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# Occupational Matching: A Test of Sorts

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This paper develops a theory of job matching in which matching information has both job-specific and occupation-specific components. If occupational matching is significant, then the theory predicts that for those who have switched jobs but remained in the same occupation, increased tenure in the previous job lowers the likelihood of separation from the current job. These predictions are tested using job tenure data from the National Longitudinal Survey's youth cohort. In general, the data are consistent with the occupational matching hypothesis.

## I. Introduction

Job shopping is a common explanation for job turnover (see Reynolds 1951). Recently, a number of studies have formalized this notion (e.g., Johnson 1978; Jovanovic 1979*a*, 1979*b*, 1984; Viscusi 1979, 1980; Wilde 1979; Lippman and McCall 1981; Miller 1984). These studies of job shopping or job matching assume that matching occurs only at the job level. It seems likely that matching also takes place at the occupational level. This is suggested by the significant fraction of people who switch occupations when switching jobs (see Miller 1984; Shaw 1987). The purpose of this paper is to see whether any empirical evidence exists to support this notion of occupational matching.[1]

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[1] For an empirical test of the job-matching hypothesis, see Flinn (1986).

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To achieve this goal, a theoretical model of job matching is developed in which matching information has both job-specific and occupation-specific components. If occupational matching is important, then the theory predicts that the likelihood of separation from the current job should decrease with increased tenure in the previous job only for those individuals who do not switch occupations when switching jobs (employers). The magnitude of this decrease increases with the arrival rate of occupation-specific matching information. In general, the effect of a change in the variance of occupation-specific matching information on this magnitude is ambiguous.

The data used in the empirical analysis were derived from the National Longitudinal Survey's youth cohort. The sample consists of those respondents who were full-time students at the start of the survey, entered the labor force, and worked for at least two different employers since leaving school. By means of a proportional hazards approach (see Kalbfleisch and Prentice 1980; Kiefer 1988), the determinants of the job separation hazard are analyzed for the second job worked after leaving school. For the full sample, an increase in tenure in the first job worked after leaving school significantly lowers the rate of separation from this second job. However, the magnitude of this effect is significantly greater for individuals whose occupation remains the same across both jobs.

The remainder of this paper is organized as follows: Section II develops the theoretical model of the paper. Under the assumption that there is no job search (i.e., no wage uncertainty) and that both job-specific and occupation-specific matching information arrive once (independently of each other) at some random time, the structural form of the job separation hazard is derived for both those working their first job and those working their second job in an occupation.

Section III describes the data, and Section IV presents the empirical results. Estimates are obtained using both parametric and semiparametric techniques (see Cox 1975; Prentice and Gloeckler 1978; Meyer 1986), for occupational switching defined at the one-digit and three-digit 1970 census classification levels, and when unobserved heterogeneity or measurement error is accounted for (see Lancaster 1979; Heckman and Singer 1984*a*).

Finally, Section V contains a summary and conclusions.

## II. Theoretical Model

This section develops a model of job matching in which matching information has both job-specific and occupation-specific components. After a simple model is formulated, the worker's optimal sampling strategy is designed. Next, the empirical implications of the theory for job turnover are discussed.

For simplicity, assume that initial wages are known with certainty. Furthermore, there is no intraoccupational variation in these wages. Let $w_i$ denote the initial wage for the $i$th occupation, $i = 1, \ldots, N$. Assume that a job (employer) switch that also results in an occupation switch is at least as costly as an intraoccupational job switch. This is reasonable whenever initial training is needed to work in an occupation. Symbolically, the cost of working the first job in the $i$th occupation is $c_{1i}$ and subsequent job switches in the $i$th occupation cost $c_i$, where $c_{1i} \geq c_i$, $i = 1, \ldots, N$. A job switch (with certain wages) entails a loss of one period's wages.

The worker receives two types of matching information. Both affect net wages additively. Job-specific information is matching information that is relevant only to the current job (this type of information is independent across *all* jobs both within and across occupations). Job-specific information for the $j$th job in the $i$th occupation is represented by $\zeta_{ij}$, a random variable identically distributed across all jobs within an occupation. Observe that $\zeta_{ij} = \zeta_i$, where $\zeta_i$ has cumulative distribution function $F_i$ with $E(\zeta_i) = 0$ and $\text{var}(\zeta_i) = \sigma^2_{\zeta_i}$ (or simply $\sigma^2_\zeta$) $> 0$. Job-specific information for the $j$th job in the $i$th occupation arrives at a geometric random time, $T_{ij}$, with parameter $r_i$.

Net wages also have an occupation-specific component. For the $i$th occupation ($i = 1, \ldots, N$), this is represented by the random variable $\omega_i$ that is distributed with cumulative distribution function $G_i$, $E(\omega_i) = 0$, and $\text{var}(\omega_i) = \sigma^2_{\omega_i}$ (or simply $\sigma^2_\omega$) $> 0$. The random arrival time, $T_i$, of the occupation-specific information is geometrically distributed with parameter $p_i$. In this case, time refers to the total time worked in the occupation.

The random variables, $\omega_i$, $\zeta_{ik}$, $T_{ik}$, $k = 1, 2, \ldots$, and $T_i$, are mutually independent for all $i = 1, \ldots, N$; for all $i \neq j$, ($\omega_i$, $\zeta_{ik}$, $T_{ik}$, $k = 1, 2, \ldots$, and $T_i$) is independent of ($\omega_j$, $\zeta_{jk}$, $T_{jk}$, $k = 1, 2, \ldots$, and $T_j$). Individuals sample jobs/occupations to maximize expected discounted income, where $\beta$ represents the discount rate, $0 < \beta < 1$.

To simplify the discussion of the worker's optimal sampling strategy, it is assumed that $P(-d_i < \zeta_i < d_i) = 0$ for some $d_i > 0$.[2] Given this

[2] Formally, the optimal intraoccupational job-switching policy must be independent of a stopping reward (call it $Z$), when one considers the sequential decision problem of working in the $i$th occupation or stopping and receiving $Z$ (see Whittle 1980). The value of $d_i$ that achieves this independence depends on the job-switching costs, $c_i$, the arrival rate of occupation-specific information, $p_i$, the discount rate, $\beta$, and the arrival rate of job-specific information, $r_i$. The empirical implications of the theory do not depend on this assumption.

assumption, the optimal sampling policy is a simple index policy.[3] An index policy is one in which an individual assigns an index to each occupation and during period $t$ works in that occupation with the highest index. Let $Z_i(x_i(t))$ denote the index for the $i$th occupation in state $x_i$ at time $t$. Then an individual will work in the $i$th occupation in period $t$ when $Z_i(x_i(t)) = \max_{j=1,\ldots,N}[Z_j(x_j(t))]$.

A complete characterization ($Z_i(x_i(t))$ for all $x_i(t)$) of the optimal sampling strategy is unnecessary for deriving an empirical test for the presence of occupation-specific information. Intuitively, when occupation-specific information is significant, individuals with previous experience in an occupation should be less likely to leave their current employer than those working in an occupation for the first time. In the former group, some occupational sorting has occurred, and departures tend to occur primarily for job-specific reasons. In the latter group, however, no occupational sorting has taken place, and job separations will occur for both occupation-specific and job-specific reasons. Before characterizing these differences in job separation rates, consider the different circumstances that induce quits.

In our model, an individual may leave a job when either job-specific or occupation-specific information arrives. For simplicity, assume that $d_i$ is sufficiently large that poor job-specific matches ($\zeta_i < 0$) always result in job switches, irrespective of whether or not occupation-specific information has arrived, and the arrival of occupation-specific information never results in an intraoccupational job switch for individuals with good job-specific matches ($\zeta_i > 0$).

Given these assumptions, an individual has two choices when occupation-specific information arrives: remain in the current occupation permanently or switch occupations. This decision depends not only on the attractiveness of alternative occupations but also on the quality of the job-specific match. Let $Z^* = \max_{j\neq i}[Z_j(x_j(t))]$ be the value of the index of the most attractive alternative occupation; let $Z_i(\omega_i)$ be the value of the index for occupation $i$ when occupation-specific but not job-specific information has arrived, and $Z_i(\zeta_i, \omega_i)$ the value of the index when both job-specific information and occupation-specific information have arrived. In the former case, an occupation switch occurs when $Z_i(\omega_i) < Z^*$, whereas in the latter case, an occupation switch occurs when $Z_i(\zeta_i, \omega_i) < Z^*$. This optimal sampling policy together with $Z_i$ increasing in $\omega_i$ *under both circumstances* gives the following proposition.

[3] The problem formulated in this section is a special case of a sequential decision problem referred to in the statistical literature as the multiarmed bandit problem (see Gittins and Jones 1974; Whittle 1980, 1982; Varaiya, Walrand, and Buyukkoc 1985). See Rothschild (1974), Viscusi (1979), McCall and McCall (1981, 1987), and Miller (1984) for some economic applications of multiarmed bandit theory.

PROPOSITION 1. (*a*) The probability that an occupation switch will occur when occupation-specific information arrives, given that it arrives after an individual receives favorable job-specific information, is $G_i(\underline{\underline{\omega}}_i(\zeta_i))$, where $\underline{\underline{\omega}}_i(\zeta_i)$ satisfies

$$\underline{\underline{\omega}}_i(\zeta_i) = Z^*(1 - \beta) - w_i - \zeta_i \tag{1}$$

and $\zeta_i$ is the realized value of the job-specific information. (*b*) The probability that an occupation switch will occur when occupation-specific information arrives, given that it arrives before job-specific information, is $G_i(\underline{\omega}_i)$, where $\underline{\omega}_i$ satisfies

$$\underline{\omega}_i = (1 - \beta)Z^* - w_i - \frac{r_i\beta \int_{d_i} \zeta_i dF_i}{1 - \beta\{1 - r_i[1 - F_i(-d_i)]\}}. \tag{2}$$

*Proof.* See Appendix A.

Define $h_{1i}(t)$ = Pr{separate from first job worked in occupation $i$ at tenure $t$ | no separation before $t$} and $h_{2i}(t|T^*)$ = Pr{separate from second job worked in occupation $i$ at tenure $t$ | no separation before $t$ and completed tenure $T^*$ for first job worked in occupation $i$}. These are the job separation hazards for individuals working in their first and second jobs in an occupation, respectively. The next theorem characterizes these job separation hazards in terms of the underlying parameters of the model.

THEOREM 1. $h_{1i}(t) = P^*(t)$ and $h_{2i}(t|T^*) = z(T^*)P^*(t) + [1 - z(T^*)]Q^*(t)$, where

$$P^*(t) = q_i^{t-1}s_i^{t-1}\left[s_i p_i G_i(\underline{\omega}_i) + r_i F_i(-d_i) + p_i r_i \int_{d_i} G_i(\underline{\underline{\omega}}_i(\zeta_i))dF_i\right]$$

$$+ (1 - q_i^{t-1})s_i^{t-1}r_i F_i(-d_i) + \frac{(1 - s_i^{t-1})q_i^{t-1}p_i \int_{d_i} G_i(\underline{\underline{\omega}}_i(\zeta_i))dF_i}{1 - F_i(-d_i)}, \tag{3}$$

$$Q^*(t) = s_i^{t-1}r_i F_i(-d_i), \tag{4}$$

and

$$z(T^*) = \frac{q_i^{T^*}}{q_i^{T^*} + [1 - G_i(\underline{\omega}_i)](1 - q_i^{T^*})}, \tag{5}$$

with $q_i = 1 - p_i$ and $s_i = 1 - r_i$.

*Proof.* See Appendix A.

On reflection, it is clear that $P^*(t) > Q^*(t)$ for all $t$, $\partial Q^*(t)/\partial t < 0$ for all $t$, and $\partial P^*(t)/\partial t < 0$ for $t$ sufficiently large. Note that $h_{2i}(t|0) = h_{1i}(t)$.

Differentiating $h_{2i}(t|T^*)$ with respect to $T^*$ gives the following proposition.

PROPOSITION 2. $\partial h_{2i}(t|T^*)/\partial T^* < 0$ for all $T^* \geq 0$.

*Proof.* See Appendix A.

Proposition 2 is the key empirical prediction of the theory. If occupational matching is significant, then tenure in the previous job should have a negative effect on the likelihood of leaving the current job for individuals who do not switch occupations when switching jobs. Trivially, if the variance of the occupation-specific component, $\sigma^2_\omega$, is zero, then the magnitude of this effect is zero.

Although the existence of occupational matching implies that an increase in tenure in the previous job should reduce the likelihood of separation from the current job for occupational stayers, the effect of "small" increases in the variance of occupation-specific information on the magnitude of this reduction is, in general, ambiguous. For example, assume that $\omega_i$ is normally distributed and that $\zeta_i$ is a random variable taking only the values $\alpha_i$ and $-\alpha_i$, each with probability ½. Then we obtain the following proposition.

PROPOSITION 3. (*a*) If $Z^*(1 - \beta) - w_i - \alpha_i > 0$, then $\partial^2 h_{2i}(t|T^*)/\partial T^* \partial \sigma^2_\omega > 0$, for $T^* > \ln\{1/[2 - G(\underline{\omega}_i)]\}/\ln q_i$. (*b*) If $\underline{\omega}_i < 0$, then $\partial^2 h_{2i}(t|T^*)/\partial T^* \partial \sigma^2_\omega < 0$, for $T^* > \ln\{1/[2 - G(\underline{\omega}_i)]\}/\ln q_i$.

*Proof.* See Appendix A.

If individuals have available attractive alternative occupations (so that $Z^*$ is high) or job-specific matching is relatively unimportant ($\alpha_i$ is low), then only those individuals whose net wages *increase* by a sufficiently large amount will remain after the arrival of occupation-specific information. Under these circumstances, an increase in the variance of occupation-specific information increases the proportion of those who have received occupation-specific information among occupational stayers with a given tenure in the previous job (so that $\partial h_{2i}(t|T^*)/\partial \sigma^2_\omega < 0$). However, the magnitude $|\partial h_{2i}(t|T^*)/\partial T^*|$ is actually reduced for $T^*$ sufficiently large.

Finally, the magnitude $|\partial h_{2i}(t|T^*)/\partial T^*|$ depends on $p_i$, the rate at which occupation-specific information arrives.

PROPOSITION 4. $\partial^2 h_{2i}(t|T^*)/\partial T^* \partial p_i < 0$ for $t < 1/p_i$ and $T^* < -1/\ln q_i$.

*Proof.* See Appendix A.

As $p_i$ increases, so does the likelihood that an occupational stayer with a given tenure in the previous job has received occupation-specific information.

## III. Data

The data used for the empirical analysis of this paper were derived from the National Longitudinal Survey's (NLS) youth cohort. This panel data set follows 12,686 youths aged 14–22 at the time of the 1979 interview. At the time of this study, panel data for the survey years 1979–85 are available.

The data set is derived from the NLS work history tape with some additional variables added from the youth cohort tapes. The tape builds a "job array" from the retrospective information included on the cohort tapes. This job array gives a week-by-week accounting of the respondent's work history over the 1979–85 period. The work history tape, however, omits some variables that are essential for the analysis that follows. Specifically, no education or marital status variables are included on this tape. In addition, it does not classify job separations by type. Hence, this information is merged to the work history tape from the youth cohort tapes.

Since the goal of the empirical work is to test for differences in job separation behavior between those switching employers, but not occupations, and those switching both employers and occupations, only respondents who have worked at, at least, two different jobs were included in the data set.[4] The two jobs used for this purpose are the first two jobs worked after leaving school. Analysis of these jobs is attractive because it is likely that new entrants into the labor market face significant occupation-specific matching uncertainty. Although some occupation-specific information may be revealed during schooling, jobs held concurrently with school attendance are not analyzed in this study. Many of these jobs involve summer employment, which necessarily terminates after 2 or 3 months. It is unlikely that jobs of this type are related to an individual's ultimate career plans.

The sample is restricted to those respondents who were full-time students as of the 1979 survey, entered the labor force, and worked for at least two separate employers. Those who held multiple jobs concurrently were omitted from the sample as were any respondents with missing values. The final sample size was 1,667. Table 1 summarizes the data.

A one-digit occupational transition matrix is given in table 2. This shows the percentage of respondents in occupation $i$ as of the end of the first job (job 1) who are in occupation $j$ as of the start of the second job (job 2). Entries along the diagonal show the percentage who have remained in the same occupation when switching jobs. As can be seen from the table, there is considerable variation in this percentage across occupations (from a low of 38.4 percent in sales to a high of 69

[4] Since job information is updated yearly for those who do not switch, it is possible to get information on occupation switching within a job. This information was not exploited for two reasons. First, most jobs did not overlap two survey periods, and so occupation status was observed only once. Second, it seems likely that switches of this type occur primarily for "stepping-stone" reasons.

TABLE 1

SUMMARY OF DATA

| | |
|---|---|
| Whites (%) | 58.1 |
| Blacks (%) | 26.3 |
| Hispanics (%) | 15.6 |
| Female (%) | 52.3 |
| High school graduates* (%) | 85.4 |
| College graduates* (%) | 15.6 |
| Average age at start of job 1 (yrs.) | 19.9 |
| Married at the start of job 2 (%) | 20.6 |
| Average tenure in job 1 (wks.) | 38.2 |
| Reasons for leaving job 1: | |
| Quit (%) | 41.4 |
| Fired or laid off (%) | 23.1 |
| Other reasons (%) | 35.5 |
| Occupation switchers job 1 to job 2 (%) | 41.0 |
| Unionized in job 2 (%) | 14.9 |
| Government in job 2 (%) | 3.7 |

* High school graduates are defined as those who have completed 12 or more years of education by the year job 2 began. College graduates are defined as those who have completed 16 or more years of education by the year job 2 began.

percent in professionals). Overall, at the one-digit level, 41.0 percent switch occupations when switching jobs.

It may be of some interest to search for any "asymmetries" in the transition matrix. This could give some rough indication of the existence of "stepping-stone" occupations. One might expect that if occupation $i$ serves as a training ground for occupation $j$, then a larger percentage would be observed to move from $i$ to $j$ than from $j$ to $i$ (the $(i, j)$ cell in table 2 would be larger than the $(j, i)$ cell). Only one notable asymmetry presents itself in the data, between clerical and sales. Approximately 28 percent of people working in the sales occupation as of job 1 switch to a clerical occupation in job 2. However, only about 6 percent move in the reverse direction. In the laborers, operatives, and craftsmen occupations, where one might suspect primarily movement from laborers and operatives to craftsmen, no clear asymmetries are present.

## IV. Empirical Results

This section reviews the empirical results of the paper. If matching at the occupational level is important, then among occupation stayers, individuals with longer tenures in their previous job should be significantly less likely to leave the current job.

Recall that $h_2(t|T^*)$ is the job separation hazard for individuals working at their second job in an occupation who had tenure $T^*$ in

TABLE 2

OCCUPATION TRANSITION MATRIX: JOB 1 BY JOB 2

| | Professionals | Managers | Sales | Clerical | Craftsmen | Operatives | Laborers | Farmers | Services | Household |
|---|---|---|---|---|---|---|---|---|---|---|
| Professionals | **69.0** | 4.1 | 4.7 | 9.4 | 1.8 | 3.5 | 1.8 | .6 | 5.3 | .0 |
| Managers | 4.6 | **56.9** | 11.4 | 15.9 | 2.3 | .0 | .0 | .0 | 9.1 | .0 |
| Sales | 8.1 | 7.0 | **38.4** | 22.2 | 4.0 | 3.0 | 4.0 | .0 | 13.1 | .0 |
| Clerical | 6.1 | 2.2 | 5.9 | **62.7** | 2.8 | 4.2 | 2.5 | .8 | 12.8 | .0 |
| Craftsmen | 1.5 | .7 | 3.0 | 8.2 | **51.9** | 10.4 | 14.1 | 1.5 | 8.9 | .0 |
| Operatives | 4.1 | 1.8 | 3.6 | 10.4 | 4.5 | **54.8** | 9.5 | 2.3 | 9.1 | .0 |
| Laborers | 3.6 | .0 | 1.2 | 12.7 | 8.5 | 12.1 | **46.1** | 1.8 | 13.9 | .0 |
| Farmers | 2.6 | .0 | 2.6 | .0 | 5.3 | 26.3 | 5.3 | **50.0** | 7.9 | .0 |
| Services | 6.0 | 2.1 | 4.2 | 13.3 | 3.1 | 7.6 | 6.0 | .5 | **57.3** | .0 |
| Household | .0 | .0 | 3.2 | 12.9 | 3.2 | 6.5 | .0 | .0 | 12.9 | **61.3** |

NOTE.—This transition matrix is read as the percentage of workers in occupation $i$ as of job 1 that are in occupation $j$ as of job 2. The diagonal gives the percentage of workers who remain in the same occupation. The total number of workers is 1,667.

their first job. A first-order expansion of $\ln h_2(t|T^*)$ around $T^* = 0$ yields

$$\ln h_2(t|T^*) = \ln h_1(t) + \alpha_{T^*}T^*, \tag{6}$$

where $\alpha_{T^*} = \partial \ln h_2(t|T^*)/\partial T^*|T^* = 0$. Proposition 2 of Section II implies that if occupation-specific information is significant, $\alpha_{T^*} < 0$. Proposition 4 suggests that $|\alpha_{T^*}|$ increases the arrival rate of occupation-specific information.

Of course, differences in job separation behavior might arise from other sources of population heterogeneity. It is assumed that this heterogeneity affects $\ln h_2(t|T^*)$ additively. Thus for the $i$th individual, $i = 1, \ldots, N$, we have

$$\ln h_2(t|T_i^*, \mathbf{x}_i) = \ln h_1(t) + \alpha_{T^*}T_i^* + \sum_{j=1,k} \beta_j x_{ji} \tag{7}$$

or, equivalently,

$$h_2(t|T_i^*, \mathbf{x}_i) = h_1(t) \exp\left(\alpha_{T^*}T_i^* + \sum_{j=1,k} \beta_j x_{ji}\right), \tag{8}$$

where $\mathbf{x}_i$ is a $k$-vector of covariates with components $x_{ji}$, $j = 1, \ldots, k$.

Recall from Section III that job 1 and job 2 are defined as the first and second jobs (employers) worked after school is left, respectively. Let $\lambda_2$ denote the separation hazard for job 2 and define SWITCH as an indicator variable that is one if the $i$th individual switches occupations between job 1 and job 2 and zero otherwise. If we denote weeks of tenure in job 1 for individual $i$ by TENURE1, then the results above imply that

$$\begin{aligned}\lambda_2(t|\text{TENURE1, SWITCH}, \mathbf{x}_i) &= \lambda_2^0(t)\\ &\times \exp\left[\alpha \times (1 - \text{SWITCH}) \times \text{TENURE1} + \sum_{j=1,k} \beta_j x_{ji}\right],\end{aligned} \tag{9}$$

where $\lambda_2^0(t)$ is the baseline hazard function. The model used for estimation purposes, however, is

$$\begin{aligned}\lambda_2(t|\text{TENURE1, SWITCH}, \mathbf{x}_i, \theta) &= \lambda_2^0(t)\\ &\times \exp\Big(\alpha_1\text{TENURE1} + \alpha_2\text{SWITCH} \times \text{TENURE1}\\ &+ \alpha_3\text{SWITCH} + \sum_{j=1,k} \beta_j x_{ji}\Big)\theta,\end{aligned} \tag{10}$$

where $\theta$ is an unobserved variable that may affect job separation ($E(\theta) = 1$). This specification allows for other possible effects of tenure in job 1 on the job 2 separation rate from which the simple theoretical model of Section II abstracted. A weaker prediction of the theory is that the effect of tenure in job 1 on the job 2 separation hazard is greater for occupation switchers ($\alpha_2 > 0$). The specification of (10) also recognizes the possibility that some sources of population heterogeneity that affect job separations may not be observed.

Before we turn to the estimates, consider the definition of an occupation switch used in the empirical estimation. An occupation switch is said to take place if the one-digit 1970 census occupation classification listed at the start of job 2 differs from *both* one-digit classifications listed at the start and end of job 1 (if they should differ). Given this definition, along with the fact that jobs 1 and 2 are the first two jobs worked after leaving school, it is unlikely that those who have switched occupations will have had any significant work experience in the occupation listed at the start of job 2. The end of this section discusses how the results change when an occupation switch is defined at the three-digit level.

A list of the other covariates ($\mathbf{x}_i$) used in the estimation of (10) is presented in Appendix B. These include the starting hourly wage for job 2, the number of weeks an individual was out of work between jobs 1 and 2, one-digit occupation and industry dummy variables for job 2, and indicator variables for whether job 2 was a union or government job as well as for race, sex, schooling, and marital status.

Four different maximum likelihood estimates of the parameters of (10), corresponding to different assumptions about $\lambda_2^0(t)$ and $\theta$, are presented in columns 1–4 of table 3.[5] The estimated baseline hazard functions are graphed in figures 1 and 2. Columns 1 and 2 of table 3 assume that the baseline hazard function is a Weibull: $\lambda_2^0(t) = \mu\phi(\mu t)^{\phi - 1}$. The estimates of column 1 do not control for unobserved heterogeneity ($\theta = 1$), whereas the estimates of column 2 assume that $\theta$ has a gamma distribution. The estimates of columns 3 and 4 are obtained by semiparametric hazard estimation techniques designed for discrete data (see Prentice and Gloeckler 1978; Meyer 1986). Unobserved heterogeneity is not controlled for in column 3 and is accounted for by a gamma mixing distribution in column 4. Estimates using Cox's partial likelihood method are presented in column 5. The estimates in columns 3–5 do not constrain the baseline

[5] The likelihood function can be derived directly from the hazard function. For example, suppose that all heterogeneity is accounted for by the covariates ($\theta = 1$). Then the unconditional probability that the $i$th individual leaves job 2 at time $t_i$ is

$$\lambda_2(t_i|\text{TENURE1, SWITCH, }\mathbf{x}_i)\exp\left[-\int_{0,t_i}\lambda_2(u|\text{TENURE1, SWITCH, }\mathbf{x}_i)du\right].$$

TABLE 3

PARAMETER ESTIMATES FOR JOB 2 SEPARATION HAZARD

| | Baseline/Mix | | | | |
|---|---|---|---|---|---|
| Parameter | Weibull/ None (1) | Weibull/ Gamma (2) | Semiparametric/ None (3) | Semiparametric/ Gamma (4) | Cox/None (5) |
| TENURE1 | −.0069*** | −.0101*** | −.0071*** | −.0086*** | −.0069*** |
| | (.0010) | (.0017) | (.0011) | (.0014) | (.0012) |
| TENURE1*SWITCH | .0025 | .0055** | .0026 | .0038* | .0025 |
| | (.0016) | (.0025) | (.0017) | (.0021) | (.0018) |
| SWITCH | −.0949 | −.3482** | −.1193 | −.2194* | −.1148 |
| | (.0787) | (.1559) | (.0852) | (.1177) | (.0844) |
| INITIALWAGE | −.0471*** | −.0554** | −.0451*** | −.0544*** | −.0436*** |
| | (.0124) | (.0197) | (.0135) | (.0166) | (.0152) |
| FULLTIME | .0750 | −.2610** | −.0683 | −.1519* | −.0616 |
| | (.0593) | (.1187) | (.0640) | (.0895) | (.0644) |
| BLACK | −.0555 | −.1921 | −.0776 | −.1414 | −.0759 |
| | (.0661) | (.1313) | (.0726) | (.1009) | (.0708) |
| HISPANIC | .0808 | .0789 | .0612 | .0686 | .0582 |
| | (.0798) | (.1578) | (.0823) | (.1193) | (.0836) |
| UNIONJOB2 | −.0728 | −.0402 | −.0610 | −.0545 | −.0531 |
| | (.0761) | (.1476) | (.0851) | (.1145) | (.0839) |
| GOVTJOB2 | .7197*** | .9723*** | .6725*** | .8898*** | .6252*** |
| | (.1437) | (.3507) | (.1415) | (.2456) | (.1441) |
| WEEKSOUTWORK | .0005 | .0014 | .0002 | .0011 | .0000 |
| | (.0010) | (.0021) | (.0011) | (.0016) | (.0011) |
| NOOUTWORKSPELL | −.3919*** | −.7604*** | −.3722*** | −.5370*** | −.3551*** |
| | (.1116) | (.2095) | (.1235) | (.1668) | (.1113) |
| HSCHGRD | −.1184 | −.0774 | −.1014 | −.0768 | −.0992 |
| | (.0781) | (.1606) | (.1264) | (.1194) | (.0824) |
| COLLEGEGRD | −.3303*** | −.3280 | −.2932** | −.2841 | −.2859** |
| | (.1198) | (.2282) | (.1289) | (.1747) | (.1316) |

| | | | | | |
|---|---|---|---|---|---|
| MARRIED*MALE | −.1571 | −.3018 | −.1459 | −.2427 | −.1489 |
| | (.1321) | (.2303) | (.1434) | (.1843) | (.1428) |
| MARRIED*FEMALE | .2689*** | .1198 | .2175** | −.1546 | .2028** |
| | (.0980) | (.1792) | (.1063) | (.1403) | (.0972) |
| NMARRIED*FEMALE | .0683 | −.0654 | .0354 | −.0286 | .0330 |
| | (.0681) | (.1352) | (.0750) | (.1028) | (.0739) |
| | | | Industry | | |
| Mining | −.0098 | −.2320 | −.0340 | −.1724 | −.0395 |
| | (.3600) | (.5992) | (.3743) | (.4683) | (.3937) |
| Construction | .6961*** | 1.0616*** | .6598*** | .8651*** | .6407*** |
| | (.1715) | (.3838) | (.1858) | (.2762) | (.2019) |
| Manufacturing | .0927 | −.1802 | .0634 | −.0620 | .0588 |
| | (.1613) | (.3285) | (.1744) | (.2397) | (.1822) |
| Transportation/communications/ | −.0083 | −.4469 | .0130 | −.1959 | .0175 |
| public utilities | (.2219) | (.4389) | (.2484) | (.3379) | (.2435) |
| Trade | −.0154 | −.2736 | −.0142 | −.1502 | −.0147 |
| | (.1545) | (.3139) | (.1670) | (.2294) | (.1736) |
| Finance/insurance/real estate | −.2611 | −.7490* | −.2271 | −.4574 | −.2204 |
| | (.2108) | (.4049) | (.2306) | (.3080) | (.2325) |
| Business/repair services | .3131* | .2345 | .2864 | .2551 | .2718 |
| | (.1841) | (.3708) | (.1978) | (.2715) | (.1984) |
| Personal services | .3866* | .1558 | .3062 | .2169 | .2984 |
| | (.1996) | (.4086) | (.2073) | (.2965) | (.2084) |
| Entertainment/recreation | .3154 | .1482 | .3362 | .2433 | .3251 |
| services | (.2480) | (.5054) | (.2742) | (.3836) | (.2518) |
| Professional services | .0848 | −.1667 | .0832 | −.0527 | .0780 |
| | (.1669) | (.3340) | (.1808) | (.2458) | (.1833) |
| Public administration | .1013 | .1138 | .1664 | .1134 | .1503 |
| | (.2265) | (.4505) | (.2450) | (.3315) | (.2553) |
| | | | Occupation | | |
| Managers | −.1744 | −.3539 | −.1464 | −.2295 | −.1468 |
| | (.2172) | (.3505) | (.2354) | (.2943) | (.2047) |

TABLE 3 (*Continued*)

| | Baseline/Mix | | | | |
|---|---|---|---|---|---|
| Parameter | Weibull/ None (1) | Weibull/ Gamma (2) | Semiparametric/ None (3) | Semiparametric/ Gamma (4) | Cox/None (5) |
| Sales | .2630* | 1.0173*** | .3323** | .6194*** | .3129* |
| | (.1515) | (.3004) | (.1703) | (.2310) | (.1715) |
| Clerical | −.1028 | .1077 | −.0653 | −.0091 | −.0609 |
| | (.1241) | (.2243) | (.1384) | (.1791) | (.1310) |
| Craftsmen | −.0190 | .0632 | .0235 | .0089 | .0222 |
| | (.1546) | (.2957) | (.1731) | (.2301) | (.1680) |
| Operatives | .0829 | .4049 | .1161 | .2361 | .1141 |
| | (.1421) | (.2684) | (.1591) | (.2110) | (.1536) |
| Laborers | .1678 | .5105* | .2045 | .3174 | .1918 |
| | (.1467) | (.2779) | (.1640) | (.2182) | (.1581) |
| Farmers | .8948*** | 2.0542*** | .9442*** | 1.5635*** | .8900*** |
| | (.2185) | (.4934) | (.2337) | (.3651) | (.2629) |
| Services | .1703 | .5332** | .1826 | .3189 | .1785 |
| | (.1254) | (.2253) | (.1399) | (.1817) | (.1316) |
| Household | .1435 | .7726* | .2443 | .5145 | .2213 |
| | (.2203) | (.4675) | (.2324) | (.3439) | (.2328) |
| $\mu$ | .0990*** | .0508*** | ... | ... | ... |
| | (.0075) | (.0054) | | | |
| $\phi$ | .8706*** | 1.5860*** | ... | ... | ... |
| | (.0231) | (.0860) | | | |
| $\sigma^2$ | ... | 1.6103*** | ... | .6331*** | ... |
| | | (.1927) | | (.1890) | |
| ln $L$ | −4,304.4 | −4,226.3 | −4,194.5 | −4,188.1 | −8,079.5 |

Note.—Job 2 is the second job worked since school has been left The number of observations is 1,667. Tenure in job 2 is measured in 4-week intervals Standard errors are in parentheses. Both occupation and industry dummies were jointly significant at the 1 percent level The variance of the gamma mixing distribution is $\sigma^2$.

* Significant at the 10 percent level.

** Significant at the 5 percent level.

*** Significant at the 1 percent level.

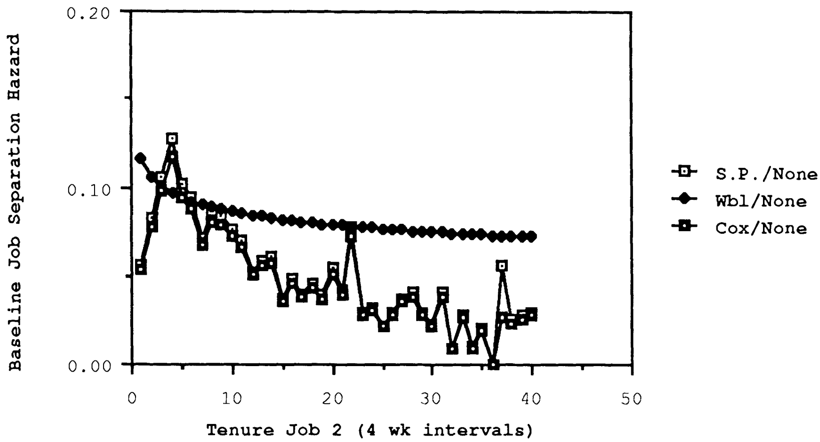


FIG. 1.—Baseline hazard estimates (no mixing distributions)

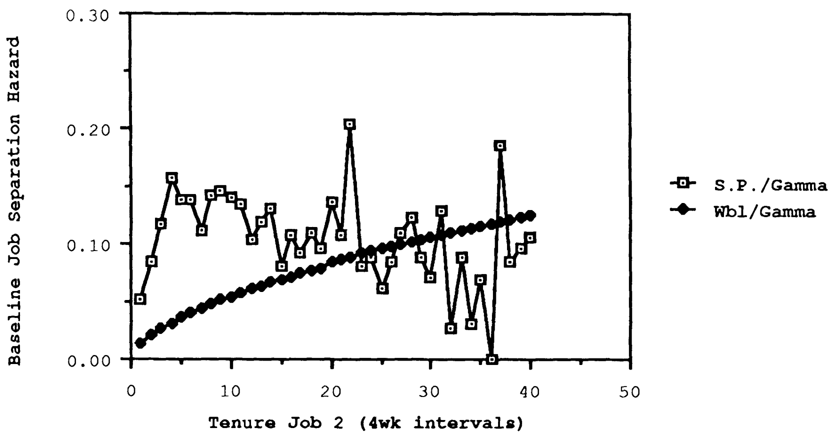


FIG. 2.—Baseline hazard estimates (with gamma mixing distribution)

hazard to any particular parametric family. To reduce numerical computations, tenure in job 2 is grouped into 4-week intervals.

The estimates of $\alpha_1$, $\alpha_2$, and $\alpha_3$ are presented in the first three rows of table 3, respectively. Turning to the estimates that do not control for unobserved heterogeneity in columns 1, 3, and 5, we see that tenure in job 1 has a significantly negative impact on the job 2 separation hazard. For those who have switched occupations, however, the magnitude of this effect is significantly smaller (at the 10 percent level on the basis of a one-sided test). The value of $\alpha_3$ is not statistically significant in all three specifications. So for a given tenure in job 1, occupation switchers are more likely to leave job 2.

When unobserved heterogeneity is accounted for, in columns 2 and 4, the estimate of $\alpha_2$ increases and is significant at the 5 percent level.[6] The estimates of $\alpha_3$, however, are significantly negative. Together, these results imply that individuals who remain in the same occupation when switching jobs are less likely to leave job 2 only if tenure in job 1 exceeds 1 year. This result is not surprising and could be due to the fact that occupation-specific matching is a more complex process than what was proposed in Section II.

Since the sample used for estimation purposes consists of only those individuals who had worked at least two different jobs in the time between leaving school and the 1985 interview date, selectivity bias may limit the usefulness of the estimates for analyzing the determinants of job separation. Nevertheless, a few observations merit mentioning. First, the estimated baseline hazards portrayed in figures 1 and 2 suggest that the Weibull distribution may be inappropriate for studying job separation. The semiparametric estimates suggest that the baseline job separation hazard increases in the first 4 months of tenure and then tends to decrease thereafter. The Weibull baseline hazard, however, is restricted to be monotonic. Second, while the semiparametric baseline estimates do not change appreciably when unobserved heterogeneity is accounted for, the estimated baseline hazard changes dramatically when it is constrained to be Weibull. Finally, for all specifications, individuals with high initial hourly wages, as well as those who have switched jobs directly without any intervening out-of-work spell, are significantly less likely to leave job 2.

Equation (5) embodies the assumption that all covariates affect the job 2 separation hazard proportionally. Some checks of this specification using methods contained in McCall (1988*a*) suggest that the proportional hazards assumption is reasonable,[7] although there is some weak indication that the one-digit occupational dummies do not affect $\lambda_2$ proportionally.[8]

[6] Heckman and Singer (1984*a*, 1984*b*) have shown that estimates can vary considerably depending on the parametric family assumed for the distribution of $\theta$. Though this result may be true when the baseline hazard is modeled as a Weibull, the semiparametric estimates should be more robust to the specification of the distribution of $\theta$ since some of the variability in the estimated effects of the covariates may be a result of misspecification of the baseline hazard (see Trussell and Richard 1985; Han and Hausman 1986). Nevertheless, a model with a Weibull baseline and two-mass-point mixing distribution was estimated to check the robustness of the estimates. The estimates of the $\alpha$'s obtained with this specification were similar to those obtained in col. 2 of table 3.

[7] To reduce computations, tests of proportionality were performed using partial likelihood methods only.

[8] Separate estimates of (5) were obtained for the services and clerical occupations. The estimate of $\alpha_1$ for services was almost double that of clericals in magnitude. This suggests that the arrival rate or variance (if the conditions of proposition 3*b* are satisfied) of occupation-specific information is greater for services.

The empirical estimates of table 3 implicitly assume that all types of job separations are alike. Although this is true under certain circumstances (see, e.g., Jovanovic 1979*b;* Borjas and Rosen 1980), in general the model of Section II is a model of quits. To control for other types of job separations, a competing risks framework is used (see Elandt-Johnson and Johnson 1980; Kalbfleisch and Prentice 1980). The results for job quitters are similar to those obtained in columns 2 and 4 of table 3. However, the overall likelihood of quitting job 2 is greater for occupation switchers only among individuals with tenures of 2 years or more on job 1.

As a final check of robustness, estimates were obtained with occupational switching defined at the 1970 census three-digit level. One might suspect more measurement error at the three-digit level, but there is no theoretical reason that the results should differ at this disaggregate level. Though not reported here, the estimates using the 1970 census three-digit classifications lend additional support to the occupational matching hypothesis. When no adjustment is made for unobserved heterogeneity, the job 2 separation behavior of those who have switched occupations and those who have remained in the same occupation is not significantly different. However, significant differences similar to those obtained at the one-digit level do emerge when unobserved heterogeneity is accounted for. Since "mixing" likelihood functions helps control for not only unobserved heterogeneity but measurement error, these results are consistent with the presumption of measurement error at the three-digit level.

## V. Summary and Conclusions

This paper developed a theory of job matching in which matching information has both occupation-specific and job-specific components. If occupational matching is significant, then the likelihood of leaving the current job will decrease with increased tenure in the previous job for those who do not switch occupations between jobs.

The National Longitudinal Survey's youth cohort was used to test these predictions. Job separation rates of youths working at their second employer since leaving school were analyzed using a proportional hazards approach. Estimates were obtained using both parametric and semiparametric hazard estimation techniques, when unobserved heterogeneity or measurement error was accounted for by mixing the likelihood function, and for occupation switches defined at both the one- and three-digit levels. In general, tenure in the previous job had a significantly negative impact on separation rate from the current job. However, for those who had switched occupations between jobs, the magnitude of this effect was significantly less. Similar results were obtained when job quits were analyzed separately using a competing risks approach.

Estimation of the underlying structural parameters of the model of Section II was not done in this paper (see, e.g., Miller 1984; Wolpin 1987). This is left to future research. The empirical results of this paper suggest that a more complex occupation-specific matching process would be necessary for any structural estimation. This estimation might best proceed by employing simulation techniques (see, e.g., Pakes 1986).

Finally, the results of this paper emphasize that, when one is modeling the job matching process, independence assumptions must not ignore the data. Of course any dynamic analysis must ultimately invoke independence. But whether this is done at an occupational or possibly industry or regional level, the data clearly indicate that at the job level the independence assumption is inappropriate.

## Appendix A

This Appendix proves the propositions and theorem of the paper.

### *Proof of Proposition 1*

*a*) By assumption, if $\zeta_i \geq d_i$, an intraoccupational job switch is not optimal when occupation-specific information arrives. So

$$Z_i(\omega_i, \zeta_i) = \frac{w_i + \zeta_i + \omega_i}{1 - \beta}.$$

Set $Z_i(\omega_i, \zeta_i) = Z^*$ and solve for $\omega_i$. Q.E.D.

*b*) By assumption, irrespective of $\omega_i$, it is assumed that if an individual is poorly matched to a specific job ($\zeta_i \leq -d_i$), then a job switch will occur. Let $V_i(\omega_i, Z)$ be the optimal value function of working at least one more period in the $i$th occupation when $\omega_i$ is known and a stopping option is available each period that pays $Z$. Whittle (1980) has shown under these circumstances that the index associated with the $i$th occupation is the value of $Z$ that makes one indifferent between continuing and stopping. Thus $Z_i(\omega_i)$ satisfies the recursive relation $Z_i(\omega_i) = V_i(\omega_i, Z_i(\omega_i))$. Writing out $V_i$, we have

$$\begin{aligned} Z_i(\omega_i) = {} & \omega_i + w_i + \beta r_i F(-d_i) Z_i(\omega_i) \\ & + \frac{\beta r_i \{(w_i + \omega_i)[1 - F(-d_i)] + \int_{d_i} \zeta_i dF_i\}}{1 - \beta} \\ & + \beta(1 - r_i) Z_i(\omega_i). \end{aligned}$$

Solving for $Z_i(\omega_i)$ and noting that an occupation switch will occur iff $Z_i(\omega_i) <$ $Z^*$ proves the result. Q.E.D.

*Proof of Theorem 1*

Before I prove theorem 1, some additional notation will be introduced. Recall from the text that $h_{1\iota}(t)$ = Pr{separate from first job worked in occupation $i$ at tenure $t|$no separation before $t$} and $h_{2\iota}(t|T^*)$ = Pr{separate from second job worked in occupation $i$ at tenure $t|$no separation before $t$ and completed tenure $T^*$ for first job worked in occupation $i$}. Also, let $oi(t-1)$ denote the event that occupation-specific information has been received by the start of the $t$th interval, $ji(t-1)$ denote the event that job-specific information has been received by the start of the $t$th interval, Bad denote the event that the information (either job-specific or occupation-specific depending on the circumstances) was sufficiently bad to induce a quit, $oi(t)$ denote the event that a person receives occupation-specific information in period $t$, $ji(t)$ denote the event that a person has received job-specific information in period $t$, and $t$ indicate the event that the person has survived up to the beginning of period $t$.

The proof follows from applying the law of iterated expectations and from noting that the probability of an event is just the expectation of the indicator function of that event. The events conditioned on at time $t$ are whether the individual has not received job- or occupation-specific information by the start of period $t$, has received one but not the other, or has received both. From this we have

$$\begin{aligned} h_{1\iota}(t) = {} & P[\sim oi(t-1) \cap \sim ji(t-1)|t] \\ & \times \{P[oi(t) \cap \sim ji(t) \cap \text{Bad}|t \cap \sim oi(t-1) \cap \sim ji(t-1)] \\ & \quad + P[ji(t) \cap \sim oi(t) \cap \text{Bad}|t \cap \sim oi(t-1) \cap \sim ji(t-1)] \\ & \quad + P[oi(t) \cap ji(t) \cap \text{Bad}|t \cap \sim oi(t-1) \cap \sim ji(t-1)]\} \\ & + P[\sim oi(t-1) \cap ji(t-1)|t] \\ & \times P[oi(t) \cap \text{Bad}|t \cap \sim oi(t-1) \cap ji(t-1)] \\ & + P[oi(t-1) \cap \sim ji(t-1)|t] \\ & \times P[ji(t) \cap \text{Bad}|t \cap oi(t-1) \cap \sim ji(t-1)], \end{aligned} \tag{A1}$$

where a tilde is the negation sign.

Now for new entrants into an occupation, the probability of not receiving occupation-specific information by the start of the $t$th interval is $q_\iota^{t-1}$. Similarly, the probability of not receiving job-specific information by the start of the $t$th interval is $s_\iota^{t-1}$. When job-specific information arrives, either before occupation-specific information or after, then the probability of a quit is $F_\iota(-d_\iota)$. If occupation-specific information arrives before job-specific information, then a quit will occur with probability $G_\iota(\underline{\omega}_\iota)$. On the other hand, if a "good" job-specific match is obtained, then the (unconditional) probability that the arrival of occupation-specific information in period $t$ will induce a job quit is $\int_{d_\iota} G_\iota(\underline{\underline{\omega}}_\iota(\zeta_\iota))dF_\iota/[1 - F_\iota(-d_\iota)]$. Substituting these facts into (A1) gives $P^*$.

To derive the expression for $h_{2\iota}(t|T^*)$, note that a person working at his or her second job in an occupation, who had tenure $T^*$ in the first job, will not have received occupation-specific information during the first job with probability $q_\iota^{T^*}$. By the memoryless property of the geometric distribution, the quit behavior of these individuals is identical to that of those first entering the occupation. The probability that an individual with tenure $T^*$ in the first job received favorable occupation-specific information during that job is $(1 - q_i^{T^*}) \times [1 - G_i(\underline{\omega}_i)]$. Individuals of this type will leave their second job only for job-specific reasons. The term $h_{2i}(t|T^*)$ is a weighted average of these two types of individuals. Q.E.D.

*Proof of Proposition 2*

From theorem 1 we have

$$h_{2i}(t|T^*) = z_i(T^*)[P_i^*(t) - Q_i^*(t)] + Q_i^*(t).$$

So

$$\frac{\partial h_{2i}(t|T^*)}{\partial T^*} = \frac{\partial z_i(T^*)}{\partial T^*}[P_i^*(t) - Q_i^*(t)]. \tag{A2}$$

But $P_i^*(t) - Q_i^*(t)$ is positive. So $\partial h_{2i}(t|T^*)/\partial T^*$ and $\partial z_i(T^*)/\partial T^*$ have the same sign. Differentiating (5) with respect to $T^*$ gives

$$\frac{\partial z_i(T^*)}{\partial T^*} = \frac{\ln q_i q_i^{T^*}[1 - G_i(\underline{\omega}_i)]}{\{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i)q_i^{T^*}\}^2} < 0. \tag{A3}$$

Q.E.D.

*Proof of Proposition 3*

*a*) Differentiating (A2) with respect to $\sigma_\omega^2$ yields

$$\begin{aligned} \frac{\partial^2 h_{2i}(t|T^*)}{\partial T^* \partial \sigma_\omega^2} &= \frac{\partial^2 z_i(T^*)}{\partial T^* \partial \sigma_\omega^2}[P_i^*(t) - Q_i^*(t)] \\ &\quad + \frac{\partial z_i(T^*)}{\partial T^*}\left\{\frac{\partial[P_i^*(t) - Q_i^*(t)]}{\partial \sigma_\omega^2}\right\}. \end{aligned} \tag{A4}$$

Since $\underline{\omega}_i > Z^*(1 - \beta) - w_i - \alpha_i > 0$,

$$\frac{\partial G_i[Z^*(1 - \beta) - w_i - \alpha_i]}{\partial \sigma_\omega^2} = \frac{\partial \Phi\{[Z^*(1 - \beta) - w_i - \alpha_i]/(\sigma_\omega^2)^{1/2}\}}{\partial \sigma_\omega^2} < 0$$

and

$$\frac{\partial G_i(\underline{\omega}_i)}{\partial \sigma_\omega^2} = \frac{\partial \Phi[\underline{\omega}_i/(\sigma_\omega^2)^{1/2}]}{\partial \sigma_\omega^2} < 0,$$

where $\Phi$ is the cumulative distribution for the standard normal. Define

$$A = s_i^t G_i(\underline{\omega}_i) + s_i^{t-1} r_i \int_{d_i} G_i(\underline{\underline{\omega}}_i(\zeta_i))dF_i + \frac{(1 - s_i^{t-1})\int_{d_i} G_i(\underline{\underline{\omega}}_i(\zeta_i))dF_i}{1 - F_i(\zeta_i)}.$$

Then

$$P_i^*(t) - Q_i^*(t) = q_i^{t-1}(1 - q_i)A.$$

So

$$\text{sign}\left\{\frac{\partial[P_i^*(t) - Q_i^*(t)]}{\partial \sigma_\omega^2}\right\} = \text{sign}\left(\frac{\partial A}{\partial \sigma_\omega^2}\right).$$

Given the assumptions of this example, $A$ reduces to

$$A = s_i^t G_i(\underline{\omega}_i) + \tfrac{1}{2} s_i^{t-1} r_i G_i[Z^*(1 - \beta) - w_i - \alpha_i] + (1 - s_i^{t-1}) G_i[Z^*(1 - \beta) - w_i - \alpha_i].$$

Hence,

$$\frac{\partial A}{\partial \sigma_\omega^2} = s_i^t \frac{\partial \Phi[\underline{\omega}_i/(\sigma_\omega^2)^{1/2}]}{\partial \sigma_\omega^2} + \tfrac{1}{2} s_i^{t-1} r_i \frac{\partial \Phi\{[Z^*(1 - \beta) - w_i - \alpha_i]/(\sigma_\omega^2)^{1/2}\}}{\partial \sigma_\omega^2} + (1 - s_i^{t-1}) \frac{\partial \Phi\{[Z^*(1 - \beta) - w_i - \alpha_i]/(\sigma_\omega^2)^{1/2}\}}{\partial \sigma_\omega^2} < 0.$$

Because $\partial^2 z_i(T^*)/\partial T^* \partial \sigma_\omega^2 = [\partial^2 z_i(T^*)/\partial T^* \partial G_i(\underline{\omega}_i)][\partial G_i(\underline{\omega}_i)/\partial \sigma_\omega^2]$,

$$\text{sign}\left[\frac{\partial^2 z_i(T^*)}{\partial T^* \partial \sigma_\omega^2}\right] = -\text{sign}\left[\frac{\partial^2 z_i(T^*)}{\partial T^* \partial G_i(\underline{\omega}_i)}\right].$$

Differentiating (A3) with respect to $G_i(\underline{\omega}_i)$ gives

$$\frac{\partial^2 z_i(T^*)}{\partial T^* \partial G_i(\underline{\omega}_i)} = \frac{\ln q_i q_i^{T^*}[1 - 2q_i^{T^*} + q_i^{T^*} G_i(\underline{\omega}_i)]}{\{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i) q_i^{T^*}\}^3}.$$

This is negative when $1 - 2q_i^{T^*} + q_i^{T^*} G_i(\underline{\omega}_i) > 0$. These results, along with (A3) and the fact that $P_i^*(t) - Q_i^*(t) > 0$, establish that (A4) is negative if $1 - 2q_i^{T^*} + q_i^{T^*} G_i(\underline{\omega}_i) > 0$ or $T^* > \ln\{1/[2 - G_i(\underline{\omega}_i)]\}/\ln q_i$. Q.E.D.

*b*) Since $Z^*(1 - \beta) - w_i - \alpha_i < \underline{\omega}_i < 0$, $\partial G_i[Z^*(1 - \beta) - w_i - \alpha_i]/\partial \sigma_\omega^2$ and $\partial G_i(\underline{\omega}_i)/\partial \sigma_\omega^2$ are both positive. Hence, $\partial[P_i^*(t) - Q_i^*(t)]/\partial \sigma_\omega^2$ is greater than zero. Substitute these results along with the fact that $\partial^2 z_i(T^*)/\partial T^* \partial G_i(\underline{\omega}_i)$ is negative when $1 - 2q_i^{T^*} + q_i^{T^*} G_i(\underline{\omega}_i) > 0$ into (A4). Q.E.D.

*Proof of Proposition 4*

It is somewhat simpler to derive $\partial^2 h_{2i}(t|T^*)/\partial T^* \partial q_i$. Since $p_i = 1 - q_i$,

$$\frac{\partial^2 h_{2i}(t|T^*)}{\partial T^* \partial p_i} = -\frac{\partial^2 h_{2i}(t|T^*)}{\partial T^* \partial q_i}.$$

Differentiating $\partial h_{2i}(t|T^*)/\partial T^*$ with respect to $q_i$ gives

$$\frac{\partial^2 h_{2i}(t|T^*)}{\partial T^* \partial q_i} = \frac{\partial^2 z_i(T^*)}{\partial T^* \partial q_i}[P_i^*(t) - Q_i^*(t)] + \frac{\partial z_i(T^*)}{\partial T^*} \frac{\partial[P_i^*(t) - Q_i^*(t)]}{\partial q_i}. \quad (A5)$$

Differentiating (A3) with respect to $q_i$ and simplifying yields

$$\frac{\partial^2 z_i(T^*)}{\partial T^* \partial q_i} = [1 - G_i(\underline{\omega}_i)] q_i^{T^*-1} \times \left(\frac{T^* \ln q_i[1 - G_i(\underline{\omega}_i) - G_i(\underline{\omega}_i) q_i^{T^*}] + [1 - G_i(\underline{\omega}_i) + G_i(\underline{\omega}_i) q_i^{T^*}]}{\{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i) q_i^{T^*}\}^3}\right). \quad (A6)$$

Recall that

$$P_i^*(t) - Q_i^*(t) = q_i^{t-1}(1 - q_i)A. \quad (A7)$$

Hence,

$$\frac{\partial P_i^*(t) - Q_i^*(t)}{\partial q_i} = q_i^{t-2}[t(1 - q_i) - 1]A. \tag{A8}$$

Substituting (A3) and (A6)–(A8) into (A5) and simplifying gives

$$\begin{aligned}\frac{\partial^2 h_{2i}(t|T^*)}{\partial T^* \partial q_i} &= Aq_i^{t-1}q_i^{T^*-1}[1 - G_i(\underline{\omega}_i)](T^* \ln q_i\{[1 - G_i(\underline{\omega}_i)] - G_i(\underline{\omega}_i)q_i^{T^*}\} \\ &\quad + \{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i)q_i^{T^*}\} + \ln q_i[t(1 - q_i) - 1] \\ &\quad \times \{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i)q_i^{T^*}\}) \div \{[1 - G_i(\underline{\omega}_i)] + G_i(\underline{\omega}_i)q_i^{T^*}\}^3.\end{aligned}$$

It is sufficient that $T^* < -1/\ln q_i$ and $t < 1/(1 - q_i) = 1/p_i$ for $\partial^2 h_{2i}(t|T^*)/\partial T^* \partial q_i > 0$. Q.E.D.

## Appendix B

This appendix describes the variables used in the empirical analysis of Section IV.

| | |
|---|---|
| INITIALWAGE | Initial hourly wage for job 2 |
| FULLTIME | Dummy variable that equals one if hours per week worked in job 2 were greater than 30 |
| SWITCH | Dummy variable that equals one if the respondent switched occupations between job 1 and job 2 |
| TENURE1 | Weeks of tenure in job 1 |
| HSCHGRD | Dummy variable that equals one if the respondent completed 12–15 years of schooling as of the start of job 2 |
| COLLEGEGRD | Dummy variable that equals one if the respondent completed 16 years (or more) of schooling as of the start of job 2 |
| WEEKSOUTWORK | Measures the number of weeks out of work between job 1 and job 2 |
| NOOUTWORKSPELL | Dummy variable that equals one if there were no weeks spent out of work between job 1 and job 2 |
| MARRIED | Dummy variable that equals one if the respondent was married as of the start of job 2 |
| NMARRIED | Dummy variable that equals one if the respondent was not married as of the start of job 2 |
| MALE | Dummy variable that equals one if the respondent was male |
| FEMALE | Dummy variable that equals one if the respondent was female |
| HISPANIC | Dummy variable that equals one if the respondent was Hispanic |
| BLACK | Dummy variable that equals one if the respondent was black |

| | |
|---|---|
| UNIONJOB2 | Dummy variable that equals one if job 2 was unionized |
| GOVTJOB2 | Dummy variable that equals one if job 2 was a government job |

The following dummy variables are industry and occupation dummies that were set to one if the respondent's classification number fell within the specified range for job 2. The numbers represent the 1970 census industry and occupation classification codes.[9]

| Industry | Classification |
|---|---|
| Mining | 047–057 |
| Construction | 067–077 |
| Manufacturing | 107–398 |
| Transportation/communications/public utilities | 407–479 |
| Trade | 507–698 |
| Finance/insurance/real estate | 707–718 |
| Business/repair services | 727–759 |
| Personal services | 769–798 |
| Entertainment/recreation services | 807–809 |
| Professional services | 828–897 |
| Public administration | 907–937 |

| Occupation | Classification |
|---|---|
| Managers | 201–245 |
| Sales | 260–284 |
| Clerical | 301–395 |
| Craftsmen | 401–575 |
| Operatives | 601–715 |
| Laborers | 740–765 |
| Farmers | 801–824 |
| Services | 901–965 |
| Household | 980–984 |

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[9] The excluded industry in the empirical analysis of the paper was agriculture/forestries/fisheries (017–028). The excluded occupation was professionals (001–195).

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