yet in 1966-69, LAYOFF during that period has no effect on contemporaneous wage growth. Thus we conclude that men who were involuntarily separated from the 1966 job do at least as well as those who did not change jobs in the 1966-71 period. This could be due to the fact that our sample of layoffs is restricted to "successful" searchers--that is, men who were laid off in 1966-67, but stayed with the new employer for the remaining four years.[35]

We also included a vector of personal and job characteristics in the regressions in Table 12. Overall, these variables are not good

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[35]If we do not restrict the sample to successful searchers, the effect of being laid off on the immediate wage gain becomes insignificant, though still positive. The coefficient is .0317 (t = .91).

predictors of an individual's wage growth.[36] The most stable results are the effects of time remaining in the labor force, REM, and changes in marital status. We find that REM has a positive and significant effect on wage growth. Theoretically, the effect of REM on wage growth is ambiguous. First, the longer the time remaining, the higher the payoff to on-the-job investment. Clearly, more investment would take place at younger ages, creating a positive relationship between REM and wage growth. On the other hand, it can be argued that the later in the life cycle the quit occurs, the more incentive there is to get as large an immediate wage gain as possible. This could be tested by examining the effect of REM on the wage growth of individuals who quit. However, the results still show a positive effect of REM on wage growth. Thus it seems that the investment hypothesis dominates. We also find that individuals who suffered a marital breakup during the period (MAR1 = 1 and MAR2 = 0), experience smaller wage growth during this period.

The 1969 NLS questionnaire provides additional information on the nature of the job change. In particular, it gives us data on whether the individual had a new job lined up prior to the separation. The data show that 47 percent of those who quit had a new job versus 12 percent of those who were laid off. Moreover, within the group of quitters, 63 percent of those who quit for job-related reasons had a new job versus 12 percent of those who quit for exogenous or personal reasons. The similarity between the latter group and the individuals who were

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[36] See Borjas and Mincer (1976) for an analysis of the determinants of individual wage growth.

laid off points out the exogenous nature of these quits. It is important to note that having a new job lined up has a strong positive effect on the 1969-71 wage gain of individuals who changed jobs during that period indicating the significance of on-the-job search. However, even when we hold having a new job constant, we still find that it is important to distinguish among types of separations.[37] Moreover interaction terms between having a job lined up and nature of the separation were generally insignificant.

Finally, we compared the effects of separating from a short job (tenure $\leq$ 3 years) and separating from a long job (tenure > 3 years). Recall that in explaining the determinants of separating from a short job we found that there was no meaningful distinction between quits and layoffs. The question arises as to whether in studying the consequences of separating from short jobs one should distinguish between quits and layoffs. We find that in comparing individuals who separated from short jobs with individuals who stayed in short jobs the results reported in Table 12 still hold, i.e., there is a meaningful distinction between quits and layoffs. In comparing individuals separating from

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[37] The regression explaining immediate wage growth between 1969 and 1971 are as follows:

$$\text{PC6971} = \underset{(-.95)}{-\,.0363}\ \text{LAYOFF} + \underset{(.67)}{.0371}\ \text{Job Related Quit} - \underset{(-2.75)}{.2346}\ \text{Personal Quit} \qquad R^2 = .006$$

$$\text{PC6971} = \underset{(-1.15)}{-\,.0440}\ \text{LAYOFF} - \underset{(-.37)}{.0228}\ \text{Job-Related Quit} - \underset{(-2.85)}{.2429}\ \text{Personal Quit} + \underset{(2.41)}{.0968}\ \text{Had} \qquad R^2 = .01$$

long jobs with individuals staying in long jobs, one important new result is obtained: being laid off from a long job has a significant negative effect on immediate wage growth. Quitting from a long job has the same effects as those reported in Table 12.

B. Future Wage Gains

We have already shown that job mobility creates discontinuous shifts in the individual's earnings profile. We now consider whether mobility in the pre-retirement years has any effect on the subsequent wage growth in the new job, i.e., on the slope of the earnings profile. Again, we consider two time periods: 1967-71 for individuals who changed in 1966-67, and 1969-71 for those who changed in 1966-69. The results are presented in Table 14.

The most striking result is the negative and significant effect of LAYOFF on future wage growth. That is, even though the immediate effect of a layoff on wage growth is positive, over the long run these individuals experience smaller wage growth than those men who stayed on the job. Generally, the effects of a quit on future wage growth are insignificant, except for the coefficient of finding a better job on 1969-71 wage growth which is negative. However, even for this group, the net gain of a quit over the five-year period, 1966-71, is positive. The fact that quitting in general has an insignificant effect on future wage growth suggests that the gain to voluntary mobility (at least for those who were "pulled" from the job) is one of an immediate wage gain rather than a continuing increase in wages. This result might be due to the age range of the sample. Clearly at older ages, the finiteness of life would imply little

TABLE 14

Job Mobility Effects on Future Wage Growth*

Change Was Due to:	14.1 b	14.1 t	14.2 b	14.2 t	14.3 b	14.3 t	14.4 b	14.4 t
	1966-67 Job Mobility				1966-69 Job Mobility			
Layoff	-.1349	(-2.15)	-.1349	(-2.15)	-.0986	(-2.33)	-.0986	(-2.33)
Personal reasons	.0066	(.07)			-.0088	(-.15)		
Found better job	-.0411	(-.40)	-.0411	(-.40)	-.1152	(-1.82)	-.1152	(-1.82)
Disliked job	.1290	(1.09)			-.0345	(-.60)		
Health			-.1253	(-.70)			-.0655	(-.61)
Disliked location			.0672	(.27)			.0223	(.17)
Disliked wages			.1467	(.82)			-.0168	(-.20)
Disliked work			.0900	(.51)			-.0673	(-.74)
Inter-personal relations			.2144	(.60)			.0003	(.00)
Other reasons			.0472	(.42)			.0217	(.25)
Family problems							-.2768	(-.82)
N	1289		1289		1383		1383	
R^2	.01		.01		.01		.01	

*"Future" wage growth is defined as:

Equations 14.1 and 14.2 = percentage wage growth in 1967-71.
Equations 14.3 and 14.4 = percentage wage growth in 1969-71.

on-the-job investment taking place. Thus these individuals undertake mobility not as a means of finding jobs which provide higher levels of job investment but as a method of obtaining an immediate increase in wages by shifting to higher, but parallel (to that of stayers), earnings profiles.[38]

C. Nonpecuniary Gains

Up to this point we have analyzed the effects of job mobility on wage growth. In this section, we explore its effects on job satisfaction. We defined an individual as "liking" his job if he indicated that he liked his job "very much" or "fairly well." Table 15 shows the percentage of individuals who liked their jobs in 1966 and 1969 by type of job separation during this period. The results are extremely interesting. About 93 percent of the individuals who stayed in the job in the 1966-69 period liked the job in 1966, while only 83 percent of those who quit their job in the next three years were satisfied with their 1966 job. It is remarkable that by 1969, the percentage of individuals who liked their jobs was 94 percent for both groups. In fact, most of the increase in job satisfaction for those individuals who quit is due to the increase attained by those individuals who were "pushed" out of the 1966 job—that is, those individuals who left the 1966 job because they were dissatisfied with a job-related characteristic such as wages, work, and interpersonal relations.

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[38] It is important, however, to note that our analysis was carried out in percentage terms and since the stayers have higher average wage levels than the quitters, those who remain in the job achieved larger absolute wage increases in the survey period.

TABLE 15

Changes in Job Satisfaction by Type of Mobility

Mobility During 1966-69:	Percent of Individuals Who Liked the 1966 Job	Percent of Individuals Who Liked the 1969 Job	Number of Observations
Stayers	93.2	93.5	1,219
Involuntary changers	88.1	94.0	67
Voluntary change due to:			
Any reason	82.5	93.8	97
Pushed	77.1	97.1	35
Pulled	89.7	96.6	29
Personal	83.3	77.7	18
Other	80.0	100.0	15

The results in Table 15 are quite important since they provide empirical evidence that an individual does not necessarily leave his job in order to achieve a money wage gain. In fact, for the groups that achieved a significant increase in job satisfaction we find insignificant money wage gains (see Table 12), while for the group that was "pulled" from the 1966 job and that achieved significant money wage gains, only a small increase in job satisfaction can be detected.

IV. Summary

This paper has analyzed the determinants and consequences of middle-age job mobility. Traditional analysis has distinguished between two types of separations: quits and layoffs. It can be argued by viewing the job as a marriage between employer and worker that this distinction has no empirical content and adds nothing to our understanding of the determinants of job separation. On the other hand, persons quitting their jobs for personal reasons may not have the same economic motivation as those who quit for job-related reasons. This latter argument would suggest an even more detailed breakdown among types of quits. By utilizing this latter breakdown of job separations we obtained several major empirical findings:

1. Theoretical models of job separation are couched in terms of a reservation wage. We took advantage of the fact that the NLS provides this information, and as expected we found that the probability of quitting for job-related reasons was significantly and negatively related to the reservation wage when job tenure was not held constant. The probability of quitting for personal reasons, however, was less

strongly related to the reservation wage since this type of quit is due to exogenous forces. Once job tenure was held constant, the effect of the wage on quit rates was diminished in all samples.

2. The availability of a pension plan had a strong negative effect on job-related quits but did not determine quitting for personal reasons. Similarly, personal characteristics such as time remaining in the labor force and the wife's labor force status had systematic effects on job-related quits and insignificant effects on exogenous quits. These findings were invariant to the inclusion of job tenure in the regression.

3. The probability of layoff was positively related to the individual's wage rate and this effect became stronger when job tenure was held constant. The positive wage effect can be explained by compensating differentials: individuals in jobs with a high degree of instability will demand higher wages. Indeed, when industry was held constant in the regression, the positive wage coefficient was diminished.

While the above results support the argument that the quit-layoff distinction as well as a more detailed breakdown of quits is meaningful, our analysis showed that the matching hypothesis has relevance as well:

4. We found strong evidence of serial correlation in job mobility. In particular, we observed that most separations occur during the first few years of job tenure. This result is evidence of a matching process between firm and worker that occurs in the early years of the job as both parties learn about each other.

5. The nature of previous job separations was not a strong determinant of the nature of future job separations. This finding conforms with the predictions of the matching hypothesis that the quit-layoff breakdown is uninformative.

6. Since it can be argued that the matching hypothesis is most relevant for short jobs, we separately analyzed the determinants of separations of short jobs versus long jobs. Once we took account of the compensating differential effect operating in the layoff equation, we found that the wage rate had no effect on the probability of separating, quitting, or being laid off from a short job. The distinction between quits and layoffs, however, remained in the analysis of long jobs.

Finally, we analyzed the consequences of job mobility during the preretirement years. Our analysis focused on the effects of job mobility on wage growth and job satisfaction:

7. We found significant evidence of the need to distinguish between types of quits. In particular, we observed that individuals who were pulled (i.e. found a better job) from their jobs had higher immediate wage gains than stayers, while individuals who were pushed (i.e. were dissatisfied with the current job) had smaller wage gains than stayers. We also found that in this age range, quitting did not affect the slope of the earnings profile in the new job.

8. Job mobility affected not only the individual's money wages, but also his degree of job satisfaction. For example, while individuals who quit because they were dissatisfied with their current job had negative or zero immediate wage gains (relative to the stayers), they experienced significant gains in job satisfaction. These individuals evidently quit not for wage gains, but for nonpecuniary aspects of the job.

The reader will recall that at the outset of this paper, we discussed several hypotheses which are useful in understanding the determinants and consequences of job mobility. The findings presented in this paper indicate that once we take account of compensatory differentials in jobs with high layoff rates, the matching view of job turnover is relevant for explaining separations from short jobs. In the case of long jobs, however, the evidence points to the relevance of specific training in explaining job turnover.

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Table A-1

OLS Regressions on the Probability of Quitting Between 1966 and 1969 using the Reservation Wage (without job tenure)

	All Quits		Quit for Job-Related Reason		Quit Because Found Better Job		Quit for Personal Reason	
	b	t	b	t	b	t	b	t
$\hat{w}$	-.0076	(-2.59)	-.0055	(-2.27)	-.0032	(-1.88)	-.0027	(-1.34)
NOTTAKE	-.0687	(-4.24)	-.0478	(-3.50)	-.0341	(-.10)	-.0277	(-2.47)
STEADY	.0400	(1.10)	.0060	(.19)	.0242	(.15)	.0446	(1.72)
ACCEPT	.0321	(.45)	-.0401	(-.63)	-.0551	(-.21)	.0703	(1.41)
OTHER	-.0537	(-1.45)	-.0536	(-1.71)	-.0452	(-.15)	-.0056	(-.22)
LIKE	-.0878	(-3.15)	-.0650	(-2.71)	.0122	(.13)	-.0405	(-2.02)
PENS	-.0592	(-3.82)	-.0514	(-3.92)	-.0263	(-.10)	-.0160	(-1.48)
PUBLIC	-.0633	(-2.45)	-.0631	(-2.89)	-.0225	(-.12)	-.0073	(-.41)
DEVP	.0014	(1.17)	.0013	(1.28)	.0008	(.03)	.0003	(.37)
DEVN	.0008	(.28)	.0004	(.16)	.0008	(.48)	.0005	(.24)
WKS	-.0002	(-.07)	.0002	(.09)	-.0036	(-2.19)	-.0004	(-.23)
SPELLS	.1127	(4.45)	.0731	(3.32)	.0496	(.13)	.0687	(3.69)
EDUC	.0018	(.74)	.0013	(.62)	.0018	(.04)	.0006	(.33)
REM	.0020	(1.26)	.0019	(1.41)	.0013	(.03)	.0002	(.20)
HLTH	-.0193	(-1.04)	-.0116	(-.74)	-.0025	(-.11)	-.0108	(-.83)
LIQ	.0001	(.12)	-.0002	(-.39)	-.0002	(-.58)	.0003	(.80)
OWN	-.0238	(-1.23)	-.0162	(-.98)	-.0027	(-.11)	-.0112	(-.82)
RES	-.0002	(-.66)	-.0002	(-.61)	.0001	(.35)	-.0001	(-.35)
MAR	-.0128	(-.42)	-.0040	(-.15)	.0032	(.14)	-.0081	(-.38)
WLFP	.0359	(1.77)	.0297	(1.73)	.0170	(.11)	.0121	(.86)
WW	-.0015	(-.22)	.0007	(.12)	.0010	(.24)	-.0030	(-.65)
N	1724		1654		1588		1608	
R^2	.071		.058		.035		.032	

Table A-2

OLS Regressions on the Probability of Quitting between 1966 and 1969 using the Actual Wage (without job tenure)

	All Quits		Quit for Job-Related Reason		Quit Because Found Better Job		Quit for Personal Reason	
	b	t	b	t	b	t	b	t
w_o	-.0079	(-1.97)	-.0071	(-2.09)	-.0045	(-1.91)	-.0018	(-.64)
NOTTAKE	-.0604	(-3.81)	-.0416	(-3.11)	-.0305	(-3.23)	-.0246	(-2.24)
STEADY	.0482	(1.33)	.0119	(.38)	.0277	(1.24)	.0478	(1.85)
ACCEPT	.0406	(.56)	-.0345	(-.54)	-.0518	(-1.15)	.0737	(1.48)
OTHER	-.0451	(-1.22)	-.0472	(-1.51)	-.0414	(-1.88)	-.0025	(-.10)
LIKE	-.0899	(-3.23)	-.0662	(-2.76)	.0117	(.66)	-.0417	(-2.09)
PENS	-.0599	(-3.86)	-.0515	(-3.93)	-.0263	(-2.83)	-.0166	(-1.53)
PUBLIC	-.0624	(-2.42)	-.0624	(-2.86)	-.0220	(-1.44)	-.0070	(-.39)
DEVP	.0013	(1.12)	.0012	(1.20)	.0007	(1.01)	.0003	(.39)
DEVN	.0001	(.03)	.0001	(.03)	.0007	(.43)	.0000	(.00)
WKS	.0000	(.00)	.0003	(.12)	-.0035	(-2.18)	-.0003	(-.17)
SPELLS	.1124	(4.43)	.0733	(3.33)	.0497	(3.10)	.0683	(3.67)
EDUC	.0015	(.60)	.0014	(.63)	.0019	(1.28)	.0002	(.12)
REM	.0019	(1.23)	.0018	(1.39)	.0013	(1.38)	.0002	(.19)
HLTH	-.0195	(-1.05)	-.0114	(-.72)	-.0022	(-.20)	-.0111	(-.86)
LIQ	.0000	(.08)	-.0001	(-.32)	-.0001	(-.48)	.0002	(.67)
OWN	-.0251	(-1.30)	-.0168	(-1.02)	-.0028	(-.24)	-.0121	(-.89)
RES	-.0002	(-.64)	-.0002	(-.62)	.0001	(.33)	-.0001	(-.31)
MAR	-.0129	(-.42)	-.0029	(-.11)	.0041	(.22)	-.0090	(-.43)
WLFP	.0379	(1.86)	.0303	(1.76)	.0170	(1.40)	.0136	(.97)
WW	-.0022	(-.32)	.0003	(.05)	.0008	(.21)	-.0034	(-.74)
N	1724		1654		1588		1608	
R^2	.070		.057		.035		.031	

Table A-3

OLS Regressions on the Probability of Quitting between 1966 and 1969 using the Reservation Wage (with job tenure)

	All Quits		Quit for Job-Related Reason		Quit Because Found Better Job		Quit for Personal Reason	
	b	t	b	t	b	t	b	t
$\hat{w}$	-.0032	(-1.14)	-.0023	(-.96)	-.0017	(-1.00)	-.0011	(-.56)
NOTTAKE	-.0372	(-2.36)	-.0248	(-1.86)	-.0230	(-2.39)	-.0163	(-1.46)
STEADY	.0316	(.91)	.0057	(.18)	.0237	(1.07)	.0410	(1.61)
ACCEPT	-.0059	(-.08)	-.0599	(-.98)	-.0633	(-1.42)	.0548	(1.11)
OTHER	-.0568	(-1.60)	-.0548	(-1.31)	-.0461	(-2.12)	-.0085	(-.34)
LIKE	-.0862	(-3.24)	-.0663	(-2.86)	.0082	(.47)	-.0425	(-2.15)
PENS	-.0300	(-2.00)	-.0285	(-2.22)	-.0158	(-1.72)	-.0060	(-.55)
PUBLIC	-.0645	(-2.61)	-.0631	(-3.00)	-.0238	(-1.59)	-.0104	(-.59)
DEVP	.0002	(.19)	.0004	(.44)	.0004	(.56)	-.0001	(-.11)
DEVN	-.0010	(-.38)	-.0009	(-.40)	.0002	(.11)	-.0002	(-.13)
WKS	-.0026	(-1.05)	-.0018	(-.84)	-.0046	(-2.84)	-.0015	(-.82)
SPELLS	.0929	(3.82)	.0597	(2.81)	.0446	(2.83)	.0637	(3.47)
EDUC	.0028	(1.16)	.0018	(.90)	.0021	(1.40)	.0010	(.56)
REM	-.0020	(-1.31)	-.0012	(-.91)	-.0002	(-.20)	-.0013	(-1.18)
HLTH	-.0154	(-.87)	-.0081	(-.53)	-.0010	(-.09)	-.0097	(-.76)
LIQ	.0003	(.67)	.0000	(.10)	-.0001	(-.25)	.0003	(1.08)
OWN	-.0158	(-.85)	-.0120	(-.75)	.0000	(.00)	-.0068	(-.50)
RES	.0003	(.77)	.0002	(.53)	.0003	(1.14)	.0001	(.49)
MAR	.0036	(.12)	.0077	(.31)	.0081	(.45)	-.0030	(-.14)
WLFP	.0362	(1.86)	.0291	(1.75)	.0170	(1.43)	.0130	(.94)
WW	-.0052	(-.82)	-.0023	(-.42)	-.0005	(-.12)	-.0045	(-.98)
TENURE	-.0089	(12.48)	-.0068	(-10.97)	-.0033	(-7.43)	-.0035	(-6.77)
N	1724		1654		1588		1608	
R^2	.149		.122		.068		.059	

Table A-4

OLS Regressions on the Probability of Quitting between 1966 and 1969 using the Actual Wage (with job tenure)

	All Quits		Quit for Job-Related Reason		Quit Because Found Better Job		Quit for Personal Reason	
	b	t	b	t	b	t	b	t
w_o	-.0020	(-.52)	-.0024	(-.71)	-.0023	(-.96)	.0004	(.14)
NOTTAKE	-.0334	(-2.18)	-.0222	(-1.70)	-.0211	(-2.25)	-.0149	(-1.36)
STEADY	.0352	(1.01)	.0082	(.27)	.0255	(1.16)	.0425	(1.67)
ACCEPT	-.0024	(-.03)	-.0575	(-.94)	-.0616	(-1.39)	.0563	(1.15)
OTHER	-.0533	(-1.51)	-.0522	(-1.73)	-.0441	(-2.04)	-.0074	(-.29)
LIKE	-.0877	(-3.29)	-.0670	(-2.90)	.0079	(.45)	-.0435	(-2.21)
PENS	-.0307	(-2.04)	-.0287	(-2.24)	-.0159	(-1.72)	-.0065	(-.60)
PUBLIC	-.0643	(-2.60)	-.0629	(-2.99)	-.0235	(-1.57)	-.0104	(-.59)
DEVP	.0002	(.21)	.0004	(.43)	.0004	(.52)	-.0001	(-.07)
DEVN	-.0015	(-.59)	-.0011	(-.51)	.0001	(.07)	-.0006	(-.34)
WKS	-.0025	(-1.01))	-.0017	(-.82)	-.0045	(-2.83)	-.0014	(-.77)
SPELLS	.0924	(3.80)	.0596	(2.80)	.0446	(2.83)	.0632	(3.44)
EDUC	.0023	(.97)	.0017	(.84)	.0021	(1.41)	.0006	(.33)
REM	-.0020	(-1.33)	-.0012	(-.93)	-.0002	(-.22)	-.0013	(-1.19)
HLTH	-.0159	(-.89)	-.0082	(-.54)	-.0009	(-.08)	-.0101	(-.79)
LIQ	.0003	(.56)	.0000	(.08)	-.0001	(-.21)	.0003	(.92)
OWN	-.0169	(-.91)	-.0125	(-.79)	-.0001	(.00)	-.0076	(-.56)
RES	.0003	(.80)	.0002	(.54)	.0003	(1.13)	.0001	(.54)
MAR	.0025	(.08)	.0077	(.31)	.0085	(.47)	-.0042	(-.20)
WLFP	.0381	(1.96)	.0297	(1.79)	.0171	(1.44)	.0144	(1.04)
WW	-.0057	(-.39)	-.0025	(-.46)	-.0006	(-.14)	-.0048	(-1.06)
TENURE	-.0090	(-12.56)	-.0068	(-10.99)	-.0033	(-7.42)	-.0036	(-6.86)
N	1724		1654		1588		1608	
R^2	.148		.122		.068		.059	

Table A-5

OLS Regressions on the Probability of Being Laid Off between 1966 and 1969

	b	t	b	t
w_o	.0040	(1.13)	.0077	(2.19)
NOTTAKE	-.0339	(-2.41)	-.0175	(-1.27)
STEADY	.0545	(1.68)	.0501	(1.58)
ACCEPT	-.0206	(-.31)	-.0360	(-.55)
OTHER	.0035	(.11)	-.0041	(-.13)
LIKE	-.0114	(-.44)	-.0193	(-.77)
PENS	-.0595	(-4.31)	-.0392	(-2.88)
PUBLIC	-.0898	(-3.84)	-.0901	(-3.96)
DEVP	.0023	(2.25)	.0016	(1.60)
DEVN	-.0005	(-.21)	-.0014	(-.62)
WKS	.0041	(1.97)	.0023	(1.10)
SPELLS	.1346	(6.59)	.1256	(6.30)
EDUC	-.0065	(-2.89)	-.0060	(-2.73)
REM	.0008	(.59)	-.0016	(-1.20)
HLTH	-.0014	(-.08)	.0016	(.10)
LIQ	-.0006	(-1.40)	-.0004	(-1.03)
OWN	.0265	(1.52)	.0308	(1.81)
RES	-.0002	(-.64)	.0001	(.36)
MAR	-.0138	(-.51)	-.0053	(-.20)
WLFP	.0011	(.07)	.0011	(.06)
WW	.0088	(1.91)	.0061	(1.35)
TENURE			-.0061	(-9.45)
N	1679		1679	
R^2	.106		.151	

TABLE A-6
Effects of the Wage Rate on the Probability of Separating from Short and Long Jobs
When Job Tenure is Held Constant*

	Separations			Quits			Layoffs		
	Short Jobs	Long Jobs	All Jobs	Short Jobs	Long Jobs	All Jobs	Short Jobs	Long Jobs	All Jobs
A. Not Holding Industry Constant									
$\hat{w}$	.0022	-.0049	-.0003	-.0030	-.0040	-.0032	.0084	-.0013	.0035
	(.29)	(-1.77)	(-.08)	(-.35)	(-1.79)	(-1.15)	(1.11)	(-.71)	(1.40)
$\bar{R}^2$	.365	.036	.200	.294	.021	.138	.393	.048	.138
w_o	.0110	-.0048	.0039	.0025	-.0050	-.0020	.0137	-.0002	.0077
	(1.05)	(-1.27)	(.87)	(.21)	(-1.63)	(-.52)	(1.28)	(-.07)	(2.19)
$\bar{R}^2$	.366	.035	.201	.293	.020	.138	.394	.047	.140
B. Holding Industry Constant**									
$\hat{w}$	.0032	-.0053	-.0013	-.0016	-.0039	-.0034	.0077	-.0018	.0026
	(.41)	(-1.92)	(-.39)	(-.19)	(-1.74)	(-1.20)	(1.13)	(-.97)	(1.04)
$\bar{R}^2$	.383	.049	.219	.313	.024	.144	.431	.059	.172
w_o	.0119	-.0055	.0014	.0049	-.0049	-.0023	.0104	-.0009	.0051
	(1.12)	(-1.46)	(.32)	(.42)	(-1.59)	(-.60)	(.98)	(-.36)	(1.46)
$\bar{R}^2$	.385	.048	.219	.313	.024	.143	.430	.058	.172
N	434	1431	1865	332	1392	1724	305	1374	1679

*The coefficients presented in this table are OLS coefficients.

**We use a set of dummies at a one-digit industry level.