# JOB MOBILITY AND GENDER-BASED WAGE GROWTH DIFFERENTIALS

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*Studies of gender differences in the returns to job mobility have yielded conflicting results. We examine whether there are gender differences in mobility patterns or in the returns to different types of mobility. Our results, based on the National Longitudinal Survey of Youth, imply that there* are *gender differences in mobility patterns, but there are* not *gender differences in the wage growth associated with different types of mobility. Therefore, it appears that empirical estimates of the gender differences in the returns to job mobility may be misleading if they do not consider the cause of separation.* (*JEL* J16, J63)

## I. INTRODUCTION

Evidence presented by Goldin [1989], O'Neill [1985] and others shows that the gender wage gap is narrowing. Much of the narrowing has been attributed to changes in female employment behavior such as the increased consistency of labor force participation and investment in education. However, because wages are not yet equal, questions remain. For example, does mobility play a role in the remaining wage gap? In this paper we examine the role of mobility, because this has been a relatively neglected area of interest for researchers concerned about the gender wage gap.

Some evidence on gender differences in mobility and the returns to mobility has been presented, but there is little consistency in the reported evidence. Keith and McWilliams [1995] demonstrate that the mobility histories of young men and women differ significantly. Loprest [1992] reports that, for a sample of individuals who worked full time for four consecutive years, young men's wages grew twice as much when they changed jobs (i.e., employers) as young women's did. However, Abbott and Beach [1994] find that the wage effects of mobility for women are similar in magnitude to the gains that Mincer [1986] and Holmlund [1984] find for men. Given the sometimes conflicting results reported so far, determining the extent and the cause of gender differences in wage-mobility effects has important implications for gender wage gap arguments.

In this study we examine the patterns and the returns to mobility for young men and women. To examine mobility patterns we disaggregate mobility into involuntary, or employer-initiated, and voluntary, or employee-initiated, separations. We further disaggregate involuntary separations into layoffs and discharges and we further disaggregate voluntary separations into family-related and non-family-related quits. By examining the impact of each of these different types of job separation on wage growth, and the difference in the relative frequency of each type of separation for men and women, we can make inferences about the impact of early career mobility on the gender wage gap.

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**ABBREVIATIONS**

AFQT: Armed Forces Qualification Test
IV: Instrumental variable
NLSY: National Longitudinal Survey of Youth
OLS: Ordinary least squares



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The remainder of this paper is organized as follows. In section II we review evidence on the returns to different types of mobility. In section III we present evidence on gender differences in mobility and wage growth. We use data from the National Longitudinal Survey of Youth (NLSY) to examine the mobility patterns of men and women and to calculate average wages and average wage growth. In section IV we describe how we selected a sample and specified a model for estimating the determinants of wage growth. In section V we examine the wage growth effects of different types of mobility, controlling for starting wages and individuals' fixed effects. In section VI, we discuss our conclusions.

## II. THE RETURNS TO DIFFERENT TYPES OF MOBILITY

Workers change jobs for different reasons, some voluntary and some involuntary. There is some evidence that the returns to different types of mobility differ. Not surprisingly, there is consistent evidence that employee-initiated separations (quits) increase wage growth, and that employer-initiated separations (layoffs and discharges) decrease wage growth, relative to staying with the same employer.[1] Additional evidence on the job mobility and wage growth of men has been offered as well. Bartel and Borjas [1981] show that job-related quits have a larger positive impact on wage growth than do non-job-related quits, with the highest wage growth resulting from those quits that are associated with successful job shopping. Parsons [1989] shows that, for young men, an employer-initiated separation with prejudice (a discharge) results in a larger wage penalty than an employer-initiated separation without prejudice (a layoff). Keith [1993] and Keith and McWilliams [1995] show that the number of previous discharges decreases, and the number of previous job-related quits increases, the current wage for samples of young men.

Mincer [1986] provides an explanation for the differential impact of employer-initiated separations (layoffs) and employee-initiated separations (quits). Mincer argues that employee-initiated separations are more likely to be associated with on-the-job or preseparation job search, while employer-initiated job separations are more likely to occur exogenously, and therefore are not likely to be associated with preseparation job search.[2] Since the reservation wage of on-the-job searchers will exceed their current wage, while the reservation wage of unemployed searchers is likely to be lower, Mincer concludes that employee-initiated separations (quits) should be associated with higher average wage growth than employer-initiated separations (layoffs).[3]

Mincer's argument can also be used to explain Parsons' evidence that the wage penalty to discharges is larger than the wage penalty to layoffs. The larger wage penalty from a discharge could reflect that the individuals who were laid off had access to better information about the likelihood of being let go than those who were fired and as a result were able to search for a new job while still employed. Parsons offers another explanation. If a discharge provides an adverse signal to potential employers, then acting on this signal the employer may penalize a previously discharged individual with a lower wage offer.

Why different types of employee-initiated separations (quits) might have a different impact on wage growth is more difficult to explain. Separating quits by motivation, as Bartel and Borjas [1981] do, may help. They suggest that quits can be separated into two types. The first type is job related. These quits result from some dissatisfaction with the current job, such as inadequate pay, poor working conditions, suboptimal hours, poor job match, etc. The second type of quit is not job related. These quits result from some non-job-related reason such as a spouse being transferred, a family member becoming ill, the worker having an additional child, etc. Mincer's [1986] argument concerning preseparation job search can be used to understand the differential im-

1. For studies of the wage effect of a quit, a layoff, or both see, for example, Blau and Kahn [1981a; 1981b], Flinn [1986], Ruhm [1987], Antel [1986; 1991] and McLaughlin [1990; 1991].

2. Studies that discuss the theoretical relationship between turnover and wage growth include Burdett [1978], Johnson [1978], and Jovanovic [1979].

3. McLaughlin [1990; 1991] offers an alternative explanation of the empirical differences between a quit and a layoff. He uses an efficient turnover model to show that the quit/layoff distinction arises out of a censoring of wage revisions. That is, if a separation results, the side initiating the wage revision determines the turnover label.

**TABLE I**
Men's and Women's Layoff, Discharge and Quit Rates

| | Men | | Women | | $t^a$ |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| *Layoff rate* | 33.2% | (0.002) | 25.9% | (0.002) | 25.8*** |
| *Discharge rate* | 6.0% | (0.001) | 4.5% | (0.001) | 10.6*** |
| *Family-related quit rate* | 1.1% | (0.001) | 8.3% | (0.001) | 50.9*** |
| *Non-family quit rate* | 59.7% | (0.002) | 61.3% | (0.003) | 4.4*** |
| *n* | 40,549 | | 34,748 | | |

Data source: The National Longitudinal Survey of Youth, 1980 to 1991 [Center for Human Resource Research 1992].

*Notes:* The standard errors are in parentheses. The layoff rate does not distinguish between temporary and permanent layoffs. Although the exact number of respondents varies from year to year, the sample contains approximately 4,545 women who had changed employers, on average, 6.21 times (SE = 0.062) and 4,466 men who had changed employers, on average, 7.24 times (SE = 0.070).

***Men/women difference is statistically significant at a 1% level (two-tailed tests).

[a] $t$ values are in absolute terms.

pact of these different types of quits on subsequent wages. If personal, or non-job-related, quits are more likely to be due to unexpected random events such as a family member's illness, the individuals who quit for these kinds of reasons may be less likely to search for a new job while still employed. Non-job-related quits would then result in lower wage growth than job-related quits—in which the individuals engaged in preseparation job search.

These arguments provide us with a rationale for analyzing the returns to the different types of mobility. Because we are interested in the possible effect of mobility on the gender wage gap, it is important that we include women in our analysis, that we disaggregate mobility by type, and that we look for possible differences in mobility patterns, as well as returns.

## III. EARLY JOB MOBILITY AND WAGE GROWTH

### *Gender Differences in Job Mobility*

To determine whether young men and women have different job mobility patterns, we identified four types of mobility decisions using information from the NLSY data files [Center for Human Resource Research 1992]. Since 1979 the NLSY panel study has collected information on approximately 12,000 young men and women who were 14 to 22 years of age in 1979. Data are collected on as many as five different jobs per year. During the annual interview, for every job separation since their last interview, respondents are asked "Which of the reasons on this card best describes why you happened to leave this job?" The possible responses to this question for the years 1980 through 1991 were:

1. a layoff, a plant closing, the end of a temporary/seasonal job, or the [job training] program ended,
2. discharged or fired,
3. quit for pregnancy/family reasons,
4. quit for other reasons.

The primary distinction between a layoff and a discharge, based on the United States Department of Labor [1971] classifications, is that a discharge implies that the terminated employee's performance was incompetent or his or her behavior was malfeasant, while a layoff occurs for exogenous (i.e., beyond the control of the employee) circumstances. Because of the wording of the responses, it is not possible to identify which quits were job related, let alone which ones were motivated by job dissatisfaction and which ones were motivated by successful job shopping. We can, however, distinguish between family-related and non-family-related (other) quits.

Table I presents the layoff, discharge, and family- and non-family-related quit rates calculated as a percentage of the total number of reported job (employer) changes between 1980 and 1991. The rates for the different types of job separation are in column 1 for the men and in column 3 for the women. An examination of the different separation rates shows that both men and women were much more likely to have reported an employee-initiated than an employer-initiated job separation. However, there were differences between men and women.

Men were more likely than women to report an employer-initiated separation, and less likely than women to report an employee-initiated separation. For example, women were twice as likely to report a family-related quit as a discharge, while men were more likely to report a discharge than a family-related quit. Although family-related quits are a small percentage of the total number of reported reasons for job change for both men and women, it was seven times more likely that a women quit for a family-related reason than it was that a man did so.

*Gender Differences in Wage Growth*

After establishing that young men and women had different separation patterns, we examined wage growth for each group using a second sample from the NLSY. This sample was limited to new entrants who entered the labor market between 1979 and 1988. We restricted our sample to new entrants to reduce the gender-correlated sample selection bias.[4] In this analysis we examined only the average annual wage growth and wage levels for our sample to determine if the different types of mobility are associated with differences in wage growth. And, although we restricted our sample to new entrants, we allowed for multiple observations of annual wage growth for these individuals. Thus, the resulting sample also includes the wage growth of individuals who had been in the labor market for several years.

Table II includes the annual average wage growth for our sample. We calculated average wage growth as the difference between the respondent's wage at $\tau+1$ and at $\tau$, where $\tau = t, t+1, ..., t+12$, and $t$ refers to the year in which the respondent entered the labor market. Column 1 contains the men's average wage growth, and column 4 contains the women's. The results indicate no significant gender difference in the average annual wage growth for all workers, a subsample of individuals who changed employers (the movers), or a subsample of individuals who did not change employers (the stayers). On average, between 1979 and 1991, the young men and women in this sample experienced real wage growth of 4% to 5% per year.

We next examined whether the "reason for job mobility" results in differential effects on wage growth. To do this, we first disaggregated mobility into employer- and employee-initiated separations. Second, we further disaggregated employer-initiated separations into layoffs and discharges, and we disaggregated employee-initiated separations into family-related and non-family-related quits. Previous studies have not disaggregated these variables to determine whether different types of mobility have different impacts on wage growth. This is unfortunate, because the aggregation could be masking important differences in the returns to different types of mobility.

An examination of the wage growth following voluntary and involuntary mobility in Table II shows that, for both men and women, the wage growth from involuntary mobility was not significantly different from zero, the wage growth associated with no mobility was approximately 4% to 5%, and that the wage growth associated with voluntary mobility was greater than 5%. Additionally, when mobility was disaggregated only into voluntary and involuntary separations, we found that the men's wage growth from voluntarily changing employers was 35% higher than the women's.

By disaggregating voluntary and involuntary mobility, we discovered more about gender differences. An examination of the average wage growth from involuntary mobility shows that the returns to layoffs are not significantly different from zero for men or women. However, while the return to discharges is not significantly different from zero for men, women suffer a wage penalty of about 8% for discharges. There was less gender difference in the returns to voluntary mobility. For both men and women, the wage growth from a non-family-related quit was 8% to 9% higher than the wage growth associated with a family-related quit. Since the data revealed no gender differences in the returns to

4. Since both male and female new entrants will have had little labor market experience or tenure, but similar levels of education, they should have similar human capital characteristics.

**TABLE II**
Average Annual Wage Growth (in natural logs) from 1979 to 1991

| | Men | | | Women | | | |
|---|---|---|---|---|---|---|---|
| | Mean (1) | SE (2) | n (3) | Mean (4) | SE (5) | n (6) | $t^a$ (7) |
| Average annual wage growth | | | | | | | |
| *All workers* | 4.7% | (0.003) | 26,889 | 4.2% | (0.003) | 24,889 | 1.18 |
| *Stayers* | 4.5% | (0.004) | 16,513 | 4.0% | (0.004) | 15,979 | 0.88 |
| *Movers* | 4.9% | (0.006) | 10,215 | 4.7% | (0.006) | 8,910 | 0.24 |
| Average annual wage growth from job change | | | | | | | |
| *Involuntary movers* | –1.1% | (0.009) | 3,830 | 0.2% | (0.012) | 2,290 | 0.87 |
| *Voluntary movers* | 8.5% | (0.007) | 6,385 | 6.3% | (0.007) | 6,620 | 2.22** |
| Average annual wage growth from disaggregated involuntary and voluntary job change | | | | | | | |
| *Layoffs* | –0.9% | (0.010) | 3,076 | 2.1% | (0.014) | 1,854 | 1.74* |
| *Discharges* | –1.8% | (0.020) | 754 | –8.1% | (0.023) | 436 | 2.07** |
| *Family-related Quits* | –5.0% | (0.049) | 114 | –2.7% | (0.025) | 749 | 0.42 |
| *Non-family-related Quits* | 8.7% | (0.007) | 6,271 | 7.5% | (0.008) | 5,871 | 1.13 |

Data source: The National Longitudinal Survey of Youth, 1980 to 1991 [Center for Human Resource Research 1992].

*Notes:* Since the sample consists of multiple observations of men and women who entered the labor market between 1979 and 1988, *n* refers to the total number of wage change observations and not the number of individuals. The following wage growth point estimates are not significantly different from zero at a 10% level: men and women's wage growth from involuntary mobility and layoffs, and the men's wage growth from discharges and family quits. All wages have been converted to real wages using the Consumer Price Index for All Urban Consumers (CPI-U, 1982–84=100).

*Men/women difference is statistically significant at a 10% level and **statistically significant at a 5% level (two-tailed tests).

[a] $t$ values are in absolute terms.

family- and non-family-related quits, we conclude that the gender difference in the return to (aggregated) voluntary mobility may be due to women's greater proportion of family-related quits. That is, it appears that the type of mobility, rather than any gender difference in the wage effect of mobility, was responsible for the difference in the aggregated employee-initiated mobility variable means.

## IV. THE DETERMINANTS OF WAGE GROWTH

Having established that the men and women in our sample have different mobility profiles, but similar average annual wage growth, we were interested in examining the determinants of wage growth for both men and women, and especially the impact of mobility on wage growth. Therefore, we estimated wage growth models in which we included the mobility variables, as well as the standard human capital and individual and job characteristic variables.

### *The Sample*

To conduct our analysis of wage growth, we constructed a third sample from the NLSY files. In order to be included in this sample a respondent had to meet two criteria. The first criterion was that the respondent had to enter the labor market between 1979 and 1988. The second criterion was that the respondent had to have two wage observations—one in the year in which (s)he entered, and the other three years later. We defined labor market entry as having been enrolled in high school or college the year prior to entry, but out of school and in the labor market the following year.

The first criterion reduced the sample by 34%. Reasons for not classifying an individual as a new entrant included: entered the labor market prior to 1979 (25%) or after 1988 (2%), and missing observations on school enrollment or labor market status (7%). The second criterion reduced the sample by an additional 21%, which is, in part, a result of the data limitations of the NLSY. Unfortunately the NLSY does not make available wage information on all jobs. Jobs that were either part-time or lasted less than nine weeks may have no wage observation. Therefore, it is possible to classify a respondent as a new entrant in the labor market, but not have any corresponding information on wages. Finally, missing observations on the other variables reduced the sample by another 9%.[5] The final sample contains 4,575 individuals of whom 49% are women.

For this analysis we did not include multiple observations on an individual. For each respondent who met both of our criteria, we looked at the wage growth between the respondent's initial wage (wage in year $t$) and his or her wage three years later (wage in $t+3$). For the mobility variable we used the reason for quitting the initial job.

Descriptions of the variables used in this analysis are presented in Appendix A. Table III presents selected descriptive statistics for the sample used. The first and third columns contain the sample means for the men and the women. There are two things worth noting. First, while the differences in men's and women's sample means are significant, most are not very large. Second, the job mobility means are similar to the ones presented in Table I.

Instead of examining turnover activity over a given period such as a year (as we did in Table II), we examined the turnover on a single job, because this was the most efficient way of matching an initial wage observation with both the reason for the job change and the subsequent wage. We first found the job that the respondent identified as his or her current or most recent job at the end of his or her first year in the labor market, i.e., the initial job, and the wage associated with this job. We then tracked the respondent's jobs at the end of each of the next three years to determine if the respondent changed jobs during the three-year period. If the respondent did change jobs, we recorded the reason for the job change. We also recorded the respondent's wage at the end of the three-year period. Descriptive statistics on this information are in Table III.

*The Specification of the Wage Growth Model*

Because in this analysis we examine the wage growth of groups that are likely to have different starting wages (e.g., union and non-union workers) and because the parameter values for new entrants could change over the wage growth period, it is important to examine the extent of these potential problems in a standard wage growth model. We do so by considering the following wage growth specifications:

$$(1) \qquad \Delta \log W_i = \beta \Delta X_i + \varepsilon_i^*,$$

and

$$(2) \qquad \Delta \log W_i = \delta \log W_{i,t} + \beta \Delta X_i + e_i^*,$$

where $W_{i,t}$ is the starting wage in the initial period for individual $i$; $X_i$ is a vector of control characteristics for individual $i$; and $\varepsilon_i^*$, and $e$ are random elements with the usual properties (i.e., where $E(\varepsilon_i^*) = 0$, $E(e_i^*) = 0$, $V(\varepsilon_i^*) = \sigma_\varepsilon^2$, and $V(e_i^*) = \sigma_u^2$). Equation (1) is a standard wage growth model. Equation (2) follows from equation B.10 in Appendix B. Equation (2) is not a standard wage growth model unless the coefficient for starting wages, ($\delta$), is equal to zero.

There are two potential problems with a standard wage growth model (equation 1). The first is that it does not control for the size of the first-period (i.e., base) wage rate. The second is that it imposes the restriction that the parameters of the independent variables remain constant over the wage growth period. For the first, if starting wage levels are likely to differ between different groups, not controlling for starting wages could mask the effect of a variable, such as union coverage, on wage growth. For example, in our sample women in unionized jobs had starting wages that were approximately 11% higher than women in non-unionized jobs (men in unionized jobs had wages that were approximately 19% higher than men in non-unionized jobs). Using the standard wage growth equation (equation 1), an identical percentage increase in wages for these women would be indicated

5. Missing observations occur because (a) the respondent was not surveyed or (b) the respondent was surveyed but either would not or could not answer the survey question.

**TABLE III**
Selected Sample Means and Standard Errors

| | Men | | Women | | |
|---|---|---|---|---|---|
| | Mean (1) | SE (2) | Mean (3) | SE (4) | $t^a$ (5) |
| *Hourly wage at t* | 5.71 | (0.072) | 4.94 | (0.058) | 8.33*** |
| *Hourly wage at t + 3* | 7.04 | (0.079) | 5.94 | (0.061) | 11.02*** |
| *Wage growth* | 20.4% | (0.010) | 17.0% | (0.011) | 2.29** |
| *Percentage who left their first CPS job* | 75.7% | (0.009) | 75.2% | (0.009) | 0.39 |
| *Reason left first CPS job* | | | | | |
| *Laid off* | 20.7% | (0.008) | 13.1% | (0.007) | 7.15*** |
| *Discharged* | 4.7% | (0.004) | 3.3% | (0.004) | 2.47** |
| *Family-related quit* | 0.5% | (0.001) | 3.8% | (0.004) | 8.00*** |
| *Non-family-related quit* | 49.8% | (0.010) | 55.0% | (0.011) | 3.50*** |
| *Received diploma or degree, t to t + 3* | 9.4% | (0.006) | 9.8% | (0.006) | 0.47 |
| *Change in years of experience, t to t + 3* | 2.79 | (0.019) | 2.52 | (0.018) | 10.32*** |
| *Change in years of tenure, t to t + 3* | 1.12 | (0.032) | 1.14 | (0.032) | 0.44 |
| *Job characteristic change, t to t + 3* | | | | | |
| *Part-to-full time* | 17.0% | (0.008) | 20.3% | (0.008) | 2.92*** |
| *Full-to-part time* | 6.4% | (0.005) | 11.4% | (0.007) | 5.81*** |
| *Full-to-full time* | 72.6% | (0.009) | 59.2% | (0.010) | 9.96*** |
| *Part-to-part time* | 4.0% | (0.004) | 9.1% | (0.006) | 7.07*** |
| *Union to non-union* | 9.9% | (0.006) | 8.8% | (0.006) | 1.30 |
| *Non-union to union* | 11.3% | (0.007) | 9.3% | (0.006) | 2.17** |
| *Union to union* | 7.1% | (0.005) | 5.4% | (0.005) | 2.40** |
| *Non-union to non-union* | 71.7% | (0.009) | 76.5% | (0.009) | 3.77*** |
| *n* | 2,330 | — | 2,245 | — | |

Data source: The National Longitudinal Survey of Youth, 1980 to 1991 [Center for Human Resource Research 1992].

*Men/women difference is statistically significant at a 10% level, **statistically significant at a 5% level and ***statistically significant at a 1% level (two-tailed tests).

[a] *t* values are in absolute terms.

only if the wages of the women in unionized jobs increased more in dollar terms than the wages of the women in non-unionized jobs. Therefore, using the standard wage growth equation would mislead us about the effect of union coverage (and other variables) on wage growth.

The second problem, that a standard wage growth model restricts the parameter values to remain constant over the wage growth period, is the same as restricting the coefficient on starting wages to equal zero. (See Appendix B for details.) If the parameter values are not constant over the two periods, the coefficient on starting wages is not zero. It seems likely that the parameter values may not be constant over time. That is, it seems likely that the effects of the independent variables could change as the new entrants continue to participate in the labor market. For example, the influence of years of education on wages may be greater in period $t$ than in $t+3$ because in $t$, by construction, education is the sample's primary type of human capital. However, by period $t+3$ the workers have had time to invest in job training and, therefore, education might no longer be the primary type of human capital.[6]

6. To determine if the parameter values *did* change over the sample periods, we estimated separate wage-level equations in $t$ and $t+3$ for separate samples of men and women. These wage-level equations included age, a measure of IQ, years of education, years of experience, years of tenure, union status, full-time work status, marital status, area of residence, and local unemployment rates. The point estimates for the human capital variables, except tenure and education (for men only), and for the union status and full-time work status variables were lower in $t+3$ than they were in $t$. Therefore, we conclude that the parameters were not constant for our sample.

To choose the most appropriate model, we estimated both. Using an $F$ test we found that we could reject the hypothesis that the coefficient on starting wages in model 2 is zero, which suggests that the coefficients are not constant over the wage growth period. Since we also have evidence that starting wages are significantly different for unionized and non-unionized workers of both sexes, we concluded that the second specification, equation 2, was the more appropriate specification for our analysis.[7]

## V. THE WAGE GROWTH EFFECTS OF JOB MOBILITY

### *The Determinants of Young Men's and Young Women's Three-Year Wage Growth*

To determine the impact of job mobility on wage growth for a three-year period, we estimate the following equation:

$$(3) \qquad \Delta \log W_i = \delta \log W_{i,t} + \beta \, \Delta X_i + \gamma R_i + e_i^*,$$

where $R_i$ is the vector of dummy variables that indicate the reason for a job change between $t$ and $t+3$, $e_i^* = \varepsilon_{i,t+3} - (1+\delta)\varepsilon_{i,t}$ with the usual properties (i.e., $E(\varepsilon_{i,t}) = 0$, $E(\varepsilon_{i,t+3}) = 0$, $V(\varepsilon_{i,t}) = \sigma_1^2$, $V(\varepsilon_{i,t+3}) = \sigma_2^2$, $E(\varepsilon_{i,t}, \varepsilon_{i,t+3}) = 0$) and $\delta$ is a proportionality factor (see Appendix B).

Using the results from a Chow test ($F(30, 4574) = 2.71$), we rejected the hypothesis of a similar structure to men's and women's wage equations at a two-tailed, 5% level. Therefore, we estimated separate models for the men and the women. Although we were primarily interested in the effect of the "different reasons for mobility" variables on wages, we estimated two models that excluded these variables. First, because the mobility variables are likely to be correlated with some of the other explanatory variables, we estimated model 1 without including any mobility variables.[8] Second, to determine if disaggregating mobility gives us any additional information, we estimated model 2, including just a single mobility variable, which was defined as one if the individual left his or her first CPS job, zero otherwise. Finally, in model 3 we disaggregated the single mobility variable into the four variables representing the different types of mobility.

The wage estimates for the three models are presented in Tables IV and V. Table IV contains the estimates of the three-year wage growth for men and Table V includes the estimates for women. Column 1 in both tables presents the results from the model in which the mobility variables were excluded; column 3 presents the results from the model in which the single mobility variable was included; and column 5 presents the results from the model that included the non-family-related quit, family-related quit, layoff and discharge variables.

There are several things worth noting. First, in model 2, when we included only a single variable for mobility, we could not reject the hypothesis that the coefficient of that variable was zero. But, in model 3, $F$-tests of the null hypothesis that the group of mobility variables' coefficients are zero could be rejected, for both men and women. Therefore, it is clear that aggregating mobility, as we did in model 2, masks the effects of the different types of mobility, as we suspected from the arguments in section II. This is evidence that it is important to disaggregate mobility when examining the wage growth effects.

Second, for the non-mobility variables, our results are consistent with earlier work. Change in education, change in years of experience, change in years of tenure, and change from non-union to union job all have a significant positive impact on wage growth for men and women. These results hold for all three models.

Third, some estimates *were* affected by the inclusion and the disaggregation of the mobility variables, however. Including the single mobility variable affected the significance of the tenure estimate for the women. Disaggregating the single mobility variable into the various reasons for mobility decreased the significance of the union-to-non-union estimates for the men. Disaggregating also affected the significance of the estimates for the increase-in-the-number-of-children and the

7. Parsons [1989] uses a similar specification.

8. For example, quitting for a pregnancy or other family-related reason will be highly correlated with having a baby. If the increased responsibility in the home caused the new parents to change jobs and accept lower wages in exchange for fewer or more flexible hours, then three variables—full to part-time employment, an increase in the number of children, and a family-related quit—will be highly correlated.

**TABLE IV**
OLS Determinants of the Three-Year Wage Growth of Young Men ($n$ = 2,330)

| | Model 1 | | Model 2 | | Model 3 | |
|---|---|---|---|---|---|---|
| | $\hat{\beta}$ (1) | *t*-value[a] (2) | $\hat{\beta}$ (3) | *t*-value[a] (4) | $\hat{\beta}$ (5) | *t*-value[a] (6) |
| *Left first CPS job* | — | — | –0.033 | (1.15) | — | — |
| *Reason left first CPS job* | | | | | | |
| *Laid off* | — | — | — | — | –0.090*** | (2.75) |
| *Discharged* | — | — | — | — | –0.155*** | (3.20) |
| *Family-related quit* | — | — | — | — | –0.333*** | (2.61) |
| *Non-family-related quit* | — | — | — | — | 0.006 | (0.20) |
| *Log wage at t* | –0.579*** | (31.00) | –0.581*** | (30.97) | –0.580*** | (31.13) |
| *Change in diploma or degree* | 0.075** | (2.54) | 0.076*** | (2.57) | 0.069** | (2.35) |
| *Change in years of experience* | 0.087*** | (8.35) | 0.086*** | (8.23) | 0.075*** | (7.04) |
| *Change in years of tenure* | 0.033*** | (5.47) | 0.027*** | (3.34) | 0.030*** | (3.75) |
| *Increase in the no. of children* | –0.019 | (0.88) | –0.017 | (0.82) | –0.012 | (0.57) |
| *Single to married* | 0.067*** | (3.12) | 0.067*** | (3.12) | 0.063*** | (2.95) |
| *Married to single* | –0.141* | (1.95) | –0.139* | (1.93) | –0.124* | (1.72) |
| *Union to nonunion* | –0.060** | (2.04) | –0.059** | (2.00) | –0.050* | (1.71) |
| *Nonunion to union* | 0.094*** | (3.43) | 0.096*** | (3.50) | 0.097*** | (3.55) |
| *Part-to-full time* | –0.051** | (2.13) | –0.048** | (2.00) | –0.051** | (2.15) |
| *Full-to-part time* | –0.030 | (0.83) | –0.028 | (0.78) | –0.029 | (0.79) |
| *Increase in unemployment rate* | –0.028 | (1.05) | –0.028 | (1.04) | –0.026 | (0.98) |
| *Decrease in unemployment rate* | –0.012 | (0.52) | –0.012 | (0.52) | 0.007 | (0.33) |
| *Urban to rural area* | –0.018 | (0.53) | 0.020 | (0.58) | 0.010 | (0.30) |
| *Rural to urban area* | 0.000 | (0.00) | –0.001 | (0.02) | 0.004 | (0.12) |
| *South to non-South area* | –0.032 | (0.54) | –0.031 | (0.52) | –0.033 | (0.55) |
| *Non-South to South area* | 0.113** | (2.25) | 0.116** | (2.30) | 0.111** | (2.21) |
| *Constant* | 0.927*** | (13.04) | 0.962*** | (12.41) | 0.991*** | (12.84) |
| *Adj. $R^2$* | 0.330 | — | 0.330 | — | 0.339 | — |

Data source: The National Longitudinal Survey of Youth, 1980 to 1991 [Center for Human Resource Research 1992].

*Notes:* The dependent variable is the difference between the log of the hourly wage at $t + 3$ and the log of the hourly wage at $t$, where $t$ refers to the year of labor market entry and $t$ = 1979 through 1988. All wages have been converted to real wages using the CPI-U (1982–84=100). Indicator variables for the year of labor market entry are included, but not reported.

*Estimate is statistically significant at a 10% level; **statistically significant at a 5% level; and ***statistically significant at a 1% level (two-tailed tests).

[a]$t$ values are in absolute terms.

moving from-a-part-time-to-a-full-time job variables for the women. These results are not surprising, because the estimates that were affected by the inclusion of the mobility variables were the ones that a priori were likely to be correlated with mobility. For example, by construction, an increase in the number of children is likely to be correlated with quitting for a family-related reason. And, if, as we suspect, layoff rates are higher in unionized jobs, moving from a union to non-union job may be correlated with a layoff. Similarly, tenure, as the sum of an individual's previous decisions not to change employers, is, by construction, correlated with the mobility variables.

The results from model 3 indicate some significant effects of mobility. Relative to staying with the same employer, men's wage growth was approximately 9% lower following a layoff, 16% lower following a discharge, and 33% lower following a family-related quit. Relative to staying with the same employer, women's wage growth was approximately 20% lower following a discharge and 17% lower following a family-related quit. We conducted $F$-tests to determine if the returns to the different types of mobility were significantly different. We found that, for both men and women, there was a significant difference between the returns to voluntary and involuntary mobility, and to the different

**TABLE V**

OLS Determinants of the Three Year Wage Growth of Young Women ($n$ = 2,245)

| | Model 1 | | Model 2 | | Model 3 | |
|---|---|---|---|---|---|---|
| | $\hat{\beta}$ | $t$-value[a] | $\hat{\beta}$ | $t$-value[a] | $\hat{\beta}$ | $t$-value[a] |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| *Left first CPS job* | — | — | −0.018 | (0.58) | — | — |
| *Reason left first CPS job* | | | | | | |
| *Laid off* | — | — | — | — | −0.042 | (1.09) |
| *Discharged* | — | — | — | — | −0.200*** | (3.33) |
| *Family-related quit* | — | — | — | — | −0.173*** | (2.97) |
| *Non-family-related quit* | — | — | — | — | 0.002 | (0.07) |
| *Log wage at t* | −0.635*** | (30.39) | −0.637*** | (30.27) | −0.636*** | (30.39) |
| *Change in diploma or degree* | 0.075** | (2.35) | 0.075** | (2.37) | 0.071** | (2.23) |
| *Change in years of experience* | 0.135*** | (10.46) | 0.135*** | (10.41) | 0.127*** | (9.73) |
| *Change in years of tenure* | 0.021*** | (3.10) | 0.017* | (1.91) | 0.017* | (1.91) |
| *Increase in the no. of children* | −0.050** | (2.05) | −0.050** | (2.06) | −0.035 | (1.41) |
| *Single to married* | −0.004 | (0.19) | −0.004 | (0.19) | 0.005 | (0.23) |
| *Married to single* | −0.075 | (1.42) | −0.074 | (1.40) | −0.072 | (1.38) |
| *Union to nonunion* | −0.026 | (0.80) | −0.026 | (0.79) | −0.025 | (0.77) |
| *Nonunion to union* | 0.084*** | (2.60) | 0.085*** | (2.62) | 0.088*** | (2.75) |
| *Part-to-full time* | −0.045* | (1.87) | −0.043* | (1.80) | −0.048** | (2.00) |
| *Full-to-part time* | −0.020 | (0.67) | −0.020 | (0.64) | −0.018 | (0.58) |
| *Increase in unemployment rate* | −0.022 | (0.79) | −0.022 | (0.78) | −0.024 | (0.87) |
| *Decrease in unemployment rate* | 0.016 | (0.67) | 0.017 | (0.69) | 0.016 | (0.67) |
| *Urban to rural area* | 0.034 | (0.90) | 0.034 | (0.89) | −0.037 | (0.97) |
| *Rural to urban area* | 0.021 | (0.56) | 0.022 | (0.60) | 0.020 | (0.53) |
| *South to non-South area* | −0.055 | (0.85) | −0.053 | (0.81) | −0.061 | (0.93) |
| *Non-South to South area* | −0.064 | (1.18) | −0.063 | (1.17) | −0.071 | (1.31) |
| *Constant* | 0.884*** | (10.93) | 0.903*** | (10.33) | 0.923*** | (10.58) |
| *Adj.* $R^2$ | 0.315 | — | 0.315 | — | 0.322 | — |

Data source: The National Longitudinal Survey of Youth, 1980 to 1991 [Center for Human Resource Research 1992].

*Notes:* The dependent variable is the difference between the log of the hourly wage at $t + 3$ and the log of the hourly wage at $t$, where $t$ refers to the year of labor market entry and $t$ = 1979 through 1988. All wages have been converted to real wages using the CPI-U (1982–84=100). Indicator variables for the year of labor market entry are included, but not reported.

*Estimate is statistically significant at a 10% level; **statistically significant at a 5% level; and ***statistically significant at a 1% level (two-tailed tests).

[a]$t$ values are in absolute terms.

types of voluntary and involuntary mobility, with the exception of layoffs and discharges for men, for which the difference was only marginally (at the 13% level) significant. This is evidence that mobility does have an impact on early wage growth and that the type of mobility matters. However, because there is no statistical difference in the magnitude of the effects for men and women, there is no evidence of gender effects to mobility.

### *Estimation Concerns*

We were concerned about two possible sources of bias in the OLS estimates in our analysis. Our first concern was that the sample restriction of wage observations in both $t$ and $t+3$ caused us to inadvertently exclude those individuals with the worst post-mobility wage prospects (relative to their reservation wage) from the sample.[9] For example, an increase in the number of children in the household may increase an individual's reservation wage. Those individuals whose subsequent reservation wage is higher than their post-separation wage offers will exit the labor market. They would not, therefore, be included in our sample because they have no wage observation in

9. We are indebted to an anonymous referee who pointed this out.

period $t+3$. This may affect the estimate of the coefficient for family-related quits because family-related quits and an increase in the number of children in the household are correlated. If the individuals who exit the market following a family-related quit are also those with the worst wage prospects, the family-related quit estimate would be upwardly biased. Because women are seven times more likely than men to quit for a family-related reason, this bias might have gender implications. Other mobility variables might be affected similarly.

To address this, we first compared the quits, layoff and discharge rates of a subsample of respondents who had a wage observation in $t$ but not in $t+3$ with the mobility rates presented in Tables I and III. We found that those individuals who did not have a wage observation in $t+3$ were more likely to have left their initial job for a type of mobility associated with the largest wage losses. Both the men and women in this subsample were more likely to have been discharged and the women (only) were more likely to have quit for a family-related reason. Because those without wage observations in period $t+3$ have different mobility patterns than those in our sample, there is a potential selectivity bias. To control for this possibility, we used the standard Heckman [1979] two-stage technique to examine the extent of sample selection bias in wage levels for the periods $t$ and $t+3$.[10]

The Heckman two-stage technique produced the following results. First, for separate samples of men and women, the inverse Mills ratio estimates were insignificant at the 5% level for both periods. However, in period $t$ the inverse Mills ratio estimate was significant at the 10% level for the women. The only estimate that was significantly different for women in period $t$ was for the "married" variable. This is understandable since having a spouse, especially an employed one, could be expected to affect the labor force participation decision. Second, including the inverse Mills ratio as a right-hand side variable in the men's and women's wage equations affected the level of significance of men's and women's experience and tenure estimates; men's unemployment rate estimates; and women's age, martial status, and non-family-related quit estimates.[11]

The weak evidence we have of sample selectivity for the women may indicate that women with the worst wage prospects, relative to their reservation wage in period $t$, are not in our sample. This implies that our sample may not be representative of all young women. The negative effect on the married estimate may be evidence that women with an employed spouse have higher reservation wages than women in general. However, the lack of evidence of selectively bias in period $t+3$ implies that there should be no impact on the mobility variables.

Our second concern is with the assumption that the starting wage and the error term are uncorrelated. If the starting wage is correlated with the error term, our estimates may suffer from simultaneous equations bias. To address this concern we estimated a variant of equation (3) in which we treated the starting wage as endogenous and, therefore, predicted by an instrumental variable. We used the predicted value of the starting wage as our instrument.[12]

The instrumental variable (IV) technique produced the following results. In the IV estimation for men, the non-family-related-quit estimate was now significant and statistically greater than it had been in the OLS estimation. In the IV estimation for women, the layoff was now significant; however it was not statistically different in magnitude from the OLS

10. In the first step we estimated probit labor force participation (LFP) equations for samples of men and women in periods $t$ and again in $t+3$. In both years the LFP equations' dependent variable equaled one if the individual had a wage observation in both $t$ and $t+3$, zero otherwise. Variables used in the first-stage probit included: age, Armed Forces Qualification Test (AFQT) score, years of education, years of experience, years of tenure, white, married, number of children, union status, urban resident, location of residence (east, central, south), five unemployment rates, non-wage income, a health-limits-work variable and the four mobility variables—layoff, discharge, family-related quit, non-family-related quit (for period $t+3$ only). Independent variables for the second-stage log wage equation included: age, married, number of children, AFQT score, white, years of education, years of experience, years of tenure, union status, location of residence (east, central, south), five unemployment rates, full-time status and the four mobility variables (for period $t+3$ only).

11. Tables with these estimates are available from the authors.

12. We estimated an equation similar to B.10 in Appendix B. The independent variables used to predict the starting wage include: AFQT score, white, years of education, union status, urban resident, location of residence, four unemployment rates, ten dummy variables for occupation, nine dummy variables for year of entry, and full-time status.

estimate. Therefore, the only evidence of simultaneous equations bias for the mobility estimates was in the non-family-related quit estimate for the men.[13]

## VI. CONCLUSIONS

This study examined gender differences in job mobility patterns and returns for a sample of new entrants in the labor market. We found significant gender differences in mobility patterns for our sample. Men were more likely to separate for an employer-initiated reason, i.e., a layoff or a discharge. Women were more likely to separate for an employee-initiated reason, i.e., a family- or a non-family-related quit. With the exception of family-related quits, the gender differences in job separations were not large. Therefore, the most significant difference in mobility patterns is associated with family-related quits. Although neither gender in our sample had quit very often for a family-related reason, women were seven times more likely than men to have done so. Although it might appear that there are gender differences in both behavior (as evidenced by women's greater proportion of family-related quits) and treatment (as evidenced by men's greater proportion of layoffs and discharges), it is not possible to determine the true extent of these differences without extending the analysis to include demand-side data.

The results of this study also reveal information about the returns to mobility. Unlike the large gender difference in wage growth that Loprest [1992] found, our wage growth means revealed no gender difference in the wage growth from mobility. However, disaggregating the wage growth into employee-initiated and employer-initiated wage growth showed that the men's wage growth from quitting (employee-initiated) was 35% higher than the women's. Further disaggregation of employee-initiated wage growth into family- and non-family-related quits showed that the return to family-related quits was significantly lower than the return to non-family-related quits for both men and women. Because women were seven times more likely to engage in family-related quits, we conclude that the gender differences in the (aggregated) employee-initiated wage growth means is most likely due to women's greater proportion of family-related quits.

Because this study examined the determinants of an *individual's* current wages and not family income, it is not clear that the individuals who quit for family-related reasons also suffered a reduction in their family income. If the post-quit wage rate is simply a poor measure of the net benefits from quitting for a family-related reason, using a better measure, such as family income, may reveal that family-related quits result in positive net benefits. One interpretation of men's relative wage gains from regional migration is that the relevant analysis may be family mobility patterns and income growth rather than individual mobility patterns and wage growth.

With regard to the gender wage gap, our results show no gender difference in the returns to (aggregated) mobility for young men and women. Our findings may differ from previous findings, such as those reported by Loprest [1992], because we used a different wage specification and because we did not impose strict work-continuity restrictions on our sample. Our results would seem to indicate that, for those individuals who do not stay out of the labor market for extended periods of time, mobility does not affect the gender wage gap, either positively or negatively.

13. There were additional non-mobility variables, for both men and women, that were significant in the IV, but not the OLS estimation. Tables with these results are available from the authors.

## APPENDIX A

Variable Definitions

| | |
|---|---|
| *Continuous Variables* | |
| log $Wt$ | The hourly (log) wage at $t$ in 1982–1984 dollars, where $t$ is the year that respondent entered the labor market, $t$ = 1979, ... 1988. |
| log $Wt$ + 3 | The hourly (log) wage at $t$ + 3 in 1982–1984 dollars. |
| Age at $t$ | The respondent's age on January 1 of the year that he or she entered the labor market. |
| Years of education at $t$ or $t$ + 3 | Years of education one or three years after the respondent entered the labor market. |
| Change in experience | The difference in the number of actual hours worked between $t$ and $t$ + 3. |
| Tenure at $t$ | Years with the respondent's initial CPS employer at $t$. |
| Tenure at $t$ + 1 | Years with current employer at $t$ + 3. |
| No. of children at $t$ | Number of children that the respondent had in the year he or she entered the labor market. |
| No. of children at $t$ + 3 | Number of children that the respondent had three years later. |
| No. of hours per week at $t$ | The number of hours worked per week at the respondent's initial CPS job. |
| No. of hours per week at $t$ + 3 | The number of hours worked per week at the respondent's current job. |
| *Dummy Variables (equal to 1 if the respondent is)* | |
| Married at $t$ or $t$ + 3 | Was married, spouse present. |
| Union at $t$ or $t$ + 3 | Wages were set by collective bargaining. |
| Received a higher degree | Received a diploma or a degree between $t$ and $t$ + 3. |
| Increase in unemployment rate | The local unemployment rate increased between $t$ and $t$ + 3. |
| Decrease in unemployment rate | The local unemployment rate decreased between $t$ and $t$ + 3. |
| Full-time to part-time | Worked more than 34 hours per week at $t$ but less than 35 hours per week at $t$ + 3. |
| Part-time to full-time | Worked less than 35 hours per week at $t$ but more than 34 hours per week at $t$ + 3. |
| Union to nonunion | Wages were set by collective bargaining at $t$ but not at $t$ + 3. |
| Nonunion to union | Wages were not set by collective bargaining at $t$ but were at $t$ + 3. |
| South to non-South | Lived in the South at $t$ but not at $t$ + 3. |
| Non-South to South | Did not live in the South at $t$ but did live in the South at $t$ + 3. |
| Urban to non-urban | Lived in an urban area at $t$ but not at $t$ + 3. |
| Non-urban to urban | Did not live in an urban area at $t$ but did at $t$ + 3. |

## APPENDIX B

Following Holmlund [1984] we derive the relationship between wage level and wage change equations as follows:

$$\text{(B.1)} \qquad \log W_{i,t} = \beta_t X_{i,t} + \varepsilon_{i,t}$$

$$\text{(B.2)} \qquad \log W_{i,t+3} = \beta_{t+3} X_{i,t+3} + \varepsilon_{i,t+3},$$

where $\varepsilon_{i,t}$ and $\varepsilon_{i,t+3}$ are random elements with the usual properties, i.e., $E(\varepsilon_{i,t}) = 0$, $E(\varepsilon_{i,t+3}) = 0$, $V(\varepsilon_{i,t}) = \sigma_1^2$, $V(\varepsilon_{i,t+3}) = \sigma_2^2$ and $E(\varepsilon_{i,t}, \varepsilon_{i,t+3}) = 0$. Note that this model does not constrain the coefficients on $X$ to be constant over time.

Subtracting equation B.1 from equation B.2 gives the following wage change equation:

$$\text{(B.3)} \qquad \log W_{i,t+3} - \log W_{i,t} = \Delta \log W_i$$

$$= \beta_{t+3} X_{i,t+3} + \varepsilon_{i,t+3} - \beta_t X_{i,t} - \varepsilon_{i,t}.$$

We then add and subtract the term, $\beta_{t+3} X_{i,t}$, to and from the right-hand side of equation B.3. This yields the following equation which contains a term reflecting the change in the value of $X$ between the two periods:

$$\text{(B.4)} \qquad \Delta \log W_i = \beta_{t+3} \Delta X_i + (\beta_{t+3} - \beta_t) X_{i,t} + \varepsilon_{i,t+3} - \varepsilon_{i,t}.$$

Again following Holmlund [1984] we assume that $\beta_{t+3}$ and $\beta_t$ are related linearly. That is,

$$\beta_{t+3} = \beta_t + \delta\beta_t, \tag{B.5}$$

where $\delta$ is a proportionality factor. If $\delta$ equals zero, then the $X$ coefficient's values stay constant over time, i.e., $\beta_{t+3} = \beta_t$. If $\delta$ is positive, then the $X$ coefficients have increased over time, i.e., $\beta_{t+3} > \beta_t$. If $\delta$ is negative, then the $X$ coefficients's values decreased over time, i.e., $\beta_{t+3} < \beta_t$. Solving equation B.5 for $\beta_{t+3} - \beta_t$, and substituting this term into equation B.4, gives us

$$\Delta \log W_i = \beta_{t+3}\Delta X_i + \delta\beta_t X_{i,t} + \varepsilon_{i,t+3} - \varepsilon_{i,t}. \tag{B.6}$$

From equation B.1 we have

$$\log W_{i,t} - \varepsilon_{i,t} = \beta_t X_{i,t}. \tag{B.7}$$

Multiplying both sides of equation B.7 by $\delta$ and substituting this term into equation B.6 yields

$$\Delta \log W_i = \beta_{t+3}\Delta X_i + \delta \log W_{i,t} + \varepsilon_{i,t+3} - (1+\delta)\varepsilon_{i,t}. \tag{B.8}$$

Thus, the wage growth equation can be thought of as the following:

$$\Delta \log W_i = \delta \log W_{i,t} + \beta_{t+3}\Delta X_i + e_i^*, \tag{B.9}$$

where $e_i^* = \varepsilon_{i,t+3} - (1+\delta)\varepsilon_{i,t}$.

Note that when $\delta$ equals zero, equation B.9 collapses to a standard wage growth equation. If $\delta$ is not equal to zero, moving $\log W_{i,t}$ from the left-hand side to the right-hand side of equation B.9 results in a wage level model that is really a wage growth model. That is,

$$\log W_{i,t+3} = (1+\delta)\log W_{i,t} + \beta_{t+3}\Delta X_i + e_i^*. \tag{B.10}$$

If $\delta$ is not equal to zero, using a standard wage growth specification instead of one similar to B.10 imposes a restriction that is false. This would have the same effect as omitting a relevant variable and the resulting estimates are likely to be biased. Our estimates of $\delta$ are negative, which implies the coefficients of the determinants of wages decreased between $t$ and $t+3$.

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