JOB MOBILITY AND THE CAREERS OF YOUNG MEN

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November 1986

RAND

The research described in this report was sponsored by the U.S. Department of Labor under Contract No. J-9-M-2-0163.

Published by The RAND Corporation
1700 Main Street, P.O. Box 2138, Santa Monica, CA 90406-2138

A RAND NOTE

N-2311-DOL

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PREFACE

American workers change jobs more than eleven times during their lives but most of these occur before the age of 30. The vast majority of these changes do not involve spells of unemployment and most are associated with increases in earnings. In this Note we study the process of job change among young men and ask about those characteristics of workers and firms that make for lasting employment relationships. The findings will be of interest to researchers and policymakers interested in employment and unemployment.

At the time this project began, Michael P. Ward was a senior economist at The RAND Corporation. He is currently a senior economist at Unicon Research Corporation. Robert H. Topel is a Professor of Business Economics and Industrial Relations at the Graduate School of Business of the University of Chicago.

This research was supported by a contract with the U.S. Department of Labor.

SUMMARY

Male workers in the United States will hold approximately 11 full time jobs during their working life. More than half of these will be held before age 30, and more than one-third before age 25. Job change patterns show that stable long-term employment is achieved only after a protracted period of job shopping. These years immediately following entry to the labor force form a critical phase of career development for young workers. During this period, the job-shopping process implies high rates of labor turnover and mobility as workers search for productive and durable employment relationships. Far from being a cause for concern, the transitory nature of jobs among young workers reflects the intensity and productivity of the search process itself.

We ask a series of questions about the factors that determine this pattern of labor mobility by focusing on job changes among younger workers between the ages of 18 and 33. We find strong evidence that declining mobility over careers is the outcome of a search process in which the workers sort themselves into better and better employment matches. On average, wages rise with each successive job. We find that as much as two-thirds of early career wage growth can be attributed to wage changes between jobs, as opposed to wage growth within-job.

Other findings include the following:

1. Among young men, more than 60 percent of all new jobs end in the first year and more than one-third end in the first three months. During the first ten years of his labor market life, the average male worker accumulates jobs at an average pace of about one job every two years.
2. The likelihood that a new job will last increases with labor market experience. While 40 percent of all first jobs end in the first three months, this frequency declines to 22 percent among new job holders with eight or more years of previous experience.

3. An important portion of rapid labor turnover among young workers is accounted for by weak labor force attachment. Early in careers, many jobs end in a transition to non-employment. More than 25 percent of all first jobs end in this way within three months. Among workers with eight years of prior experience, however, the corresponding figure is only four percent.
4. Young blacks hold jobs that are significantly less stable than those of whites. Shorter job durations among blacks are due both to higher rates of mobility between employers and more frequent transitions to non-employment. Among blacks, the likelihood of a non-employment spell declines with experience, but much less rapidly than for whites: Blacks with eight years of experience are twice as likely as whites to enter non-employment.
5. Jobs in large firms are more stable than in small ones. This finding remains even after controlling wages, which are also higher in large firms, and for unobserved differences in the characteristics of the workers in these forms. Within-career moves from small to large employers significantly reduce the likelihood of future mobility.

ACKNOWLEDGMENTS

For their comments on this and earlier drafts we thank John Abowd, Gary Becker, Kevin Murphy, Melvin Reder, Sherwin Rosen, and David Ross. RAND reviewer James Dertouzos gave us a careful and useful critique. We have also benefited from comments from participants at seminars given at the University of Chicago, Northwestern University, National Bureau of Economic Research, University of Florida, New York University and Yale University. David Ross provided very capable assistance.

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I. INTRODUCTION

Among male workers in the United States, the average completed length of a job is less than four years. Nevertheless, most employment occurs in jobs that last much longer: The typical male worker is now holding a job that can be expected to last 18 years (Akerlof and Main, 1982). The major feature of the labor market that supports these disparate facts is that the distribution of job-changing is heavily skewed toward young workers. More than half of the approximately 11 full-time jobs that the typical worker will hold during his labor-market life occur before age 30, and more than one-third before age 25 (Hall, 1982). Evidently, quite durable jobs are a prime characteristic of employment relations for the majority of the prime-aged labor force, yet the early phase of a worker's career is commonly one of rapid turnover and only weak attachment to particular jobs.

These facts are of vital importance to a number of modern theories of employment relations and labor market dynamics. For example, the relevance of popular "implicit contract" models of the labor market, which proceed from the assumption of durable employment relationships, would be called into question if long-term employment with a single firm is the exception rather than the rule for most workers.[1] The description applies to the many "lifetime" jobs that are important components of older workers' careers, yet the contract paradigm is clearly less attractive when applied to younger workers who are in the process of sorting themselves into durable jobs. For them the labor market is dynamic, characterized by intensive search for a productive employment relationship, while the expected length of any new job is quite short.

In this Note we begin to study the determinants of early career mobility and the process by which workers sort themselves into durable employment relationships. We study three related questions. First,

[1]Early examples are Azariadis (1975) and Hall and Lilien (1979). For contract models that rely explicitly on long job durations, see Lazear (1979, 1981).

what do careers "look like" in terms of job mobility and productivity growth? These basic facts about the market for young workers and the extent of their mobility are largely unknown. Second, what factors determine patterns of labor mobility? Here we are interested in effects that operate across individuals at a point in time--for example, sectoral or demographic differences in mobility--as well as those that determine the timing of job changes within individual careers. Finally, what theory of individual decisionmaking underlies observed patterns of job changing? Is the transitory nature of young people's job a cause for concern, or is rapid turnover merely the consequence of productive search activity?

To answer these questions, we study a large sample of individual labor-market histories drawn from a unique and rich source of panel data, the Longitudinal Employee-Employer Data (LEED) file, which is based on the Social Security Administration's Continuous Work History Sample. These data provide quarterly mobility, job, and earnings histories for over 1.3 million individuals during the 16 years 1957-1972, including firm-specific information for each employer that made OASDI contributions to a worker's account during any three-month period. Unlike BLS surveys or other sources of panel data that have been used to study job mobility, the LEED file is a complete longitudinal history without loss of memory. In combination with its size and the length of the panel, this feature makes LEED a unique source of information on the short-term nature of job-shopping and labor turnover.

In what follows we focus exclusively on young workers for two reasons. First, the processes of acquiring information about jobs and human capital investment are most important for recent entrants to the labor force. Thus, to understand the transition to "permanent" employment relationships we want to follow the sorting process that leads to stable matches, beginning with market entry. Second, the analysis of job changes among older workers is subject to inference problems associated with "left censoring" of employment spells and careers. In order to describe the likelihood of an older worker changing jobs, we must understand the process that brought him to his current position--but employment events that occurred before the beginning of most panel data sets are unrecorded. In contrast, the labor market histories that we analyze are complete. They follow young individuals (between 18 and 33 years old) from their entry into the market. Our sample represents a random drawing from the population of entrants, and so we view our results as being typical of the market for young workers. In contrast to other sources of panel data, this restricted focus on uncensored careers is not a meaningful limitation on the available sample size from the LEED file.

In contrast to early, nonbehavioral models of labor mobility such as the "mover-stayer" model,[2] our evidence points to job-shopping (search) and the accumulation of job-specific human capital as dominant elements affecting the mobility decision of young workers. Specifically, we find strong evidence that declining average mobility over careers is the outcome of a search process as workers sort themselves into "good" employment matches. Average match quality rises with experience as new jobs are sampled and, within an individual's career, better employment matches are significantly less likely to terminate. Within a particular worker-firm pairing, we find that the probability of termination declines with current job tenure (though not as rapidly as aggregate tabulations of labor mobility would suggest) implying the importance of specific capital and information accumulation on the job. In contrast, controlling for job-specific wage growth, we find that both current job tenure and labor market experience *increase* mobility. As we demonstrate, this sign reversal is a key implication of human capital and learning models. The key econometric evidence on this points relies crucially on (i) observable wages on jobs as indicators of changing match quality during careers, and (ii) the length of the panel, which allows us to follow the typical worker over several job spells.

Other results of our analysis include the following:

1. Job mobility among young men is remarkably fluid. Among the young workers we study, more than 60 percent of all new jobs end in the first year, and more than 30 percent end in the

[2]See Blumen, Kogan, and McCarthy (1955) or Singer and Spilerman (1976) for a survey.

first three months. During the first 10 years of his labor market life, the average male worker accumulates jobs at an average pace of about .5 jobs per year.

2. New entrants to the labor force are significantly more mobile than experienced workers. While 40 percent of all first jobs end in the first three months, this frequency declines to (a still high) 22 percent among new job holders with eight or more years of previous experience.
3. An important portion of rapid labor turnover among young workers is accounted for by weak labor force attachment. Early in careers, many jobs end in a transition to non-employment. More than 25 percent of all first jobs end in this way within three months. Among workers with eight years of prior experience, however, the corresponding figure is only four percent.
4. Young blacks hold jobs that are significantly less stable than those of whites. Shorter job durations among blacks are due both to higher rates of mobility between employers and more frequent transitions to non-employment. Among blacks, the likelihood of a non-employment spell declines with experience, but much less rapidly than for whites: Blacks with eight years of experience are twice as likely as whites to enter non-employment.
5. Jobs in large firms are more stable than in small ones. This finding remains even after controlling wages, which are also higher in large firms, and for unobserved differences in the characteristics of the workers in these firms. Within-career moves from small to large employers significantly reduce the likelihood of future mobility.

The Note is organized as follows. Section II briefly reviews theories of labor turnover and their implications for observed patterns of labor mobility. It then describes the LEED sample and presents our main tabulations of the extent of job changing among young workers. Following this documentation of main "facts", Sec. III provides more formal econometric evidence on the decisions that determine mobility within individual careers. Section IV summarizes our findings and discusses further research in this area.

II. MODELS OF LABOR TURNOVER

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 job tenure.[1] Most theories of turnover are designed to accommodate these crude facts, though they differ in fundamental details. For our purposes, we may categorize these theories into three broad groups.

Nonbehavioral models have the longest history (Blumen, Kogan, 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 tenure acts as filter, systematically selecting on individuals who have low probabilities of terminating a worker-firm pairing. Thus any tabulation of the frequency of mobility in such a sample will show a negative relationship between the probability of moving and current job tenure, since "stayers" are more likely to survive, though for any individual the probability of moving may be independent of tenure. Statistical methods of accounting for this type of bias in duration models due to unobserved heterogeneity are the subject of a growing literature (e.g., Heckman and Singer, 1982, 1984; Elbers and Ridder, 1982).

A closely related source of bias is generated in optimizing models of individual search for a good employment match (e.g., Jovanovic 1979b). Following the terminology of Nelson (1970), 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, so with information only on the duration of job spells, this source of unobserved heterogeneity generates the same types of bias as above. The aggregate frequency of job mobility will decline with tenure, even if tenure exerts no direct influence on an individual's probability of

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

changing jobs. In addition, holding tenure constant, this type of search implies that mobility should 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 (the current job) is higher for workers who have searched longer. A key issue addressed below is the extent to which this 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," so the worker learns about match quality by observing his productivity over time (Jovanovic, 1979a). Uncertainty about the true quality of a match then 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.[2] 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 is independent of time in the market (experience), holding tenure constant.

None of these extreme models is likely to completely describe mobility data; all will play a role. Yet except under extreme circumstances, these models are observationally equivalent in data on job durations alone. This fact suggests that additional information on the productivities of particular job pairings would provide important identifying leverage of distinguishing competing theories. The observability of wages, in conjunction with the long-panel aspects of the LEED file, plays this role in Sec. 3 below. First, however, we concentrate on describing the most prominent features of labor mobility among young workers.

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[2]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.

III. JOB MOBILITY AMONG YOUNG WORKERS

THE DATA

The LEED file contains quarterly employee-employer records for each individual, including Social Security earnings credited to the worker's account during each quarter, a unique employer identifier, 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. So as to measure labor market events from the beginning of careers, we selected only individuals who were born after 1938. Thus the oldest person in our data is 18 in 1957. All others have their histories measured starting before their 18th birthday, or when they are first recorded in the data. Consequently, the oldest person in these data is 34 years old in 1972, having accumulated 64 quarters of experience during the panel.

Among the young men who compose our sample, multiple job holding, rapid turnover, and return to past employers are common. Transitory or part-time jobs 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. The break is not as dramatic as full-time schooling models suggest, but rather seems best characterized as a gradual switch from school or leisure to full-time employment along a path of high turnover and intermittent employment.

In light of these facts, our definition of the length of a job was influenced primarily by our initial inspection of the data. We first 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. Job tenure is accumulated continuous quarters with a single employer.[1]

Three final requirements were imposed on individual records that were selected for analysis. First, because of ambiguities about the meaning of jobs held while very young, we begin each career with the first "full-time" job that is held on or after the person's 18th birthday. Second, because of the possibility that individuals may enter uncovered employment, for example, the military, we exclude records with continuous non-employment gaps of two years or more.[2] Finally, we require that each record analyzed have at least six years of potential market experience after initial entry, so individuals who entered the labor force after 1966 are excluded.

MOBILITY

All of the following results are based on the occurrence or termination of a worker-firm pairing as identifiable economic events, ignoring such subcategories as geographic or sectoral mobility for which these data are also suited. We confine most of our analysis to white males, of which there are 9527 individuals with 51,331 full-time jobs in our data.[3] We begin with Table 1, which shows the age distribution of first entry to the labor market under various definitions of labor force attachment, which we take to depend on the length of the initial

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[1]In cases 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.

[2]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 limit on non-participation.

[3]Sample size could be expanded, but we sought approximately 10,000 "clean" individual histories for computational convenience.

employment spell.[4] Over 80 percent of young men have held a substantial job of some sort by age 20 and over half by age 18, yet weak attachment to the labor force is a prime characteristic of new entrants: The distribution of first employment spells lasting a year or more is shifted sharply to the right relative to the age distribution of first employment. Less than one-third of spells lasting a year or more occur by age 18, implying that early spells are likely to end with a transition to non-employment.

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 our data, showing the distribution of actual labor force experience for a sample of individuals with exactly 10 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 nearly one-third of these workers are continuously employed over the first 10 years of their careers, and almost 60 percent have spent less than one

Table 1

AGE DISTRIBUTIONS OF ENTERING THE LABOR FORCE: VARIOUS DURATIONS OF FULL-TIME EMPLOYMENT

	Age at Beginning of Employment Spell							
Length of Spell	18	19	20	21	22	23	24	> 25
> 1 quarter	54.39	17.55	9.29	6.28	4.76	3.68	2.00	2.05
> 2 quarters	40.89	19.48	12.17	9.06	7.89	4.92	2.21	3.68
> 3 quarters	34.21	19.88	12.81	10.18	8.53	6.35	3.74	4.31
> 1 year	30.16	18.98	13.42	11.10	9.16	7.15	4.38	5.55

NOTE: Reported figures are proportions of individuals born between 1938 and 1943 who first entered the labor force for the indicated period of time at the indicated age. Sample is 3888 white males.

[4]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.

cumulative year without a job. Nevertheless, panel B shows that the distribution of non-employment time is heavily skewed toward the earliest part of careers. During the first few years in the market, nearly a 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.

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 five or more jobs, and nearly 20 percent have held eight jobs or more. Only one worker in 20 has held a single job for ten years, though we saw in Table 2 that six times this many are continuously employed. Thus, even for continuously employed individuals, rapid job changing is an important component of early careers. The average number of jobs by this career point is 5.54, which is very close to Hall's (1982) estimate for similarly aged workers, which was derived by a far different sampling procedure.[5] Note, however, that our definitions preclude multiple job holding as contributing to this count, so we probably 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 between two jobs. The average pace of sampling jobs then declines fairly smoothly as experience accumulates.[6] 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 model of the pure

[5]Hall estimates that workers aged 25-29 have held 5.5 jobs and that workers aged 20-24 have held four. During the age interval 20-24, Hall estimates that the average worker holds 2.1 jobs. Our corresponding estimate is 2.4. The close correspondence between his estimates and ours is support for the representativeness of our sample.

[6]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 2

A. DISTRIBUTION OF ACTUAL LABOR MARKET EXPERIENCE FOR PERSONS WITH 10 YEARS OF POTENTIAL EXPERIENCE (N = 3397)

	Years					Quarters							
	< 4	5	6	7	8	33	34	35	36	37	38	39	40
Percent with indicated experience	0.12	0.56	2.35	6.27	11.48	4.53	5.80	4.30	6.12	8.27	10.01	8.24	31.94

B. ACTUAL LABOR MARKET EXPERIENCE BY YEARS OF POTENTIAL EXPERIENCE (Full Sample)

	Potential Market Experience (Years)									
	1	2	3	4	5	6	7	8	9	10
Actual market experience	.77	1.53	2.35	3.21	4.12	5.10	6.10	7.07	8.03	8.98
Average additional experience	.77	.76	.82	.86	.92	.98	1.00	.97	.96	.95

Table 3

A. DISTRIBUTION OF CUMULATIVE FULL-TIME JOBS FOR PERSON WITH 10 YEARS OF POTENTIAL EXPERIENCE (N = 3397)

	Cumulative Full-Time Jobs												
	1	2	3	4	5	6	7	8	9	10	11	12	13
Percent with indicated number of jobs	5.27	10.51	13.37	14.31	13.81	10.89	9.01	6.06	5.15	3.12	6.71	1.21	0.15

B. ACTUAL 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	1.51	2.10	2.74	3.34	3.83	4.27	4.54	4.87	5.20	5.54
new jobs	1.51	.59	.64	.60	.49	.44	.27	.33	.33	.34
By actual experience	2.00	2.79	3.36	3.80	4.13	4.41	4.70	4.95	5.09	5.22
new jobs	2.00	.79	.57	.44	.333	.28	.29	.25	.14	.13

mover-stayer type, in which new jobs within careers would form a renewal process. That is, 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 in Fig. 1, which shows the conditional relative frequency of job terminations given current tenure--the empirical exit-hazard function. 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 out of every three end within a year. The probability of moving declines dramatically over the early periods 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 data in figure 1 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. (Table 5 translates these data into average durations of jobs.) Even after eight full-time years in the market, the quarterly rates imply that 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 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. On average, new jobs are more durable the longer the duration of prior job

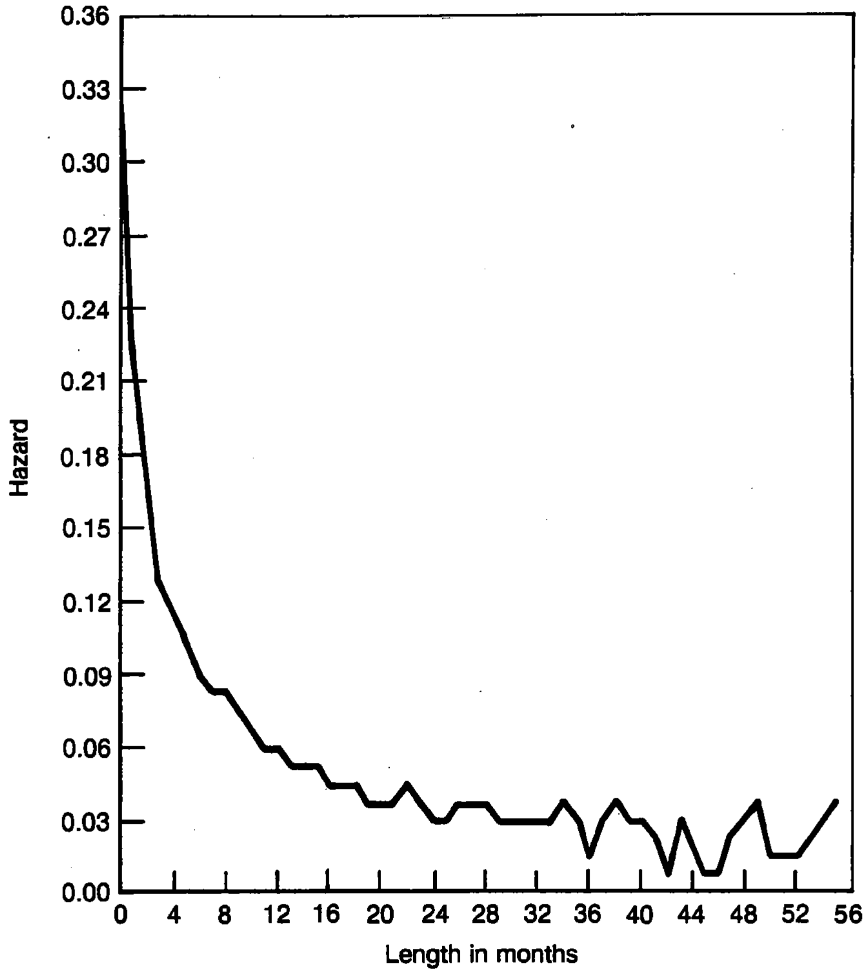

Fig. 1—The aggregate exit-hazard function

Table 4

EMPIRICAL MOBILITY FUNCTIONS BY PRIOR MARKET EXPERIENCE, NUMBER OF PRIOR JOBS, AND INDUSTRY

	Current Job Tenure (Quarters)												
	1	2	3	4	5	6	7	8	9	10	11	12	13
A. Prior Experience													
1 year or less	.44	.27	.17	.13	.13	.11	.09	.09	.09	.08	.08	.07	.05
1-2 years	.33	.23	.17	.14	.11	.11	.11	.09	.09	.08	.07	.06	.06
2-4 years	.29	.22	.16	.13	.11	.10	.09	.08	.07	.07	.06	.06	.07
4-8 years	.26	.19	.14	.11	.11	.09	.08	.07	.07	.06	.05	.04	.05
More than 8 years	.22	.16	.11	.10	.08	.06	.08	.06	.06	.07	.07	.06	.07
B. Prior Jobs													
None	.40	.22	.15	.14	.12	.11	.10	.09	.09	.09	.08	.07	.07
1	.35	.22	.15	.10	.11	.10	.10	.08	.07	.08	.06	.07	.05
2	.33	.21	.14	.12	.11	.10	.09	.08	.08	.08	.06	.06	.06
3 or 4	.30	.21	.14	.12	.10	.09	.09	.08	.07	.06	.06	.05	.05
5-7	.29	.22	.16	.13	.12	.10	.09	.09	.09	.08	.06	.06	.08
8 or more	.39	.30	.22	.16	.14	.15	.13	.10	.10	.12	.06	.06	.06
C. Industry													
Finance	.25	.18	.14	.10	.10	.13	.08	.06	.10	.09	.04	.06	.05
Construction	.44	.33	.21	.19	.16	.14	.13	.11	.11	.09	.10	.08	.07
Wholesale Trade	.30	.19	.14	.11	.10	.10	.10	.08	.09	.09	.08	.07	.07
Retail Trade	.36	.26	.19	.16	.15	.13	.12	.10	.10	.10	.08	.07	.08
Manufacturing	.30	.19	.13	.10	.10	.09	.08	.08	.07	.07	.06	.05	.06
Aggregate Hazard	.34	.22	.15	.13	.11	.10	.09	.08	.08	.08	.07	.06	.06

search. Of course, among these new job pairings those of lower quality are more likely to end early, which partly accounts for the shape of the hazard.

Panels B and C of Table 4 show empirical hazards by number of prior jobs and selected industries in order to illustrate two key points 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 very large number of prior jobs. The decline is less prominent than in panel A, and it is mainly concentrated in the first quarter. 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 job. This may be because some people are "movers" or because of differences in job search histories. For example, if workers search systematically for a productive match, then a sequence of low-productivity prior job matches implies a lower expected quality of current job, raising average turnover in the population of new jobs. The second point is related. Industries and occupations may differ in the firm-specificity of human capital that workers acquire, or in other factors that affect mobility. Mobility may then be higher in sectors that offer more general (or sector specific) skills. This effect is illustrated in panel C for selected major industries. Over 75 percent of all new jobs in construction end in the first year, and the expected duration of these jobs is only one year. In contrast, jobs in the financial sector and in manufacturing industries are much more stable, lasting more than twice as long on average (see Table 5). To the extent that workers invest in skills that are specific only to such sectors, this heterogeneity in mobility across industries implies the existence of (endogenous) "movers" and "stayers" at the individual level.

Table 5

EXPECTED REMAINING DURATION OF A JOB BY YEARS OF CURRENT TENURE, YEARS OF PRIOR EXPERIENCE, NUMBER OF JOBS, AND INDUSTRY

Item	Current Job Tenure (Years)					
	0	1	2	3	4	5
A. Prior Experience						
1-4 quarters	1.38	3.72	4.98	5.93	6.51	6.88
2-4 years	1.86	3.95	5.15	5.95	6.51	6.87
5-7 years	2.15	4.14	5.35	6.06	6.59	6.87
B. Prior Jobs						
None	1.78	3.97	5.20	6.67	7.05	7.33
3 or 4	2.02	4.24	5.49	6.36	6.90	7.36
C. Industry						
Construction	0.99	2.81	3.83	4.46	4.71	4.76
Manufacturing	2.56	5.25	7.00	8.20	9.27	10.20
Finance	2.17	4.07	5.26	6.22	6.96	7.42

LABOR FORCE ATTACHMENT AND TYPES OF JOB ENDINGS

In discussing Table 2, we noted that labor force attachment tends to be weak in the early phases of careers. Table 6 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, fully 63 percent of all first-year job endings result in a transition to non-employment. 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 non-employment in the first year is .10.[7] It is impossible for

[7]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.

Table 6

MOBILITY FROM CURRENT JOB BY TYPE OF TRANSITION
AND EXPERIENCNE AT START OF JOB
(White Males)

Prior Experience (x)		Current Job Tenure (Quarters) 1	2	3	4	5	6	7	8	9	10	11	12	13
$x = 0$	$j \to j$	.105	.090	.070	.074	.056	.061	.055	.053	.041	.045	.043	.039	.036
	$j \to n$	.262	.112	.062	.051	.050	.038	.028	.019	.032	.022	.018	.012	.022
$1 \leq x \leq 4$	$j \to j$	.137	.117	.093	.071	.069	.069	.052	.054	.042	.052	.036	.034	.024
	$j \to n$	.260	.126	.061	.043	.047	.031	.023	.019	.027	.013	.019	.019	.019
$5 \leq x \leq 8$	$j \to j$	.173	.125	.102	.084	.068	.053	.059	.050	.044	.049	.038	.033	.033
	$j \to n$	.134	.078	.042	.033	.029	.032	.021	.018	.017	.015	.011	.008	.009
$9 \leq x \leq 16$	$j \to j$	.171	.126	.094	.069	.059	.060	.049	.040	.040	.041	.031	.029	.028
	$j \to n$	.092	.064	.037	.026	.025	.020	.021	.015	.011	.014	.011	.007	.012
$17 \leq x \leq 32$	$j \to j$	.152	.110	.080	.058	.057	.048	.042	.037	.031	.027	.025	.018	.018
	$j \to n$	.054	.039	.027	.024	.016	.014	.014	.011	.010	.011	.007	.009	.004
$32 \leq x$	$j \to j$	.124	.078	.049	.040	.031	.030	.033	.018	.020	.013	.020	.019	.012
	$j \to n$	.041	.035	.023	.018	.015	.008	.014	.012	.016	.021	.015	.013	.008

NOTE: $j \to j$ denotes job to job transitions (a change of major employer)
$j \to n$ denotes job to "not employed" transitions

us to tell in these data whether these high "non-employment" flows represent withdrawals from the labor force or transitions into unemployment. Nevertheless, these data suggest that a prime component of well-known higher unemployment rates among young workers is their vastly higher flows out of jobs rather than the durations of their spells. In fact, related evidence on unemployment transitions in Topel (1984) indicates that unemployment durations are substantially shorter among workers with little labor market experience, though their flows into unemployment are much higher. A plausible interpretation of these facts is that the costs of unemployment (or non-employment) are quite low for young workers, so changes in employment status are frequent.

Note also that the absolute importance of job-to-job transitions actually 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.

Black-White Differences

We defer a detailed examination of demographic differences in labor mobility to a companion paper. Here, we touch briefly on black-white differences because of the role that job stability may play in generating racial differences in human capital growth, and also because transition to non-employment may be a prime factor in generating higher unemployment among young blacks. A number of studies (Hall, 1971; Ehrenberg, 1980; Topel, 1984) have found that higher unemployment rates among blacks are mainly due to more frequent transitions into unemployment; the durations of their spells are not materially different from those of other groups. Yet Hall (1982) presents evidence that black *jobs* are no less stable.[8] The implication is that job-to-job

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If anything, we probably understate non-employment transitions by our method.

[8]Hall (1982) concludes that "the vastly higher incidence of unemployment among blacks is not at all the result of higher flows of workers out of jobs." This assumes that all job endings result in unemployment. For the opposite conclusion, see Hall, (1971).

transitions must be lower for blacks if their job durations are the same.

Our evidence contradicts Hall's. Blacks' jobs are less stable than whites'. Table 7 reports empirical mobility functions for young blacks, stratified by years of experience and type of transition. Comparison with the conformable tabulations for whites in Table 6 reveals that black jobs are more likely to terminate at all levels of experience and tenure. Their jobs are substantially shorter. This occurs both because blacks are more likely to transit to non-employment *and* because they are more likely to change employers. For example, among blacks with eight or more years of experience, only 31 percent survive the first year on a new job, which is less than half the corresponding figure for whites (65 percent). The probability of a first-year non-employment transition among blacks with eight years of experience is .20, also double the white rate. Importantly, the sharp decline in mobility with experience that was observed among young whites is not nearly so apparent among blacks. Among blacks with eight years in the market, the probability of surviving one year in a new job (.31) is almost the same as for completely inexperienced workers (.35), while experienced white workers are much more stably employed than new entrants (.65 and .38 survival probabilities). In fact, the job-to-job mobility rate *rises* across all experience levels, while initial jobs are fairly similar to those of young whites. If the pattern for whites indicates a "settling in" to a good employment match over their careers--as our evidence below indicates--then a possible implication is that "good" worker-firm pairings are much rarer in the types of jobs held by blacks. The productivity growth that is an outcome of the job search process may then be much less prominent in the careers of blacks, with obvious implications for subsequent earning capacity.

Firm Size and Sectoral Differences

As we illustrated in Table 4, there are important sectoral differences in rates of job mobility and job duration. While these between-industry differences in job durations are interesting in their own right, they may also exert important influence on the shape of the aggregate mobility functions tabulated above. To investigate these

Table 7

MOBILITY FROM CURRENT JOB BY TYPE OF TRANSITION
AND EXPERIENCE AT START OF JOB
(Black Males)

Prior Experience (x)		Current Job Tenure (Quarters) 1	2	3	4	5	6	7	8	9	10	11	12	13
x = 0	j j	.103	.081	.072	.082	.069	.062	.065	.053	.054	.055	.047	.036	.055
	j n	.283	.140	.104	.081	.070	.066	.047	.042	.057	.042	.035	.040	.030
1 < x < 4	j j	.150	.122	.092	.088	.085	.087	.069	.066	.053	.060	.067	.043	.065
	j n	.259	.149	.090	.070	.066	.062	.051	.057	.043	.042	.034	.036	.031
5 < x < 8	j j	.175	.154	.105	.105	.093	.079	.083	.065	.075	.076	.086	.079	.073
	j n	.203	.113	.088	.063	.058	.061	.043	.029	.049	.042	.026	.027	.037
9 < x < 16	j j	.204	.160	.126	.094	.108	.090	.080	.100	.124	.100	.109	.105	.099
	j n	.150	.102	.074	.065	.051	.051	.047	.037	.029	.026	.024	.018	.032
17 < x < 32	j j	.234	.184	.152	.142	.135	.129	.098	.116	.126	.110	.110	.120	.115
	j n	.106	.073	.047	.037	.031	.040	.032	.029	.030	.021	.023	.019	.015
32 < x	j j	.254	.176	.165	.177	.175	.126	.111	-*	-	-	-	-	-
	j n	.082	.072	.044	.028	.024	.041	.018	-*	-	-	-	-	-

NOTE: j j denotes job to job transitions (a change of major employer).
j n denotes job to "not employed" transitions.
* less than 200 jobs

Sample consists of 5286 black males, holding 28,566 jobs. Selection criteria are the same as for the white sample. See text for details.

issues, we estimated by maximum likelihood "proportional hazards"[9] models of the form:

$$\lambda(t;x,I) = \exp\{h(t) + x\beta + \alpha(I)\}, \quad (1)$$

where $\lambda(t)$ is the conditional probability of a job spell ending at tenure t given survival to t, h(t) is a specified function of tenure, and $\alpha(I)$ refers to a shifting intercept that applies to each of 64 two-digit (SIC) industry classifications. Thus, the estimated hazard is allowed to shift by industry. Parameter estimates for Eq. (1) for the full sample of 51,331 job spells are shown in Table 8, where we report specifications estimated with and without the industry effects. Since we allow for no regressors other than job tenure and experience (and firm size), the first specification is simply fitting Eq. (1) to the tabulated data reported above.

Somewhat surprisingly, controlling for heterogeneity across detailed industry in this way has only a small effect on the estimated shape of the empirical hazard function. The function flattens with respect to tenure (as it must), but only slightly. This is not because the estimated industry differences are trivial. The (unweighted) mean industry effect is -2.51 with a standard deviation of 0.25. Jobs are least durable in construction (-2.07) and gasoline stations (-2.18), while mobility is lowest in communications (-3.13) and in banking (-3.00). Note that the fact that the experience effect in column (2) is only slightly different than in column (1) implies that the observed relation between mobility and experience is not caused by systematic migration of workers to more stable sectors of the economy over their careers. Within-industry accumulation in experience also stabilizes employment.

The table also reports effects of firm size on mobility.[10] It is well known that workers in large firms earn higher wages, other (observable) things equal. Our evidence indicates that their jobs are

[9]See Kalbfleish and Prentice (1980) for an introduction to statistical models for duration data. Details of the likelihood function for Eq. (1) are appended.

[10]The omitted size category is firms with 2500 or more employees.

Table 8

PROPORTIONAL HAZARDS MODELS OF JOB MOBILITY: CONTROLLING FOR SECTORAL DIFFERENCES

Item	Without Industry Effects	With Industry Effects	Distribution of Industry Effects (Unweighted)	
Quarter 1	1.3336	1.255		
	(.0367)	(.0317)		
Quarter 2	.8843	.8342	mean:	-2.51
	(.0407)	(.0433)		
Quarter 3	.5705	.5407	St. Dev:	0.25
	(.0479)	(.0527)		
Quarter 4	.3979	.3766	Maximum	-2.07
	(.0532)	(.0597)		
Quarter 5	.2787	.2637	Minimum	-3.13
	(.0583)	(.0751)		
Job tenure	-.0404	-.0376		
	(.0031)	(.0033)		
Experience	-.0168	-.0171		
	(.0012)	(.0009)		
Firm size:				
1-9	.4557	.2546		
	(.0593)	(.0711)		
10-99	.5966	.3547		
	(.0609)	(.0658)		
100-499	.4252	.2629		
	(.0639)	(.0629)		
500-999	.3216	.2019		
	(.0782)	(.0916)		
1000-2499	.2266	.1647		
	(.0783)	(.0962)		
Intercept	-2.7470	--		
	(.0600)			
ℓn likelihood		-64276.6		

NOTE: For specification of the sample likelihood function and method of calculating fixed effects, see the Appendix. Figures in parentheses are asymptotic standard errors. Sixty-four industry-specific intercepts not reported.

also significantly more stable. This may reflect Oi's (1983) notion that human capital is more job-specific in large firms, or more simply that mobility takes place within firms when scale permits the operation of a large internal labor market. Though average firm size differs substantially across industries, the effect is only partially reduced when detailed industry of employment is controlled for, indicating that within-industry job changes from small to large employers also result in more durable employment relations.

These facts imply that sectoral differences in labor turnover, while numerically important, are not prime factors in affecting the career patterns of job changing that we observed above. But these tabulations offer little guidance on the decisions that actually generate job changing activity. We turn to more formal econometric evidence on the job shopping process.

IV. ECONOMETRIC EVIDENCE ON THE JOB-SHOPPING PROCESS

THEORIES AND MODELS OF JOB-SHOPPING

As we noted in Sec. 1, there are a number of more or less complementary theories of the job-shopping process that seek to rationalize known facts about labor turnover. In Nelson's (1970) terminology, these models treat sampled jobs as some combination of "search" or "experience" goods, possibly also with opportunities for on-the-job accumulation of job-specific human capital (Jovanovic, 1979a,b; 1984). To fix ideas on the issues raised by these models, consider the following simple model of jobs as search goods with on-the-job accumulation of match-specific productivity. Let

ϕ = initial match-specific capital (productivity), known at t=0,

$G(\phi)$ = cumulative distribution function of offers ϕ,

$w(t)$ = $\phi + y(t)$ = match-specific productivity at job tenure t, assumed equal to the wage,[1]

δ = Poisson arrival parameter for job offers drawn from $G(\phi)$,

$F(t)$ = probability that completed tenure is less than t, and

$\lambda \equiv (f(t)/[1-F(t)]$ = conditional failute time density (the hazard function) at tenure t.

Following Jovanovic (1979b), an optimal job-shopping policy involves the choice of a reservation wage $\phi^*(t)$ that maximizes the discounted value of continued search on the current job. Since ϕ and t summarize all information about the job, this value is:

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[1]The identity between wages and productivity implies that separations are entirely worker-initiated and efficient. This assumption is not crucial to what follows. What is required is that wages index the value of jobs to workers, as in Eq. (2). Note that in this model y(t) (learning by doing) is deterministic and identical across jobs.

(2) $v[\phi,t] = \underset{\phi^*(\tau)}{\mathrm{Max}} \; {}_t\int^{\infty} \exp -{}_t\int^{\tau}[\lambda(s)+r]ds \quad \{w(\tau)+\delta\int_{\phi^*(\tau)} \; v(x,0)g(x)dx\}d\tau$

Job offers that satisfy $\phi > \phi^*(\tau)$ are accepted at τ. The reservation wage offer must satisfy

(3) $\quad v[\phi^*(\phi,t),0] = v[\phi,t].$

That is, the value of a new job, paying ϕ^* at zero tenure, must yield equal value to remaining on the old job and earning $\phi + y(t)$. If $y'(t) > 0$, so productivity and wages grow with tenure, then

$$\frac{\partial v(\phi,t)}{\partial t} > 0$$

so the reservation wage defined by (3) rises with current job tenure. Since the hazard rate is the probability of obtaining an acceptable offer, $\lambda(\phi,t) = \delta[1 \; G(\phi^*(\phi,t))]$, we have

(4) $\quad$ (a) $\dfrac{\partial\lambda(\phi,t)}{\partial\phi} < 0,\qquad$ (b) $\dfrac{\partial\lambda(\phi,t)}{\partial t} < 0.$

Both of these say that jobs with greater match-specific productivity are less likely to end, given ϕ and current tenure t.

Implications like (4) are common to most models of labor turnover, yet as we indicated above they are not readily testable from information on job durations only. The reason is that ϕ--match "quality"--is normally treated as unobservable. Job tenure then acts as a filter that selects on good matches so the aggregate mobility function declines with tenure even if $\partial\lambda(\phi,t)/\partial t = 0$.[2] As it is written, however, (4) ignores other potential observables than t. In particular, note that any two of $w(t)$, t, and ϕ are sufficient information to describe the current value of a job (2), so an equivalent expression for $v[\phi,t]$ is $v[w(t),t]$. Indeed, in this model $v[w(0),t] \equiv v[\phi,t]$.

[2] Heckman and Singer (1982) discuss the issue of heterogeneity in duration models at length. Normally in economics these models have been applied to the durations of unemployment spells where theory offers little guidance as to sources of heterogeneity.

Conditioning the value of on-the-job search on the (potentially observable) wage has special implications. In particular, suppose that y(t) is concave as almost all studies of the relation between wages and job tenure indicate (e.g., Mincer and Jovanovic, 1981; Bartel and Borjas, 1981). Then since productivity grows more rapidly in the early periods of a job it must be true that

$$v[w(t),t] < v[w(t),0].$$

That is, a new (t=0) job that pays starting wage equal to the current wage on the current job, w(t), offers higher lifetime wealth because of its greater potential for wage growth. This implies that workers are willing to accept a reduced wage on a new job as a form of investment, so the reservation wage is *less* than the current wage: $\phi^*(w(t),t) < w(t)$ for $t > 0$. Put differently, an increase in t holding w(t) fixed must *reduce* the value of the current job because the potential for future productivity growth is diminished when y(t) is concave. Therefore

$$(5) \qquad \frac{\partial v(w(t),t)}{\partial w(t)} > 0; \quad \frac{\partial v(w(t),t)}{\partial t} < 0$$

Using the definition of $\phi(w(t),t)$ these imply

$$(6) \qquad \frac{\partial \lambda(w(t),t)}{\partial w(t)} < 0; \quad \frac{\partial \lambda(w(t),t)}{\partial t} > 0$$

so tenure *increases* the probability of moving, given the current wage, and the wage reduces the probability of leaving the current job. Alternatively, had we conditioned on the initial wage on the current job, w(0), then

$$(7) \qquad \frac{\partial \lambda(w(0),t}{\partial w(0)} < 0; \quad \frac{\partial \lambda(w(0),t)}{\partial t} < 0$$

Implications (6) - (7) are robust to a broad class of mobility models. For example, in Jovanovic's (1984) setup, workers gather information about the quality of a new match from an initial unbiased test plus repeated observations on productivity (wages) as tenure accumulates. With such learning, uncertainty about match quality declines with t, since more readings are available. Again, w(t) and t are sufficient statistics to describe the value of a job. Here, however, $v_2[w(t),t] > 0$ because of the option value of accepting an alternative: tenure reduces uncertainty about current match quality, so the probability that the current match will turn out to be of particularly high quality, given current $w(t)$, is declining in t. This makes alternatives more attractive, so mobility increases with tenure and (6) and (7) hold here as well.[3]

The common feature of these models is that knowledge of the current wage and current job tenure are sufficient to describe the present value of holding a particular job at tenure t. No unobserved matching heterogeneity remains in the model. These represent an important class of models, though they are not exhaustive.[4] Nevertheless, conditioning the hazard on wages, as in (6) and (7), raises other difficult issues for empirical analysis since the wage variations that underlie these effects describe changes *within* an individual career as tenure accumulates and better job matches are located. In actual data, other sources of wage variation *across* individuals and their jobs contaminate the experiment described by these restrictions on the hazard function.

To illustrate, suppose the wage offer distribution for person i with market experience x is given by

$$(8) \qquad w^o_{i\tau} = x_{i\tau}\beta + \mu_i + \phi$$

where $\beta > 0$ describes the growth of general human capital with experience and μ_i is an individual constant representing ability, education, and the like. Both shift the offer distribution. As above, ϕ is a matching productivity component. The reservation wage for person i on the job j is described by the current wage and tenure:

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[3]A difference is that the hazard may actually rise when in the early period of a job if we condition on only the initial wage. For a proof, see Jovanovic (1979a).

[4]In particular, if investment in job-specific human capital is endogenous then incentives to invest are greater in a good match that is more likely to survive. Initial wages may then be reduced to the extent that workers share in costs and returns, breaking the connection between wages, productivity, and match quality. Moreover, if the accumulation function y(t) differs across jobs then the maximum acceptable offer is not single-valued, i.e., it depends on the initial wage and the rate of growth.

$$\phi^*(w,t) = w_{ij\tau} + t_{ij}\theta, \quad \theta < 0.$$

A new job offer is accepted if $w_{i\tau} > \phi^*(w,t)$ or if

$$(9) \qquad \phi > w_{ij\tau}\theta - \tau_{ij\tau}\theta - x_{i\tau}\beta - \mu_i.$$

Equation (9) illustrates two key points. First, though conditioning on the wage eliminates the unobservable match effect from the right side of (9), a new form of individual-specific heterogeneity is introduced by the presence of the unobservable μ_i. Persons whose wages would be increased by reasons apart from the job-matching process are *more* likely to move for any given level of the wage and tenure. Thus, there are "movers" and "stayers" in the model once we condition on the wage, even if this form of heterogeneity was absent from the original model. Second, time varying observable factors such as experience that are associated with increased wages must also *increase* mobility once we condition on the wage. The reason is that an increase in market experience, wage and tenure constant, implies a lower current match quality so long as general productivity rises with experience. The probability of an acceptable outside offer is therefore higher. This sign reversal relative to the unconditional hazard is a strong implication of the search model. Finally, note that the effect of experience in (9), β, is the true experience effect on wages in (8), so β is identified from the mobility rule.

To implement these ideas empirically, we specify the hazard function for person i on the job j in the proportional hazards form[5]

$$(10) \qquad \begin{aligned} \lambda(w_{ij\tau}, t_{ij\tau} | z_{ij\tau}, \mu_i) &= \exp\{y_{ij\tau}\} \\ y_{ij\tau} &= h(t_{ij\tau}) + \beta_1 w_{ij\tau} + \beta_2 z_{ij\tau} + \mu_i \end{aligned}$$

where z is a vector of observables that shift the hazard. According to (6) and (7), $h'(t_{ij\tau}) > 0$ if the current wage is used in (9), while $h'(t_{ij\tau}) < 0$

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[5]If the distribution of match-specific productivities is exponential, $G(\phi) = 1 - \exp(-\backslash{*}Gad)$, the hazard is $\lambda(\ ) = \mathrm{Prob}\ (\phi > z_\gamma) = \exp\{-z\alpha\gamma\}$ which is exactly the proportional hazards model applied to (9).

if the initial wage on job j is used. In both cases $\beta_1 < 0$ while $\beta_2 > 0$ for factors in z that increase wages but do not otherwise affect mobility.

Given the nonlinear model (10), if the individual effects μ were known then the likelihood function for a sample of job durations has a well-known form (e.g., Kalbfleisch and Prentice, 1980). Here, however, the μ_i are unobservable and the model implies that they are correlated with other observed regressors, in particular the wage. This fact implies that random effects methods of integrating out the unobservables are inappropriate for this model.[6] This leaves individual effect estimators that treat the μ_i as estimable parameters and that rely directly on the panel aspect of our data. It is well known, however, that estimates from such models are biased in short panels. So an important issue is whether this bias is small in panels of reasonable length like those that may be extracted from the LEED file.

Monte Carlo evidence for the performance of the fixed-effect estimator, in data with the turnover characteristics contained in LEED, indicates that the bias is quite small in panels of over eight years. This evidence can be obtained on request.[7] For the empirical results that follow, we select a subsample of 864 white males for whom at least 13 years of post-18 labor market history were available, beginning with entry into the labor market. Summary statistics for the variables used in estimation are reported in Table 9.

Details of the likelihood function corresponding to (10) and techniques for maximizing it are reported in the Appendix.[8] In light of the prediction embedded in (6), (7), and (9) we report estimates that successively (i) ignore wages, (ii) condition on the initial wage on the

[6]For random effects models of duration under the assumption of a known mixing distribution, see Flinn and Heckman (1982). Nonparametric methods are developed in Heckman and Singer (1984).

[7]For example, we estimate fixed-effect models of the form $\lambda_{it} = \Theta_i \exp\{\alpha t\}$ where $\Theta_i \sim u[.1, 1.3]$. For various length panels with 100 individuals each and 25 samples generated in each case, the bias in the fixed-effect estimator of α was 10 percent or less, in panels of 30 quarters or more, with true $\alpha = -.33$.

[8]There are 864 incidental parameters, μ_i, to be estimated. Fortunately, the likelihood function determined by (10) may be concentrated on these fixed effects, solving for each μ_i as a function of the data and other parameters of the model. See the Appendix for details.

Table 9

SUMMARY STATISTICS
PERSONS WITH 13 OR MORE YEARS OF POTENTIAL EXPERIENCE (N = 864)

Variable	Definition	Mean	Standard Deviation
Job tenure	Accumulated full-time quarters with current employer	8.07	0.36
Quarter 1	= 1 if first quarter of job, zero otherwise	.268	.443
Quarter 2	" "	.139	.346
Quarter 3	" "	.086	.280
Quarter 4	" "	.651	.246
Quarter 5	" "	.046	.210
Low wage	Natural logarithm of real (1967 dollars) quarterly earnings	7.135	0.525
Experience	Years of full-time labor market experience	7.58	3.90
First job	= 1 if the individual has no prior experience, zero otherwise	.157	.41
Prior job	= 1 if the current job began with a transition from a prior job, zero if transition was from non-employment	.751	.21
Firm size	Number of employees in reporting unit		
1-9		.50	.50
10-99		.27	.44
100-499		.12	.33
500-999		.04	.19
1000-2499		.03	.17
> 2500		.04	.19
Number of jobs		6.40	4.04

current job, and (iii) condition on the current wage. The estimated form for h(t) in these models allows the hazard to have an arbitrary shape over the first 15 months of a job, after which it is linear and splined to the fifth quarter. Experiments with this form produced only trivial differences in the results.

Column (1) of Table 10 reports estimates that do not condition on the observed wage. Thus, we expect $h'(t) < 0$ due to accumulation of specific capital on a particular job and because better jobs are more likely to survive. In fact, the hazard is fairly steep, especially during the first 15 months. At the sample mean value of μ_i, the hazard for a new entrant falls from .26 to .13 over this length of tenure, but it is fairly flat beyond this point. Note also that labor market experience reduces mobility, as implied by the job search model: Expected quality of the current job match increases with time spent searching. Consistent with this prediction, jobs that began from a spell of non-employment are significantly more likely to end, which would occur if expected quality is lower for jobs that are simply revealed to be preferred to non-employment. As in the aggregate tabulations above, we also find first jobs are much less stable than subsequent ones.

We also find that firm size significantly affects labor turnover. Since the estimates account for individual differences in propensities to move, they imply that career moves from small to large employers result in much more durable employment relations. The estimated difference between the largest and smallest size classes implies that individuals in small firms are almost 84 percent more likely to move. Even the second largest category (1000-2499) shows 30 percent greater mobility than in the largest firms.

Column (2) of the table reports estimates that condition on the initial level of (log) quarterly wages on the job. Because of the structure of the LEED file, jobs that last two or fewer quarters are unlikely to have a complete quarter from which we may gauge earnings: Jobs lasting two quarters or less can be expected to begin and to end during successive quarters. We therefore define the initial wage to be the maximum quarterly earnings for an individual over the first three quarters of a job. To account for the short job bias in measured wages, we interact the wage with dummy variables for the first two quarters. Thus, the estimated main effect of the wage is from "clean" jobs lasting three or more quarters.

Table 10

PROPORTIONAL HAZARDS MODELS WITH INDIVIDUAL EFFECTS:
PERSONS WITH 13 OR MORE YEARS OF POTENTIAL EXPERIENCE (N = 864)

(asymptotic standard errors in parentheses)

Item	Initial Wage on Job				
	(1)	(2)		(3)	
Quarter 1	.834	3.451	.582[a]	3.431	.604[a]
	(.049)	(.591)	(.054)	(.589)	(.053)
Quarter 2	.599	.916	.538	.828	.600
	(.055)	(.716)	(.055)	(.071)	(.055)
Quarter 3	.377	.343		.353	
	(.067)	(.067)		(.067)	
Quarter 4	.259	.232		.242	
	(.073)	(.073)		(.073)	
Quarter 5	.140	.120		.126	
	(.086)	(.087)		(.086)	
Job tenure	-.005	-.010		-.012	
	(.003)	(.003)		(.003)	
Wage	-	-.948		-.976	
		(.058)		(.057)	
Wage* quarter 1	-	-.402		-.396	
		(.084)		(.083)	
Wage* quarter 2	-	-.053		-.032	
		(.100)		(.100)	
Years of experience	-.052	.030		.036	
	(.006)	(.007)		(.008)	
First job	.532	.309		.209	
	(.059)	(.063)		(.056)	
Prior job	-.167	-.047		-.051	
	(.040)	(.044)		(.042)	
Firm size:					
1-9	.838	.652		-	
	(.107)	(.112)			
10-99	.689	.566		-	
	(.106)	(.108)			
100-499	.506	.414		-	
	(.111)	(.114)			
500-999	.487	.433		-	
	(.132)	(.133)			
1000-2499	.332	.306		-	
	(.135)	(.141)			
Log likelihood	-16955.15	-16688.55		-16720.97	

[a] Wage-quarter interactions evaluated at the sample mean log wage.

Since these estimates condition on the initial wage, theory still predicts that $h'(t) < 0$, which the estimates show. However, the initial wage will subsume the statistical role of experience as a predictor of current job quality, so by our arguments above $\partial\lambda/\partial x > 0$. This sign reversal for experience when the wage is included is a strong implication of the matching model, and it is firmly supported by the data. We also confirm that *within* an individual's career, higher-quality job matches are less likely to end ($\partial\lambda\partial w < 0$). Evaluating the short job interactions at the mean wage, we find some flattening of $h(t)$ that is concentrated in the first six months of a job.

Are these estimates reasonable? As a rough check, note that once we condition on the wage, experience enters the model because of its impact on career wage growth. This implies that we may estimate β from the ratio of the experience and wage effects on mobility. The resulting estimate of the impact of an extra year's experience on log earnings is $(.030/.948) = .032$. This compares favorably with other estimates of this effect from panel data, which typically lie between 3.0 and 4.0 percent, though these estimates are for workers substantially older than those in our data.[9] In fact, when wage equation (8) is estimated by least squares in our sample of young workers, we obtain $\beta = .09$ per year. Note that this estimate is biased up because the search process implies that experience and unobserved match quality, ϕ, are positively correlated. Comparison of these estimates implies that as much as two-thirds of the observed relationship between experience and earnings is attributable to search and the intense job-changing activity of young workers, and only one-third to the smooth accumulation of productivity with experience indicated by β in (8). At this stage of the life cycle, the search process is *the* major contributor to career productivity growth.[10]

[9]For example, using a sample of young men from the National Longitudinal survey, Mincer and Jovanovic (1981, Table 1.6) estimate an experience effect on log wages of .0338, evaluating their estimate at the mean level of experience in our data. Altonji and Shakotko (1984) find a similar effect of about .035 using data from the Panel Study of Income Dynamics.

[10]A more detailed study of the effects of experience, tenure, and job mobility on observed wage growth is a topic of our current research.

Other estimates also change when we control for wages. The higher mobility attributable to first jobs falls by more than half, indicating that instability of these jobs relative to later ones is due in large part to lower average match quality and earning power. The previously strong effect of entering the current job from previous employment falls by even more, and is no longer statistically different from zero. The implication is that jobs commencing from a spell of non-employment are of lower average quality than those for which the workers alternative was a prior job. Finally, we find that the firm-size effect on mobility falls when we control for the job-specific wage, especially for the smallest size categories. This change is consistent with the well-documented finding that large firms pay higher wages (e.g., Brown and Medoff, 1984), yet a strong size effect on mobility remains. As we noted previously, this may reflect Oi's (1983) notion that human capital is more job-specific in large firms, or more simply that mobility takes place within firms when scale permits a large internal labor market.

Table 11 reports the results for specifications that condition the hazard on the current wage, so the wage is a time-varying regressor within a job spell. Because we condition on the current wage, we retain a free form for h(t) only in the first two quarters.[11] The search model predicts a negative effect of within-career wage growth on mobility, while both experience *and* job tenure should increase mobility in this case. The evidence favors each of these predictions. In particular, the partial effect of job tenure changes from a negative and statistically significant effect in Table 10, when initial wages are controlled for, to a positive effect when we condition on the current wage. Taken literally, the estimates in column (1) imply that the wage and reservation wage diverge at a rate of (.005/1.43) x 4 = .014 per year. In other words, a worker with five years of current job tenure would be willing to accept up to a seven percent (.014 x 5) initial wage cut in moving to a new job because of increased investment opportunities.

[11]That is, h(t) is subsumed by conditioning on the current wage. The first two quarter dummies are retained because of the problems in measuring the wage during those quarters. A χ^2 test for the inclusion of the dummy variables for quarters 3-5 had a value of only .87. Other estimates are almost completely insensitive to the exclusion.

Table 11

PROPORTIONAL HAZARDS MODELS WITH INDIVIDUAL EFFECTS:
PERSONS WITH 13 OR MORE YEARS OF POTENTIAL EXPERIENCE (N = 864)

(asymptotic standard errors in parentheses)

Item	Current Wage on Job			
	(1)		(2)	
Quarter 1	.401	.287[a]	.360	.288[a]
	(.610)	(.060)	(.615)	(.060)
Quarter 2	1.966	.260	-1.99	.260
	(.720)	(.060)	(.725)	(.060)
Job tenure	.005		.005	
	(.003)		(.003)	
Initial wage	-		-.032	
			(.091)	
Current wage	-1.43		-1.41	
	(.059)		(.089)	
Wage* quarter 1	-.016		-.010	
	(.087)		(.087)	
Wage* quarter 2	.312		.316	
	(.101)		(.102)	
Years of experience	.044		.046	
	(.008)		(.007)	
First job	.271		.271	
	(.062)		(.619)	
Prior job	-.050		-.049	
	(.043)		(.043)	
Firm size:				
1-9	.625		.624	
	.109)		(.109)	
10-99	.569		.568	
	(.114)		(.106)	
100-499	.411		.410	
	(.112)		(.112)	
500-999	.445		.445	
	(.133)		(.133)	
1000-2499	.311		.311	
	(.139)		(.140)	
Log likelihood	-16607.33		-16607.20	

[a]See notes to Table 10. Current wage refers to measured log earnings during the last full quarter of employment on the job. Thus wages are always measured from the quarter prior to termination, unless the job ends in the first quarter.

As final evidence, the second set of estimates in Table 11 conditions the hazard on both initial and current wages on the job. The evidence strongly supports the "sufficiency" on the current wage: Given current earnings, the initial wage is irrelevant to current mobility decisions. Stated differently, within-job wage growth (w(t) - w(0)) and initial earnings on the current job have nearly identical effects on current mobility. Since these effects of wage growth within and between jobs are net of the effect of experience on earnings, they represent elements of earnings capacity that are specific to the current job. The evidence indicates that workers are indifferent between the sources of these rents, so on-the-job earnings growth adds no information about the value of continuing the match once the level of wage is known.

CONCLUSIONS

The years immediately following entry to the labor force form a critical phase of career development for young workers. During this period, the job-shopping process implies high rates of labor turnover and mobility as workers search for productive and durable employment relationships. Far from being a cause for concern, the transitory nature of jobs among young workers reflects the intensity and productivity of the search process itself. The well-known decline in mobility with age or experience is the result of this sorting, in conjunction with the growth of job-specific capital that tightens the bond between worker and firm.

This view of the search process has important implications for the interpretation of life-cycle productivity or earnings growth. Following models of life-cycle investment in human capital, the standard specification of "human capital earnings functions" decomposes individual wages into components due to personal characteristics and smooth growth components attributed to labor market experience and current job tenure (e.g., Mincer and Jovanovic, 1981). Job shopping activity offers a competing explanation: At least among young workers, life-cycle productivity growth is a discrete process associated with job-changing as a form of human capital investment. Thus at this stage of the life-cycle, job matching activity alone may account for the observed positive relation between earnings, general market experience and tenure. We presented preliminary evidence that as much as two-thirds of the early career "experience effect" on earnings can be explained in this way, but further conclusions await a detailed analysis of the relation between mobility and earnings growth.

APPENDIX

The Sample Likelihood Function

For each job spell in our data, we observe the (possibly truncated) length of the 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 at t_{ij},
= 0 otherwise (censored observations),
J_i = number of job spells for person 1,
$z_{ij}(\tau)$ = regressors at job tenure τ, and
μ_i = fixed individual effect.

Let $F(s_; Z, \mu)$ be the cumulative distribution function of completed spell lengths s. Then the likelihood contribution of the jth spell for person i is

$$\text{(A1)} \qquad L_{ij} = f(t_{ij}, \ )^{ij}[1-F(t_{ij}, \ )]^{1-d_{ij}}$$

$$= \lambda(t_{ij}, \ )^{d_{ij}}[1-F(t_{ij}, \ )]$$

$$= \lambda(t_{ij}, \ )^{d_{ij}}\exp[-\int^{t_{ij}} \lambda(\tau, \ )d\tau]$$

where $\lambda(s, \ ) = f(s; \ )/[1-F(s; \ )]$ is the conditional density of spell endings at s, the hazard function. In the proportional hazards case given by (10) in the text, the log-likelihood contribution for person i is

$$\text{(A.2)} \qquad L_i = \sum_{j=1}^{J_i} (Z_{ij}(t_{ij})\delta + \mu_i)d_{ij} - \sum_{j=1}^{J_i} \int_o^{t_{ij}} \exp[Z_{ij}(\tau)\delta + \mu_i]d\tau$$

Summing (A2) over individuals yields the sample likelihood function. Note that the integral on the right of (A2) involves the entire history over the job spell $[0, t_{ij}]$ of all time-varying regressors in $Z_{ij}(\tau)$. For example, the full history of the wage enters

(A2) in calculating the estimates in Table 11. In practice, we assumed that $Z_{ij}(\tau)$ was constant within quarters in calculating the integral. The fixed-effect estimates in Tables 10 and 11 take advantage of the fact that the likelihood contribution (A2) can be concentrated on μ_i:

$$\text{(A3)} \qquad \mu_i = \log \left\{ \frac{\sum_{j=1}^{J_i} d_{ij}}{\sum_{j=1}^{J_i} \int_0^{t_{ij}} \exp[Z_{ij}(\tau)\delta]d\tau} \right\}$$

Substituting (A3) into (A2) yields a sample likelihood that does not depend on μ_i:

$$\text{(A4)} \qquad L_i = \sum_{j=1}^{J_i} d_{ij}Z_{ij}(t_{ij})\delta - N_i \log\{ \sum_{j=1}^{J_i} \int_0^{t_{ij}} \exp[Z_{ij}(\tau)\delta]d\tau\}$$

where $N_i = \Sigma d_{ij}$ is the number of completed spells for person i. Thus persons who never change jobs in he panel contribute zero weight to the likelihood.

Likelihood function (A4) was maximized using the scoring method described in Berndt et al. (1974). The estimated asymptotic covariance matrix is

$$\text{(A5)} \qquad V(\delta) = \frac{1}{n} \sum_{i=1}^{n} \frac{\partial L_i}{\partial \delta} \frac{\partial L_i}{\partial \delta'}$$

The derivations are the same for the likelihood that calculates industry differentials, except that μ_i is a fixed effect for all individuals in industry i and J_i is the total number of spells in industry i.

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