Job Duration, Seniority, and Earnings
Author(s): Katharine G. Abraham and Henry S. Farber
Source: *The American Economic Review*, Vol. 77, No. 3 (Jun., 1987), pp. 278-297
Published by: American Economic Association
Stable URL: http://www.jstor.org/stable/1804095
Accessed: 23-10-2015 11:51 UTC

Job Duration, Seniority, and Earnings

By KATHARINE G. ABRAHAM AND HENRY S. FARBER*

An important stylized fact about labor markets is that workers with longer seniority with their current employer have higher earnings than other workers with the same total labor market experience. This study shows that the measured positive cross-sectional return to seniority is largely a statistical artifact due to the correlation of seniority with an omitted variable representing the quality of the worker, job, or worker-employer match. The implication is that earnings do not, in fact, rise very much with seniority.

It is a commonly accepted empirical finding that workers with more seniority on their current job earn more than other workers with the same total labor market experience. The standard explanations for a positive correlation between seniority and earnings are based on the existence of implicit employment contracts under which earnings grow with time on the job in order to provide workers with appropriate incentives regarding turnover and/or effort. For example, if a job involves investment in firm-specific training, then it may be optimal for workers and employers to structure implicit employment agreements such that compensation is deferred until late in the job so that workers will not quit (taking their specific capital with them).[1] Another possible motivation for such a deferral arrangement exists where effort is important. The promise of eventual compensation in excess of the opportunity value of time provides the worker with an incentive to exert the appropriate level of effort on the job. A worker who left or whose performance fell below agreed-upon standards and in consequence was fired would lose the opportunity to enjoy the benefits of the high deferred wages.[2]

The firm-specific human capital explanation and effort-incentive wage deferral explanations differ in their implications regarding the relationship between earnings and productivity growth over the work life.[3] However, they share the implication not only that the seniority-earnings profile will be upward sloping, but also that there will be a positive return to seniority even after controlling for total labor market experience.

The empirical support for these views of the labor market rests entirely on the positive cross-sectional association between seniority and earnings, but this is not sufficient evidence to establish that earnings rise with seniority. An alternative interpretation of the

*The Brookings Institution, 1775 Massachusetts Avenue, NW, Washington, DC 20036, and Department of Economics, E52–252f, MIT, Cambridge MA 02139, respectively. We thank Joe Altonji for providing us with the data, Jim Powell for providing valuable econometric advice, and Leslie Sundt for able research assistance in preparing the data for analysis. Useful comments on an earlier version of this paper were received from Altonji, Jerry Hausman, Steve Nickell, Bob Topel, and participants in numerous workshops. Farber received support from the National Science Foundation. The views expressed here are our own and should not be attributed to the trustees, officers, or staff members of the institutions with which we are affiliated.

[1]See Gary Becker (1964), Jacob Mincer (1974), and Dale Mortensen (1978) for discussions of investment in firm-specific training. Mincer and Boyan Jovanovic (1981) present an analysis of the relationships among seniority, mobility, and earnings that relies on investment in specific human capital.

[2]See W. Kip Viscusi (1980), Becker and George Stigler (1974), and Edward Lazear (1979) for models in which wage deferral provides this sort of incentive for workers. George Akerlof and Lawrence Katz (1986) argue that such deferral arrangements cannot, in fact, yield efficient effort levels unless the present value of promised earnings at the start of the job exceeds the present value of earnings on the next-best alternative job.

[3]James Medoff and Abraham (1980, 1981) and Abraham and Medoff (1982) offer evidence on the relationship between seniority-related earnings growth and seniority-related productivity growth.

cross-sectional evidence is based on the idea that workers who are 1) better workers, 2) in better jobs, or 3) in better worker-employer matches earn more throughout their jobs and also stay on their jobs longer. It is straightforward to show that the distribution of seniority in a cross section has a higher mean for workers on longer jobs. Thus, as long as those workers who earn more from the start have longer average completed job durations, the ordinary least squares estimate of the return to seniority in a cross-section earnings function is biased upward.

The key testable link in this argument is that workers in long jobs earn more from the start than observationally equivalent workers in short jobs. Why might this be true? First, some workers may be both more stable and more productive than others. To the extent that there are turnover and training costs, stability per se can raise an employee's value to the firm. Second, some employers may choose to pay higher wages than others (some jobs are "better" than others), so that workers are unlikely either to shirk or to quit.[4] Finally, some worker/job matches may be better than others, in the sense that the worker is more productive in those matches than in other possible matches, so that the specific value of the match is shared between the employer and the worker, and the match is less likely to be broken off.[5] To the extent that worker, job, and/or match quality are unmeasured, they represent omitted variables in a cross-section earnings regression.

Here we develop a simple stochastic model of earnings determination that illustrates the upward bias in the estimated return to seniority just described and provides the basis for two related approaches to correcting this bias. These approaches are implemented using data from the *Panel Study of Income Dynamics*. The results of this empirical analysis reveal 1) that workers on longer jobs earn more in every year on the job, and 2) that much of the apparent cross-section return to seniority reflects omitted variable bias.

I. Completed Job Duration in the Earnings Function

Suppose that the earnings of a particular worker i on job j in year t can be written

$$(1)\quad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \mu_{ij} + \eta_{ijt},$$

where W = hourly earnings, S = current seniority (tenure), β_1 = return to seniority, EXP = pre-job experience, β_2 = return to pre-job experience, μ = a person/job-specific error term representing the excess of earnings enjoyed by this person on this job over and above the earnings that could be expected by a randomly selected person/job combination, and η = a person/job/time-period-specific error term.

For simplicity of exposition, other factors that might influence earnings are omitted from the theoretical discussion and all variables are assumed to be measured as deviations from their means. In this formulation, μ_{ij} captures the net influence of three unobservables on hourly earnings: unobserved person quality, unobserved job quality, and unobserved match quality. The error μ_{ij} is assumed to be fixed over the course of a job and may be correlated with S and EXP. The error η_{ijt} is assumed to be orthogonal to S, EXP, and μ.

In equation (1), β_2 represents the returns to experience per se, including the returns to general human capital and any other growth in earnings that occurs automatically with time in the labor market. Additionally, earnings are likely to grow with experience because more experienced workers typically end

[4] The efficiency wage literature suggests various possible reasons why some employers might pay higher wages than others to workers of equal quality. These include differences in the costs of turnover, differences in the costs of monitoring worker shirking, and differences in the value of worker loyalty. Carl Shapiro and Joseph Stiglitz (1984) and Jeremy Bulow and Lawrence Summers (1986) present formal efficiency wage models. Janet Yellen (1984), Stiglitz (1984), and Katz (1986) survey the literature.

[5] Mortensen and Jovanovic (1979) both present theoretical analyses of the effects of heterogeneous match quality that have these implications.

up in better jobs and/or better matches.[6] This will be reflected in higher values of μ_{ij} as EXP_{ij} increases. Let the relationship between pre-job experience and μ be approximated by

$$\mu_{ij} = \alpha EXP_{ij} + \phi_{ij}, \tag{2}$$

where α captures the growth in μ_{ij} with experience and ϕ_{ij} is the component of μ_{ij} that is uncorrelated with EXP_{ij}. Substituting into equation (1) yields

$$\ln W_{ijt} = \beta_1 S_{ijt} + (\beta_2 + \alpha) EXP_{ij} + \phi_{ij} + \eta_{ijt}, \tag{3}$$

where β_1 is the total return to seniority and $\beta_2 + \alpha$ is the total return to pre-job experience including both the return to experience per se and systematic returns to search. The net return to seniority is appropriately defined as the excess of the growth in earnings on a given job over and above the total returns to general labor market experience, or $\beta_1 - (\beta_2 + \alpha)$.

In practice, earnings functions are generally estimated using cross-section data and ϕ_{ij} is not observable. The standard cross-section earnings equation is

$$\ln W_{ijt} = b_1 S_{ijt} + b_2 EXP_{ij} + v_{ijt}, \tag{4}$$

where v is the estimating equation error. Omission of ϕ_{ij} from equation (4) means that b_1 and b_2 may be biased estimators of the return to seniority, β_1, and the return to experience, $\beta_2 + \alpha$. Deriving the expected values of b_1 and b_2 requires knowing more about the relationship between ϕ_{ij} and S_{ijt}. In particular, we need to know the partial relationship between ϕ_{ij} and S_{ijt}, holding EXP_{ij} constant.[7]

It was argued above that good workers and workers in good jobs or good matches are likely to stay on their jobs longer. Formally, this implies that the completed duration of jobs is positively related to μ_{ij}. Let this relationship be expressed as

$$D_{ij} = \gamma \mu_{ij} + \varepsilon_{ij}, \tag{5}$$

where D_{ij} is the completed length of the current job, γ is a parameter that summarizes the relationship between D and μ, and ε_{ij} captures the variation in completed job duration that cannot be linked to the earnings advantage associated with worker, job, and/or match quality.[8] Substituting from equation (2),

$$D_{ij} = \alpha\gamma EXP_{ij} + \gamma\phi_{ij} + \varepsilon_{ij}. \tag{6}$$

Holding initial experience constant, completed job duration is positively related to ϕ_{ij}.

What does this tell us about the relationship between ϕ_{ij} and S_{ijt}? In a cross section of individuals, those with longer current seniority are likely to be on longer-lasting jobs. More formally, if each year of any given job is equally likely to be represented in the cross-section sample of observations used to estimate the earnings function, then on average the observed seniority on the job will be halfway through the job so that

$$\begin{aligned} S_{ijt} &= 1/2 \times D_{ij} + \xi_{ijt} \\ &= 1/2 \times \alpha\gamma EXP_{ij} + 1/2 \times \gamma\phi_{ij} \\ &\quad + 1/2 \times \xi_{ij} + \xi_{ijt}, \end{aligned} \tag{7}$$

where ξ_{ijt} is a random variable with zero mean. Thus, the existence of a positive relationship between ϕ and D, holding pre-job experience constant, implies the existence of a positive relationship between ϕ and S,

[6] See Kenneth Burdett (1978), Jovanovic, and Robert Topel (1986).

[7] By construction, ϕ_{ij} is uncorrelated with EXP_{ij}. However, if S_{ijt} and EXP_{ij} are correlated, omitting ϕ_{ij} may bias both b_1 and b_2.

[8] There is a large literature that estimates turnover probabilities as a function of wage rates. Donald Parsons (1977) and Richard Freeman and Medoff (1984) present surveys of some of this work. Job duration can be viewed as an inverse turnover measure. The prediction that job duration should be positively related to μ is consistent with evidence that, *ceteris paribus*, turnover rates are negatively related to wages.

holding pre-job experience constant. The distribution of ξ_{ijt} will vary depending upon the completed length of the job. However, its mean is always zero, as are its covariances with μ_{ij}, ϕ_{ij}, D_{ij}, and EXP_{ij}.

Using the relationships in equations (4) to (7), it can be shown that the expected value of the seniority coefficient, b_1, in equation (3) is

$$(8)\quad E(b_1)=\beta_1+\left[\gamma\times\text{var}(EXP_{ij})\times\text{var}(\phi_{ij})\right] \Big/ \left[2\times\left[\text{var}(S_{ijt})\times\text{var}(EXP_{ij}) - \text{cov}^2(S_{ijt},EXP_{ij})\right]\right].$$

$E(b_1)$ is larger than β_1 provided γ (the coefficient summarizing the relationship between D and μ) is positive. The expected value of the pre-job experience coefficient, b_2, in equation (4) is

$$(9)\qquad E(b_2)=\beta_2+\alpha K_1,$$

where the multiplicand of α is

$$(10)\quad K_1=\left[\text{var}(EXP_{ij})\times\left\{\text{var}(S_{ijt}) - 1/4\times\text{var}(D_{ij})+1/4\times\text{var}(\varepsilon_{ij})\right\}\right] \Big/ \left[\text{var}(S_{ijt})\times\text{var}(EXP_{ij}) - \text{cov}^2(S_{ijt},EXP_{ij})\right].$$

It is straightforward to show that K_1 is positive but less than one so that the experience coefficient (b_2) is a downward-biased estimate of $\beta_2+\alpha$. Therefore, b_1-b_2 has an expected value greater than the true net return to seniority, $\beta_1-(\beta_2+\alpha)$, both because b_1 is an upward-biased estimate of β_1 and, secondarily, because b_2 is a downward-biased estimate of $\beta_2+\alpha$.

One approach to correcting the estimates of b_1 and b_2 from equation (4) is to find an instrument for the seniority variable.[9] Equation (7) suggests a suitable instrument: ξ_{ijt}, which equals $S_{ijt}-1/2\times D_{ij}$. By construction, this instrument is correlated with seniority but orthogonal to everything else in the equation. The vector of parameters estimated using ξ as an instrument for S equals

$$(11)\qquad \bar{b}=(Z'X)^{-1}Z'W,$$

where $\bar{b}$ is a 2×1 vector containing $\bar{b}_1$ and $\bar{b}_2$, X is the $n\times2$ matrix containing values of S_{ijt} and EXP_{ij}, Z is the $n\times2$ matrix containing values of ξ_{ijt} and EXP_{ij}, and W is the $n\times1$ matrix of observations on the log wage. The instrumental variables estimator of the seniority coefficient, $\bar{b}_1$, has expected value β_1. In addition, the instrumental variables estimator of the experience coefficient, $\bar{b}_2$, has expected value $\beta_2+\alpha$. Finally, $\bar{b}_1-\bar{b}_2$ is an unbiased estimate of the net return to seniority.

The key assumption underlying the proposed instrumental variables estimator is that observed seniority in a cross section is, on average, half completed job duration. This assumption may not hold exactly. For example, a sample might include a disproportionate number of workers near the start of their work lives who are, on average, closer to the start of their jobs than they are to the end. In this case, the deviation of seniority from one-half completed duration may be correlated with duration and thus not be a valid instrument for seniority. However, if some more general relationship between seniority and completed job duration holds, such as

$$(12)\qquad S_{ijt}=\omega_0+\omega_1 D_{ij}+\xi_{ijt},$$

then, regardless of the values of ω_0 and ω_1, the residual from this regression is an appropriate instrument for seniority. Using estimates of these residuals to calculate the instrument yields consistent estimates of the returns to experience and seniority.

An alternative approach to removing the upward bias in the estimated total return to seniority is to control explicitly for completed job duration in the earnings equation.

[9]Joseph Altonji and Robert Shakotko's (1987) analysis of the return to seniority uses this approach, though with a somewhat different instrument than ours.

Intuitively, the tenure coefficient is biased only because S_{ijt} is associated with D_{ij}, which in turn is correlated with ϕ_{ij}. This suggests that controlling for D_{ij} should eliminate the upward bias in the estimated tenure coefficient. Augmenting the standard cross-section earnings equation by adding D_{ij} as an explanatory variable yields

$$(13)\quad \ln W_{ijt} = \hat{b}_1 S_{ijt} + \hat{b}_2 EXP_{ij} + \hat{b}_3 D_{ij} + \pi_{ijt},$$

where π is the estimating equation error.

It can readily be shown that $E(\hat{b}_1)$ equals β_1 so that $\hat{b}_1$ is an unbiased estimator of the gross return to seniority. The experience coefficient in equation (13) has expected value:

$$(14)\quad E(\hat{b}_2) = \beta_2 + \alpha K_2,$$

where

$$(15)\quad K_2 = \left[\mathrm{var}(EXP_{ij}) \times \mathrm{var}(\varepsilon_{ij})\right] / \left[\mathrm{var}(D_{ij}) \times \mathrm{var}(EXP_{ij}) - \mathrm{cov}^2(D_{ij}, EXP_{ij})\right].$$

The value of K_2 is positive but less than one, which means that $\hat{b}_2$ is an underestimate of $\beta_2 + \alpha$. Thus, so long as α is positive, $\hat{b}_1 - \hat{b}_2$ is an upward-biased estimate of the net return to seniority ($\beta_1 - [\beta_2 + \alpha]$). None of these results are sensitive to the precise form of the linear relationship between seniority and completed duration, and this approach can be used to bound the true return to seniority.

A major attraction of the augmented OLS approach is that it provides a direct estimate of the relationship between completed job duration and earnings, $\hat{b}_3$. This is an indicator of the importance of the relationship of individual, job, and/or match heterogeneity with earnings through job duration. In addition, it serves as a useful device for investigating the underlying hypothesis that better workers, jobs, or matches are associated with higher earnings *throughout* the job. In terms of the earlier analysis, it can be shown that

$$(16)\quad E(b_3) = (1 - K_2)/\gamma,$$

where K_2 is as just defined. Thus, $\hat{b}_3$ is a downward-biased estimate of $1/\gamma$, the coefficient of the regression of μ_{ij} on D_{ij}.

The preceding discussion assumes that individual-, job-, and match-specific earnings components all have the same incremental association with job duration. A more general specification would allow for separate earnings increments associated with unobserved individual characteristics and job/match characteristics. Given that any bias in the estimated return to seniority attributable to omitted earnings components is mediated through the relationship of those omitted earnings components with completed job duration, all of the results just described for the simple model are unaffected by a generalization to multiple unobserved earnings components.

Using the identity that links total experience with pre-job experience and seniority ($EXP_{ijt} = EXP_{ij} + S_{ijt}$) and ignoring second-order terms, valid inferences regarding the net return to seniority can be drawn from an earnings function specification that includes either pre-job experience or total experience along with seniority. Where pre-job experience is used, the net return to seniority ($\beta_1 - [\beta_2 + \alpha]$) is calculated as the difference between the coefficient on seniority and the coefficient on pre-job experience. Where total experience is used, the net return to seniority is simply the coefficient on seniority. To facilitate discussion regarding the net return to seniority, the empirical analysis proceeds using total experience.[10]

[10] For the instrumental variables estimator, this requires the minor modification that total experience (and its square, if included) must be instrumented for along with seniority. The natural additional instruments are pre-job experience (and its square). In a model without squared terms, the IV estimator of the specification using pre-job experience and the IV estimator using total experience with pre-job experience as an ad-

II. Estimating Completed Job Duration

The first step in implementing the analysis described in the preceding section is to derive a measure of completed job duration. The *Panel Study of Income Dynamics* (*PSID*) is used in the empirical analysis. Unfortunately, there is a general problem with virtually all longitudinal data sets, including the *PSID*, when information is required on the completed duration of a spell of any kind. This is that the individuals are followed for only a limited period of time so that there are likely to be many jobs which do not end by the date at which the individual is last observed. Some procedure must be used to impute completed durations to these jobs.

We take the approach of estimating a parametric model of job duration that accounts for the censoring of duration in those jobs for which the end is not observed. This model is then used to compute an estimate of expected completed job duration *conditional* on the job lasting at least as long as the last observed seniority level. This estimate is used as the measure of completed job duration for the censored spells. The actual completed job duration is used for jobs for which the end is observed. This procedure has the advantage of using all available information on duration.

A. *The Jobs Sample*

All of the subsequent analysis is performed using data for male household heads aged 18–60 who participated in the *PSID*.[11] We used only observations from the random national sample portion of the *PSID* (the so-called Survey Research Center or SRC subsample). Persons who were retired, permanently disabled, self-employed, employed by the government, or residents of Alaska or Hawaii were excluded from the sample. Because we were concerned that different processes might govern tenure attainment and earnings in the union sector than in the nonunion sector, we also excluded observations on unionized jobs.[12] We were also concerned about differences across occupations in the processes determining job duration and earnings. In what follows, we therefore focus our discussion on results for two occupational subgroups. The first is a subset of white-collar occupations including nonunion professional, technical, and managerial employees.[13] The second subsample is comprised of nonunion blue-collar employees. In each year from 1968 through 1981 in which those individuals satisfying our selection criteria were household heads, information was available on number of years they had

ditional instrument yield identical estimates of the underlying parameters.

[11]Altonji kindly provided us with an extract containing the variables which we used in performing our analyses. The procedures followed in creating this extract are described in detail in an appendix to Altonji and Shakotko. In order to delete non-SRC subsample observations, we added information on whether a given individual was part of the SRC subsample or the nonrandom continuation subsample from the *Survey of Economic Opportunity*. We also smoothed the tenure variable in instances where a given individual had been assigned the midpoint value of a tenure interval. If the individual on a given job changed tenure intervals in succeeding years, we computed a smooth tenure variable forward and backward from the change point. If all observations for the individual on a given job were in the same tenure interval, we computed a smooth tenure variable forward and backward from the middle observed year on the job assuming that tenure in that year was equal to the midpoint of the interval.

[12]In some years, unionization refers to union membership, and in other years, to working on a job covered by a collective bargaining agreement. Where both measures were available, collective bargaining coverage was used. Observations on jobs for which the worker changed union status during the course of the job do not appear in the sample. In 371 jobs, workers were coded nonunion in some years and union in others. If 1) at least two-thirds of the observed years on one of these jobs were coded nonunion, 2) there were no runs of three or more years coded union, and 3) the first and last years observed on the job were coded nonunion, then the entire job was considered a nonunion job and was included in our sample.

[13]Excluded from the analysis were the approximately 16 percent of all nonunion jobs that were clerical and sales jobs. This is too few to support separate job duration models for these groups. At the same time, these jobs seemed likely to differ significantly from other white-collar jobs.

TABLE 1—SELECTED CHARACTERISTICS OF JOBS SAMPLES FOR OCCUPATIONAL SUBGROUPS[a]

	Managerial and Professional Nonunion			Blue Collar Nonunion		
	All	Complete	Censored	All	Complete	Censored
Proportion with Years of Tenure at Last Date Job Observed in Range:						
$T \leq 1$	.280	.483	.147	.483	.639	.294
$1 < T \leq 3$	.235	.251	.224	.215	.216	.213
$3 < T \leq 10$	.257	.205	.291	.109	.115	.268
$T > 10$	.228	.0614	.338	.119	.0310	.224
Mean of:[b]						
Years of Tenure at Last Date Job Observed	6.76 [8.05]	2.96 [4.03]	9.25 [9.00]	4.07 [6.43]	1.9 [3.2]	6.7 [8.1]
Years of Pre-Job Experience	9.76 [8.09]	10.00 [7.76]	9.61 [8.30]	10.68 [9.49]	9.8 [9.0]	11.8 [10.]
(Years Pre-Job Experience)^2	160.8 [253.7]	160.2 [230.1]	161.1 [268.1]	204.2 [348.6]	175.7 [325.1]	238.5 [372.0]
Years of Education	14.6 [2.09]	14.5 [2.1]	14.7 [2.1]	11.3 [2.43]	11.4 [2.3]	11.3 [2.6]
Proportion:						
Nonwhite	.0416	.0537	.0337	.138	.139	.137
Married	.862	.854	.867	.845	.823	.872
Disabled	.0528	.0563	.0505	.0932	.0761	.114
Managerial	.483	.517	.461	–	–	–
Prof., Tech.	.517	.483	.539	–	–	–
Foreman, Craft	–	–	–	.439	.421	.461
Oper., Labor	–	–	–	.561	.579	.539
No. of Observations	985	391	594	1417	775	642

[a]Except for tenure and years of previous experience, all variables are reported as of the first year the job was observed. Previous experience was computed as the difference between reported experience in the first year the job was observed and seniority at that point.
[b]Standard deviations are shown in brackets.

held their current job, number of years they had worked prior to taking the current job, years of education, race, marital status, disability status, occupation, industry, region, and earnings.

There are 985 jobs held by 706 individuals represented in the nonunion white-collar subsample, and 1417 jobs held by 831 individuals represented in the nonunion blue-collar sample. Our concern at this point is with ascertaining how long each of these jobs ultimately lasted.

Various characteristics of the jobs in each of the samples are reported in Table 1. Variables that can change over time in an unpredictable fashion (for example, marital status, occupation) are assumed constant and measured at the first point the job is observed in the sample. Any jobs for which there are some blue-collar years and some white-collar years appear in both occupational subsamples. The last observed seniority on a job is always considered to be the seniority at the last date the person is observed with an employer, whether or not there has been a change of occupation during the course of the job. There were 87 cases in which an individual reported moving from blue-collar status to white-collar status, and 83 cases in which an individual reported moving from white-collar status to blue-collar status while a job was in progress.[14]

We observe the actual completed duration for 391 of 985 jobs in the white-collar sam-

[14]It is likely that these numbers overstate the true number of changes since there are undoubtedly some errors in classification that produce spurious movements between the two broad occupational groups.

ple and for 775 of 1417 jobs in the blue-collar sample. Not surprisingly, a large proportion of the completed jobs are relatively short. However, in both samples, there are a sizable number of completed jobs lasting 3 to 10 years and over 10 years. Longer jobs are more common among the still-in-progress jobs.

B. *Specification and Estimation of the Job Duration Model*

The proportional hazard Weibull specification serves as the basis of the estimation reported here. In that specification, the probability that a job has completed duration (D) greater than or equal to T is

$$(17)\qquad Pr(D \geq T) = \exp[-\lambda T^{\tau}],$$

where τ is a positive parameter. The proportional hazard assumption implies that

$$(18)\qquad \lambda = e^{-Z\Gamma},$$

where Z is a vector of observable individual characteristics hypothesized to affect job duration and Γ is a vector of parameters.

If the parameters of this distribution are estimated, there is some ambiguity in the interpretation of the estimate of τ. The obvious interpretation is that the estimated value of τ indicates "true" duration dependence in the hazard of a job ending. An alternative interpretation is that the estimate of τ is biased downward by unmeasured heterogeneity in match quality. For the purposes of this study, we are interested only in an accurate estimate of completed duration, and we proceed with the simple Weibull specification.[15]

The contribution to the likelihood function made by a completed job is the probability-density that the job lasted *exactly* S_f years given that the job lasted at least S_0 years.[16] Given a Weibull distribution for duration, this is

$$(19)\qquad Pr(D = S_f | D \rangle S_0) = \lambda\tau S_f^{\tau-1}\exp\left[-\lambda\left(S_f^{\tau} - S_0^{\tau}\right)\right].$$

Similarly, the contribution to the likelihood function made by a job with a censored duration is the probability that the job lasted *more than* S_f years given that the job lasted at least S_0 years. This is

$$(20)\qquad Pr(D > S_f | D \rangle S_0) = \exp\left[-\lambda\left(S_f^{\tau} - S_0^{\tau}\right)\right].$$

The log-likelihood function is formed from these probabilities as

$$(21)\qquad \ln(L) = \sum_j \left\{C_j \ln Pr\left(D_j > S_{fj} | D_j \rangle S_{0j}\right) + (1 - C_j)\ln Pr\left(D_j = S_{fj} | D_j \rangle S_{0j}\right)\right\},$$

where j indexes jobs and C_j is an indicator variable that equals one if the completed job duration is censored (i.e., the job does not end during the sample period) and equals zero otherwise (i.e., the completed job duration is observed).[17]

Table 2 contains estimates of the Weibull job duration model estimated over the subsamples of 985 white-collar jobs and 1417 blue-collar jobs, respectively. These estimates were derived by maximizing the likelihood function defined above with respect to the parameters Γ and τ.[18] In interpreting the

[15] See Tony Lancaster (1979) for a parametric approach to the problem of estimating unmeasured heterogeneity in a Weibull model of unemployment duration. James Heckman and Burton Singer (1984) present a nonparametric approach to estimating duration models with unmeasured heterogeneity.

[16] It is important to condition on the length of the job as of the date it is first observed because the sampling scheme is such that jobs will not be observed unless they last long enough to make it to the start of the sample period.

[17] Note that this specification of the likelihood function assumes that unmeasured factors affecting completed job durations are independent across spells. However, within each sample, there are multiple observations on job durations for some individuals. Given the nonlinear nature of the model, an appropriate tractable procedure for accounting for the induced correlation is not obvious.

[18] The algorithm described by Ernst Berndt et al. (1974) was used to find the maximum.

TABLE 2—SELECTED COEFFICIENTS FROM FINAL TENURE MODELS[a]

	Managerial and Prof. Nonunion (1)	Blue Collar Nonunion (2)
Γ (Inverse Baseline Hazard, $\lambda = e^{-z\Gamma}$)		
Years of Experience	$-.0288$	.0611
	(.0204)	(.0010)
(Years of Experience)^2	.00123	$-.00113$
	(.00071)	(.00027)
Years of Education	.0699	.0243
	(.0244)	(.0136)
Nonwhite (yes = 1)	$-.412$	$-.0387$
	(.224)	(.0878)
Married (yes = 1)	.270	.464
	(.135)	(.073)
Disabled (yes = 1)	.0768	.180
	(.218)	(.113)
Manager (yes = 1)	$-.0764$	–
	(.1121)	
Foreman, Craftworker (yes = 1)	–	.0656
		(.0623)
"Duration" Parameter		
τ	.380	.394
	(.028)	(.017)
Log-Likelihood	-900.4	-1097.9
Sample Size	985	1417

[a]These coefficient estimates are from a Weibull proportional hazards model implemented using the jobs samples described in Table 1. All explanatory variables are reported as of the start of the job. Professional/technical employees are the omitted occupational group in the col. 1 model and operatives/laborers are the omitted occupational group in the col. 2 model. The numbers shown in parentheses are asymptotic standard errors.

estimates of the determinants of the baseline hazard (λ), recall that the hazard rate was specified such that $\lambda = e^{-Z\Gamma}$. Thus, an increase in a variable with a positive coefficient reduces λ and increases the expected duration of the job. The hypothesis that the models of completed job duration for the two occupational groups are the same is strongly rejected.[19] The marginal effect of pre-job experience on job duration for white-collar workers is never statistically significant at the .05 level, while among blue-collar workers, having more pre-job experience has a significant positive association with completed job duration. The estimates also suggest that education has a stronger positive relationship with job duration in white-collar occupations than in blue-collar occupations.

C. *Prediction of Job Duration for Incomplete Jobs*

We used the parameter estimates from the appropriate column of Table 2 to predict the expected completed job duration of each of the incomplete jobs in each sample. This expectation is computed conditionally on the job lasting longer than the last observed seniority. Note that the job duration model we have estimated is based on data for the preretirement period. It will capture the net effects of quit and layoff processes on job duration, but it will not capture the effect of the competing retirement process which comes into play for older workers. If we predicted job durations without taking retirement into account, some would be implausibly long. We therefore assume that all jobs that are in progress when the worker reaches age 65 end at that point. For an individual/job match with observable characteristics Z that has lasted S_f years as of the last date we observe it, the conditional expected completed job duration is

$$(22)\quad \hat{E}(D|D\rangle S_f)$$
$$= \frac{1}{Pr(D>S_f)} \int_{S_f}^{S_{65}} \lambda\tau t^{\tau} e^{-\lambda t^{\tau}} dt + \frac{Pr(D>S_{65})}{Pr(D>S_f)} * S_{65},$$

where S_{65} represents the seniority attained if a match lasts until the worker turns 65,

$$(23)\qquad Pr(D\rangle S_f) = \exp[-\lambda S_f^{\tau}],$$
$$Pr(D \geq S_{65}) = \exp[-\lambda S_{65}^{\tau}],$$
and $$\lambda = e^{-Z\Gamma}.$$

[19]The hypothesis that the parameters of the models for the two subgroups are identical except for a constant shift and the occupation dummies in $Z\Gamma$ can be rejected at any reasonable level of significance. The test statistic, distributed as χ^2 with 22 degrees of freedom, is 52.3 (p-value < .001). The independence assumption of this test is not strictly satisfied, since the two samples contain some jobs in common.

TABLE 3—DISTRIBUTION OF COMPLETED JOB DURATIONS[a]

Proportion with Completed Job Duration in Range:	Managerial and Professional Nonunion			Blue Collar Nonunion		
	All	Complete	Censored	All	Complete	Censored
$D \leq 1$	.193	.483	.00168	.354	.639	.0109
$1 < D \leq 3$	.105	.251	.00842	.155	.216	.0826
$3 < D \leq 10$	.182	.205	.167	.253	.115	.419
$D > 10$	.521	.0614	.823	.238	.0310	.488
No. of Observations	985	391	594	1417	775	642

[a]For the completed job subsample, the distribution of actual completed job duration is shown. For the censored job subsample, the distribution of predicted completed job duration is reported.

With an appropriate change of variables, this expected duration can be computed numerically using incomplete *gamma* functions. We also use the square of completed job duration in the earnings function estimation. For incomplete jobs, this is estimated analogously to completed job duration.

As noted earlier, actual job duration was observed for a substantial fraction of the jobs in both samples, and this measure was used in these cases. For the jobs with censored duration, the predicted values were used. Table 3 contains the breakdown of the completed durations of the jobs in the two occupational samples. As expected, the censored jobs are generally longer than the completed jobs.

III. Is Seniority Half Completed Job Duration?

Samples of individual-year observations from the two *PSID* jobs samples just discussed are used to estimate the earnings functions which constitute the core of our analysis. Recall that the samples consist only of nonunion male heads of households. There are 3493 individual-year observations on the 706 workers in the 985 white-collar jobs, and 3554 individual-year observations on the 831 workers in the 1417 blue-collar jobs. The means and standard deviations of the central variables for each of the samples are contained in the first column of Tables 4a and 4b.

Two alternative approaches to removing the bias in the estimated return to seniority have been suggested: 1) using the residual from a regression of seniority on completed duration as an instrument for seniority, and 2) including completed duration as a regressor in the earnings function. Both approaches started from the recognition that a true random sample of years from jobs would have the property that, on average, seniority equals one-half completed duration. This is equivalent to the hypothesis that in a regression of seniority on completed duration the constant term is zero and the coefficient on completed duration is .5.

In order to investigate this directly, we regressed seniority on the completed duration for the two occupational subsamples. The results are

$$S_{ijt} = \underset{(.18)}{-2.24} + \underset{(.007)}{.534} \cdot D_{ij} \qquad R^2 = .61 \quad \text{(white collar)}, \tag{24}$$

$$S_{ijt} = \underset{(.097)}{-1.31} + \underset{(.005)}{.550} \cdot D_{ij} \qquad R^2 = .75 \quad \text{(blue collar)},$$

and the numbers in parentheses are OLS standard errors.[20] Clearly, the hypothesis

[20]Given that D_{ij} is a predicted value for the observations from censored jobs, the OLS standard errors are not appropriate. However, the inferences are unchanged by use of (computationally tedious) standard errors that account for the fact that D_{ij} is predicted as well as general heteroskedasticity of the form analyzed by Halbert White (1980). See Whitney Newey (1984) for computational details on the correct standard errors.

TABLE 4a—SELECTED COEFFICIENTS FROM ln (AVERAGE HOURLY EARNINGS) MODELS MANAGERIAL AND PROFESSIONAL NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	IV (2)	OLS (3)	OLS (4)
Years of Experience	18.14 [10.08]	.0349 (.0027)	.0392 (.0058)	.0288 (.0027)	.0263 (.0031)
(Years of Experience)^2	430.77 [407.84]	−.00062 (.00006)	−.00077 (.00014)	−.00048 (.00007)	−.00043 (.00007)
Years of Current Seniority	8.88 [8.34]	.0106 (.0011)	.00585 (.00128)	.00548 (.00178)	.00520 (.00256)
E(Completed Job Duration)	20.83 [12.18]	–	–	.0198 (.0024)	.0265 (.0050)
{E(Completed Job Duration)}^{2}	631.55 [505.56]	–	–	−.00035 (.00006)	−.00059 (.00016)
E(Job Duration) $\times[=1$ if $3<$ Seniority $\leq 10]$	6.02 [10.46]	–	–	–	−.00094 (.00432)
{E(Job Duration)}^{2} $\times[=1$ if $3<$ Seniority $\leq 10]$	165.4 [325.3]	–	–	–	.00009 (.00015)
E(Job Duration) $\times[=1$ if Seniority $>10]$	11.09 [15.78]	–	–	–	−.00798 (.00455)
{E(Job Duration)}^{2} $\times[=1$ if Seniority $>10]$	380.1 [572.9]	–	–	–	.00030 (.00015)
R^2	–	.3696	.3575	.3871	.3883

[a]All models also include controls for education, race, marital status, disability, occupation, industry, region, and year. E (Completed Job Duration) is computed using the estimates in col. 1 of Table 2. The numbers shown in parentheses are standard errors. Sample size = 3493.

TABLE 4b—SELECTED COEFFICIENTS FROM ln (AVERAGE HOURLY EARNINGS) MODELS BLUE-COLLAR NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	IV (2)	OLS (3)	OLS (4)
Years of Experience	17.34 [11.14]	.0205 (.0024)	.0173 (.0040)	.0117 (.0026)	.0120 (.0026)
(Years of Experience)^2	424.70 [470.81]	−.00045 (.00006)	−.00042 (.00009)	−.00026 (.00006)	−.00028 (.00006)
Years of Current Seniority	6.31 [7.46]	.0142 (.0011)	.00290 (.00172)	.00241 (.00213)	−.00054 (.00302)
E(Completed Job Duration)	13.86 [11.75]	–	–	.0154 (.0021)	.0381 (.0057)
{E(Completed Job Duration)}^{2}	362.44 [444.45]	–	–	−.00014 (.00006)	−.00104 (.00024)
E(Job Duration) $\times[=1$ if $3<$ Seniority $\leq 10]$	4.57 [8.27]	–	–	–	−.00592 (.00538)
{E(Job Duration)}^{2} $\times[=1$ if $3<$ Seniority $\leq 10]$	102.13 [211.7]	–	–	–	.00031 (.00024)
E(Job Duration) $\times[=1$ if Seniority $>10]$	6.45 [12.90]	–	–	–	−.0241 (.0055)
{E(Job Duration)}^{2} $\times[=1$ if Seniority $>10]$	215.1 [461.9]	–	–	–	.00103 (.00024)
R^2	–	.3878	.3513	.4041	.4098

[a]All models also include the controls listed in Table 4a, fn. a. E (Completed Job Duration) is computed using col. 2, Table 2. Standard errors are shown in parentheses. Sample size = 3554.

that the coefficient on completed duration is .5 can be rejected in both samples at conventional levels of significance. This means that completed duration is not orthogonal to the deviation of seniority from one-half completed duration so that this deviation is not a valid instrument for seniority. However, the coefficient of completed duration is not far from one-half, which suggests that the general approach is valid.

On the basis of these results, the IV estimation of the earnings function proceeds using the residuals from the regression of seniority on completed duration to instrument seniority. We also present OLS estimates of earnings functions augmented with our measure of completed duration.

IV. Earnings Function Estimates

Consider an earnings function for individual i in job j in year t of the form:

$$(25)\quad \ln(W_{ijt}) = \theta_0 + \theta_1 S_{ijt} + \theta_2 E_{ijt} + \theta_3 E_{ijt}^2 + X_{ijt}\Omega + \varepsilon_{ijt},$$

where $\ln(W_{ijt})$ is the logarithm of real average hourly earnings, S_{ijt} is seniority, E_{ijt} is total experience, X_{ijt} is a vector of other individual characteristics, and ε_{ijt} represents unmeasured factors affecting earnings.[21] The coefficient θ_1 is the net return to seniority and corresponds to $\beta_1 - (\beta_2 + \alpha)$ in the model of Section I. The coefficient θ_2 is the return to general labor market experience and corresponds to $\beta_2 + \alpha$ in the model of Section I.

Tables 4a and 4b contain estimates of earnings functions for the samples of professional, technical, and managerial workers and of blue-collar workers, respectively. In addition to the regressors shown, all models include controls for education, race, marital status, disability, occupation, industry, region, and year. The standard errors presented are the "simple" standard errors computed from $\hat{\sigma}^2(X'X)^{-1}$ for the OLS models and from $\hat{\sigma}^2(Z'X)^{-1}$ for the IV models.[22]

In both tables, column 1 contains a standard OLS earnings equation that neither instruments for seniority nor includes completed job duration as a regressor. These estimates suggest that there are sizable returns both to general labor market experience and to seniority with a particular employer for workers in both occupational groups. The estimated net return to seniority is on the order of 1 to 1.5 percentage points per year.

The column 2 models were estimated by IV using pre-job experience, the square of pre-job experience, and the residual from the regression of seniority on completed job duration as instruments for total experience, the square of total experience, and seniority. The IV coefficient estimates provide consistent estimates of the return to general labor market experience and the net return to seniority. While instrumenting has relatively little effect on the estimated return to experience, the estimated net return to seniority falls substantially. The return for white-collar workers falls from 1.1 to 0.6 percent per year. The return for blue-collar workers falls from 1.4 to 0.3 percent per year; moreover, the corrected estimate is not significantly different from zero at conventional levels. This suggests that most of the cross-sectional correlation between earnings and seniority controlling for experience reflects the influence of omitted variables.

The coefficients in column 3 were estimated using ordinary least squares, but with completed job duration and its square added to the list of explanatory variables. The first

[21] We have estimated all of the models in this section with pre-job experience and its square in place of total experience and its square, and the qualitative conclusions emerging from the analysis do not change.

[22] These standard errors are not strictly appropriate for the estimation here because they do not account for the fact that the measure of completed job duration is predicted for the observations on censored jobs. Standard errors that are corrected both for this fact and for general heteroskedasticity (see Newey and White) were computed for a number of specifications. These were uniformly very close to the simple standard errors, and in no case was any inference resulting from a statistical test changed.

thing to note about these augmented OLS models is that the estimated returns to seniority are virtually identical to those obtained using the IV approach. This confirms that a substantial portion of the usual cross-sectional return to seniority is due to an omitted worker, job, and/or match quality measure. The virtual equality of the results using the augmented OLS approach and the IV approach also suggests that the bias in the augmented OLS estimates that we discussed in Section I is not a problem in practice.[23]

Perhaps the most interesting aspect of the augmented OLS estimates is that there is a *very* strong positive association between completed duration and earnings in both occupational groups. Consider two otherwise equivalent workers, one of whom holds a job that will eventually last 20 years and the other of whom holds a sequence of two 10-year jobs. In a white-collar occupation, the worker in the single 20-year job is estimated to earn 9.3 percent (standard error = 0.8 percent) more *in each year* than the worker in the sequence of 10-year jobs. In a blue-collar occupation, the worker in the single 20-year job is estimated to earn 11.2 percent (standard error = 1.0 percent) more *in each year* than the worker in the sequence of 10-year jobs.

The finding that workers in longer jobs earn more in every year on the job than workers in shorter jobs is verified by the results in column 4. These results are based on specifications that allow the effect of completed duration on earnings to differ by seniority level. Specifically, completed duration and its square are included in the regression along with interactions of these two variables with two dummy variables indicating seniority of 3 to 10 years and seniority greater than 10 years. The hypothesis that these additional four variables have zero coefficients cannot be rejected for the white-collar sample, but can be rejected for the blue-collar sample.[24] Closer examination of the estimated parameters reveals that, consistent with the results of the formal test, the four interaction terms have inconsequential coefficients in the white-collar sample. For blue-collar workers, the wage advantage of long jobs falls with seniority after starting at a much higher level than suggested by the results in column 3 of Table 4b. Thus, there is still a positive wage advantage to being in longer blue-collar jobs at all levels of seniority. It simply is not uniform in magnitude.

Overall, the results in this section strongly confirm our expectations. Rather small estimates of the return to seniority are found using either the IV or the augmented OLS approach relative to the standard cross-sectional OLS regression. The corrected estimate of the net return to seniority is 0.5 to 0.6 percent per year for white-collar workers and a statistically insignificant 0.2 to 0.3 percent per year for blue-collar workers. In addition, our results provide direct evidence that workers in longer jobs earn substantially more throughout the job than workers in shorter jobs.

V. Some Issues of Specification and Estimation

Probably the most significant potential problem with the analysis just described is that the key job duration variable had to be estimated for jobs whose end was not observed. This introduces three conceptually distinct sources of measurement error that could affect our estimates.

The first source of measurement error is that the expected value is used in place of the actual realization of completed job duration. This is not a problem so long as the correct parameters and the correct model have been used in computing expected job duration. In this case, the measurement error

[23]Recall that the theoretical discussion implied that the IV estimate of the net return to seniority is a consistent estimate of $\beta_1 - [\beta_2 - \alpha]$ while the augmented OLS approach yields an upward biased estimate. In fact, we find that the augmented OLS estimate of the net return to seniority is slightly, though not significantly, *smaller* than the IV estimate.

[24]The test statistic for the white-collar sample is 1.69 and that for the blue-collar sample is 8.48, both distributed as F with 4 and approximately 3500 degrees of freedom. The critical value of this distribution is 2.37 at the 5 percent level and 3.32 at the 1 percent level.

is exactly the deviation between expected and actual job duration. This is uncorrelated with expected job duration (the included regressor) by construction. Thus, there is no bias from this source in our estimated earnings function coefficients.

The second source of measurement error is that the parameters of the job duration model are only estimates, so that the predictions of expected job duration are themselves subject to error. However, it can be shown that in large samples, the estimation error in the parameters of the job duration model is of small enough order that coefficient estimates in equations which use the derived measure of duration as an explanatory variable are consistent.

The third source of measurement error is that the job duration model may be misspecified. If misspecification results in random errors in expected completed duration then classical measurement error is introduced. In the context of the augmented OLS estimates, this measurement error will tend to 1) bias the coefficient on completed duration toward zero, and 2) result in an estimated return to seniority that is higher than it would be in the absence of measurement error. The estimated return to seniority would, however, still be useful as an upper bound. If misspecification results in systematic measurement error, the coefficient on completed duration and the estimated return to seniority will be biased in an unknown way.

Given the finding of a large positive return to completed duration and the sharp reduction in the return to seniority using the augmented OLS estimates, it is unlikely that random measurement error is a serious problem. In any case, since the effects of random measurement error are predictable, our findings provide bounds on the "true" effects. We can do nothing about the potential of systematic measurement error except to note that our results are conditional on the Weibull specification of completed job duration.[25]

The appropriateness of the method used to impute completed durations for the censored jobs clearly merits careful investigation. One obvious question is whether the particular function of the explanatory variables and last-observed seniority that we use in creating our measure of completed duration is appropriate or whether it is contributing to the fit of the model simply because it incorporates the information on last-observed seniority. One way to investigate this is to reestimate the augmented OLS model including last-observed seniority and its square directly as regressors. If our measure of completed job duration has a significant effect on earnings even after controlling for last-observed seniority, then we have added support for our measure.

The first columns of Tables 5a and 5b contain estimates of earnings functions that include measures of completed job duration and its square, but not last-observed seniority (repeated here for comparison purposes). The second columns of these tables contain estimates of earnings functions that include last-observed seniority and its square, but no completed duration measure. When entered separately, both completed job duration and last-observed seniority contribute significantly to the fit of the model. Do these relationships hold up if we control for both simultaneously? The third column of Tables 5a and 5b contain estimates of earnings functions for the two occupational groups that include measures of both completed duration and last observed seniority. After controlling for completed duration, last-observed seniority is not a significant determinant of earnings. However, even after controlling for last observed seniority, completed duration adds significantly to the models' explanatory power.[26] Thus, our

[25]While the first two sources of measurement error do not induce inconsistency in the earnings function parameter estimates, they do affect the estimates of the standard errors of the coefficients. As discussed in fn. 22, appropriate standard errors that are corrected for the effects of these errors and for general heteroskedasticity are almost identical to the usual standard errors.

[26]This is verified by F-tests of the general specification in col. 3 against the restricted specifications in cols. 1 and 2. The test statistics for the hypothesis that

TABLE 5a—SELECTED COEFFICIENTS FROM ALTERNATIVE ln (AVERAGE HOURLY EARNINGS) MODELS
MANAGERIAL AND PROFESSIONAL NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	OLS (2)	OLS (3)	OLS (4)
Years of	18.14	.0288	.0272	.0280	.0288
Experience	[10.08]	(.0027)	(.00280)	(.0028)	(.0027)
(Years of	430.77	−.00048	−.00042	−.00046	−.00051
Experience)^2	[407.84]	(.00007)	(.00007)	(.00007)	(.00007)
Years of Current	8.88	.00548	−.00472	.00186	.00865
Seniority	[8.34]	(.00178)	(.00311)	(.00345)	(.00244)
E(Completed Job	20.83	.0198	–	.0167	.0220
Duration)	[12.18]	(.0024)		(.0041)	(.0025)
{E(Completed Job	631.55	−.00035	–	−.00030	−.00044
Duration)}^{2}	[505.56]	(.00006)		(.00009)	(.00007)
Years Last Observed	12.08	–	.0275	.00682	–
Seniority	[9.52]		(.0031)	(.00567)	
{Years Last Observed	236.8	–	−.00043	−.00007	–
Seniority}^{2}	[304.9]		(.00008)	(.00012)	
E(Job Duration)	12.15	–	–	–	.00067
×[=1 if Uncensored]	[34.45]				(.00592)
{E(Job Duration)}^{2} ×	133.41	–	–	–	−.00030
×[=1 if Uncensored]	[629.26]				(.00028)
R^2	–	.3871	.3841	.3874	.3885

[a]See Table 4a, fn. a. Sample size = 3493.

TABLE 5b—SELECTED COEFFICIENTS FROM ALTERNATIVE ln (AVERAGE HOURLY EARNINGS) MODELS
BLUE-COLLAR NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	OLS (2)	OLS (3)	OLS (4)
Years of	17.34	.0117	.0135	.0118	.0119
Experience	[11.14]	(.0026)	(.00256)	(.00259)	(.0027)
(Years of	424.70	−.00026	−.00028	−.00026	−.00027
Experience)^2	[470.81]	(.00006)	(.00006)	(.00006)	(.00006)
Years of Current	6.31	.00241	−.00304	.00375	.00412
Seniority	[7.46]	(.00213)	(.00322)	(.00352)	(.00307)
E(Completed Job	13.86	.0154	–	.0175	.0167
Duration)	[11.75]	(.0021)		(.0052)	(.0022)
{E(Completed Job	362.44	−.00014	–	−.00017	−.00020
Duration)}^{2}	[444.45]	(.00006)		(.00012)	(.00008)
Years Last Observed	8.80	–	.0255	−.00377	–
Seniority	[8.82]		(.0031)	(.00730)	
{Years Last Observed	155.2	–	−.00031	.00005	–
Seniority}^{2}	[263.8]		(.00008)	(.00017)	
E(Job Duration)	12.92	–	–	–	.00556
×[=1 if Uncensored]	[33.92]				(.00627)
{E(Job Duration)}^{2}	13.77	–	–	–	−.00053
×[=1 if Uncensored]	[60.21]				(.00030)
R^2	–	.4041	.4001	.4042	.4051

[a]See Table 4b, fn. a. Sample size = 3554.

measure of completed duration contains information on earnings well beyond the information contained directly in the variables that are used in its computation, including the last-observed seniority.

Perhaps the most important question concerning our completed job duration variable is whether the relationship between completed job duration and earnings differs between the observations that come from jobs where actual completed durations are observed and from jobs where completed durations are predicted. To answer this question, we reestimated the augmented OLS earnings functions with two additional variables: 1) the interaction between completed job duration and a dummy variable that equals one if the job duration is uncensored and equals zero otherwise; and 2) the interaction between the square of completed job duration and the same dummy variable. This allows completed duration and its square to have different effects where they are actually observed (uncensored jobs) and where they are predicted (censored jobs). These estimates are contained in column 4 of Tables 5a and 5b.

For white-collar workers, the hypothesis that the effects are the same (that the two new variables have zero coefficients) can be rejected at the 5 percent level, but not at the 1 percent level. However, the interaction terms have rather small coefficients relative to the coefficients on the basic variables. For blue-collar workers, the hypothesis that the effects are the same cannot be rejected at either the 5 or 1 percent level, and the point estimates of the interaction terms' coefficients are insubstantial.[27] In sum, observed completed duration in uncensored jobs and our estimate of completed duration in censored jobs have almost identical relationships with earnings.

The last potential issue we consider here is related to the fact that, because we use pooled time-series cross-section data to estimate our earnings functions, there are repeated observations on particular individuals. If the earnings function residuals are correlated across observations within individuals, our standard errors are likely to be understated. We are reluctant to present estimates of a fixed-effect model because all of the variation in the measure of completed duration in such a model comes from those workers who change jobs within the sample period. These job changes will be dominated by short jobs which are not representative of the sample of jobs as a whole.[28] A *very* conservative alternate approach to this problem is to reestimate the key specifications of the model on a single-year cross section. If the inferences from such a specification are similar to those from the pooled model, then we can have more confidence in the overall validity of the results. The estimation of a single-year cross section also has the advantage that it provides results directly comparable to much of the existing literature on earnings functions.

In order to carry out this analysis, 1973 was selected as a representative year, and the analyses of Tables 4a and 4b were repeated on 1973 cross sections of 244 white-collar workers and 240 blue-collar workers. The single-year estimates are contained in Tables 6a and 6b, and two things are clear from a comparison of these results with the results in Tables 4a and 4b. First, as we expected, the results are less precisely estimated due to the much smaller sample sizes. Second, the results are very similar in character those derived using the pooled samples. The "standard" OLS results in column 1 of Tables 6a and 6b show statistically significant net returns to seniority. The IV estimates of the return to seniority are much

last-observed seniority adds explanatory power to the model are 0.845 for white-collar workers and 0.295 for blue-collar workers. The test statistics for the hypothesis that completed job duration adds explanatory power are 9.30 for white-collar workers and 12.1 for blue-collar workers. All the test statistics are distributed as F with 2 and approximately 3500 degrees of freedom. The critical value of this distribution is 4.61 at the 1 percent level of significance.

[27] The test statistics are 3.95 for the white-collar sample and 2.95 for the blue-collar sample, both distributed as F with 2 and approximately 3500 degrees of freedom. The critical value of this distribution is 3.00 at the 5 percent level and 4.61 at the 1 percent level.

[28] Note that year effects are removed in all specifications through the use of a complete set of year dummies.

TABLE 6a—SELECTED COEFFICIENTS FROM 1973 CROSS SECTION ln (AVERAGE HOURLY EARNINGS) MODELS MANAGERIAL AND PROFESSIONAL NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	IV (2)	OLS (3)	OLS (4)
Years of	18.38	.0349	.0514	.0273	.0278
Experience	[10.12]	(.0104)	(.0199)	(.0106)	(.0126)
(Years of	440.06	−.00066	−.00111	−.00049	−.00050
Experience)^2	[402.58]	(.00025)	(.00049)	(.00025)	(.00029)
Years of Current	9.00	.0135	.00653	.00721	.0109
Seniority	[7.97]	(.0043)	(.00524)	(.00684)	(.0104)
E(Completed Job	20.99	–	–	.0230	.0432
Duration)	[13.00]			(.0090)	(.0181)
{E(Completed Job	642.94	–	–	−.00041	−.00095
Duration)}^{2}	[534.27]			(.00023)	(.00054)
E(Job Duration)	6.40	–	–	–	−.0180
$\times[=1$ if $3<$ Seniority $\leq 10]$	[10.92]				(.0159)
{E(Job Duration)}^{2}	174.22	–	–	–	.00044
$\times[=1$ if $3<$ Seniority $\leq 10]$	[341.65]				(.00051)
E(Job Duration)	11.19	–	–	–	−.0231
$\times[=1$ if Seniority $>10]$	[16.01]				(.0173)
{E(Job Duration)}^{2}	385.70	–	–	–	.00060
$\times[=1$ if Seniority $>10]$	[585.00]				(.00054)
R^2	–	.4318	.4105	.4566	.4640

[a]See Table 4a, fn. a. Sample size = 244.

TABLE 6b—SELECTED COEFFICIENTS FROM 1973 CROSS SECTION ln (AVERAGE HOURLY EARNINGS) MODELS BLUE COLLAR NONUNION SAMPLE[a]

	Mean [s.d.]	OLS (1)	IV (2)	OLS (3)	OLS (4)
Years of	18.47	.0340	.0133	.0205	.0188
Experience	[11.32]	(.0100)	(.0149)	(.0099)	(.0099)
(Years of	468.83	−.00075	−.00036	−.00046	−.00046
Experience)^2	[484.73]	(.00020)	(.00033)	(.00022)	(.00022)
Years of Current	6.50	.0102	−.00453	.00063	−.00781
Seniority	[7.60]	(.0038)	(.00646)	(.00840)	(.0122)
E(Completed Job	14.57	–	–	.0194	.0502
Duration)	[12.88]			(.0075)	(.0208)
{E(Completed Job	402.81	–	–	−.00028	−.00161
Duration)}^{2}	[491.00]			(.00024)	(.00088)
E(Job Duration)	4.86	–	–	–	.00108
$\times[=1$ if $3<$ Seniority $\leq 10]$	[8.95]				(.0199)
{E(Job Duration)}^{2}	114.07	–	–	–	.00031
$\times[=1$ if $3<$ Seniority $\leq 10]$	[237.2]				(.00086)
E(Job Duration)	7.05	–	–	–	−.0381
$\times[=1$ if Seniority $>10]$	[13.81]				(.0203)
{E(Job Duration)}^{2}	245.9	–	–	–	.00171
$\times[=1$ if Seniority $>10]$	[508.6]				(.00089)
R^2	–	.4958	.4048	.5182	.5344

[a]See Table 4b, fn. a. Sample size = 240.

smaller than the OLS estimates, and they are *not* significantly different from zero in either occupational group. The augmented OLS estimates of the return to seniority, contained in column 3 of Tables 6a and 6b, are quite similar to those derived with the IV estimator. In addition, the augmented OLS estimates imply that workers in longer jobs earn significantly more throughout the job than workers in shorter jobs. We conclude that the general findings of the previous section are not simply due to an exaggerated precision that comes from ignoring the error component structure in a pooled sample.

VI. Concluding Comments

The basis for considering implicit contracts under which compensation is deferred from early until late in workers' time with their employers to be an important feature of the labor market has been the simple cross-sectional evidence that long seniority workers have higher wages, even taking their total labor-market experience into account. The evidence presented in this study seriously undermines the empirical foundations of this sort of implicit contract. Contrary both to the conventional wisdom and to our own prior expectations, there seems to be only a small average return to seniority in excess of the average return to general labor market experience. For the nonunion professional, technical, and managerial sample, the corrected estimates of the seniority coefficient suggest that the true return to seniority is approximately half a percent per year, rather than the approximately 1 percent per year suggested by the standard cross-section model. For the nonunion blue-collar sample, the corrected estimates yield a statistically insignificant return to seniority of approximately one-quarter of 1 percent per year, rather than the statistically significant 1.5 percent per year suggested by the standard model.[29]

This evidence does not imply that implicit contracts entailing the posting of a bond by workers through a deferral of compensation are never important. Indeed, they could be very important for some subgroups of workers and even a small return to seniority could translate into a substantial cumulative contribution to annual earnings over a period of time. It is also possible that parts of the total compensation package other than earnings, such as fringe benefits or other perquisites, are structured so as to reward longevity with a particular employer.[30] However, earnings deferral under implicit contracts appears to be a much less important factor in both white- and blue-collar labor markets than has generally been believed.

The other result of our study that is potentially very important for the understanding of the nature of the long-term employment relationship is that workers in long jobs earn substantially more throughout their jobs than do workers in short jobs. Whether this is due to individual differences, inter-job differences, or match-specific differences, this finding has important implications for the decisions of workers and employers as they affect investment in match-specific capital and the dynamics of the employment relationship. Note that our finding of a correlation between completed duration and earnings does *not* imply that it is the length of the job that *causes* earnings. It is likely that, at least to some extent, the higher earnings throughout the job provide an incentive for workers to remain on their job. Viewed in this light, our results are consistent with the currently popular efficiency wage models.

The finding of a strong positive relationship between job duration and earnings in both of the occupational subsamples provides strong evidence against another view of long-term employment relationships based on models of incomplete information where firms offer workers insurance regarding their unknown abilities. In particular, Milton Harris and Bengt Holmstrom (1982) argue

[29] Our findings regarding the returns to seniority are consistent with those obtained by Altonji and Shakotko.

[30] Freeman and Medoff provide evidence that the value of nonwage benefits such as vacations and pension plans rise with seniority. There may also be less tangible advantages that accrue with seniority.

that the positive return to experience found in most data sets could reflect such insurance contracts. A simple version of this model considers the case where a worker and all firms are initially uncertain about the worker's productivity, and where that uncertainty is reduced over time as the worker and all firms learn about the worker's productivity from the worker's stream of output. The optimal contract for the firm to offer the worker specifies an initial wage equal to the expected value of the worker's productivity minus an insurance premium, and the employer guarantees not to reduce the initial wage if the worker is revealed to be relatively unproductive. Since all new information about productivity is common knowledge, workers revealed to be relatively productive receive wage increases either from their original employer or by taking a new job with another employer. Workers revealed to have low productivity cannot duplicate their original wage elsewhere and for that reason are more likely to stay with their original employer. Thus, this simple insurance model predicts a *negative* correlation between job duration and earnings. Our results indicate that this correlation is strongly positive.

The recognition that the direction of causality between earnings and job duration is ambiguous highlights the difficulty of using our results to make definitive statements on the existence of particular types of labor contracts. Observed job durations are generated as the result of mobility decisions that are poorly understood. While we conclude tentatively that the evidence from earnings functions is not consistent with simple earnings deferral models of incentive contracts or with the simple model of insurance contracts, it is clear that further analysis of the joint determination of job duration and earnings is necessary for a full understanding of long-run employment relationships.

REFERENCES

Abraham, Katharine G. and Medoff, James L., "Length of Service and the Operation of Internal Labor Markets," *Proceedings of the Thirty-Fifth Annual Meeting of the Industrial Relations Research Association*, December 1982, 308–18.

Akerlof, George A. and Katz, Lawrence F., "Do Deferred Wages Dominate Involuntary Unemployment as a Worker Discipline Device?," NBER Working Paper No. 2025, September 1986.

Altonji, Joseph and Shakotko, Robert, "Do Wages Rise with Job Seniority?," *Review of Economic Studies*, forthcoming 1987.

Becker, Gary S., *Human Capital*, NBER, Ann Arbor: University Microfilms, 1964.

______ **and Stigler, George J.,** "Law Enforcement, Malfeasance and Compensation of Enforcers," *Journal of Legal Studies*, January 1974, *3*, 1–18.

Berndt, Ernst K. et al., "Estimation and Inference in Nonlinear Structural Models," *Annals of Economic and Social Measurement*, 1974, *3/4*, 653–65.

Bulow, Jeremy and Summers, Lawrence, "A Theory of Dual Labor Markets with Application to Industrial Policy, Discrimination, and Keynsian Unemployment," *Journal of Labor Economics*, July 1986, *4*, 376–415.

Burdett, Kenneth, "A Theory of Employee Job Search and Quit Rates," *American Economic Review*, March 1978, *68*, 212–20.

Freeman, Richard B. and Medoff, James L., *What Do Unions Do*?, New York: Basic Books, 1984.

Harris, Milton and Holmstrom, Bengt, "Ability, Performance and Wage Differentials," *Review of Economic Studies*, July 1982, *49*, 315–33.

Heckman, James and Singer, Burton, "A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data," *Econometrica*, March 1984, *52*, 271–320.

Jovanovic, Boyan, "Job Matching and the Theory of Turnover," *Journal of Political Economy*, October 1979, *87*, 972–90.

Katz, Lawrence F., "Efficiency Wage Theories: A Partial Evaluation," in Stanley Fischer, ed., *NBER Macroeconomics Annual 1986*, Cambridge: MIT Press, 235–76.

Lancaster, Tony, "Econometric Methods for the Duration of Unemployment," *Econometrica*, July 1979, *47*, 939–56.

Lazear, Edward, "Why Is There Mandatory Retirement?," *Journal of Political Econ-*

omy, December 1979, *87*, 1261–84.

Medoff, James L. and Abraham, Katharine G., "Experience, Performance and Earnings," *Quarterly Journal of Economics*, December 1980, *95*, 703–36.

______ **and** ______, "Are Those Paid More Really More Productive: The Case of Experience," *Journal of Human Resources*, Spring 1981, *16*, 186–216.

Mincer, Jacob, *Schooling, Experience and Earnings*, NBER *Studies in Human Behavior and Social Institutions*, No. 2, New York: Columbia University Press, 1974.

______ **and Jovanovic, Boyan,** "Labor Mobility and Wages," in Sherwin Rosen, ed., *Studies in Labor Markets*, NBER Universities-National Bureau Conference Series, No. 31, Chicago: University of Chicago Press, 1981, 21–64.

Mortensen, Dale T., "Specific Capital and Labor Turnover," *Bell Journal of Economics*, Autumn 1978, *9*, 572–86.

Newey, Whitney K., "A Method of Moments Interpretation of Sequential Estimators," *Economics Letters*, 1984, *14*, 201–06.

Parsons, Donald O., "Models of Labor Market Turnover: A Theoretical and Empirical Survey," in Ronald G. Ehrenberg, ed., *Research in Labor Economics*, Vol. 1, Greenwich: JAI Press, 1977, 185–223.

Shapiro, Carl and Stiglitz, Joseph, "Involuntary Unemployment as a Worker Discipline Device," *American Economic Review*, June 1984, *74*, 433–44.

Stiglitz, Joseph, "Theories of Wage Rigidity," mimeo., Princeton University, 1984.

Topel, Robert, "Job Mobility, Search, and Earnings Growth: A Reinterpretation of Human Capital Earnings Functions," in Ronald G. Ehrenberg, ed,. *Research in Labor Economic*, Vol. 8, 1986, 199–223.

Viscusi, W. Kip, "Self-Selection, Learning-Induced Quits and the Optimal Wage Structure," *International Economic Review*, October 1980, *21*, 529–46.

White, Halbert, "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," *Econometrica*, May 1980, *48*, 817–38.

Yellen, Janet, "Efficiency Wage Models of Unemployment," *American Economic Review Proceedings*, May 1984, *74*, 200–08.