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JOB MOBILITY AND THE CAREERS OF YOUNG MEN

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ABSTRACT

We study the joint processes of job mobility and wage growth among young men drawn from the Longitudinal Employee-Employer Data. Following individuals at three month intervals from their entry into the labor market, we track career patterns of job changing and the evolution of wages for up to 15 years. Following an initial period of weak attachment to both the labor force and particular employers, careers tend to stabilize in the sense of strong labor force attachment and increasing durability of jobs. During the first 10 years in the labor market, a typical young worker will work for seven employers, which accounts for about two-thirds of the total number of jobs he will hold in his career. The evolution of wages plays a key role in this transition to stable employment: we estimate that wage gains at job changes account for at least a third of early-career wage growth, and that the wage is the key determinant of job changing decisions among young workers. We conclude that the process of job changing for young workers, while apparantly haphazard, is a critical component of workers' move toward the stable employment relations that characterize mature careers.

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> The wastefulness of this "try and try again" process of advancing to a better position is self-evident. The worker does not know in detail the nature of the job which he is obtaining nor does he know his own capacities. Nevertheless it is the principal method by which workers at the present time improve their condition on their own initiative.
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> Sumner H. Slichter, 1919, p. 218.

1. Introduction

New entrants to the labor force can look forward to about 40 years of work. Over their careers their wages will about double (in the cross-section) and they will change jobs ten times. But the pace of these changes is far from even. The first ten years of a career will account for 66 percent of lifetime wage growth for male high school graduates and almost exactly the same fraction of lifetime job changes. As the career progresses both turnover and wage growth subside.[1]

In this paper we ask how these features of the early career, high turnover and rapid wage growth, are related. That the two should coincide during the beginning of the career is not surprising. Tenuous attachment to both the labor force, and to any given employer, is combined with rapidly growing generalized work experience among young workers. Most modern theories of job mobility attach a structural interpretation to the process of job change and wage growth (Burdett, 1978; Jovanovic, 1979b). In particular, models of job change offer an addition to the traditional schooling and training explanations for why earnings rise with experience. Investments here take the form of learning about one's comparative advantage by sampling and experiencing a variety of jobs; that wages should grow as an outcome of this process is an observation traceable to Stigler's (1962) seminal article.

The most prominent and widely documented facts about labor mobility are that average rates of job changing decline with age or experience and, especially, with current

[1]Hall (1982) estimates the number of lifetime jobs; Murphy and Welch (1988) estimate the structure of wage growth over the career.

job tenure (duration).[2] When changing jobs, the wages of young people increase, but these gains decline with age (Bartel and Borjas, 1981; Mincer, 1986). Most theories of turnover are designed to accommodate these crude facts, though they differ in fundamental details.

Nonbehavioral models have the longest history (Blumen, Kogen, and McCarthy, 1955; Singer and Spilerman, 1976). The key idea is that individuals may differ for unobservable reasons in their propensities to leave a job. This presence of "movers" and "stayers" in turnover data implies that observed job duration acts as a filter, selecting individuals who have low probabilities of terminating a worker-firm pairing. Thus a tabulation of the frequency of mobility will show a negative relationship between the probability of leaving a job and its current duration, though for any individual the probability of leaving may be independent of tenure. If "stayers" are fertile for firm-specific training, while "movers" are not, then average earnings will rise faster within a job than between jobs.

A related source of heterogeneity is generated in optimizing models of individual search for a good employment match (e.g., Jovanovic 1979b). Following the terminology of Nelson (1974), suppose that jobs are pure "search goods," the quality (productivity) of any worker-firm pairing being known upon initial inspection. Rational search then implies that good pairings are more likely to survive. This source of heterogeneity may also bias inferences about the relation of job duration to mobility, though the direction of bias depends on details of the search technology in this case.[3] Mobility will also decline with time in the labor market for purely statistical reasons: given some process by which job offers are generated, the expected value of the maximum offer received (the current job) is higher for workers who have searched longer (Burdett, 1978). For exactly the same reason,

[2]See Mincer and Jovanovic (1981) and Parsons (1978).

[3]Though good matches tend to survive, persons with low tenure in survey data are well matched by virtue of the fact that they have recently accepted a new job. In general, this bias means that tabulations of the frequency of mobility based on survey data will understate any negative relation between mobility and tenure (Topel, 1986).

earnings grow with labor market experience as a return to search, even without accumulation of general training. A key issue addressed below is the extent to which this type of matching process is descriptive of actual careers.

In contrast to these models where the true effect of tenure on mobility may be zero, the probability that an *individual* changes jobs will decline with tenure if either job-specific human capital or information about the quality of a match accumulates with time on the job. Suppose that jobs are like "experience goods," in that a worker learns about match quality by observing his productivity over time (Jovanovic, 1979a). Uncertainty about the true quality of a match declines with tenure, and poor matches are more likely to end. A key difference is that survivors have learned that they are well matched, so that their own probabilities of moving have declined with tenure after some critical amount of information has accumulated.[4] At the extreme where there is no ex ante information about the quality of a new job, jobs during a worker's career form a renewal process and mobility (and earnings) is independent of time in the market (experience), holding current job duration constant.

None of these extreme models of labor turnover is likely to completely describe mobility data and earnings dynamics. Yet except under extreme circumstances, they are observationally equivalent in data on job durations alone. This suggests that additional information on the productivities of particular job pairings may provide important identifying leverage for distinguishing the importance of competing theories.

Our analysis of these issues is based on the Longitudinal Employee-Employer Data (LEED) file--a large sample of individual labor market histories taken from Social Security earnings records. The observability of job specific wages, in conjunction with the long

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[4]Jovanovic (1979a) demonstrates that the probability of leaving a new job may initially rise with tenure in this case. The reason is that it pays to remain and collect information on a new job, especially if ex ante information is scarce. Eventually, however, the probability of changing jobs must decline with tenure.

panel aspects and size of the LEED file provides important information for distinguishing among models of job mobility and earnings growth. Using these data, we follow approximately 10,000 young men from entry into the labor market, and we observe up to 15 years of quarterly post-entry labor market experience. Because we observe careers from their inception, common issues of "left censoring" of labor market histories do not arise. The data are unique in that their basis in administrative records allows us to view fast turnover jobs and transitions between them with a fidelity that is absent in all self-reported longitudinal samples, while closely tracking the evolution of wages within careers.

The paper is organized as follows. The next section describes patterns of job mobility and labor force attachment among recent entrants to the labor force. The data indicate that job attachments after entry are extremely fragile: two-thirds of all new jobs among young workers end in the first year, with both transitions to new employers and exit from the labor force playing important roles. The importance of job-to-job transitions as a source of turnover increases with experience, and jobs become more durable as the career progresses. The typical young worker holds seven full-time jobs during his first ten years in the labor market. Section 3 relates this evidence of intense job shopping to early career wage growth. Wages grow extremely rapidly during this career phase, averaging over eleven percent annually in the period of our data. Wage gains at job changes average about 10 percent, and account for about one third of total wage growth during the first ten years in the market. We also find that the evolution of wages within jobs closely approximates a random walk, so that the wage, conditioned on market experience and tenure, is a sufficient statistic for the value of a job.

We follow this evidence with an explicit model of mobility decisions that emphasizes both the evolution of wages within jobs and the arrival of external offers in affecting mobility. The main empirical result is that the job-specific wage is a key determinant of mobility: controlling for individual heterogeneity, a ten percent within-career increase in the wage reduces the probability of changing jobs by about 20 percent, and jobs with more rapid wage growth are also more stable. This connection between the evolution of wages and job changing behavior is strong evidence in favor of job shopping as a determinant of mobility and early career development.

The epigraph, taken from Slichter's dissertation, reflects a view that has all but disappeared from contemporary discussion of the mobility process.[5] He deplored the high rate of turnover that he observed among young factory workers as a hinderence to career development and an important contributor to the idleness of nonhuman capital.[6] In contrast, our evidence is that the process of job changing among young workers, while apparantly haphazard, is a critical phase in workers' move toward the long-term, stable employment relations that characterize mature careers.

2. Job Mobility Among Young Men

A. The Data

Our data source is the LEED, the longitudinal employee-employer data. The complete LEED file contains quarterly employee-employer records for over one million individuals, including Social Security earnings credited to the worker's account from each

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[5] It is still the case, however, that frequent job change may inhibit investment in firm-specific human capital and contribute to flatter wage growth profiles (Mincer and Jovanovic, 1981). Our point is that frequent job change is the rule rather than the exception and that wage growth may well be accommodated rather than hindered by the process.

[6] Reynolds (1951), in his careful study of the work history of manual workers in New Haven in the 1940s, was the first to coin the term "job shopping" in his description of the early years in the labor market. (See also Ginsberg, 1951.) Parnes (1954), in a comprehensive review of exisiting studies of labor mobility, relates a pessimistic appraisal of this mobility as a market mechanism. He concludes, for example, that workers rarely consider wages or even make job comparisons in deciding on a move (p. 188) and that manual workers only slightly more often than not improve their wages when they change jobs voluntarily (p. 190). With regard to the mobility differences by age, Parnes emphasized institutional factors (seniority rules) and maturation in explaining declining turnover with age.

employer during each quarter, the number of employees in the firm, and the detailed industry (4-digit SIC) and location (county) of the employer. Each employee and each employer has a unique identifier in the file. For individuals, information on personal characteristics is limited to age, race, and sex. The most prominent nonreported items are the person's schooling and dimensions of labor supply (hours and weeks worked).

The panel begins in the first quarter of 1957 and ends in 1972. So as to measure labor market events from the beginning of careers, we selected only individuals who were born after 1938.[7] Thus the oldest person in the data is 18 in 1957. Because of ambiguities about the meaning of jobs held while very young, all others have their histories measured starting with the first job held on or after their eighteenth birthday. Consequently, the oldest person is 34 years old in 1972, having accumulated 60 quarters of experience during the panel.[8]

Among the young men who comprise our sample, multiple job holding, rapid turnover, and return to past employers are common. Transitory jobs and employment spells followed by a gradual move toward stable employment characterize the prototypical career sequence. Unfortunately, it is not obvious how one should weight employment experience in transient jobs, or even how one distinguishes them except ex post. A revealing feature of the data is that it is extremely difficult to tell when individuals leave school to enter full-time work. In some cases the break is not as dramatic as full-time schooling models

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[7]In the succeeding empirics the selection criteria are restricted further to illustrate characteristics of the early career and, later, to enable us to analyze wage change.

[8]We generated samples of careers starting at age 16 in 1957 (N = 8,102) and, atlernatively, at age 18 (N = 9,919). The samples differed only in that job events before the respective birth quarters were ignored. The younger sample allowed us to examine labor force entry and job change prior to age 18 and to study the effect of "left-censoring" of job histories. We found that rates of job transition were essentially identical for the two samples. Table 1 below uses the younger sample while the remaining tables use job histories which begin on the eighteenth birth quarter. Versions which use the younger sample are available from the authors on request.

suggest, but rather seems best characterized as a gradual switch from nonparticipation to full-time employment along a path of high turnover and intermittent work.

In light of these facts, our definition of the length of a job was influenced primarily by inspection of the data. We define the major employer as the one who contributes the most to total earnings in a quarter. We smooth over periods of part-time or short-time employment by considering a person to be participating on a "full-time" basis if he earned at least 70 percent of the minimum quarterly wage during that quarter, assuming full-time work. This limit is gauged against the sum of earnings on all jobs, and only these quarters of full-time work are used to accumulate measured work experience.[9] Quarters in which the earnings limit is not reached are treated as "nonemployment" periods, though these may actually include employment in the uncovered sector or part-time work. Job tenure is accumulated continuous quarters with a single employer.[10] Finally, we restrict ourselves to jobs that begin after age 18 (except for Table 1). Jobs in progress at age 18 are measured from age 18.

Two final requirements were imposed on career histories that were selected for analysis. First because of the possibility that individuals may enter uncovered employment, for example the military, we exclude records with continuous nonemployment gaps of two years or more.[11] Finally, for our initial tabulations, we require that each record analyzed

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[9]These criteria for accumulating jobs and experience are more stringent than in other studies of job durations and earnings. For example, using the PSID Cline (1979), Altonji and Shakotko (1987), and Abraham and Farber (1987) include all individuals with positive hours during a calendar year. The latter two studies accumulate a year of experience if an individual reported 100 hours worked during a year. In our data, these criteria imply increased participation and job changing.

[10]In case of transitory (one quarter) changes in the identity of the main employer, we smoothed over the break and treated the employment spell as continuous. This has the effect of eliminating terminations due to such factors as temporary layoffs, which we do not view as a severance of the employment relationship.

[11]Results are not materially different when these records are maintained. However, for such workers it is not clear how to evaluate the accumulation of experience, since market work is defined by recorded earnings. We took two years to be a reasonable upper

have at least six years of potential market experience after initial entry. More stringent requirements are imposed later. Further details on sample selection are appended.

B. Mobility

All of the following results are based on the occurrence or termination of a worker-firm pairing as identifiable events, ignoring such subcategories as geographic or sectoral mobility for which these data are also suited. We confine the analysis to white males, of which there are 9,919 individuals with 58,181 "full-time" jobs in our data. Table 1 shows the age distribution of first entry to the labor market under various definitions of

TABLE 1

AGE DISTRIBUTIONS AT ONSET OF CONTINUOUS WORK
16 YEARS OF AGE AND OLDER (N = 8,102)

	Age at Beginning of Employment Spell							
Length of Spell	≤ 18	19	20	21	22	23	24	≥ 25
≥ 1 Quarter	46.0	25.6	14.7	7.6	4.1	1.5	0.5	0.1
≥ 2 Quarters	29.6	24.9	18.8	11.4	8.1	4.8	1.7	0.7
≥ 3 Quarters	24.2	22.7	18.6	12.9	10.1	7.0	2.8	1.5
≥ 1 Year	21.6	21.0	17.3	13.6	11.8	8.3	3.9	2.6

labor force attachment, which we take to depend on the length of the initial employment spell.[12] The data show that over 86 percent of young men have held a substantial job of some sort by age 20 and over 46 percent by age 18, yet weak attachment to the labor force is a prime characteristic of new entrants: the distribution of first employment spells lasting

limit on nonparticipation.

[12]These are not *job* spells. Thus an employment spell of one year or more (in the fourth row of the table) may include many individual jobs. For this table only, we consider jobs which begin as early as age 16.

a year or more is shifted sharply to the right relative to the age distribution of first employment. Only about 20 percent of spells lasting a year or more occur by age 18, implying that early spells are likely to end with a transition to nonemployment.

These tabulations suggest that young individuals spend a significant portion of their post-entry, potential labor force time without a job. To illustrate, table 2 exploits the panel aspects of the data, showing the distribution of actual labor force experience for a sample of individuals with exactly ten years of potential experience (years since first entry), and also the rate at which experience accumulates post-entry. Panel A of the table shows that only 16.2 percent of these workers are continuously employed over the first ten years of their careers, and only 44 percent have spent less than one cumulative year without a job. Nevertheless, panel B shows that the distribution of nonemployment time is heavily skewed toward the earliest part of careers. During the first few years in the market, nearly one-fourth of potential market time is spent without a job. The rate at which actual experience accumulates accelerates rapidly, however, so that strong labor force attachment is the rule after five or six years.

In part this weak attachment to the labor force reflects summer employment and other temporary or part-time jobs among persons in school, as well as transition between school and full-time work. We see now reason to exclude this type of employment spell, either in terms of accumulated experience or the job shopping process. Yet it is important to realize that school would be the main activity of only a small portion of individuals in these data: only 58 percent of the potential high school class of 1958 actually graduated, and only 30 percent of these graduates went on to attend college.[13] Thus less than 18

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[13]U.S. National Center for Education Statistics, *Digest of Education Statistics*, annual. These figures are higher today: roughly 75 percent of each cohort graduates from high school and 46 percent go on to college. Among 19- and 20-year-olds who reported their major activity as school in 1970, 57 percent worked fewer than 13 weeks in the previous calendar year and 21 percent did not work at all (March Current Population Survey).

Table 2A

Distribution of Actual Labor Market Experience for Persons with 10 Years of Potential Experience (18 Years of Age and Older)

	Years					Quarters							
	≤4	5	6	7	8	33	34	35	36	37	38	39	40
Percent with Indicated Experience	0.8	1.7	4.6	8.4	15.8	6.2	5.9	6.0	6.7	8.2	8.8	10.8	16.2

Table 2B

Actual Labor Market Experience by Years of Potential Experience

	Potential Market Experience (Years)									
	1	2	3	4	5	6	7	8	9	10
Actual Market Experience	.70	1.36	2.10	2.89	3.73	4.61	5.49	6.38	7.29	8.19
Additional Experience	.70	.66	.74	.79	.84	.88	.88	.89	.91	.90

Note.--Actual experience is cumulative quarters spent working by the definitions in the text. Potential experience is time since first entry to the labor market.

percent of the relevant post-18 population would be in school, and a smaller fraction would meet the entry criteria we impose. Because of this, we view the gradual transition to full-time participation as a prime characteristic of young workers' careers.

Table 3 shows corresponding data for the accumulation of "full-time" jobs. By the tenth year after entry, more than half of young workers have held six or more jobs, and over a third have held eight jobs or more. Only one worker in 20 has held a single job for ten years. The average number of jobs by this career point is 6.96, which is slightly higher than Hall's (1982) estimate for similarly aged workers, which was derived by a far different sampling procedure.[14] Note, however, that our definitions preclude multiple job holding as contributing to this count, so these tabulations underestimate the total number of jobs actually sampled in the job-shopping process.

Panel B of table 3 reports the flow of new jobs with experience. On average, the first full year of actual employment is divided among almost three jobs. The average pace of sampling jobs then declines fairly smoothly as experience accumulates.[15] This is our first real evidence on the sorting process; as experience accumulates, the frequency of job changing declines. This fact is inconsistent with a pure mover-stayer model, as well as with the pure "experience goods" version of job matching, in which new jobs within careers would form a renewal process. If all jobs were identical ex ante, then new jobs would not be systematically more stable than past ones.

Underlying many of these results is a prime feature of mobility data that was referred to above: the average frequency of job mobility is a declining function of current job tenure. The strength of this association in the market for young workers is illustrated

[14]Hall estimates that workers aged 25-29 have held 5.5 jobs and that workers aged 20-24 have held four. Our estimates for these age ranges are 5.9 and 3.8, respectively.

[15]Note that the tenth year of actual experience is associated with fewer new jobs than the tenth year of potential experience. This is due to sample selection: persons who change jobs frequently have spent more time out of the labor force on average, and so they are selected out of the sample of persons with high actual experience.

Table 3A

Distribution of Cumulative Full-Time Jobs for Persons with 10 Years of Potential Experience

	Cumulative Full-Time Jobs												
	1	2	3	4	5	6	7	8	9	10	11-15	16-20	≥21
Percent with Indicated Number of Jobs	4.3	7.0	9.9	11.1	11.6	9.5	9.0	8.2	6.9	4.8	13.0	3.2	1.3

Table 3B

Average Cumulative Full-Time Jobs by Years of Labor Market Experience

	Years of Experience									
	1	2	3	4	5	6	7	8	9	10
By Potential Experience	1.60	2.47	3.24	3.92	4.56	5.14	5.66	6.09	6.54	6.96
New Jobs	1.60	.87	.77	.68	.64	.58	.52	.43	.45	.42
By Actual Experience	2.78	3.77	4.50	4.99	5.37	5.62	5.83	6.11	6.18	6.22
New Jobs	2.78	.99	.73	.49	.38	.25	.21	.28	.07	.04

Note.--Potential experience is time since entry, actual experience is time employed. "Jobs" refers to t number of new employers during the indicated time span. Multiple employers at any date are not counted.

in figure 1, which shows the conditional relative frequency of job terminations given current tenure--the empirical exit-hazard function.[16] The shape of this function is a main component of the rapid accumulation of jobs among young workers. More than one-third of all new jobs among young workers end within three months, and two of every three end within a year. The probability of moving declines dramatically over the first year of a job (by a factor of three), yet even after a full year's tenure more than one-third of all remaining jobs will end in the next twelve months. After about four years on the job this annualized probability of job terminations stabilizes at about .20. An alternative representation of these data is the expected remaining duration of a job given current tenure. The estimates imply that a typical new job among young men can be expected to last only about 1.5 years, but having reached that point the job can be expected to last an additional four years.

This high-speed turnover among young workers is not simply an artifact of planned short periods of participation by new entrants to the labor force. Panel A of table 4 illustrates this fact, showing quarterly mobility functions for various levels of previous full-time experience at the start of the job. Even after eight full-time years in the market, half of all new job holders will move again within one year. This estimate may overstate the extent of mobility for the typical worker, since the fact that a new job has begun after eight years in the labor market may select individuals who are more likely to move on subsequent jobs. Even so, the uniform decline in the hazard rate with experience suggests the importance of job shopping.

Panel B of table 4 shows empirical hazards by number of prior jobs in order to illustrate a key point about observed mobility and the search process. First, there is a tendency for mobility to decline as jobs accumulate, except for individuals who have had a

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[16]For computing exit hazards we restrict ourselves to jobs held after age 18, including jobs in progress at age 18.

Table 4

Empirical Mobility Functions by Prior Market Experience, Number of Prior Jobs

	Current Job Tenure (Quarters)												
	1	2	3	4	5	6	7	8	9	10	11	12	13
A. Prior Experience:													
None	.39	.20	.15	.15	.17	.13	.10	.12	.12	.09	.08	.08	.08
1 Year or Less	.41	.25	.17	.14	.16	.11	.11	.11	.11	.11	.09	.08	.07
1-2 Years	.34	.24	.17	.13	.13	.13	.11	.11	.09	.09	.08	.08	.07
2-4 Years	.30	.23	.16	.13	.12	.11	.10	.10	.09	.08	.07	.06	.07
4-8 Years	.27	.20	.15	.13	.10	.10	.09	.09	.08	.08	.05	.06	.05
More than 8 Years	.22	.19	.14	.12	.10	.09	.08	.06	.05	.05	.05	.06	.05
B. Prior Jobs:													
None	.39	.20	.15	.15	.17	.13	.10	.12	.12	.09	.08	.08	.08
1	.34	.21	.16	.13	.13	.11	.10	.10	.10	.10	.08	.07	.07
2	.32	.21	.14	.12	.12	.10	.10	.10	.09	.08	.07	.07	.07
3 or 4	.29	.21	.15	.12	.12	.10	.09	.09	.09	.09	.08	.07	.07
5-7	.30	.22	.17	.13	.11	.12	.10	.10	.09	.08	.06	.07	.06
8 or More	.37	.30	.22	.18	.15	.14	.14	.11	.09	.12	.05	.06	.08
C. Aggregate Hazard	.33	.22	.16	.14	.13	.11	.10	.10	.10	.08	.07	.07	.07

Note.--Figures refer to the percentage of the relevant group that change major employer between successive quarters. N = 58,181.

Table 5

Mobility from Current Job by Type of Transition and Experience at Start of Job

		Current Job Tenure (Quarters)												
		1	2	3	4	5	6	7	8	9	10	11	12	13
Prior Experience:														
None	*j-j*	.093	.078	.059	.070	.068	.066	.054	.072	.065	.050	.047	.053	.045
	j-n	.297	.118	.091	.080	.099	.059	.050	.047	.054	.035	.036	.023	.032
1 Year or Less	*j-j*	.150	.122	.094	.083	.083	.064	.078	.075	.068	.076	.058	.055	.047
	j-n	.262	.125	.073	.062	.075	.050	.036	.039	.045	.034	.028	.028	.029
1-2 Years	*j-j*	.183	.144	.109	.089	.084	.079	.083	.078	.067	.063	.056	.050	.050
	j-n	.156	.094	.060	.040	.044	.048	.025	.031	.026	.023	.020	.033	.018
2-4 Years	*j-j*	.186	.151	.120	.096	.086	.078	.074	.070	.067	.060	.054	.048	.049
	j-n	.109	.075	.043	.036	.036	.029	.025	.027	.025	.021	.018	.017	.021
4-7 Years	*j-j*	.193	.145	.114	.098	.080	.067	.071	.069	.054	.060	.039	.048	.041
	j-n	.078	.057	.035	.029	.121	.029	.020	.017	.022	.018	.015	.015	.012
More than 8 Years	*j-j*	.175	.140	.109	.092	.070	.075	.065	.049	.042	.044	.041	.048	.043
	j-n	.048	.049	.032	.025	.026	.019	.011	.014	.008	.008	.013	.016	.011

Note.--Estimates are the percentage moving within one quarter for the indicated reason. *j-j* represents job-to-job transitions, *j-n* represents job-to-"nonemployment" transitions. See text for definitions. Sample is 9,919 white males.

very large number of prior jobs. The decline is less prominent than in panel A, and it is mainly concentrated in the first quarter of new jobs. Thus, while these facts suggest that later jobs are more stable, they also imply substantial heterogeneity among individuals in probabilities of leaving a job. Persons with many prior jobs are also more likely to leave a new one. This may be because some people are "movers," say, due to occupational differences in specificity of human capital or because of differences in job search histories. For example, if workers search systematically for a productive match, then a sequence of low-quality prior jobs implies a lower expected quality of the current job, raising average turnover in the population of new jobs.

C. Labor Force Attachment and Types of Job Endings

In table 2, we noted that labor force attachment tends to be weak in the early phases of careers. Table 5 shows the surprising importance of this fact as a determinant of turnover and the unstable jobs of young workers. For workers on their initial jobs, half of all first-year job endings result in a transition to nonemployment.[17] Both the relative and absolute importance of this transition decline with experience, however, though even among workers with eight full years of prior experience the proportion who leave new jobs for nonemployment in the first year is .15.[18] It is impossible for us to tell in these data whether these high "nonemployment" flows represent withdrawals from the labor force or transitions into unemployment. They are consistent with related evidence on unemployment flows, which shows that a prime component of higher unemployment rates among young workers is their vastly higher flows out of jobs (Topel, 1984). A plausible interpretation of

[17]A transition to a quarter with earnings below 70 percent of the minimum wage.

[18]Some of this decline may be due to our sample selection criteria. For any fixed level of labor supply, high-wage workers are less likely to cross the earnings threshold that determines participation. However, we think this cutoff is sufficiently low that the bias is not serious. If anything, we probably understate nonemployment transitions by our method.

these facts is that the costs of unemployment (or nonemployment) are relatively low for young workers, so changes in employment status are frequent.

Note also that the absolute importance of job-to-job transitions rises over the first few years in the market before turning down, though the total frequency of job endings declines throughout. The implied low intensity of search and job sampling during the earliest phase of market participation is another dimension of the weak labor force attachment of young workers. Because we restrict ourselves to full-time jobs after age 18 the group with zero experience may be anomalous. These are individuals who have never worked full-time until after their eighteenth birthday.

3. Wage Growth Within and Between Jobs

To this point we have ignored the role of job changing as a determinant of wage growth, yet in terms of theory the connection is obvious. Empirically, profiles of earnings growth mirror the pattern of early career mobility. For the typical male worker in the U.S., over two-thirds of total lifecycle earnings growth occurs during the first ten years of labor market experience (Murphy and Welch, 1988).

We examine two important aspects of the evolution of earnings for young workers. First we analyze wage growth within jobs and the time series properties of wage innovations. The data indicate that earnings evolve within jobs as approximately a random walk with drift. This result is important for our subsequent analysis of workers' mobility decisions. We follow with direct evidence on the gains from mobility, based on observed changes in earnings at job transitions. The evidence is that more than a third of early career wage growth is associated with job changing, while larger wage gains at transitions are associated with a corresponding decline in subsequent mobility.[19] Because we wish to

[19]These results differ from those of Mincer (1986) and Bartel and Borjas (1982) by a factor of almost two. They find smaller wage changes at job changes in large part (we theorize) because of the inability to follow intra-year job and wage changes in the popular

track the evolution of wages and jobs over the career, in what follows we restrict attention to individuals with at least 13 years of potential labor market experience in the panel. Further, so as to focus on workers with relatively permanent attachments to the labor force, "entry" is defined to occur when (i) earnings in a quarter exceed 70 percent of the minimum wage for full-time work, and (ii) earnings over the subsequent four quarters also satisfy this criterion.[20] There were 872 persons satisfying these requirements, representing 44,089 job-quarters.

A. Within-Job Growth

The common form of "human capital earnings functions" applied to micro data expresses the natural logarithm of earnings as a function of labor market experience (X_{jt}) and current job tenure (T_{jt}) on job j at time t. Suppressing other regressors for ease of exposition, a prototype model of individual log earnings is:

$$w_{jt} = H(X_{jt}, T_{jt}) + \phi_j + \epsilon_{jt} . \tag{1}$$

The most common specification of $H(\cdot)$ is quadratic in its arguments (e.g., Mincer and Jovanovic, 1981). To maintain the notion of heterogeneity among jobs in earning capacity for an individual worker, we have specified the unobservables ϕ_j and ϵ_{jt} to denote a fixed effect specific to a particular job, j, and a time varying random component of measured earnings.

Sample selection induced by job search and matching implies that $E(\phi + \epsilon | X, T) \neq 0$, so standard regression techniques are inappropriate for (1). Our interest is in the

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longitudinal surveys, and because of measurement error in identifying true job changes. In addition, for the "movers" in these data, recall of past job changes in an interview setting is understandably incomplete.

[20]The second condition eliminates individuals whose careers start with summer jobs. Tabulations of the March CPS indicate that only 17 percent of 19- and 20-year-olds enrolled in school would meet the second condition on annual earnings.

determinants of wage growth, however, so fixed effects can be eliminated by differencing (1) within jobs:

$$\Delta w_{jt} = \Delta H(X_{jt}, T_{jt}) + \Delta\epsilon_{jt} \ . \tag{2}$$

If $E(\Delta\epsilon|X, T) = 0$ then least squares will yield unbiased estimates of the parameters of $\Delta H(\cdot)$, though separate effects of experience and tenure on y_{jt} are not identified since $\Delta X = \Delta T = 1$ within jobs.[21]

Estimates of equation (2) are reported in panel A of table 6 under various functional form assumptions for $H(X_{jt}, T_{jt})$. Inspection of the quarterly data revealed a strong tendency for within-job earnings changes to occur at annual intervals. Because of this, the reported estimates are based on annual differences of (log) real ($1967) quarterly wages. In the absence of job mobility, the average rate of annual wage growth was about seven percent ($\bar{X} = 7.1$ years), but among new entrants to the labor force it was about 14 percent.[22] This illustrates the decline of within-job wage growth with experience, and growth also declines (at a decreasing rate) with job tenure. Rows (iv) and (v) of table 6 control for the *completed* duration of the current job and for whether the job ends within one year of the current period. We find that more durable jobs offer slightly greater wage growth--an extra five years of completed tenure raises average annual growth by about 1.5 points--though jobs that will end "soon" are characterized by lower growth in periods preceding the end of the job. Both of these effects suggest that the evolution of wages on a job plays a key role in influencing mobility decisions. Jobs that yield higher wage growth

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[21] For $E(\Delta\epsilon|X,T) = 0$ we require that current wage innovations do not generate sample selection. For example, if low values of $\Delta\epsilon$ immediately induce mobility, then these values will be not observed in the sample of individuals who do not change jobs. We abstract from this type of selection in estimating (2).

[22] Because wage changes are annual, to be included in the sample a job must have survived for at least six quarters. This in itself focuses the analysis on relatively stable jobs. Note that the estimates include aggregate wage growth, so comparison to cross-sectional estimates is not appropriate.

TABLE 6

DETERMINANTS OF WITHIN-JOB WAGE GROWTH

A: Wage Growth Models

($\Delta \bar{W} = .071$)

	Intercept	ΔX^2	ΔT^2	ΔT^3	Completed Tenure	Job Change	R^2
(i)	.136 (.006)	-.005 (.0004)	-	-	-	-	.025
(ii)	.138 (.006)	-.004 (.0004)	-.002 (.0005)	-	-	-	.026
(iii)	.151 (.007)	-.004 (.0004)	-.006 (.0015)	.0003 (.0001)	-	-	.027
(iv)	.129 (.009)	-.004 (.0005)	-.009 (.0017)	.0003 (.0001)	.004 (.0010)	-	.029
(v)	.139 (.0104)	-.004 (.0005)	-.008 (.0017)	.0004 (.0001)	.003 (.0012)	-.013 (.007)	.030
Mean (S.d.)		13.29 (6.67)	6.66 (5.49)	56.15 (85.93)	7.00 (3.67)	0.28 (0.45)	

B: Residual Autocovariances for $\Delta\epsilon_{jt}$: Model (v)

Lag	0	1	2	3	4
Covariance	.0434 (.0020)	-.0132 (.0014)	-.0002 (.0009)	-.0003 (.0008)	-.0007 (.0009)
Autocorrelation	1.00	-.3269	-.0043	-.0088	-.0173

C: Estimated Error Structure (θ)

ρ	σ_{vv}	$\sigma_{\eta\eta}$
.9701 (.0436)	.0173 (.0011)	.0128 (.0012)

Notes.--Standard errors are in parentheses. The dependent variables is the annual change in log quarterly earnings, beginning with the sixth quarter on a job. Experience and tenure are measured in years. The data included 6,698 job years on 872 individuals.

tend to survive, and sluggish wage growth is associated with impending mobility. Since mobility decisions are endogenous, however, these empirical relationships have no structural interpretation at this stage.

For purposes of our subsequent analysis of mobility decisions, the autocovariance structure of the residuals, $\Delta\epsilon$, is of greater interest than the form of $H(X, T)$. For example, if $\Delta\epsilon$ is i.i.d., then the evolution of wages within a job is a random walk with drift. In this case past wage innovations do not predict future growth, so the current quarterly wage, experience, and tenure are sufficient statistics for the distribution of future wages on a job. Estimates of the autocovariances of the residuals from model (v) are reported in panel B of table 6.[23] The standard deviation of annual within-job wage changes (conditioning on experience and tenure) is about 19 percent. This estimate is substantial, but it is smaller than corresponding estimates from other sources of panel data based on surveys. This difference reflects both the smaller importance of reporting errors in Social Security data and the fact that we eliminate between-job wage changes from the calculations.

More interesting is the strong negative autocorrelation in $\Delta\epsilon$ at lag one, followed by small, though uniformly negative correlations at higher lags.[24] This pattern suggests the following, quasi-ARMA model for wage innovations. Decompose ϵ as $\epsilon_{jt} = e_{jt} + \eta_{jt}$ where e is a systematic shock to wages and η represents a purely transitory disturbance. For example, η may be generated by transitory changes in hours or wages, or by coding and

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[23]Corresponding estimates for other models are virtually identical to these.

[24]There is weak evidence of nonstationarity in the sense of declining variance of $\Delta\epsilon_{jt}$. For the first ten years of a job the estimates of $C_0(T)$ are .056, .045, .034, .037, .038, .045, .041, .025, .021, .034. Thus any decline in the variance of wage innovations (as implied by learning models, for example) must be concentrated in the early periods of jobs.

measurement errors.[25] We allow e, the "permanent" component, to follow an AR(1) process with parameter $\rho > 0$. The theoretical autocovariances of $\Delta\epsilon$ are then:

$$\begin{aligned} C_0 &= 2\sigma_{vv}/(1+\rho) + 2\sigma_{\eta\eta} \\ C_1 &= -\sigma_{vv}(1-\rho)/(1+\rho) - \sigma_{\eta\eta} < 0 \\ C_2 &= -\sigma_{vv}\rho(1-\rho)/(1+\rho) < 0 \\ C_k &= -\sigma_{vv}\rho^{k-1}(1-\rho)/(1+\rho) < 0 \end{aligned} \tag{3}$$

where $C_k \equiv E(\Delta\epsilon_t\Delta\epsilon_{t-k})$ and σ_{vv} is the variance of innovations to e; that is, $e_{jt} = \rho e_{jt-1} + \eta_{jt}$. The three parameters $\theta = (\rho,\sigma_{vv},\sigma_{\eta\eta})$ yield a parsimonious description of the empirical covariance structure. Note that if $\rho = 0$ the theoretical autocorrelation in $\Delta\epsilon_t$ is -.5. Thus the estimated first-order autocorrelation of -.33 in table 6 strongly suggests $\rho > 0$. The very small covariances at higher lags suggest that ρ is close to 1.0.

To estimate θ, denote the right-hand side of (3) as $F(\theta)$, so (3) is of the form $C = F(\theta)$ where $C = (C_0, C_1, ..., C_k)'$. The estimated covariances then satisfy $\hat{C} - F(\theta) \sim N(0, \Sigma)$, and we seek an estimate of θ, $\hat{\theta}$, that minimizes the quadratic form $S = (\hat{C} - F(\hat{\theta}))'\Sigma^{-1}(\hat{C} - F(\hat{\theta}))$. Minimizing S and expanding $F(\hat{\theta})$ about an initial consistent estimate θ_0 leads to the iterative formula $\hat{\theta} - \theta_0 = [F'(\theta_0)^T\Sigma^{-1}F'(\theta_0)]^{-1}F'(\theta_0)\Sigma^{-1}(\hat{C} - F(\theta_0))$, which is a method of moments estimator for θ. The last panel of table 6 reports estimates of θ based on the first four covariances of ϵ and a consistent estimate, $\hat{\Sigma}$, of the covariance matrix of $\hat{C}$.[26]

The key feature of these estimates is that the evolution of earnings within jobs is approximately a random walk with drift: the estimate of $\hat{\rho} = .97$ is not materially different

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[25]The decomposition is similar to a permanent income model, where e and η represent permanent and transitory shocks. To fit the empirical covariances we require only that one component of ϵ be white noise. An alternative model for the data is MA(1), though we know of no plausible economic rationale for this model.

[26]Inclusion of higher-order lags had no appreciable impact on the results. Our estimate of Σ is derived from the fourth moments of the data.

from unity.[27] This is important, since it implies that a good predictor of *future* wages on a particular job is the current wage, adjusted for predictable growth due to accumulated experience and seniority. Past wage changes are not informative. Thus heterogeneity among jobs in predictable wage growth is not an important feature of the data. In conjunction with this estimate, the estimated variances of e and η imply that about 95 percent of the within-job residual variance in measured quarterly wages is associated with the systematic component, e; the remainder is accounted for by measurement error or other transitory shocks to measured earnings.[28]

B. Between-Job Wage Growth

The LEED file provides a more complete record of earnings growth than is available in other sources of panel data. Yet even with quarterly observations it is difficult to uniquely assign a quarterly wage to each job. For example, since the beginning and end of each job are known to occur within particular three-month intervals, recorded earnings during the first and last periods of a job will not provide valid estimates of the wage. Because short jobs are common, it is difficult to reliably track earnings growth associated with every job change in a worker's career.

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[27]Estimates of the covariance structure of earnings within jobs are rare. Altonji and Shakotko (1987) discuss the point, but they assume that within-job wage disturbances are purely transitory. Our results are inconsistent with that assumption. Abowd and Card (forthcoming) report estimates of the autocovariances of reported earnings and hours changes for single job spells in the NLS and PSID samples. Their results for earnings are very similar to ours, though their estimated variances are greater because of greater measurement error in survey data. We obtain nearly identical parameter estimates to those in panel C using PSID hourly earnings data.

[28]One caveat to these results is that sample selection may cause the assumption $E(\Delta\epsilon|X, T) = 0$ to be violated. If current innovations to within-job wage offers, $\Delta\epsilon$, affect current mobility decisions, then we observe wage outcomes only for individuals who chose not to change jobs. Wage growth in the sample of stayers will overestimate potential growth available to a random worker. The estimated residuals from (2) are affected by this selection bias, which may affect our inferences about the error structure.

To get around this problem we study transitions to jobs that survive for at least one quarter. Consider the transition from job j-1 to job j. Denoting the first period of job j by t, we know that w_{jt+1} and $w_{j-1,t-2}$ are valid wage observations. We seek an estimate of between-job growth, given by

$$E(w_{j,t} - w_{j-1,t-1} | w_{j,t+1}, w_{j-1,t-2}) = w_{j,t+1} - w_{j-1,t-2} \tag{4}$$
$$- E(w_{j,t+1} - w_{j,t} | \cdot) - E(w_{j-1,t-1} - w_{j-1,t-2} | \cdot)$$

The last two terms on the right of (4) denote expected wage growth on the new (j) and old (j - 1) jobs, respectively. We estimate these terms from the within-job earnings growth model reported in row (v) of table 6. Note that by considering only those new jobs that survive more than one quarter we treat temporary jobs lasting a quarter or less as elements of a single transition.

Panel A of table 7 reports estimates of wage growth at job transitions during the first ten years of labor market experience, based on (4). The first row shows mean values of wage growth at transitions, broken out by 2.5-year experience intervals. The typical job change during this career phase is associated with a 12 percent increase in an individual's quarterly wage, though for experienced workers with 7.5-10 years in the labor market the average change is only half this large. These estimates should be compared to average quarterly wage growth *within* jobs of only 1.75 percent (table 6). Adjusting the estimates for wage growth that would have occurred in the absence of mobility (again using (4)) shows an average wage *gain* at transitions, w_{jt} - $w_{j-1,t}$, of about 10 percent. These wage changes account for a significant proportion of early-career wage growth: Average cumulative (log) wage growth during the first ten years of labor market experience is about .95, of which 40 percent (.380/.947) is accounted for by job transitions. In terms of wage gains at job transitions (row 4), the corresponding estimate is that about one-third of total

TABLE 7

WAGE CHANGES AT JOB TRANSITIONS FOR YOUNG MEN

A: Average Wage Changes at Job Transitions as a Component of Wage Growth

	Experience Interval (Years)				
	0-2.5	2.5-5	5-7.5	7.5-10	0-10
Average wage change at job transitions	.171 (.015)	.119 (.016)	.079 (.015)	.057 (.016)	.114 (.007)
Average wage gain at job transitions	.145 (.015)	.099 (.016)	.064 (.015)	.046 (.016)	.094 (.007)
Cumulative wage change at transitions	.169 (.014)	.118 (.015)	.054 (.015)	.039 (.015)	.380 (.022)
Cumulative wage gains at transitions	.143 (.013)	.091 (.015)	.039 (.015)	.032 (.015)	.313 (.021)
Cumulative wage growth	.316 (.012)	.286 (.013)	.204 (.012)	.141 (.013)	.947 (.015)

B: Determinants of Wage Changes at Job Transitions

Dependent Variable	Intercept	ΔX^2	T_{j-1}	T_{j-1}^2	T_{j-1}^3	Δ (Completed Tenure)	R^2
$Y_{jt} - Y_{j-1,t-1}$	.1992 (.0168)	-.0013 (.0003)	-.0257 (.0042)	.0016 (.0003)	-2.4^{-5} (5.4^{-6})	.0028 (.0007)	.037
	.2305 (.0151)	-.0015 (.0003)	-.0272 (.0042)	.0016 (.0030)	-2.5^{-5} (5.4^{-6})		.032
$Y_{jt} - Y_{j-1,t}$	.1684 (.0168)	-.0010 (.0003)	-.0249 (.0042)	.0016 (.0003)	-2.5^{-5} (5.4^{-6})	.0029 (.0007)	.031
	.1998 (.0151)	-.0013 (.0003)	-.0264 (.0042)	.0016 (.0003)	-2.5^{-5} (5.4^{-6})		.026

Notes.--Standard errors are in parentheses. Dependent variable is the change in log quarterly earnings at job transitions. The estimates are calculated from job transitions that result in a new job lasting more than one quarter. There were 3,367 such transitions in the data. See text for description of procedures. In panel B job tenure and experience are measured in quarters.

earnings growth during the first ten years of labor market experience is attributable to job changing activity.

Panel B of table 7 reports estimated determinants of wage changes at job transitions. The underlying specifications are the same as for within-job wage growth in table 6 with two differences. First, since current tenure on the new job is zero, the between-job change in tenure is just $T_{j-1,t-1}$. Second, to pursue the point that completed job duration is an (endogenous) indicator of the relative quality of the job via mobility decisions, we include the between-job change in completed tenures as a regressor.[29] As expected, between-job wage gains decline with experience, and also with prior job tenure. More importantly, average wage gains are largest in transitions to more durable jobs: An increase of one year in completed job duration is associated with a one percentage point increase in the *initial* wage on a new job. Again, since mobility decisions are endogenous no causality is implied by this relationship. Yet the relationship suggests that workers' mobility decisions are strongly affected by the job-specific wage, and that these wage gains are a key element generating workers' sorting to stable employment relations. We next provide a more formal analysis of these decisions.

4. Econometric Evidence on the Job Shopping Process

Our model of mobility decisions is based on wealth maximizing on-the-job search.[30] In light of our preceding evidence we allow job changing decisions to be affected by the (stochastic) evolution of wages *within* a job--which affects the value of remaining with the current employer--as well as by the arrival of external wage offers from new employers.

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[29] Abraham and Farber (1987) report corresponding results for wage levels.

[30] For earlier treatments of this problem see Burdett (1978), Jovanovic (1979a,b; 1984), Topel (1986), and Mortensen (forthcoming).

The discussion that follows outlines the main empirical implications of the model. Details are appended.

A. A Model of Mobility Decisions

We assume that wage offers from potential new employers are generated by a known offer distribution. The location of this distribution differs among individuals due to differences in talent or other characteristics, and it may also shift within a worker's career as general human capital accumulates. Suppressing individual differences for the moment, we account for the latter effect by assuming that the location of the external wage offer distribution depends on an individual's cumulative labor market experience, X:

$$\text{prob}(w^0 < z;\ X) = G(z;\ X),\ G_x(z;\ X) \leq 0\ . \tag{5}$$

Experience increases wage offers if the last inequality in (5) is strict, though recall that *observed* wages will increase with experience due to search, even if general productivity is independent of experience ($G_x(\cdot) = 0$).[31] An important implication of the model is that the hypothesis $G_x(\cdot) = 0$ is empirically testable. We assume that the occurrence of offers from (5) is Poisson with parameter π, so the probability of obtaining a new offer during a short period Δt is $\pi\Delta t$. Since these offers represent starting wages on new jobs, they are observed in our data only if a worker changes jobs; that is, if an offer is accepted.

Our previous evidence showed that the evolution of wages within jobs closely approximates a random walk with drift. To encompass this case we assume that the

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[31] Equation (5) subsumes the standard specification of earnings equations as a special case. For example, let the log wage offer for person i on job j be $\ln w_{ij}^o = A(X_i) + \varphi_{ij} + \mu_i$ where φ_{ij} is a job-specific effect and μ_i is a person-specific intercept. Then $G_x(\cdot) < 0$ is equivalent to $A'(X) > 0$; expected wage offers rise with experience.

probability distribution of a new *internal* wage offer (w^i) from the current employer depends on the current wage (w), experience (X), and tenure (T):[32]

$$\text{prob}(w^i < y;\ w,\ X,\ T) = F(y;\ w,\ X,\ T)\ . \tag{6}$$

Thus within-job wage growth is stochastic, and the triplet (w, X, T) is a sufficient statistic for the distribution of current and future wages on a job. Note that past wages do not enter (6), so we implicitly assume that past wage growth is not a predictor of future growth on a job. This is implied by our results above. We assume that a higher current wage increases the distribution of future offers ($F_w(\cdot) < 0$). If expected wage growth is nonincreasing with experience and tenure (wage levels are concave) then $F_X(\cdot) \geq 0$ and $F_T(\cdot) \geq 0$. As with outside offers, we assume that the occurrence of within-job wage changes is Poisson.

Distributions (5) and (6) are the sources of observed wage changes within and between jobs, so the wage, experience, and tenure summarize all relevant information about the current job and the distribution of alternatives. This implies the existence of a function $V(w, X, T)$ that gives the expected present discounted value of lifetime wealth from searching optimally on a job that currently pays wage w. Given $V(\cdot)$ and an external wage offer w^o, a job change occurs if $V(w, X, T) < V(w^o, X, 0)$; that is, if the new job with zero tenure offers greater expected wealth than the current job. This decision rule defines a "reservation offer," $R(w, X, T)$, satisfying $V(R(w, X, T), X, 0) = V(w, X, T)$ such that any external offer exceeding $R(\cdot)$ is acceptable. The probability of receiving a new job offer in a short interval of time is π, so the probability (density) of leaving a job at tenure T, given that the worker has not left before T, is:

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[32] In an earlier version of this paper we treated within-job earnings growth as nonstochastic. Discussions with Dale Mortensen improved our treatment of this issue, and added generality. See Mortensen (forthcoming) for an elegant treatment.

$$\lambda(w, X, T) = \pi \text{ Prob}(w^o > R(w, X, T)) \tag{7}$$
$$= \pi\{1 - G(R(w, X, T); X)\}.$$

We seek empirically verifiable restrictions on the *hazard function*, (7), which is to be estimated.[33] For purposes of exposition treat $R(\cdot)$ as differentiable. Then the effects observables on mobility are

$$\lambda_w(w, X, T) = -\pi g(R, X)R_w(w, X, T)\ , \tag{8a}$$

$$\lambda_T(w, X, T) = -\pi g(R; X)R_T(w, X, T)\ , \tag{8b}$$

$$\lambda_X(w, X, T) = -\pi g(R; X)R_X(w, X, T) \tag{8c}$$
$$- \pi G_X(R; X)\ ,$$

where $g(z; X) = G_z(z; X)$ is the density of wage offers. It is fairly obvious that $\lambda_w(\cdot) < 0$. A higher wage increases the value of the current job, and hence the reservation offer, so the job is less likely to end. Similarly, the sign of $\lambda_T(\cdot)$ depends on whether $V(w, X, T)$ is increasing or decreasing with tenure. If expected wage growth is larger at the beginning of jobs ($F_T(\cdot) > 0$) then $V(w, X, T)$ is monotonically decreasing in T and $R(w, X, T) < w$ for $T > 0$. At the same wage, new jobs are more valuable because they offer higher expected wage growth. Workers would therefore accept a wage cut to obtain them. Thus, conditional on the current wage and experience, mobility *increases* with tenure if expected on-the-job wage growth is declining.[34]

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[33]The model can be extended to allow transitions to nonemployment, which would lead to a competing risk specification with two types of job endings. Let $V(w^n,X)$ be the value of search for a jobless person with nonwage income w^n. If a within-job wage change causes $V(w,X,T) < V(w^n,X)$ then the individual will quit in order to search from nonemployment. Recall that the distinction in our data is only between *covered* employment and all alternatives, including employment in the uncovered sector. Thus the extension offers few additional restrictions in terms of analyzing the data. We therefore focus on a single type of job ending.

[34]If workers learn about match quality as tenure accumulates, then the variance of wage innovations also declines with tenure. Variance increases the option value of the current job, so the probability of leaving the job increases with tenure in this case as well.

The effect of labor market experience on mobility (8c) is determined by the effect of experience on the reservation offer and on the distribution of alternatives. If wage offers grow with experience ($G_X(\cdot) < 0$) the latter effect is positive. The effect on the reservation offer is unsigned because additional experience affects wealth on both the current and alternative jobs. Note, however, that $R(w, X, 0) = w$: at the beginning of a job the least acceptable offer is the current wage, so $R(\cdot)$ is independent of experience at this time. Thus, at $T = 0$ (8c) implies that mobility is *increasing* with experience ($\lambda_X(w, X, 0) > 0$). The reason is that additional experience improves outside wage offers if $G_X(\cdot) < 0$, so the current job is *more* likely to end as experience accumulates. Put differently, of two identical workers with identical wages, the one with greater experience is employed in a poorer match relative to his alternatives. Thus a direct test of the hypothesis that wage offers rise with experience is that mobility must rise with experience, conditional on current tenure and the wage.

These predictions for the impact of experience and tenure are the opposites of what we found in the simple tabulations of job mobility in section 2. There we argued that the value of the current job tended to increase with both experience and tenure because of sample selection induced by search and the accumulation of job-specific capital. The key difference is that the hazard function, (7), controls for the current wage, which theory implies is a main factor affecting job-changing decisions. Here the wage, experience, and tenure are sufficient statistics for the value of the job.[35] Thus a key implication of the

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See Jovanovic (1984) or Mortensen (forthcoming).

[35]This feature of the model is implied by our specification of within-job wage changes, (6). It need not hold in alternative models. For example, if investment in job-specific human capital is endogenous, then incentives to invest will be greater in a good match. If workers finance investment through reduced initial wages, the connection between the current wage and the value of the match is broken. "Bonding" models of within-job wage profiles (Lazear, 1981) have the same effect. The brevity of the typical job in our data, and our finding of no serial correlation in within-job wage growth, suggest that these models are not important for very young workers.

theory is that the *unconditional* exit hazard may decline in both labor market experience and current job tenure, as in table 4, but these effects are reversed when mobility decisions are properly conditioned on the current wage.

A final point is important. Recall that we suppressed individual differences in earning capacity when specifying the wage offer distribution, (5). To explicitly account for these differences, let the offer distribution for person i be $G(w^o; X, \mu_i)$ where μ_i is an unobservable person-specific shifter such as ability. Assume that $G_\mu(\cdot) < 0$; greater ability increases wage offers. Then the hazard is of the form $\lambda(w, X, T, \mu_i)$, and the effects of observables on the probability of changing jobs refer to comparisons between otherwise identical individuals, or to variations within a person's career. Parameter μ_i enters as an unobservable that increases mobility for any values of (w, X, T). A high-μ person is more likely to leave a job for any wage because his offer distribution dominates that of a low-μ person. Therefore, conditioning on the current wage serves to control for job-specific heterogeneity in match qualities, but the theory also implies a form of "mover-stayer" heterogeneity due to fixed individual differences in mobility decisions. We explicitly account for the presence of these unobservables in estimating the model.[36]

B. Estimation

Equation (7) gives the conditional failure time density for job spells, expressed in continuous time. Since the LEED data are quarterly, the recorded information is that a job transition did or did not occur within a particular three-month period. Denote the continuous time hazard rate for individual i on job j at tenure T as $\lambda(T; Z_{ij}(T), \mu_i)$ where Z is a vector of variables that affect mobility decisions. Then by well-known results (e.g.,

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[36]This heterogeneity cannot be ignored. For instance, suppose that more able (high wage) persons are less likely to change jobs for reasons not accounted for by the model. Then a cross-sectional comparison of mobility behavior would show a negative relation between wages and mobility even if there is no effect of the wage on individuals' decisions.

Kalbfleisch and Prentice, 1980), the probability that job j ends within an interval $(t, t+\Delta t)$ given that it did not end prior to t is:

$$\bar{\lambda}(t, t+\Delta t; Z, \mu_i) = 1 - \exp\{-\int_t^{t+\Delta t}\lambda(u; Z_{ij}(u), \mu_i)du\}. \tag{9}$$

The empirical model is closed by choosing a functional form for $\lambda(\cdot)$. We adopt the "proportional hazards" specification,

$$\lambda(T; Z_{ij}(T), \mu_i) = \exp\{h(T) + Z_{ij}(T)\beta + \mu_i\}, \tag{10}$$

where $h(T)$ is an estimable form for the effects of job tenure on the hazard and μ_i is a person-specific effect that is fixed across jobs within a career, as indicated by the theory.[37]

If the individual effects were known, the likelihood function implied by the nonlinear model (9) and (10) for a sample of job durations would have a straightforward form. Here, however, the μ_i are unobservable. Further, the theory implies that they are correlated with other observed regressors, the wage in particular. Thus random effects methods of integrating out unobservables, which assume orthogonality to observed regressors, are inappropriate for this model.[38] This leaves true fixed effect estimators that treat the μ_i as estimable parameters and that rely directly on the panel aspect of the data. However, the fixed effect estimator applied to nonlinear models is known to be biased in short panels (Anderson, 1973; Chamberlain, 1980). Thus, an important issue is whether the panel lengths derivable from the LEED file are long enough that the estimator can be applied with confidence.

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[37]In estimating the model we assume that the regressors $Z_{ij}(T)$ are fixed within the three-month intervals available to us.

[38]See Kiefer and Neumann (1981) or Heckman and Singer (1984) for applications of random effects methods.

Monte Carlo evidence for the performance of the fixed effect estimator is reported in Topel (1986) (also see Heckman [1981] for evidence on the fixed effects probit model). The key result is that, in data with the turnover characteristics of the LEED, the short panel bias is quite small in panels of over eight years.[39] To be conservative, as noted above, we restricted the analysis to persons with minimum panel lengths of 13 years of potential post-entry labor market experience. There were 872 persons satisfying this restriction. We estimate a separate intercept, μ_i, for each of these persons, so our results on mobility refer to the effects of changes that occur *within* individual careers.[40] Details of the likelihood function are appended.

C. Empirical Results

Before turning to details of parameter estimates, figures 2 and 3 illustrate the shape of the hazard function (10) when experience and tenure effects are unrestricted by functional form. In each case, the effects of experience (figure 2) or tenure (figure 3) are represented by a sequence dummy variables. Because theory predicts that conditioning the hazard on the current wage will change the effect of experience or tenure on mobility, we show these effects for models that both include and omit the current wage as a control variable.[41]

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[39] In general the bias causes an underestimate of the importance of negative duration dependence ($\partial\lambda/\partial T < 0$). In panels of over eight years, this bias was always smaller than 10 percent. Heckman's (1981) evidence for the fixed effect probit model also showed a critical panel length of about eight periods, though the critical panel length must depend on the frequency of the event being studied. Very rare or very common events require longer panels.

[40] The estimated individual effects reflect factors that shift the offer distribution, as above, as well as other unobserved differences that affect mobility but that are fixed within a person's career.

[41] Other regressors for these models are reported in table 8, as are summary statistics. The dummies for experience are at 0-6 months, 6-12 months, followed by annual dummies out to 12 years. The tenure dummies are quarterly for four years, after which the hazard is assumed to be constant.

Figure 2

Estimated Probability of Changing Jobs
by Years of Labor Market Experience

Figure 3

Estimated Probability of Changing Jobs
by Quarters of Current Job Tenure

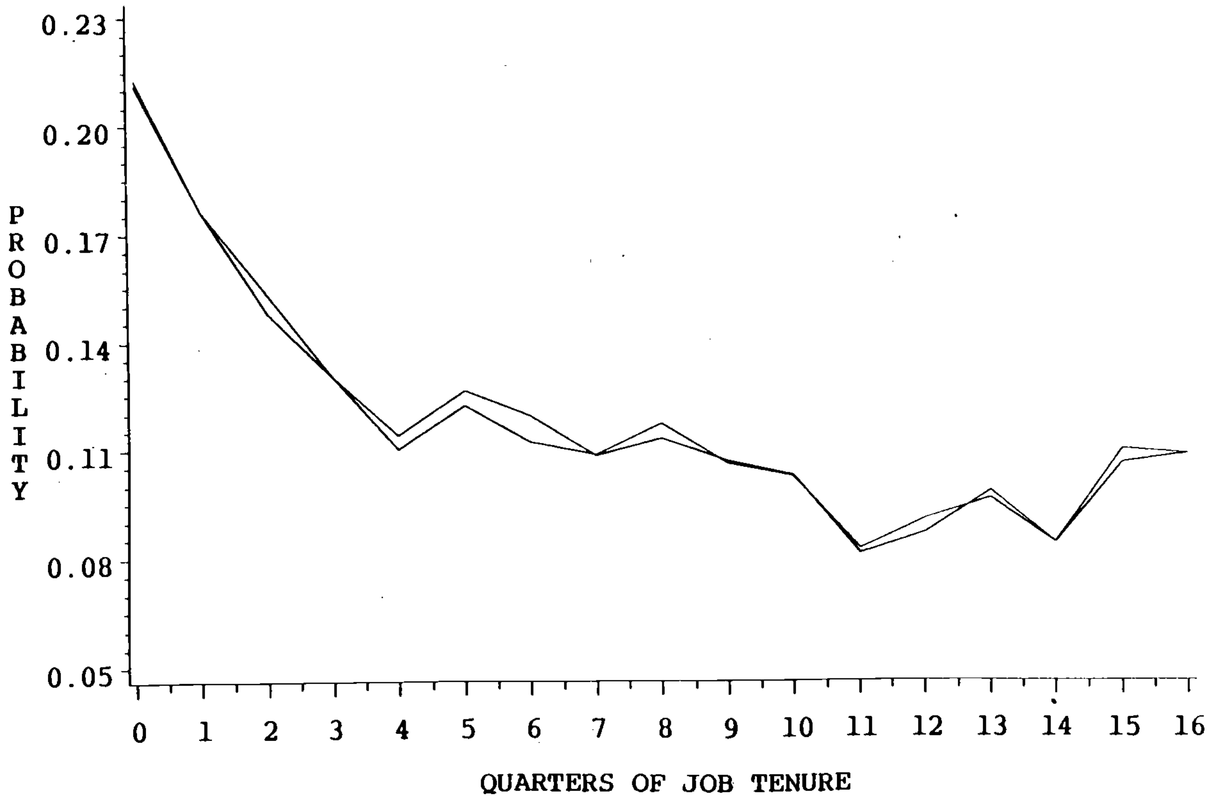

Figure 2 shows that in the unconditional model mobility declines with experience.[42] A worker with 12 years of labor market experience is about half as likely to leave a new job as is a new entrant. This decline conforms with our earlier tabulations. When mobility is conditioned on the wage, however, this pattern is reversed: wage constant, a worker with 12 years of experience is three times as likely to leave to leave his current job as an otherwise identical new entrant. This reversal in the experience profile of job mobility is what the theory predicts when wage offers increase with experience. More formally, the mobility data support the hypotheses of a positive return to search combined with $G_x(\cdot) < 0$.[43]

Figure 3 is a corresponding plot for the effects of job tenure. Conditioning on unobserved person effects and other regressors, mobility declines with seniority. Most of the decline occurs in the first year of a job, and the hazard is fairly flat beyond this point. Further, this shape is virtually unchanged when the estimates are conditioned on the wage; it is not even worth identifying which curve conditions on the wage. Unlike the pattern for experience effects in figure 2, there is no evidence that mobility increases with tenure when the wage is held constant. In retrospect this may not be surprising: the predicted sign reversal for experience is a direct implication of human capital growth over the career, while the prediction for the effect of job tenure relies on concavity of wage growth profiles on succesive jobs. While these and other data sources favor concavity, the evidence is not strong among these young workers (see table 6). Even so, if mobility declines with tenure

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[42]For the estimates in the figure, the model is evaluated at sample means of all regressors except tenure, which is set to zero for internal consistency (tenure cannot exceed experience).

[43]If mobility declined with experience simply because older workers are more stable, conditioning on the wage would not have the indicated effect. Thus the pattern illustrated in figure 2 is also evidence that the decline in the unconditional hazard is caused by the sorting of workers to successively more productive jobs.

because of accumulating specific capital reflected in the wage, conditioning on the wage should eliminate the relationship. It does not.

Table 8 reports parameter estimates for various forms of (10). For these models, inspection of the unconstrained estimates led us to specify experience and tenure effects as three-piece linear splines with break points at one and two years. In accord with the theory, which requires that the effect of experience be evaluated at the beginning of a job, we also allow for interactions between experience and tenure effects over each of these intervals. The specifications contain shifters for calendar quarter (summer is omitted), six firm size categories (the largest is omitted), and for whether the current job began with a direct transition from another job as opposed to nonemployment. The models also include 872 person-specific intercepts, for which we report the mean and standard deviation of the estimates. Because of the proportional hazards assumption (10), the parameter estimates in table 8 are approximate elasticities of the quarterly probability of leaving a job with respect to a unit change in the indicated variable.[44]

For reference the model in column (1) omits the wage and interactions between experience and tenure. In addition to the previously illustrated pattern of experience and tenure effects, we find that jobs are more stable in large firms. At means, the turnover rate in the smallest size class is about double that of the largest class. Since the estimates account for individual differences in propensities to change jobs, the interpretation is that transitions from small to large employers result in much more durable employment relations. There are several possible interpretations of this fact; the one we prefer is that large organizations encompass transitions that would otherwise occur between smaller ones. This "internal labor market" means that careers develop within the firm, though there may be no less mobility among tasks in large organizations. We also find lower turnover in jobs that

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[44]The interpretation is most accurate for small probabilities of leaving a job. For probabilities in the neighborhood of .1 per quarter, the approximation is very good.

TABLE 8

ESTIMATED PROPORTIONAL HAZARDS MODELS FOR JOB MOBILITY
LEED WHITE MALES WITH 13+ YEARS OF EXPERIENCE

	Mean (s.d.)	(1)	(2)	(3)	(4)
1. Current Wage	2.780 (.472)	-	-1.700 (.059)	-1.637 (.060)	-1.77 (.059)
× Quarter 1	.338 (.877)	-	-.236 (.048)	-.211 (.052)	-.303 (.052)
× Quarter 2	.254 (.785)	-	-.093 (.032)	-.125 (.033)	-.128 (.033)
2. Prior Job	.623 (.484)	-.221 (.040)	-.012 (.042)	.042 (.043)	.032 (.043)
3. Experience Spline:					
a. 0-1 Year	2.88 (.506)	.045 (.028)	.081 (.029)	.038 (.036)	.035 (.036)
b. 1-2 Years	6.847 (2.507)	-.016 (.007)	.053 (.008)	.051 (.013)	.055 (.013)
c. > 2 Years	15.989 (13.673)	-.014 (.0016)	.013 (.002)	.022 (.003)	.023 (.003)
4. Tenure Spline:					
a. 0-1 Year	2.33 (1.105)	-.189 (.016)	-.343 (.047)	-.448 (.072)	-.437 (.072)
b. 1-2 Years	3.892 (3.571)	-.039 (.008)	-.028 (.008)	-.092 (.020)	-.104 (.020)
c. > 2 Years	3.915 (7.601)	.016 (.005)	.014 (.005)	.015 (.008)	.015 (.008)
5. Experience × Tenure Interactions?	-	No	No	Yes	Yes
6. Firm Size:					
a. 1-9	.063 (.243)	.863 (.095)	.566 (.096)	.571 (.097)	-
b. 10-99	.321 (.467)	.670 (.091)	.473 (.091)	.475 (.092)	-
c. 100-499	.232 (.422)	.442 (.096)	.331 (.096)	.333 (.097)	-
d. 500-999	.055 (.227)	.099 (.116)	.053 (.116)	.056 (.116)	-
e. 1000-2499	.066 (.249)	-.186 (.126)	-.231 (.126)	-.223 (.123)	-
f. Not Reported	.212 (.409)	.978 (.092)	.534 (.093)	.533 (.093)	-

TABLE 8
(continued)

	Mean (s.d.)	(1)	(2)	(3)	(4)
7. Season					
Winter	.240	.267	.285	.288	.292
	(.427)	(.041)	(.041)	(.041)	(.041)
Spring	.244	.170	.172	.173	.174
	(.430)	(.039)	(.039)	(.039)	(.039)
Fall	.260	.272	.292	.293	.297
	(.439)	(.040)	(.040)	(.040)	(.040)
8. Individual Effects					
Mean	-	-2.206	-2.227	-2.328	-2.957
(s.d.)		(1.238)	(.825)	(.845)	(.868)
9. Log Likelihood	-	-12951	-12351	-12321	-12370
10. Observations	44089				

NOTE.--Figures in parentheses are asymptotic standard errors. Estimates for the tenure spline are evaluated at 12 years of experience. Six tenure-experience interactions are not reported. 872 individual effects are not reported.

began with a transition from a previous (covered) employer. This effect is implied by search: jobs that begin with a transition from non-employment are less valued, on average, because they need only dominate non-employment rather than a previous job.

Column (2) conditions on the log quarterly wage from the previous period. Because jobs may begin or end at any time, jobs lasting two or fewer quarters do not provide a complete quarter from which we may guage earnings. To account for this short job bias, we interact the quarterly wage with dummy variables for the first two quarters of a job. Thus the main effect of the wage refers to jobs that survive more than two quarters. We find that the job-specific wage is a key determinant of mobility: At means, the estimate implies that a 10 percent *within career* wage gain reduces the quarterly probability of leaving a job by about two percentage points. Since experience and tenure (and other variables) are held constant, this estimate is net of any systematic growth associated with these variables. We take this as strong evidence in favor of a search-based model of job changing.

Conditioning on the wage affects the other estimates. In column (1) we interpreted the effect of "Prior Job" in terms of selection on more valuable jobs. Consistent with this argument, the effect is eliminated by conditioning on the wage, which appears to be the key indicator of the value of a job. Firm size differentials also decline by roughly a third, indicating that size is in part a predictor of match quality. Nevertheless, the fact that substantial size differences in turnover remain after controlling for the wage also supports the scale interpretation offered above.

The estimates confirm the sign reversal in the effect of experience on mobility, which was illustrated in figure 2. The theory on this point refers to the effect of experience at the start of a job, so the specification in column (3)--which includes interactions between tenure and experience--is the most appropriate test of the hypothesis that experienced workers are more poorly matched, wage held constant. At means, the estimates imply that an extra five years of experience would raise the quarterly probability of changing jobs by about 50 percent (six percentage points).[45] As in figure 3, the tenure profile of job changing declines sharply during the first two years, after which the hazard is fairly flat.[46] And as in the figure, this pattern is not strongly affected by conditioning on the wage.

Table 9 reports variants of the basic mobility model that control for the starting wage on the job and for the rate of job-specific wage growth. In column (1), the only wage control is the starting wage, which sharply reduces mobility. This evidence ignores the evolution of wages within a job, but it indicates the important role of heterogeneity in available wage offers: *transitions to higher wage jobs stabilize employment*. This confirms the heuristic evidence from table 7, above, where we found that between-job wage growth was larger in transitions to more durable jobs. Notice that since within-job wage gains are ignored in this model, the interpretation of the tenure effect is altered so long as the value of the job grows with tenure. Consistent with this, the tenure profile of mobility is marginally steeper than in models that control for the current wage.

Column (2) allows the initial and current wages to have separate effects. If the current wage is fully informative about the future value of a job--an assumption of our model based on the random walk result in table 6--the starting wage should have no effect on behavior. This hypothesis is rejected ($t = 3.15$): Though the current wage clearly dominates in mobility decisions, a higher starting wage *increases* mobility, conditional on the current wage. This suggests a role for job-specific wage growth in affecting workers'

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[45]As a check on whether the estimate is reasonable, assume that experience has a linear effect, β, on the mean of the log wage offer distribution. Then the hazard is $\lambda = \{1 - G(R - \beta X)\}$. At $T = 0$, $R = w$. Thus $\beta = - \lambda_x/\lambda_w|_{T=0} = .022/1.64 = .0134$ per quarter, or 5.4 percent annual growth in wage offers due to the accumulation of general experience. This estimate is well within the range of commonly estimated experience effects on earnings.

[46] The reported estimates evaluate the experience-tenure interactions at 12 years of experience, because of the restriction that job tenure can be no larger than total experience.

TABLE 9

Estimated Hazard Functions for Job Mobility: Initial Wages and Wage Growth

	1	2	3
1. Initial Wage	-1.475	.473	-
	(.061)	(.150)	
2. Wage Growth	-	-	-5.318
			(1.276)
3. Current Wage	-	-2.09	-1.616
		(.143)	(.061)
× Quarter 1	-.387	-.330	-.331
	(.051)	(.052)	(.052)
× Quarter 2	-.185	-.147	-.148
	(.033)	(.034)	(.034)
4. Prior Job	.032	.034	.035
	(.043)	(.043)	(.043)
5. Experience Spline:			
a. 0-1 Year	.032	.038	.038
	(.036)	(.036)	(.036)
b. 1-2 Years	.047	.050	.050
	(.013)	(.013)	(.013)
c. > 2 Years	.019	.021	.021
	(.003)	(.003)	(.003)
6. Tenure Spline:			
a. 0-1 Year	-.541	-.476	-.483
	(.072)	(.072)	(.072)
b. 1-4 Years	-.094	-.091	-.087
	(.021)	(.020)	(.020)
c. > 4 Years	-.009	.024	.018
	(.008)	(.009)	(.008)
7. Experience × Tenure Interactions?	Yes	Yes	Yes
8. Individual Effects			
Mean	-2.099	-2.247	-2.249
(s.d.)	(.859)	(.845)	(.845)
9. Log Likelihood	-12418	-12311	-12307
10. Observations	44089		

NOTE.--Figures in parentheses are asymptotic standard errors. Estimates for the tenure spline are evaluated at the mean of experience. Other regressors from table 8 are included in the model, but not reported separately. 872 individuals are not reported.

mobility decisions. We test for this in column (3), which controls for the average rate of past wage growth on the job. We find that jobs offering higher wage growth are significantly less likely to end, holding the current wage fixed. At means, a five percent increase in annual job-specific wage growth would reduce the quarterly probability of changing jobs by about 0.8 points. This finding is reasonable if jobs systematically differ in their prospects for earnings growth, though it is somewhat puzzling in light of our previous evidence that within-job wage growth approximates a random walk. In that case, past wage growth on a job is not informative about the future value of the job; so it should not affect workers' mobility decisions.

5. Conclusions

We have documented the important role of job mobility as an element of career development among young workers. This career phase is characterized by a transition to relatively stable employment relationships along a path of high turnover and rapid wage growth. During the first ten years of labor force participation the typical young worker holds seven jobs, and our estimates imply that over one third of average wage growth during this period is attributable to job changing. This finding alone calls for a reevaluation of the standard human capital investment model of lifecycle earnings growth. Consistent with this, we have also shown that the job changing activities of young workers are strongly consistent with matching models of on the job search: controlling for unobserved heterogeneity, the key element leading to the eventual durability of jobs is the wage, growth of which is largely an outcome of the search process itself. Good matches tend to survive, and the decline in average mobility as experience accumulates is mainly attributable to locating such a match.

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APPENDIX A

JOB SEARCH WITH HUMAN CAPITAL GROWTH

In this appendix we verify claims made in text regarding the optimal mobility decisions of workers. The model is similar to the matching models of Jovanovic (1979a,b), though our analysis allows for (stochastic) growth of external wage offers--due to accumulation of general human capital--as well as stochastic growth of wage offers within a job. Specifically, we assume that the offer distribution facing an individual with labor market experience X is given by equation (5) in the text, where w^0 denotes a randomly drawn wage offer. If the location of the distribution is increasing in X, then $G_x(z; X) < 0$. This is assumed. We also assume that distribution (5) has uniformly bounded support on $X \geq 0$, and that the occurrence of offers from (5) is generated by a Poisson process with parameter π_1.

Internal wage growth is also stochastic. Recall that our evidence on internal wage growth is that wages evolve as a random walk with drift. To reflect this structure, we assume that the probability distribution of a new wage offer from the current employer is given by (6), which depends on the current wage, w, experience, X, and tenure, T. Thus on-the-job wage growth is a submartingale. We assume that a higher current wage increases the distribution of future offers ($F_w(\cdot) < 0$). If expected wage growth is nonincreasing in X and T then $F_X(\ ) \geq 0$, and $F_T(\ ) \geq 0$. The Poisson arrival rate of new offers on the current job is π_2, so wage growth is a 'jump process."

Let w be the current wage and t be the random arrival time of the first wage offer, either external (from (5)) or internal (from (6)), measured from the current period. The arrival rate of offers is $\pi = \pi_1 + \pi_2$, so t is exponentially distributed with parameter π. We assume for simplicity that leisure has no value, so workers will never quit in order to search from unemployment. For a generalization, see Mortensen. With a constant rate of time preference, r, the value of proceeding optimally on a current job paying w at experience X and tenure T is

$$V(w,X,T) = \frac{w}{r+\pi} \tag{A1}$$

$$+ \int_0^\infty \Big[\pi_1 \int \max[v(z;\ X+t,\ 0),\ v(w,\ X+t,\ T+t)]\Big] g(z;\ X+t)dz$$

$$+ \pi_2 \int_w v(y;\ w,\ X+t,\ T+t)f(y;\ w,\ X+t,\ T+t)dy]e^{-(r+\pi)t}dt.$$

Our boundedness assumptions on the distributions (5) and (6) guarantee that $V(\cdot)$ is a contraction from the space of bounded functions into itself (see Ross, 1970).

The properties of the mobility rule that were stated in the text depend on the effects of w, X, and T on the value function (A1). That $V(\cdot)$ is monotonically increasing in w is obvious: a higher current wage raises the value of the current job. The effects of tenure on experience are more complicated, since the value function is not generally differentiable in these arguments. Note, however, that the right side of (A1) maps the space of continuous functions that are increasing in w into itself. Thus if $F_T(\cdot) \geq 0$ then $V(\cdot)$ is nonincreasing in T because a leftward translation of the density $f(y;\ w,\ X,\ T)$ must decrease the expected value of a function that is increasing in w. Therefore $V(w,\ X,\ T) \geq V(w,\ X,\ T+\Delta T)$, with strict inequality if $F_T(\cdot) > 0$. Expected future wage growth declines with tenure, so an increase in current job tenure reduces the value of the job. In contrast, the effect of additional experience on $V(w,\ X,\ T)$ is unsigned: greater experience increases the distribution of external offers ($G_X(\cdot) < 0$), but reduces wage growth on the current job ($F_X() \geq 0$).

These effects on the value of a particular job are observable only through mobility decisions. Since $V(\cdot)$ is increasing in w, the worker's search policy will possess the reservation wage property. Thus there is a reservation offer, $R(w, X, T)$, defining the minimum acceptable outside wage offer, that satisfies

$$V(R(w, X, T), X, 0) = V(w, X, T). \tag{A2}$$

Clearly $R = w$ at zero tenure. Given $R(\cdot)$, the exit hazard function for job mobility is

$$\lambda(w, X, T) = \pi_1(1 - G(R(w, X, T); X)). \tag{A3}$$

Effects of w, X, T on mobility that are stated in the text follow from the effects of these variables on $V(\cdot)$, and hence $R(\cdot)$, through (A2) and (A3).

APPENDIX B

THE SAMPLE LIKELIHOOD FUNCTION

For each job spell in our data, we observe the (possibly truncated) length of each spell and a vector of constant-within-job or time-varying regressors that affect mobility. For individual i, let

t_{ij} = maximum observed duration of job j

d_{ij} = 1 if job ended in the interval $(t_{ij}, t_{ij}+1)$

= 0 otherwise (censored observations)

J_i = number of job spells for person i

$Z_{ij}(s)$ = regressors at job tenure s

μ_i = fixed individual effect.

Let $F(s; Z, \mu)$ be the cumulative distribution function of complete spell length s. Let $\bar{F}(\cdot) = 1 - F(\cdot)$ be the survivor function. Then the probability that spell j ends in the interval $(t_{ij}, t_{ij}+1)$ is

$$\bar{\lambda}(t_{ij}; Z, \mu_i) = 1 - \frac{\bar{F}(t_{ij}+1; Z, \mu_i)}{\bar{F}(t_{ij}; Z, \mu_i)}$$

$$= 1 - \exp\left\{- \int_{t_{ij}}^{t_{ij}+1} \lambda(s, Z(u), \mu_i) ds\right\} \qquad \text{(B1)}$$

where $\lambda(\cdot) = F(\cdot)/\bar{F}(\cdot)$ is the conditional density of failure times; the hazard function. The log likelihood contribution for the j^{th} spell is then

$$L_{ij} = d_{ij}\log \bar{\lambda}(t_{ij}, Z, \mu_i) + \sum_{r=0}^{t_{ij}} \log(1 - \bar{\lambda}(r, Z, \mu_i)). \tag{B2}$$

Letting $\lambda(s; Z, \mu_i) = \exp(Z(s)\beta + \mu_1)$, the log likelihood contribution of the i^{th} person is obtained by summing over j in (B2). In practice, we assume that the regressors $Z(s)$ are fixed within individual quarters of a job spell.

In maximizing (B2), we treat the 872 individual effects, μ_i, as estimable parameters. To calculate these, we modify the scoring method of Chamberlain (1980) for updating the parameter vectors β and μ. The procedure is as follows. Define the discrete hazard in period r of a worker's career, from (B1), as $\bar{\lambda}_{ir} = \bar{\lambda}(Z_{ir}\beta + \mu_i)$, which is a cumulative probability. Define $\ell_{ir} = \bar{X}_{ir}$, and let $\delta_{ir} = 1$ if a job ending occurred in period r, $\delta_{ir} = 0$ otherwise. Note that the likelihood is concave, so the Newton-Raphson method is appropriate for updating the parameter vector. At each iteration, let

$$h_{ir} = -\left[1 - \frac{\ell_{ir}}{\bar{\lambda}_{ir}}\right]\left[\frac{1-\delta_{ir}}{\bar{\lambda}_{ir}} - \frac{\delta_{ir}}{1-\bar{\lambda}_{ir}}\right]\ell_{ir} - \left[\frac{(1-\delta_{ir})}{\lambda_{ir}^2} + \frac{\delta_{ir}}{(1-\bar{\lambda}_{ir})^2}\right]\ell_{ir}^2 \tag{B3}$$

$$k_{ir} = \left[\frac{(1-\delta_{ir})}{\lambda_{ir}^2} - \frac{\delta_{ir}}{(1-\lambda_{ir})}\right]\ell_{ir}^2 \tag{B4}$$

$$h_i = \sum_r h_{ir} \tag{B5}$$

$$k_i = \sum_r k_{ir} \tag{B6}$$

$$Z_i = \frac{1}{h_i} \sum_r h_{ir} Z_{ir} \tag{B7}$$

$$Z = [Z_i] \tag{B8}$$

$$m = [k_i / h_i] \tag{B9}$$

Applying the partitioned inverse rule to the matrix of second derivatives of the likelihood, the formulae for updating β and μ at each iteration are:

$$\Delta\beta = - \left[\sum_{ir} h_{ir} Z'_{ir} Z_{ir} - \sum_i h_i Z'_i Z_i \right]^{-1} \left[\sum_{ir} k_{ir} Z_{ir} - \sum_i k_i Z_i \right] \tag{B10}$$

$$\Delta\mu = - Z\Delta\beta - m.$$

The asymptotic covariance matrix for β is given by the first matrix in (B10). In practice, this procedure converges in about seven iterations from $(\beta, \mu) = 0$. A computer program that performs the calculations is available on request.

APPENDIX C

THE DATA

The LEED file begins in the first quarter of 1957 and ends in 1972. Our basic criteria for extracting records from the file were:

1. The quarter of the individuals 18th birthday must occur in the data.
2. Each record must yield potential, post-18 experience of at least six years.
3. A record can contain no gaps in earnings that exceed two years.
4. Quarters of experience are accumulated when cumulative earnings from all reported employers exceeds 70 percent of the minimum wage.

Reported earnings in the data are actual earnings reported by an employer up to the Social Security limit in any calendar year. In some cases the limit was reached, in which case we imputed earnings from past and future earnings with the same employer. For example, for persons who broke the limit in the fourth quarter, imputed earnings are the average of third quarter earnings from year t and first quarter earnings from t+1 if no change of employer occured. For persons with a change of employer, third quarter earnings are used. In 50 cases out of 1,103,980 an individual broke the limit in the first quarter. These cases were deleted from the analysis of wage growth.

In some cases, persons showed temporary changes in the identity of the major employer, or temporarily fell below the 70 percent earnings cutoff though the identity of the employer did not change. We did not treat these cases as a break on the employment relationship; rather we smoothed over the temporary change in the identity of the major employer, so the periods are treated as a single job spell. Thus temporary changes in earnings do not generate job changes unless a change in employer also occurred.

These selection criteria resulted in 9919 individual records, and approximately 296,000 job quarters of data. Further selection for the analysis of wage growth and the impact of wages on mobility restricted the sample to persons with 13 or more years of labor market experience. For these persons "entry" was deemed to occur when earnings for a quarter first exceeded the cutoff, and subsequent earnings for the following year also exceeded the cutoff calculated on an annual basis. The resulting sample consisted of 872 persons who entered the market in or before the first quarter of 1959, with 44089 job quarters in the data.

The programs used in creating the working data files may be obtained from the authors on request. These files have been sent to the University of Michigan archives.