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WAGE CHANGES IN JOB CHANGES

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ABSTRACT

This is a study of short and longer-run wage gains observed in moving from one job (firm) to the next. Short-run wage gains are defined as wage changes over the survey year bracketing the move minus the opportunity cost of moving. The latter is measured by wage growth of a subgroup of stayers whose mobility behavior and other characteristics are the same as of the current period movers. Longer-run wage gains are defined as the difference in wages between two successive jobs at the same tenure levels, net of experience, again net of opportunity costs.

Wage gains of movers are generally positive, except for layoffs of older workers. A large part of the gain is due to the lesser wage growth on the job of movers compared to (all) stayers. This is consistent with below average amounts of on the job training observed for movers compared to all workers.

Wage gains of quits exceed those of layoffs, despite similar wage levels and wage growth on the preceding job. Wage gains of older movers are smaller compared to gains of younger movers, both in quits and in layoffs. Differences in search conditions and in the nature of separations help to explain these findings.

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WAGE CHANGES IN JOB CHANGES

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I. INTRODUCTION AND SUMMARY

In this study we estimate short and longer-run wage changes observed in the period of moving from one job to the next. Short-run wage changes are defined as the difference between the starting wage on the new job and the wage observed a year before on the old job. Longer-run changes are defined as the difference in wages between the two jobs at the same tenure levels, net of experience. In effect, this change in wages is measured by the shift of the tenure-wage profile in the successive jobs. Wage gains of movers are usually estimated by the difference between wage growth of movers and wage growth of all stayers during the observation period. Heterogeneity in wage profiles of workers creates a selectivity bias in using all stayers as the control group. We try to reduce the bias by using a more appropriate control group, namely of those stayers whose mobility behavior, in addition to other observable characteristics, is similar to that of current period movers. Their wage growth on the job proxies for the wage growth foregone by movers.

Wage changes of movers were estimated in the period 1970-1981, as well as in the more recent subperiod 1976-1981, for white male workers, non-students, up to age 60. We distinguished subgroups of young (first decade of work experience) and of all other, more experienced workers. We also distinguished movers by type of separation, quit and layoff. Other characteristics, such as education, marital status and union membership were used as independent variables in the statistical regression equations.

Findings in both the shorter and longer run reveal similar facts: Wage gains of movers are generally positive,[1] except for layoffs among the older

workers. Wage gains in quits exceed wage gains in layoffs. Wage gains of older movers are smaller (or even negative) than wages of young movers. This is partly due to the greater frequency of layoffs (compared to quits) among older workers.

These systematic patterns are clearly observed, when the mobility wage gains are related to the group of next-period movers rather than to the less appropriate control group of all stayers. Indeed, a large part of the gain is due to the lesser wage growth on the job of movers compared to stayers. This is consistent with findings in our previous studies, according to which less frequent movers receive more job training and grow more rapidly on the job.[2] This is one reason for the flatter life-time trajectories of frequent movers. The other, according to our present findings is that, with the exception of young workers (especially quits), movers, despite average gains in moving, do not catch up with wage levels of stayers. This is true mainly of the more experienced movers.

The adverse effect of layoffs compared to quits on mobility wage gains is not traceable to differential behavior on the prior job, insofar as it is reflected in wage levels and wage growth: Both are about the same prior to separation, and lower than among (the average) stayer. However, search behavior both on the job and off-the-job is apparently different. Unemployment encountered by most layoffs reduces wage gains, especially if it is prolonged. Fewer quits enter unemployment which also tends to be shorter than in layoffs.

A search model which focuses on search efficiency, defined both as search effort and as personal or environmental (state of the market) search productivity is capable of explaining differential mobility wage gains by layoffs relative to quits and by age, as well as by other characteristics such

as education, marital status, national unemployment, and probably other criteria which we have not studied. The model suggests that differences in search efficiency create differences in duration of search (unemployment) and in acceptance wages, as well as a negative relation between the two. Differences in costs of search would, of course, create a positive correlation. In our findings differences in search efficiency appear to dominate.

Although the decline in wage gains of older movers is largely due to the adverse effects of layoff unemployment (quit unemployment is not as, or not at all, deleterious), wage gains decline with age in quits as well. The work of Bartel and Borjas (1981) with NLS data indicates that quits which are exogenous (for family, health, and other reasons) result in wage losses, and so do, to a lesser extent, quits which represent trade-offs of wages for preferred other working conditions. We find one important example of the latter in our data: Reduced wage gains are traded off for preferred changes in hours. The phenomenon is significant in the experienced labor force, not so among the young. We may conclude that the reduction of wage gains in quits of older workers is due primarily to the greater prevalence of the exogenous and trade-off categories among them.[3]

II. Measuring Wage Changes in Transitions

In estimating returns to inter-firm job changes we should distinguish short-term wage changes obtained in the transition from longer-run changes represented by a shift in the tenure profile of wages after the move relative to that profile in the previous job. Job change decisions of workers are motivated by both kinds of gains (or losses). Their relative importance depends on the worker's discount rate. More generally, it is the present

value of the net gain that matters. Although present value effects of mobility are not estimated in this study, some inferences about comparative magnitudes (as between younger and older workers) are discussed in Section IV. It should also be noted that the wage gains we observe are net of costs of foregone wages, but not of direct and other costs. While such costs may be negligible in job change within the local labor market which is preponderant, some 65% in our data, they are surely important in geographic mobility. Data limitations prevented us from properly distinguishing between local and geographic mobility in the present report.

(A) Short Run Wage Gains

Most studies of longitudinal data contain estimates only of short-run wage transitions,[4] but the quality of available findings is not secure. Wage transitions observable in panel data cover an interval of at least one year[5] between reports of the first wage on the new job and the last wage on the old. Since wages change (usually grow) for stayers as well as for movers, a correct estimate of the wage gain from moving is the difference between the actual wage gain of movers over the interval and the unobserved but expected wage gain of movers had they not moved over the same interval. In the usual procedure, the coefficient of a job change dummy (S_t) in a wage growth equation is used to estimate the wage transition. What it really measures is the difference between the wage gain of movers and wage gain of stayers both defined over this particular time interval (t).

Define the wage gain of movers over the interval t as $w_{m,t}$ and the wage gain of stayers over this interval as $w_{s,t}$. Thus, the coefficient on S_t is $w_{m,t} - w_{s,t}$. What is known as the "selectivity problem" is that wage growth of stayers ($w_{s,t}$) is not likely to be the same as the expected wage growth of movers, call it $w_{ms,t}$, had they stayed. Put more strongly, the coefficient

on S_t tells us how much better (or worse) movers fare compared to stayers, but this is an irrelevant and even faulty question. It is prima facie faulty, because any answer would suggest that one or the other group is acting irrationally: if $w_{s,t}$ is the opportunity cost of movers in moving (we are ignoring other costs here), then $w_{m,t}$ must be the opportunity cost of stayers in staying. Hence if $w_{m,t} > w_{s,t}$, the stayers are irrational, and conversely if the inequality sign is reversed. It is irrelevant because economic optimization means that movers are doing their best by moving and stayers by staying. Strictly speaking this is true ex ante, as well as, on average, ex post, so long as most people are not misled by incomplete information. We must replace the incorrect opportunity cost of movers $w_{s,t}$ by an estimate of the correct one, $w_{ms,t}$.

We should be careful to note, however, that the dichotomy of movers and stayers which bears on the selectivity problem was defined solely for the particular interval t. Some workers who moved in t may otherwise move very infrequently, while some of those who did not may otherwise move quite frequently. In a longer-term perspective, there is no dichotomy between movers and stayers, there is a spectrum of workers ranging from those who move rarely to those who move very frequently. Since average job tenure in our sample is about 7 years, movers observed in a one year interval must have an above-average probability of moving, that is of being repeat movers over their working lives. Indeed, their job tenure is on average closer to 3 years. As was shown in our previous study, workers who tend to invest more heavily in human capital formation on the job, are likely to move less frequently and to have steeper wage growth while on the job.[6] Consequently, movers in period t should have systematically weaker on-the-job growth than stayers in period t. The data on job training are consistent: while all workers had an average

training period of 2.5 years, movers had training of about 1.5 years (in the 1976-1981 annual periods). Hence the "selectivity" bias. The correct opportunity cost of movers $w_{ms,t}$ is smaller than $w_{s,t}$. But it can be estimated on a subgroup of stayers in period t. Since the time period t is arbitrary, workers who have similar personal characteristics including mobility behavior should have similar wage growth on the job. If so, we can approximate the unobserved growth of movers had they not moved ($w_{ms,t}$) in period t by the observed wage growth of stayers in period t who are otherwise similar to our movers, and who are observed to move in period (t+1), call it $w^{m,t+1}_{s,t}$. The only difference is that in period t the "future" movers (those moving in t+1) stay on the job.[/]

The wage growth equation now contains two change dummies, S_t, denoting separation in period t, and S_{t+1}, denoting separation in period t+1, while the dependent variable is wage growth of all in period t. S_t is 1 if a move takes place in period (t) and 0 if no move occurs in it, nor in the next, while S_{t+1} = 1 if a move occurs in period (t+1), but not in (t), and 0 if it does not occur in the second (or preceding) period. The coefficients on S_1 and S_2 are, respectively.

$$D_t = w_{m,t} - w_{s,t} = G(s) \quad (1)$$

$$D_{t+1} = w_{m,t+1} - w_{s,t}. \quad (2)$$

Hence the corrected estimate of the relevant wage gain is

$$D_t - D_{t+1} = w_{m,t} - w_{m,t+1}. = G(m)$$

The basic idea is that $w_{m,t+1}$, the wage growth of stayers observed in the period preceding their move (in t+1) is a good approximation for the unobserved wage growth of movers in period t, had they stayed on the job ($w_{ms,t}$). This procedure may not eliminate the bias entirely, but we expect it to provide a much better control than $w_{s,t}$ in the naive procedure, which

ignores the selectivity issue.

(b.) Longer-Run, Job Gains in Mobility

Longer-run changes in wages resulting from an inter-firm change in employment are estimated as upward (or downward) parallel shifts in the tenure-wage profile in the new relative to the preceding job. Since we observe only initial wages on the new job, the shift is basically an estimate of the difference between the starting wages on the new and the preceding job, net of wage levels resulting from accumulated experience. The same issue of selectivity arises here as it did in estimating short-run changes. Our procedure to deal with this problem is essentially the same. It permits us to compare wage gains of movers with wage gains of all stayers, as well as with wage gains of comparable stayers during the year or tenure which was foregone by the movers.

Both wage level and wage change equations are used in the estimation of job gains. Wage level equations are used separately for prior and subsequent surveys bracketing the move.[o] Both equations contain the same separation dummies. In time t, the coefficient on the separation dummy indicates the cross-section wage difference between otherwise similar movers and stayers prior to the move, while in time (t+1) it measures this differential after the move. Separation dummies one year forward are added in each equation to provide for the control group of future movers. While the prospective equation (at t) indicates the wage level selectivity of movers, the difference between coefficients on the same separation in retrospective at (t+1) and the prospective equation measures the gain from moving relative to all stayers.

The gain relative to comparable (next-period) movers is obtained by subtracting the difference in coefficients on the future move from the difference in the coefficients of the current move. The prospective equation

(P) containing levels at time t has (in addition to other variables) dummies on S_t, the move in the subsequent period and on S_{t+1}, the move in the following period. The dummy on S_t denotes the differential between wages (at t) of next period movers before they moved and stayers (at t), $D^P_t = w^t_{m,t} - w^t_{s,t}$. The dummy on S_{t+1} denotes the differential between wages (at t) of those who will move two periods ahead, and of stayers, $D^P_{t+1} = (w^t_{m,t+1} - w^t_{s,t})$. In the retrospective equation (R) containing levels at t+1 the dummy on S_t denotes the differential between wages (at t+1) between those who moved in the preceding period and those who stayed, $D^R_t = (w^{t+1}_{m,t} - w^{t+1}_{s,t})$. Again, the dummy on S_{t+1}, the move in the subsequent period, denotes the differential between wages (at t+1) of such movers and of stayers, $D^R_{t+1} = (w^{t+1}_{m,t+1} - w^{t+1}_{s,t})$.

The wage gain of movers relative to stayers is therefore measured by the difference between the dummies on S_t in the two equations,

$$G(S) = D^R_t - D^P_t = (w^{t+1}_{m,t} - w^{t+1}_{s,t}) - (w^t_{m,t} - w^t_{s,t}) = w_{m,t} - w_{s,t} \qquad (1')$$

The difference between the wage growth of next period movers and of stayers is measured by the difference between the dummies on S_{t+1} in the two equations

$$G(mS) = D^R_{t+1} - D^P_{t+1} = (w^{t+1}_{m,t+1} - w^{t+1}_{s,t}) - (w^t_{m,t+1} - w^t_{s,t}) = w_{m,t+1} - w_{s,t} \qquad (2')$$

hence the corrected estimate of wage gains of movers is

$$G(m) = G(S) - G(mS) = (w_{m,t} - w_{s,t}) - (w_{m,t+1} - w_{s,t}) = w_{m,t} - w_{m,t+1} \qquad (3')$$

As (2') suggests, instead of taking differences between coefficients in two level equations, we may use a wage change equation which is derived from the two level equations. This alternative procedure is, in principle, equivalent to the first. However, estimates in the two procedures may differ

insofar as wage change equations require panels in neighboring years, which reduces the sample size somewhat, and insofar as the error structure is altered. Although the estimate of short-term gains in (2) is algebraically the same as the estimate of longer-run "job gains" in (2'), the estimates are derived in a different manner: The short-term estimates are derived from a wage change equation in which all variables (other than the separation dummies) are levels prior to moving. In the job-gain estimates these variables are in form of first difference over the relevant periods. Note that in differencing the independent variables of the level equations, the experience variable becomes $X=1$ for all, and its coefficient enters the intercept, but X^2 differs with the level of experience. The tenure variable T equals 1 for job stayers but becomes a negative-T, where T is length of job tenure on the preceding job. Similarly, T^2 is positive for stayers, but is negative and equals $-T^2$ for movers. The use of T is the key to estimating "job" gains. When tenure <u>levels</u> before separation were held fixed in the equations, the mobility dummies indicated the more immediate wage change in moving from one job to another. To estimate "job gains" T is "held constant."[9] The present specification estimates the wage change from the prior to the current job at comparable tenure levels, in effect at starting wages, assuming the same shapes (but not level) of the tenure curve for both jobs. This is corrected for selectivity by the dummies on future movers.

III. <u>Empirical Findings</u>

Tables 1, 2, and 3, show real wage gains[10] from job changes. Table 1 shows the short-run gains of movers over the year the move took place, Table 2 shows the longer-run gain we call "job gain" obtained from a wage change specification, and Table 3 measures the same using pairs of wage levels

equations "prospective" and "retrospective" bracketing the move. In our (PSID) data the period 1976-1981 contains a comprehensive wage and salary coverage for straight-time real wages on the main job. The earlier period (1970-75) data are restricted to wage earners (hourly rated).[11] Tables 1, 2, and 3, which extend the period back to 1970, utilize a dummy for the 1976-81 period to distinguish the subperiods both by coverage and possible historical differences.[12]

Tables 1-3 utilize coefficients of dummy variables for separations in one and two adjacent periods, after inclusion in wage regressions of a number of relevant independent variables. In effect, we are asking how much do movers gain relative to comparable workers who did not move over the period, and relative to a more similar subgroup of stayers, namely those who move in the subsequent period.

In Table 1 short-run ("transition") wage gains were estimated in pooled wage growth equations. Year to year wage growth measured as $\ln(w)$ is the dependent variable. We look at net dollar gains in some of the subsequent tables (4,6,8) when we try to rank present values of gains from moving in different age groups. The independent variables, other than the separation dummies, are shown in Table 1A.

Table 1 shows coefficients on dummies for separation (S), quits (Q), and layoffs (L) for periods t alone (col. 1), and of pairs of dummies in (t) and (t+1) in col. (2) and (3) respectively. Col. (1) measures the gain of movers relative to all stayers G(s), while col. (4) = col. (2) - col. (3), measures the gain G(m) relative to movers in (t+1) who stayed on the job in period t. G(s) is the coefficient of S_t, when S_{t+1} is excluded, while G(m) is the difference between the coefficients of S_t and S_{t+1}, shown in col. (2) and (3).

The separation dummies are added to a list of other, standardizing

variables measured in the survey prior to the move. The variables include education (Ed), experience (X), and tenure (Ten) in linear and quadratic terms, marital and health status (Mar) and (Disab), location variables (city size, whether in SMA, geographic area). Also, the percent-point change in the national unemployment rate (U) of adult males (ages 35-54)[13] over the interval t, and an interaction of the separation with union membership in any of the periods t and (t+1). Statistically significant effects were observed on several of the standardizing variables, shown in Table 1A. Wage growth of all workers (mainly stayers in any given year), diminishes with experience and with tenure in the firm, both in a decelerating manner -- features similar to those in cross-section wage profiles. Changes in the adult male unemployment rate are significant, indicating a pro-cyclical fluctuation in real wages of homogenous labor. The response of real wage changes to unemployment changes is stronger in the more recent period (1976-1981) than in the longer period back to 1970.

Table 1 shows wage gains between jobs calculated from coefficients of mobility dummies. The effects of separations, quits, and layoffs are shown for all workers in the sample, young workers (defined by at most one decade of work experience), and the older, more experienced workers (defined by $x > 10$).

Findings in both the shorter and longer period show similar patterns: (1) Starting with short-run gains in Table 1: Wage growth between jobs does not exceed wage growth on-the-job for the average separation (col. 1). It is even less than the wage growth of stayers, if the separation is due to layoff among the more experienced workers. However, the gain in wages between jobs relative to comparable movers who stayed on in period t, shown in col. (4), is positive -- again, with the exception of layoffs among the more experienced workers. This positive gain is, therefore, largely due to the fact that

movers have weaker wage growth on the job than stayers.

(2) Gains of quitters are positive and exceed gains from layoffs which are small or negative, especially for the more experienced workers. This despite the fact, evident in col. (3) that the flatter on-the-job growth of movers, which represents the opportunity cost of moving, is about the same whether the following separation is a quit or a layoff. The adverse effect of layoffs on wage growth between jobs has been noted before, albeit in terms similar to our col. (1) than to the more appropriate col.(4). The finding appears natural to some and puzzling to others.[14]

(3) Gains from separations, as shown in Table 1 (col. 4) decline with age (experience), both in quits and in layoffs. Once again, the facts are familiar, although based on diverse methodologies. Numerically, the gains from separations range from -1.8% in layoffs of older workers to 7.4% in quits of young workers in the longer period.

Before we proceed to interpretations of the findings, we turn to estimates of longer-run effects of mobility shown in Tables 2 and 3, since they are qualitatively similar to those in Table 1. The layout of Table 2 is the same as that in Table 1. The coefficients on separation dummies come from wage change regressions in which the dependent and independent variables are the same as in those underlying Table 1, except that levels of experience and of tenure are now replaced by annual changes in them.

The job gains (tenure-wage profile shifts) estimated in Table 2 are numerically larger than the "transition" gains shown in Table 1. This is pronounced for the younger movers, suggesting perhaps that beyond the immediate gain, young movers gain also in wage growth on the new job. Otherwise the pattern of profile shifts by type of separation and by age is similar to that found for short-run gains in Table 1.

Table 2A shows the significant coefficients of the standardizing variables: Thus the coefficient of ΔX^2 is negative. This is consistent with a negative sign on X in Table 1A, since $\Delta X^2 = 2X+1$. The sign of ΔTen^2 is similarly negative, but positive for ΔTen, suggesting a negative effect of longer tenure on subsequent gain from moving. Increases in unemployment reduce wage growth as before, though the effects are smaller than in the shorter-run. Interactions of union membership with separations are generally positive, probably reflecting moves from non-unionized to unionized firms.

Table 3 shows the job gain (profile shift) estimates obtained by using prospective (P) and retrospective (R) wage functions in successive cross-sections.[15] G_s, the gain relative to all stayers, is the difference between separation coefficients $(R_t - P_t)$, in row (3) minus row (1); G_{ms} -- the gain of future movers relative to current stayers is $(R_{t+1} - P_{t+1})$ row (4) minus row (2); and G_m -- the gain of movers relative to comparable (next period) movers is $(R_t - P_t) - (R_{t+1} - P_{t+1})$, i.e. rows (3 - 1) - (4 - 2).[16]

It is reassuring to find that the numerical estimates of the relative gains G_m in Table 3 are quite close to those in Table 2. The conclusions are therefore the same as in Table 2. There is, however, additional information in Table 3, namely observations on wage levels of movers and stayers before and after the move. Thus the fact that all coefficients in rows (1) through (4) are negative means that movers have lower wages than stayers both in the old and in the new job, although the discrepancy is reduced by moving. Also, prior to moving the wage disadvantage (relative to stayers) is about the same for quits and for layoffs (row 1 in Table 3). In conjunction with the findings in Tables 1 and 2 to the effect that wage growth on the job is about the same for quits and layoffs, it would seem on the whole that there is no significant difference in on-the-job wage experience of quits and layoffs

prior to separation.

By moving young quits make up the bulk of the deficit, while older quits cut it in half only. Young layoffs remove less than half of the deficit, while older layoffs worsen it if they change it at all. Apparently, it is behavior during the transition rather than on the job that distinguishes quits from layoffs. Since mobility does not compensate for the slower wage growth on the job of frequent movers, their wage trajectories are flatter, and even flatter for those movers whose separations are dominated by layoffs. However, in the long run more frequent movers are not distinguishable by quits or layoff. No significant correlation between frequency of moving and relative frequencies of quits to layoffs was found among movers in the PSID.

Interaction Effects

Mobility wage gains differ by type of separation and by age, according to Tables 1-3. Do mobility wage effects differ also by other characteristics, such as education, experience, tenure, and so on? To detect these interactions we restrict our observations to episodes of moves. In Table 4, in the upper panel we show relative wage gains as before. In the lower panel we show dollar wage gains. In both panels we look at wage changes during the move (in period t) and before the move (at t-1). The latter are substitutes for the wage change of stayers who will move in t+1.

We find that both the relative and dollar gains differ by education, tenure, by health, by location (in SMA), and less significantly by experience and by marital status. These results are qualitatively similar in the percent and dollar measures, and are stronger when lagged wage effects (coefficients in (t-1)) are subtracted from current effects (in t) which refer to current movers without reference to any comparison group.

Gains from job mobility increase with education, mainly because gains from quits do. In relative terms the effect of education is negative only in layoffs of young workers. Note also, in col. (t-1), that more educated workers tend to quit jobs which are relatively inferior (in terms of wage growth), and gain more by moving, the more inferior the previous job was. Gains decline with tenure in a decelerating fashion. They are reduced in ill health, and they are smaller in SMA's. They are larger (certainly in dollar terms) for married than for single men. Experience effects are negative and near significance when health variables are left out.

Recessions (increases in national unemployment) have a stronger negative effect on wage growth of movers (-3.7% in Table 4 as compared to -2.5% in Table 1A) than on wage growth of stayers. Thus, the cyclical variability of real wages is greatest for movers and declines with experience (and/or tenure) among stayers (Table 1A). Lesser cyclical fluctuations in wages may be expected of workers who have accumulated more firm-specific capital than others, or, more generally, carry greater fixed labor costs to employers. The comparative findings on movers and stayers, young and old, and early vs. later tenure are consistent with this hypothesis.

IV. Wage Changes and Unemployment in Job Transitions.

1. Why do gains from quits exceed gains from layoffs?

As was noted in Tables 1-3, gains from quits exceed gains from layoffs in all groups and periods, both in the transition and in the job sequence. The differential is rather stable. Gains from quits exceed gains (or losses) from layoffs by about 4% points. A dummy for (layoffs=1, quits=0) included in the wage change equations for moves only, indicates the difference. It is shown in Col. 1 of Table 5.

Explanations of the disadvantage of layoffs in the transition are put forward by several analysts. Cline (1980) argues that quitters move from inferior jobs when better jobs become available, while involuntary separations (layoffs) affect workers in both inferior and superior jobs. The implication is that the pre-separation wage levels are, on average, higher for layoffs, while post-separation wages are about the same. Our findings in the prospective and retrospective regressions of Table 3 show exactly the opposite: Pre-separation wages of both quits and layoffs are about the same, both significantly lower than wages of stayers, while post-separation wages, although not equal to wages of stayers, are lower for layoffs than for quits. Rosen argues against the asymmetry between layoffs and quits in his preface to NBER volume (1981): "Who initiates the turnover decision should be irrelevant, since job separations should occur if and only if productivity on the current job is less than productivity on an alternate job." Rosen adds the speculation that perhaps layoffs are more heavily selected from unstable sectors, hence a greater average loss (or lesser gain) may be expected for layoffs than for quits. This assumes higher (compensatory) wages in unstable sectors and (some) moves from unstable to stable sectors: While our prospective regressions are consistent with the symmetry argument, they reject the unstable sector hypothesis.

A hypothesis that may explain the larger gains of quitters, relies on job search behavior: Most quitters change jobs directly without intervening unemployment, while most layoffs are unemployed between jobs. The implied on-the-job search of quitters carries a reservation wage which exceeds the wages on the old job (abstracting from non-wage components of the real wage package), while the reservation wage of the laid-off unemployed searchers is lower:[17] The starting wage on the new job may, but need not, be higher than

on the old job. In our data 2/3 of quits changed jobs without unemployment, while about the same proportion of layoffs became unemployed. If reservation wages are indeed lower for unemployed than for employed searchers we should find that the reduction in wage gains due to unemployment is greater than that due to layoff. This is shown in Table 5, comparing col. 2 with col. 1.

The disadvantage of searching while unemployed is less significant for quits than for layoffs (compare col. 4 with col. 6), especially among the more experienced workers. Quit into unemployment is deliberate -- it is likely to be prompted by high costs of searching on-the-job. It therefore indicates the intent to search more intensively. Moreover, the intensity of search by unemployed quitters is likely to be strengthened by lack of unemployment compensation for which layoffs are eligible. Indeed, average duration[18] of layoff unemployment is almost twice as long as of quit unemployment (Shown in Table 7).

If intensity of search of unemployed quitters is greater than that of those on layoff, it should also follow that <u>for the same length of search</u> quitters should be more successful in locating a better (higher wage) job. This is confirmed in our regression (Table 5) in which a duration variable is added to the unemployment dummy: The coefficient on duration is not significant for quits, but is negative and significant on layoffs. That is to say, that each additional month of unemployment reduces the acceptance wage of unemployed layoffs by an additional 2.8%. Apparently, longer duration reflects lesser or decreasing efficiency of search[19] of those laid off, but not of those who quit.

2. <u>Why do gains from moving decline with age?</u>

(a) Incidence and Duration of Unemployment

According to Table 1, short run gains from separations of experienced

workers (X > 10) are much smaller (about a third) than gains of young workers. The reduction is somewhat less in quits (about a half), but goes from positive to negative in layoffs. The use of percent gains may, of course, be misleading. It is possible for the net dollar gain (the difference between the dollar gains of movers and the dollar gains of the control-group) to increase when the percent gain declines. Table 6 below shows that the gross dollar gain from moving of young workers was larger than the corresponding gain of older workers. This implies that the net dollar gain declines as well, given the decline in the percent measures[20] (G_m).

As Tables 2 and 3 indicate, the decline in G_m with age is even greater in our job gain estimates, which are conceptually closer to present values.

One reason for the decline in the wage gain of the older job changers is that a greater proportion of them experience layoff. The ratio of layoffs to separations is close to 70% among the older workers, compared to about 50% for the younger ones. And, as was shown in Table 5, the effects of unemployment are especially severe for laid off older workers, and more so in prolonged unemployment, which is more typical of laid off workers. It follows that the differential effects of unemployment, especially in layoffs explains both the wage gain differences between quits and layoffs, and the declines of gains for older job changers.

An explanation of all these patterns, already alluded to, may be found in differential efficiency of search by type of separation, age, education or other characteristics of workers or labor markets. Greater efficiency in search may be a matter of personal efficiency, or greater intensity of search, or of a more favorable environment. We define it as the probability (p) of finding a vacancy (whether or not the job is ultimately acceptable) per unit of time. Of course, p is a function of resources invested in search and of

the environment, but it will suffice for our purposes to treat it as a parameter:[21]

In the terminology of search models, we argue that, on average, older workers who separate from jobs have a lesser probability of finding a job per unit of search time, not because they are holding out for a higher acceptance wage within the relevant wage offer distribution (though it is true of some), but because the probability of getting any offer, that is, the probability of finding a vacancy, is smaller. On this assumption we can show that older workers who separate will search longer when unemployed, and quit less frequently, while their acceptance wage will be relatively lower, so the wage gain will be smaller for older job movers than for younger ones.

In the initially standard search model, the individual samples from his wage offer distribution f(w) receiving one offer per unit of time. The worker decides on an optimal wage floor which equates the gain from an additional unit of search to the cost of it. The resulting rule is:

$$(1) \quad P_a(\overline{W}_a - W_a) = c$$

where W_a is the lowest acceptable wage, P_a is the probability of getting an acceptable wage offer, that is, of $W > W_a$, $\overline{W}_a$ the mean of all acceptable wage offers; c is the (marginal) cost of search which includes opportunity and other costs. Income offsets z which are contingent on a continued search, such as unemployment compensation, enter costs with a negative sign. Expected duration of search D is inverse to P_a. In this model search is longer the higher the acceptance wage, which is higher the lower cost of search.

However, the probability of accepting a wage offer must be redefined given that the probability of finding any offer in a unit period can be less than 1. A lesser frequency of vacancies may be a result of depressed business conditions in general, or depressed markets for a particular type of labor, or

a function of lesser efficiency or intensity of search. The optimum condition becomes:

$$(2) \quad p \cdot P_a(\bar{W}_a - W_a) = c$$

Here p is the probability of finding a job offer, P_a the probability of finding an acceptable job conditional on finding a vacancy, and $p \cdot P_a$ is the probability of finding an acceptable job. D is now the inverse of the product $p \cdot P_a$. As before, changes in c produce a positive relation between W_a and D. However, changes in p over the business cycle, or differences in p across people, are likely to produce a negative correlation between W_a and D.

A reduction in p leads to a downward revision of W_a, hence to an increase in P_a. The question is whether $p \cdot P_a$ will rise or fall when p declines. No perfectly general answer can be given to this question, but a most plausible answer is that $(p \cdot P_a)$ will fall, hence the duration of search will lengthen even though W_a is revised downward in consequence of a fall in p. It is easy to see that the difference $(\bar{W}_a - W_a)$ increases as W_a is lowered in a uniform or triangular wage offer distribution.[21] When W_a is reduced, $\bar{W}_a$ is reduced by a smaller amount, so that $p \cdot P_a$ must fall if c is fixed or reduced.

The conclusion that a lower p is very likely to produce longer search and lower acceptance wages holds both for unemployed and for employed searchers. An increased duration of search on the job, of course, means a reduction in the frequency of quit.

In sum, workers facing fewer vacancies in their search may be expected to have a longer duration of search and a lesser wage gain when unemployed, and to inhibit their job change (quitting) when employed.[22] These conclusions are consistent with worker behavior during the business cycle: duration of unemployment increases and quits decline while layoffs increase, partly because employment demand declined and partly to substitute for a decline in

attrition (quits).

Applying the same model to the life-cycle, we may argue that either p or c declines at older ages. A decline in c is not plausibe except very early when labor market entrants become eligible for unemployment compensation. A decline in c would lead to increases in W_a and in wage gains, but the opposite is implied by a fall in p and is observed. The implications that older men have a lesser tendency to quit, a reduced Q/L ratio, and a lower W_a when changing jobs are confirmed in our data.

The longer duration of unemployment of older workers is a well known feature of national statistics. In our data this fact is observed in Table 7, in which duration of unemployment of unemployed job changes is regressed on a number of variables. In col. (1) and (3) where the variables are restricted to education and experience, duration initially decreases (over the first half a dozen years of experience) then increases with experience. When tenure and other variables are added in Col. (3) and (5), it is lengthening of tenure that appears to be responsible for the increased duration of unemployment by age. Apparently, unemployed older movers with longer tenure face greater hardships in job search -- a fact consistent with the smaller wage gains of longer tenured movers, observed in Table 4. It is also consistent with losses of firm and industry specific capital due to structural changes affecting older, longer tenured workers in manufacturing industries especially. Sectoral information as well as that on plant closings is required to probe these matters more deeply.

Looking at the other variables, Table 7 also shows that duration of unemployment diminishes with education. The hypothesis that the efficiency parameter (p) increases with education is therefore consistent with both implications of the search model: greater wage gain (Table 4) and shorter

duration of unemployment. Similarly duration increases with the level of the national unemployment rate (Table 7) and wage gains decline (as we saw in Table 4). Here p reflects the decline of vacancies in recessions. Effects of health disabilities and of marital status similarly fit the search model: longer duration of unemployed and smaller wage gain of the disabled, and the converse for married compared to single men. Finally, if unemployment results from layoff, its duration is 70% longer than if it results from quit, and as shown before, the wage gain is far smaller.

(b) Why do wage gains from quits decline with age?

Although the lesser gains from mobility can be in large measure attributed to unemployed search, wage gains decline also in quits where unemployment plays a minor and not necessarily deleterious role. In their study of NLS data, Bartel and Borjas (1981) were able to classify quits into 3 categories: (1) for personal or family non-market reasons, such as change in health, in family status, and so on), (2) for reasons of dissatisfaction with working conditions, which they call "push", and (3) "job-related" real wage maximizing reasons which they call "pull". Their findings are shown in Table 8. Quits in the first 2 categories led to lesser wage gains or to losses, and the incidence of them was greater among old workers (age over 45 in the NLS) than among the younger ones: 75% vs. 57%. Although the gains estimated by Bartel and Borjas are calculated relative to all stayers, therefore probably understated, the differential patterns of gains by age are consistent with our estimates.

The distinctions within quit categories are not available in our data. However, the role of changes in non-wage conditions can be explored in one rather important case, which fits the Bartel and Borjas category (2), namely

when the job change contains a significant change in hours per week. The different wage-hours schedules offered in different jobs, firms, or industries are due to various technologically and administratively conditioned systems of coordination of the production process, as well as to the presence of fixed labor costs which differ among firms. Consequently, desired changes in hours are more likely to be accomplished in job transitions than within the firm.[25] Where the change is in the direction opposite to the desired one, workers must be compensated by an increase in the wage. When the change in hours is in the desired direction, workers are willing to trade off some of the acceptance wage for the preferred change in hours. Note that the negative effect of wages applies whether or not the desired change is negative or positive, and that the trade-off applies to search and match mobility rather than to moves in response to wide-spread shifts in demand. In the latter case wages would rise with increases in hours and decline with decreases. Thus, if most of job changing which is associated with significant changes in hours, is not in response to demand fluctuations and is in the direction of preferred hours, we should expect to see a negative effect of absolute (positive and negative) changes in hours schedules on the wage gain.

In Table 9 we find that coefficients on the interaction of separations[26] and change of hours in wage gains is negative, suggesting that trade-offs dominate in moves with sizable changes in hours (Δ hrs $>$ 5 per week, in absolute value).

We find also that changes in hours do not significantly affect wages of young workers. Apparently job matching and experimentation applies to work schedules and other working conditions during the early years of work experience.[26] Since reports on hours worked are conceptually less reliable for salaried than for hourly rated workers, we show separately effects of

changes in hours in the sample of wage earners. The results are equally significant, despite the smaller sample size. More importantly, the effects are observed not merely on starting wages. The effects on the longer run ("job") wage gains are at least as clear and significant.

In demand fluctuations hours decline in downturns and so do real hourly wages, but the opposite is true in upswings. To test whether we were not confounding effect of utility trade-offs with demand fluctuations, we also distinguish increases from decreases in hours. If a negative sign on the latter dominates, we may have misinterpreted our findings as a utility trade-off. We find, however, that each of the changes carries a negative coefficient, not significant for the young, but significant for workers with more than 10 years of experiences. The trade-off is a 3 to 6% reduction in wage for over 5 hours change in work schedule for wage and salary earners. It is between 5 and 8% for wage earners alone.

The illustration of effects of changes in hours on wage gains fits category (2) of the Bartel and Borjas classification of quits shown in Table 8. The preponderance of "personal" and "push" types of quits among older workers explains the anomaly that gains from quits decline with age in both relative and dollar terms. It is an apparent anomaly, because a declining payoff period would require increases in gains to induce workers to move. Note that this requirement is indeed fulfilled in the "pull" category of quits. Here returns to strictly wage motivated quits are twice as high in dollar terms for the old movers than for the young ones. The percent gain is about the same. If the gains are assumed to be permanent, the present value of older quits is probably not smaller than that of the younger movers.[28]

Footnotes

[1]Of course, the dispersion is not small. Hence significant numbers of movers incur losses.

[2]Mincer and Jovanovic (1981) provide the human capital explanation and limited empirical evidence. Much stronger evidence is shown in Mincer (1984).

[3]According to Table 8, about half of quits of young workers are mainly wage maximizing ("pull"). This is true of only one quarter of older quits.

[4]The exceptions are Borjas and Rosen (1980), Cline (1980), Mincer (1983) and Polachek and Horvath (1977). These studies differ both in methodology and in population coverage.

[5]One exception is the DIME-SIME data set, in which intervals are as short as one month (Mortensen and Neumann, 1985).

[6]Mincer (1984).

[7]Although one might use the wage growth of movers in the year before they moved, $W^{s,t-1}_{m,t}$ instead of $W^{m,t+1}_{s,t}$, there are several disadvantages due to differences in (calendar) period, non-linearity in tenure-wage profiles, and significant loss of observations, given short tenure of movers.

[8]This approach was used also in Mincer (1983) in the study of union effects on wages.

[9]Among the other independent variables listed in Table 1A.

[10]Wages and salaries on the current, main job, were divided by scheduled hours and deflated by the CPI.

[11]Hourly rated wages were truncated at $9 before 1978, causing some (small) biases.

[12]In fact, most of the estimates in the shorter period are similar to those in the longer period.

[13] This index of unemployment is not affected by compositional changes in the labor force, and it is less likely to reflect supply responses to the business cycle, as do some of the other demographic components of the aggregate unemployment rate.

[14] Bartel and Borjas (1981) term the differential effect of quits and layoffs "not surprising," while Rosen in his introduction to the 1981 volume calls it "a puzzle".

[15] These equations exhibit the usual coefficients found in cross-section wage functions.

[16] See p. 8 above for derivation.

[17]Recall unemployment is excluded from our data, by definition.

[18]Our measures of duration of unemployment in Tables 5 and 7 are weeks unemployed during the year when job change took place. This may represent more than one spell, and may create some inaccuracy.

[19]Although job changers do not, by definition, return to their previous jobs, some of the unemployed may have expected recall. These expectations may wane as duration lengthens, resulting in a drop in the reservation wage.

[20] Let the gross dollar gain of movers be g and of the control group (movers who stay) g_{ms}. The net dollar gain is therefore $g_m = g - g_{ms}$. But

$g_{ms} = g/1+G_m$, so $g_m = g(1-1/1+G_m) = g \times G_m/1+G_m$. It follows that

the net dollar gain g_m declines with age if g does, even if G_m is the same. For g_m to remain the same or to increase, g must rise at least as fast as G_m declines. The decline in G_m is therefore not merely a matter of arithmetic. It does indicate a decline in the gross dollar gain. Indeed, the relative measures (G_m) shown in Tables 1-3 would have to increase with age to keep the net dollar gain from falling.

[21]This parameter is called the "arrival rate" (of offers) in the mathematical search literature. In the version that follows, the model was described in Mincer and Leighton (1982).

[22]Barron (1975), Feinberg (1977), and Nickel (1979) analyze wider classes of wage offer distributions, with similar results. These distributions belong to a more general class of log-concave probability distributions, including uniform, triangular, normal, and exponential among others. Proofs that such wage offer distributions generate a negative correlation between our p and D are given by Flinn and Heckman (1983) and by Burdette and Ondrich (1985).

[23]We need not assume that p is exogenous. It may decline as a result of the search process (e.g. Salop, 1973). The distinction is immaterial for our purposes.

[24] See Altonji and Paxson (1985).

[25] Altonji and Paxson have similar findings for quits of workers who previously expressed the desire to change hours.

[26] Also, moves toward <u>higher wages</u> in preference to other components of the job dominate the mobility of young compared to older workers according to Table (8) below, taken from Bartel and Borjas (1981).

[27] Let i be the discount rate, g_o and g_y the dollar gains of old and young movers respectively, and R the remaining payoff period in years. R=10 for the older movers in the NLS, while R=40 for the young, which is almost infinity for discounting purposes. For the present values of older quits to be no less than for the younger, the following inequality must hold:

$$1/i[1 - (1/1+i)^R]g_o \geq (1/i)g_y$$

$$[1 - (1/1+i)^{10}] \geq 1/2, \text{ hence } (1/1+i)^{10} \leq 1/2, \text{ so } (1+i)^{10} \geq 2$$

The inequality holds for $i \geq 7\%$, a quite realistic condition.

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Table 1

Short Run Wage Gains in Job Change
Annual, 1970-1981

	Separations			
	(1)	(2)	(3)	(4)
	G(s)	S_t	S_{t+1}	G(m)
All	-.005	.0034	-.035	.039
n=7,246	(.6)*	(.4)	(3.9)	
Young	.024	.031	-.033	.064
n=2,689	(1.8)	(2.3)	(2.7)	
Older	-.032	-.022	-.043	.021
n=4,571	(2.4)	(1.5)	(3.1)	

	Quits			
	G_s	Q_t	Q_{t+1}	G(m)
All	.010	.019	-.034	.053
n=7,246	(.9)	(1.7)	(3.1)	
Young	-.006	.024	-.050	.074
n=2,689	(.4)	(1.4)	(3.1)	
Older	.020	.027	-.019	.046
n=4,571	(1.3)	(1.7)	(1.3)	

	Layoffs			
	G(s)	L_t	L_{t+1}	G(m)
All	-.036	-.026	-.044	.018
n=7,246	(2.4)	(1.6)	(3.0)	
Young	.009	.008	-.049	.057
n=2,689	(.4)	(.8)	(2.6)	
Older	-.071	-.059	-.041	-.018
n=4,571	(3.2)	(2.6)	(1.8)	

* Conventional t-statistics in parentheses.

(1) Gains relative to stayers (S_{t+1} not in the equation)

Young = Experience $\leq$ 10 yrs
Older = Experience $\geq$ 10 yrs

(2) Coefficient on separation in t

(3) Coefficient on separation in t+1

(4) Gains relative to movers. (col. 2 - col. 3)

Table 1A

Other Variables in Short-run Wage Change Regressions

Variables	All		Young		Old	
	1976-81	1970-81	1976-81	1970-81	1976-81	1970-81
C	.097*	.097*	.130*	.104*	.079	.096*
	(4.4)	(5.8)	(2.6)	(2.7)	(2.1)	(3.6)
Ed	-.0008	-.0002	.00002	.0016	-.0008	-.0005
	(.6)	(.2)	(.01)	(.8)	(.5)	(.5)
X	-.0037*	-.0035*	.0021	-.0020	-.0032	-.0029
	(2.6)	(3.4)	(.1)	(.2)	(1.2)	(1.5)
X^2	.00004	.00005*	-.0009	-.0004	.000026	.00003
	(1.2)	(2.0)	(.7)	(.5)	(.5)	(.8)
CS	.005	.005	.011	.009	.0025	.004
	(.6)	(.9)	(.7)	(.8)	(.2)	(.5)
Dsab	-.008	-.00008	-.020	-.012	-.005	.003
	(.6)	(.1)	(.7)	(.6)	(.4)	(.3)
ΔU	-.024*	-.014*	-.029*	-.016*	-.021*	-.012*
	(5.2)	(4.7)	(3.4)	(2.9)	(3.9)	(3.6)
Mar	.0008	.0002	.008	.011	.0007	-.004
	(.1)	(.02)	(.5)	(.8)	(.04)	(.4)
SMA	.006	.006	.0026	.005	.006	.005
	(.8)	(1.0)	(.2)	(.5)	(.7)	(.8)
Ten	-.006*	-.0050*	-.036*	-.032*	-.003*	-.0026*
	(4.0)	(4.4)	(3.9)	(4.2)	(1.8)	(2.4)
Ten^2	.0002*	.00015*	.003*	.0027*	.0001*	.00008*
	(3.8)	(4.0)	(2.6)	(2.7)	(2.0)	(2.3)
D(76-81)		-.017*		N.S.		-.024*
		(3.2)				(4.1)

t - statistics in parentheses

C = intercept
Ed = education
X = experience
CS = city size
Dsab = health impairment
ΔU = annual change in the unemployment rate
Mar = married
SMA = standard metropolitan area = 1
Ten = tenure in firm
D(76-81), 1 if in 76-81
n.s. = not significant

Table 2

Job Gains in Mobility (Δln wage) 1970-1981

	Separations			
	(1)	(2)	(3)	(4)
	G(s)	S_t	S_{t+1}	G(m)
All	.032	.040	-.032	.072
	(3.0)	(3.3)	(3.5)	
Young	.089	.099	-.044	.143
	(4.2)	(4.4)	(3.1)	
Older	.035	.048	-.030	.078
	(2.4)	(2.8)	(2.4)	
	Quits			
	G_s	Q_t	Q_{t+1}	G(m)
All	.044	.056	-.030	.086
	(3.6)	(4.0)	(2.8)	
Young	.113	.130	-.052	.182
	(5.0)	(5.4)	(3.2)	
Older	.028	.038	-.016	.054
	(1.6)	(1.9)	(1.0)	
	Layoffs			
	G(s)	L_t	L_{t+1}	G(m)
All	.007	.008	-.041	.049
	(.5)	(.5)	(2.7)	
Young	.059	.057	-.039	.096
	(2.4)	(2.1)	(1.7)	
Older	.012	.029	-.004	.033
	(.6)	(1.2)	(2.4)	

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* Conventional t-statistics in parentheses.
(1) Gains relative to stayers
(2) Coefficient on separation in t
(3) Coefficient on separation in t+1
(4) Gains relative to movers.

Table 2A

Variables in Job Gains (Δ) Regressions

	All		Young		Old	
	1976-81 (1)	1970-81 (2)	1976-81 (3)	1970-81 (4)	1976-81 (5)	1970-81 (6)
C	.64 (2.7)	.63 (3.8)	n.s.	n.s.	.066 (22)	.067 (3.4)
ΔX^2	-.001 (5.0)	-.0008 (6.2)	-n.s.	-.0033 (2.4)	-.001 (3.7)	-.0007 (4.0)
ΔU	-.018 (3.3)	-.014 (4.5)	-.018 (1.7)	-.015 (2.5)	-.016 (2.6)	-.012 (3.5)
ΔTen	.017 (2.1)	.011 (3.8)	.072 (4.9)	.061 (5.7)	n.s.	.004 (1.4)
ΔTen^2	-.0004 (2.4)	-.0004 (3.1)	-.005 (3.0)	-.004 (3.6)	n.s.	n.s.
D(76-81)		-.018 (3.2)	n.s.			-.025 (4.0)

n.s. = not significant

Table 3

Job Gains in Mobility
Prospective (P) and Retrospective (R) Cross-Sections
1970-1981

	All		
Equation	S	Q	L
Q_t	-.131 (10.5)	-.13 (8.8)	-.128 (6.6)
Q_{t+1}	-.085 (6.7)	-.0982 (5.4)	-.087 (3.9)
R_t	-.109 (8.6)	-.082 (5.5)	-.148 (7.7)
R_{t+1}	-.121 (10.1)	-.118 (8.1)	-.126 (6.7)
G(s)	.022	.047	-.020
G(m)	.058	.085	.019

	Young		
Equation	S	Q	L
Q_t	-.125 (8.1)	-.131 (7.1)	-.122 (5.0)
Q_{t+1}	-.062 (3.9)	-.050 (2.6)	-.075 (2.9)
R_t	-.067 (3.5)	-.033 (1.6)	-.096 (3.7)
R_{t+1}	-.125 (8.3)	-.117 (6.6)	-.133 (5.6)
G(s)	.060	.098	.026
G(m)	.123	.167	.083

Table 3 (continued)

	Older		
Equation	S	Q	L
P_t	-.122 (6.2)	-.119 (4.9)	-.112 (3.6)
P_{t+1}	-.100 (5.0)	-.111 (4.3)	-.088 (2.9)
R_t	-.083 (3.9)	-.057 (2.3)	-.135 (4.2)
R_{t+1}	-.109 (5.7)	-.111 (4.6)	-.105 (3.6)
G(s)	.039	.062	-.023
G(m)	.048	.062	-.006

S = separations, Q = quits, L = layoffs
Other independent variables listed in Table 1A

Table 4

Effects of Mobility on Wage Gains
by Characteristics of Movers, 1976-1981

Relative Wage Change:	n=763 Separations		n=352 Quits*	
	At(t)[a]	At(t-1)[b]	At(t)	At(t-1)
Ed	.068	-.048	.0767	-.088
	(1.9)	(2.1)	(2.0)	(3.2)
Ed^2	-.0024	.002	-.0028	.0036
	(1.7)	(2.1)	(1.8)	(3.1)
Ten	-.028	-.011	-.037	n.s.
	(2.2)	(1.8)	(3.4)	
Ten^2	.0017	.0005	.002	n.s.
	(2.7)	(1.8)	(3.4)	
ΔU	-.037	n.s.	-.027	n.s.
	(1.8)		(1.6)	

Dollar Wage Change:	Separations		Quits*	
	At(t)	At(t-1)	At(t)	(At(t-1)
Ed	.85	-.66	1.002	-1.071
	(4.5)	(4.5)	(4.1)	(4.9)
Ed^2	-.032	.026	-.037	.041
	(4.1)	(4.3)	(3.7)	(4.6)
Ten	-.095	-.063	-.127	n.s.
	(1.8)	(1.6)	(1.8)	
Ten^2	.006	.002	.008	n.s.
	(2.1)	(1.4)	(2.2)	
ΔU	-.15	n.s.	-.13	n.s.
	(2.4)		(.9)	

Note: Of the additional variables (listed in Table 1A) Disability was negative, Marital Status positive, both nearly significant. Experience variables had the same signs as Tenure (at t) but were not significant when Tenure was included.

*Layoff variables had the same signs as separations, but were less significant.

[a]At (t) wage growth in move
[b]At (t-1) wage growth a year before move

Table 5
Effects of Layoffs and of Unemployment
on Wage Gains

Sample of Job Changes, 1970-1981

Sample	Separations			Quits		Layoffs	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	L	U	D	U	D	U	D
All	-.046	-.070	-.011	-.050	-.008	-.064	-.014
n=1,082	(2.1)	(3.2)	(3.1)	(1.6)	(1.3)	(1.8)	(3.0)
Young	-.044	-.060	-.008	-.038	-.014	-.036	-.007
n=675	(1.5)	(2.1)	(1.7)	(1.1)	(1.2)	(.6)	(1.1)
Older	-.041	-.085	-.016	-.064	.007	-.119	-.025
n=407	(1.3)	(2.7)	(3.2)	(1.3)	(.6)	(2.3)	(4.3)

Col. (1) Coefficient of dummy for Layoff = 1, Quit = 0.
(2) Coefficient on dummy for Unemployment in Transition = 1, No Unemployment = 0.
(3) Coefficient on duration of unemployment measured in 2-week intervals, linear terms only, in addition to dummy as in (2).

Table 6

Average Relative and Dollar Wage Gains
of Subgroups, by Age and Separation

1970-81	Separations		Quits		Layoffs	
Wage Gains	%	$	%	$	%	$
Young	4.7	.129	5.3	.154	1.6	.090
Older	3.5	.104	4.2	.121	.9	.030
1976-81						
Young	3.7	.134	4.5	.160	2.6	.091
Older	2.7	.096	3.0	.103	.9	.038

Table 7

Duration of Unemployment Between Jobs

Unemployed Movers, 1970-1981

Variables	Coefficients	
Ed	-.56	-.54
	(2.0)	(2.0)
Ed^2	-.018	.017
	(1.6)	(1.5)
X	-.032	.028
	(.9)	(.6)
X^2	.003	.0014
	(2.2)	(1.0)
U		1.40
		(8.1)
Disab.		.72
		(1.8)
Mar.		-.72
		(2.1)
Ten		-.45
		(4.5)
Ten^2		.020
		(3.8)
Layoff		1.71
		(6.3)

Table 8

Wage Changes by Type of Quit in NLS Data

Type of Quit:	Personal			Push			Pull		
Measure	%	$	P[c]	%	$	P	%	$	P
Young[a]	12.8	-.365	.15	.6	.054	.42	6.9	.30	.43
Old[b]	-19.5	-.46	.27	-2.8	-.097	.48	7.1	.60	.25

[a]At most 30 years old.

[b]At least 45 years old, not retired.

[c]Proportion of quits in each category.

Source: Bartel and Borjas (1981), p. 68, Table 2.1

Table 9

Interaction Effects of Separation and Changes in Hours on Wage Gains (1970-1981)

(A) Wage and Salary Earners

	All	Young	Older
Short Run Gains	-.012 (1.0)	n.s.	-.04 (2.1)
Job Gains	-.016 (1.1)	n.s.	-.042 (2.1)
$\Delta H > 5^a$	n.s.	n.s.	-.057 (2.3)
$\Delta H < -5^a$	-.024 (1.5)	-.029 (1.1)	-.028 (1.7)

(B) Wage Earners Only

	All	Young	Older
Short Run Gains	-.022 (1.5)	n.s.	-.058 (2.9)
Job Gains	-.026 (1.7)		-.061 (2.9)
$\Delta H > 5^a$	-.026 (1.4)	n.s.	-.076 (2.9)
$\Delta H < -5^a$	-.026 (1.5)	n.s.	-.047 (1.8)

[a]In job gain equations.