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# JOB DURATION, SENIORITY, AND EARNINGS

Katharine G. Abraham

Henry S. Farber

Number 407

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massachusetts
institute of
technology

50 memorial drive
cambridge, mass. 02139

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ABSTRACT

Job Duration, Seniority, and Earnings

Katharine G. Abraham

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Massachusetts Institute of Technology

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The stylized fact that seniority and earnings in a cross-section are positively related, even after controlling for total labor market experience, has served as the basis for theoretical analyses of implicit labor contracts suggesting that workers post bonds in the form of deferred compensation in order to ensure their continued performance at an adequate level. An alternative interpretation of the cross-section evidence is that good workers or workers in good jobs or good matches both earn more throughout the job and have longer job durations. Another stylized fact, that labor market experience and earnings in a cross section are positively related, has been taken as evidence of the importance of general human capital accumulation. An alternative interpretation of this evidence is that workers with more experience have had more time to find good jobs and/or good matches, resulting in higher earnings.

A measure of the completed duration of jobs is developed, based on a Weibull model of job duration using a longitudinal sample from the Michigan Panel Study of Income Dynamics, in order to distinguish between the competing hypotheses regarding both seniority and experience. Earnings functions using this measure of job duration as a regressor are estimated. These yield three main results. First, workers in longer jobs earn significantly more in every year of the job than do workers in shorter jobs. Second, controlling for completed job duration eliminates most of the apparent return to seniority found in standard cross-section models. Thus, it appears that implicit contracts that provide for workers posting bonds through deferred wage payments are less important than has been believed. Third, for blue collar workers there is evidence that a part of the small observed (cross-sectional) return to labor market experience is due to sorting of workers into better jobs over time. There is no evidence of sorting for white collar workers.

"It is better to know nothing than to know what ain't so,"
- Josh Billings, Proverb [1874]

## I. Introduction and Theoretical Background

An important stylized fact of the labor market is that workers with longer seniority have higher earnings even after controlling for total labor market experience. The empirical support for this stylized fact is that standard earnings functions estimated using cross-section data produce significant positive seniority coefficients even when a measure of total labor market experience is also included in the regression.

The standard explanations for this positive correlation posit 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] In effect, the worker is posting a bond to ensure continued employment. Another possible motivation for such a deferral arrangement exists where effort is important. Having the worker post a bond through a deferral of compensation 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

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1. See Becker (1964) and Mincer (1974) for discussions of investment in firm specific training. Mincer and Jovanovic (1981) present an an analysis of the relationships among seniority, mobility, and earnings that relies on investment in specific human capital.

would lose the bond that he or she had posted under a deferred wage contract.[2]

The firm-specific human capital explanation and 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 that there will also be a positive return to seniority even after controlling for total labor market experience. In other words, the theories imply that seniority produces earnings growth in excess of the earnings growth associated with labor market experience.

As stated above, the empirical support for these views of the labor market rests entirely on cross-sectional earnings function estimates. However, the positive cross-sectional association between seniority and earnings does not in fact imply that earnings rise with seniority. An alternative interpretation of the cross-sectional evidence is that workers who have been on their jobs for a long time started out earning more. In this case, despite the fact that earnings are fixed over the course of a job, a standard earnings function regression will yield a positive seniority coefficient.

It is in fact reasonable to suspect that the cross-sectional evidence is contaminated in this way. Some workers may be both more stable and more productive than others.[4] Some employers may choose to pay higher wages than

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2. See Viscusi (1980), Becker and Stigler (1974) and Lazear (1979) for models in which wage deferral provides this sort of incentive for workers.

3. 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.

4. Indeed, to the extent that there are turnover and training costs, stability per se can raise an employee's value to the firm.

others, so that workers are unlikely to leave them.[5] 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 match is less likely to be broken off.[6] It is quite plausible, then, that high quality workers, jobs, and matches will be associated with both higher wages and longer job duration. To the extent that worker, job, and/or match quality are unmeasured, they represent omitted variables in a cross-section earnings regression. Since those spells in any cross-section sample of spells in progress that have long durations to date are more likely to be spells that will ultimately last a long time (so that seniority will be positively correlated with completed job duration) and completed job duration is positively correlated with any omitted worker, job, and/or match quality measures, seniority in a cross-section will also be positively correlated with these omitted quality measures. The result is an upward bias in the estimated seniority coefficient in cross-section earnings functions.

A primary goal of the empirical analysis in this study is to distinguish between the competing hypotheses regarding the relationships among job seniority, completed job duration, and earnings that imply a positive cross-sectional relationship between seniority and earnings. The results of this analysis shed light on the various theoretical models that have been adduced to explain the pattern of earnings over the worklife.

A second and related stylized fact of the labor market is that in a

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5. 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. Shapiro and Stiglitz (1984) and Bulow and Summers (1985) present formal efficiency wage models; Yellen (1984) and Stiglitz (1984) survey the literature.

6. Jovanovic (1979) presents a theoretical analysis of the effects of heterogeneous match quality.

cross-section of individuals, workers with more total years of labor market experience have higher earnings. The standard explanation for this positive association is that general labor market experience raises workers' on-the-job productivity, which translates into higher earnings (Becker, 1964; Mincer, 1974; Brown, 1983). An alternative view of the positive association between experience and earnings stresses the fact that those with more experience have had more time to find a good job and/or a good match.[7] If sorting into better jobs and/or better matches over time is an empirically important phenomenon, it could lead to a strong positive association between experience and earnings even in the absence of any experience-related growth in productivity.

Note that if sorting over time of workers into better jobs is an important phenomenon, then there ought to be a positive correlation between labor market experience at the start of a job and completed duration of the job. In order to make this clear, consider a simple model where workers sample at set intervals from a stable wage offer distribution. Workers change jobs when the wage associated with the latest draw from the wage offer distribution exceeds their current wage. It is straightforward to show that, as time goes by and workers are in successively higher paying jobs, the probability of receiving an offer that dominates the best offer to date will fall. The result will be a positive relationship between initial experience and job duration.

On the basis of these arguments, another goal of the empirical analysis is to distinguish between the competing hypotheses regarding the relationships among experience, job duration, and earnings that imply a positive cross-sectional relationship between experience and earnings. This analysis will

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7. This argument was pointed out to us by Robert Topel. See Burdett (1978) and Jovanovic (1979).

shed light on the extent to which there is support for the general human capital accumulation theory as opposed to the job sorting theory.[8]

An intuitively appealing approach to dealing with bias in the estimated return to seniority in the cross-section is to control for the completed length of the job in the earnings equation.[9] After all, by the argument given above, the seniority coefficient is biased only because seniority is correlated with job duration, which in turn is correlated with omitted worker, job and/or match quality. Section II develops a simple stochastic model of earnings determination that provides a basis for this approach. Moreover, if completed job duration is a proxy -- even an imperfect proxy -- for job and/or match quality and more experienced workers do sort themselves into better jobs and/or better matches, one might expect that controlling for completed job duration in the earnings equation should remove at least part of the positive search component in the estimated return to labor market experience. The model contained in section II provides a more formal basis for this intuition as well.

In section III we construct a completed job duration measure for use in cross-section earnings equations, using a pooled time-series cross-section sample of nonunion males from the Michigan Panel Study of Income Dynamics (PSID). The analysis requires a panel data set, since information on how long people end up staying on their jobs is needed. While not all of the jobs we observe end during the period over which the panel is followed, there is some information available on ultimate durations of all jobs. Even for jobs still in progress at the end of the sample period, we know that completed job

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8. Our results also have implications for a third model of the return to experience presented in a recent paper, Harris and Holmstrom (1982).

9. Altonji and Shakotko (1985) address the issue of bias in cross-section estimates of the return to seniority empirically, using an instrumental variables approach.

duration is at least as long as the last observed tenure on the job. This information is used to predict the completed duration of these matches in the context of a parametric model of job duration estimated separately for nonunion professional, technical and managerial employees and for nonunion blue collar employees; results for a nonunion sample spanning all occupations are presented for comparison purposes. One interesting finding emerging from this piece of the analysis is that additional prior experience increases the expected length of a new blue collar job, but not the expected length of a new professional, technical or managerial job.

In section IV we augment standard earnings functions with our job duration measure. Because there is reason to think that the wage determination process operates differently for workers in different labor market segments, these earnings equations are estimated separately for nonunion professional, technical and managerial employees and for nonunion blue collar workers. Results for a nonunion sample spanning all occupations are again presented primarily for comparison purposes. Three issues are addressed with these estimates. The first issue is the extent to which earnings are higher in jobs that end up lasting longer. Our results on this issue are quite straightforward: workers who are in long jobs earn higher pay in every year on the job. The second issue is the relative importance of the "direct" effects of seniority versus omitted person, job and/or match quality in producing the observed cross-sectional positive correlation between seniority and earnings. We find that most of the apparent cross-section return to seniority reflects omitted variable bias. The final issue is the extent to which the observed positive correlation between experience and earnings reflects the sorting of more experienced workers into better jobs and/or better matches. Our results suggest that more experienced blue collar workers earn more in part because they are in better jobs and/or better matches. There is no evidence of such sorting for professional, technical, and managerial workers.

Section V contains an overview of the results and offers a few concluding observations.

## II. Completed Job Duration in the Earnings Function

We turn now to a simple model of the relationships among current job tenure, pre-job experience, individual-, job- and match-specific earnings components, and completed job duration. This model illustrates two major points. First, under the assumptions of the model, the estimated return to tenure in standard cross-section earnings equations is biased upwards and the estimated return to experience reflects both returns to experience *per se* and returns to search. Second, again under the assumptions of the model, controlling for completed job duration eliminates the bias in the estimated return to seniority and removes part of the search component in the estimated return to experience.[10]

Suppose that the earnings of a particular worker on a particular job at a point in time can be written:

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

where

$W$ = hourly earnings,

$S$ = current seniority (tenure),

$EXP$ = pre-job experience,

$\mu$ = a person/job specific error term representing the excess of earnings enjoyed by this person on this job over and

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10. A more detailed derivation of the results reported in this section of the paper can be found in Appendix A.

above the earnings that could be expected by a randomly selected person/job combination,

$\eta$ = is a person/job/time-period specific error term,

i = index of individuals,

j = index of jobs,

t = index of years within jobs, and

$\beta$'s = parameters of interest.

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, $\mu_{ij}$ captures the net influence of three unobservables on hourly earnings: unobserved person quality, unobserved job quality and unobserved match quality.[11] The error $\mu_{ij}$ is assumed to be fixed over the course of a job and may be correlated with S and EXP. The error $\eta_{ijt}$ is assumed to be orthogonal to S, EXP and $\mu$.

In practice, earnings functions are generally estimated using cross-section data and $\mu_{ij}$ is not observable. The earnings function contained in equation (1) as ordinarily implemented in a cross-section can be rewritten as

$$(2) \qquad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \nu_{ijt}.$$

The error term ($\nu_{ijt}$) is defined in terms of the underlying stochastic structure as

$$(3) \qquad \nu_{ijt} = \mu_{ij} + \eta_{ijt}.$$

Thus, the person/job specific component of the error ($\mu_{ij}$) is omitted from the equation. Note that there is only a single cross-sectional observation on any individual i or job j in a cross-section.

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11. A brief discussion of a more general model that allows for separate person and job/match components is presented below in the text; a more detailed discussion can be found in Appendix A.

How does omitting $\mu_{ij}$ from equation (2) affect the estimated values of $\beta_1$ and $\beta_2$? This depends on the correlations between $\mu_{ij}$ and the variables $S_{ijt}$ and $EXP_{ij}$. As was argued above, good workers or workers in good jobs or good matches are likely both to stay on the job longer and to earn more money. In a cross-section of workers, those with high current seniority are more likely to be in long jobs, so that observed seniority and completed job duration will be positively correlated; there will thus be a positive correlation between observed seniority and $\mu_{ij}$. One would therefore expect $\hat{\beta}_1$ estimated from equation (2) to be an upward biased estimate of $\beta_1$. In addition, because those workers with more pre-job experience have had more chances to find a good job or good match and because good jobs or good matches are likely to pay higher wages, there will be a positive correlation in a cross-section between $\mu_{ij}$ and $EXP_{ij}$. This means that $\hat{\beta}_2$ estimated from equation (2) is likely to be an upward biased estimate of $\beta_2$, the returns to experience <u>per se</u>.

These potential biases can be stated more precisely. Suppose that the completed duration of jobs is positively correlated with $\mu_{ij}$ so that the following relationship holds:

(4) $D_{ij} = \gamma\mu_{ij} + \varepsilon_{ij}$,

where $D_{ij}$ is the completed length of the current job, $\mu_{ij}$ is as defined above, $\gamma$ is a parameter which summarizes the relationship between D and $\mu$, and $\varepsilon_{ij}$ captures the variation in completed job duration that cannot be linked to the earnings advantage associated with worker, job and/or match quality. Suppose further that workers with more pre-job experience are likely to have found better jobs and/or better matches. One simple approximation to such a relationship is

(5) $\mu_{ij} = \alpha EXP_{ij} + \phi_{ij}$,

where $\mu$ and EXP are as defined above, $\alpha$ is a parameter that summarizes the relationship between $\mu$ and EXP and $\phi_{ij}$ is the variation in $\mu$ not systematically related to EXP.[12] We assume that pre-job experience is not correlated with completed job duration except insofar as more experienced workers are more likely to have found a good job and/or a good match, so that $E(EXP_{ij}*\varepsilon)$ is zero.[13] 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. More formally,

$$(6) \quad E(S_{ijt}) = 1/2\ D_{ij}$$

and

$$(7) \quad S_{ijt} = 1/2\ D_{ij} + \xi_{ijt},$$

where $\xi_{ijt}$ is a random variable with zero mean. The distribution of $\xi_{ijt}$ will vary depending upon the completed length of the job. However, its mean is always zero, as is $cov(\mu_{ij},\xi_{ijt})$, $cov(D_{ij},\xi_{ijt})$, and $cov(EXP_{ij},\xi_{ijt})$. Applying the standard bias formula to the seniority coefficient in equation (2) estimated from cross-sectional data and simplifying yields

$$(8) \quad E(\hat{\beta}_1)-\beta_1=(1/2)*(1/DET_1)*\gamma*var(EXP_{ij})*var(\phi_{ij})$$

which is positive provided $\gamma$ (the coefficient summarizing the relationship between D and $\mu$) is positive.

The expected value of the pre-job experience coefficient in equation (2) estimated from cross-sectional data exceeds $\beta_2$, the return to experience _per se_. The difference between the two is

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12. It should be emphasized that equations (4) and (5) are not structural equations but simply summarize the relationships expected in cross-section data.

13. The consequences of relaxing this assumption are discussed briefly at

(9) $E(\hat{\beta}_2)-\beta_2=(1/DET_1)*$

$$[\alpha*var(EXP_{ij})*(var(S_{ijt}) - 1/4*var(D_{ij}) + 1/4*var(\varepsilon_{ij}))],$$

which is also positive so long as $\alpha$ (the coefficient summarizing the relationship between $EXP_{ij}$ and $\mu_{ij}$) is positive.

An intuitively appealing approach to removing the just-noted upward biases is to control explicitly for the completed length of jobs.• After all, the tenure coefficient in equation (2) is biased only because $S_{ij}$ is associated with $\mu_{ij}$ indirectly through $D_{ij}$. Moreover, equation (4) suggests that $D_{ij}$ may be useful more generally as a proxy for $\mu_{ij}$, so that introducing $D_{ij}$ into the earnings equation might remove at least part of the search component in the experience coefficient estimated from a cross-section. Augmenting equation (2) by adding $D_{ij}$ as an explanatory variable in a cross-section yields the following specification:

(10) $\ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \beta_3 D_{ij} + \omega_{ijt}$,

where $\omega$ is the estimating equation error. Using the relationships in equations (3) and (4), this error is

(11) $\omega_{ijt} = -(1/\gamma)*\varepsilon_{ij} + \eta_{ijt}$ .

How does introducing $D_{ij}$ into the earnings equation affect the estimated tenure and experience coefficients? The bias in the tenure coefficient in the augmented equation (10) model estimated using a cross-section can be shown to equal zero. Intuitively, even though completed job duration ($D_{ij}$) is an imperfect proxy for the omitted job/person/match quality variable ($\mu_{ij}$) in equation (2), seniority ($S_{ijt}$) is correlated with $\mu_{ij}$ only through its correlation with $D_{ij}$, so that controlling for $D_{ij}$ eliminates the bias in the estimated tenure coefficient.

The difference between the experience coefficient in the augmented

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the end of this section and in greater detail in Appendix A.

equation (10) model estimated using a cross-section and $\beta_2$ equals

$$(12)\quad E(\hat{\beta}_2)-\beta_2=(1/DET_2)*\alpha*var(EXP_{ij})*var(\varepsilon_{ij})*(var(S_{ijt})-1/4*var(D_{ij})).$$

This is still positive so long as $\alpha$ is positive, though comparison of equation (12) against equation (9) reveals that introducing $D_{ij}$ into the earnings equation reduces the difference. Note that the expected value of the search component in $\beta_2$ falls to zero as $var(\varepsilon_{ij})$ falls to zero, that is, as $D_{ij}$ becomes a better proxy for $\mu_{ij}$.

There are at least two obvious ways in which the model presented here might be generalized. First, one might believe that more experienced (older) workers are simply more stable than less experienced (younger) workers and allow experience to have a direct positive relationship with completed job duration that is not reflected in higher earnings once experience is controlled for. In more formal terms, the expected value of $EXP_{ij}*\varepsilon_{ij}$ could be positive. This generalization in no way affects our conclusions concerning the seniority coefficient; the standard cross-section estimate of $\hat{\beta}_1$ is still biased upward, and this bias is eliminated by including a measure of completed job duration in the regression. The condition

$$(13)\quad cov(D_{ij},\varepsilon_{ij})/var(D_{ij}) > cov(EXP_{ij},\varepsilon_{ij})/cov(EXP_{ij},D_{ij}),$$

is sufficient for the experience coefficient in the standard cross-section earnings equation to be an upward biased estimate of the returns to experience *per se* and for introducing completed job duration into the earnings equation to reduce but not eliminate this bias.

A second generalization of the base model is to relax the assumption that individual-, job-, and match-specific earnings components all have the same incremental association with job duration. In a more general specification in which earnings increments associated with unobserved individual characteristics have a different relationship to job duration than those attributable to job/match characteristics, equation (4) would be rewritten as

$$(14)\quad D_{ij} = \gamma_1\mu_i + \gamma_2\theta_{ij} + \varepsilon_{ij},$$

where $\mu_i$ is the earnings increment attributable to unobserved individual differences, $\theta_{ij}$ is the earnings increment attributable to unobserved job/match characteristics, $\varepsilon_{ij}$ captures the variation in completed job duration that cannot be linked to the earnings advantage associated wiith either worker or job/match heterogeneity, i index individuals, and j indexes jobs. The parameters $\gamma_1$ and $\gamma_2$ represent the relationships of individual and job/match characteristics respectively with job duration, and they are both assumed to be positive. Experience would affect job/match specific attributes but not (fixed) individual attributes, so equation (5) would be rewritten in terms of $\theta_{ij}$. Once again, generalizing the model in this way does not affect our conclusions concerning bias in the seniority coefficient; the standard cross-section estimate of $\hat{\beta}_1$ is still biased upward, and this bias is eliminated by including a measure of completed job duration in the regression. With regard to the estimated effect of experience, as long as the condition

$$(15)\ \mathrm{var}(\varepsilon_{ij}) + \gamma_1(\gamma_1 - \gamma_2)*\mathrm{var}(\mu_i) \geq 0$$

is satisfied, the experience coefficient in the standard cross-section earnings function is an upward biased estimate of $\beta_2$ and introducing completed job duration into the model reduces the upward bias.

In sum, our theoretical analysis suggests that, under a reasonable set of assumptions, controlling for the completed duration of the job in a standard cross-section earnings equation can: 1) eliminate the upward bias in the tenure coefficient associated with the failure to consider individual, job and/or match specific earnings components and 2) remove at least part of the search component in the experience coefficient arising from the same source. On this basis, the approach we adopt in the empirical analysis is to develop a measure of completed job duration using longitudinal data and use this measure in estimating a standard earnings function.

As a final note, it should be stated that we put no structural interpretation on the estimated completed job duration coefficient in earnings functions like equation (10). The relationship between completed job duration and earnings is certainly simultaneous. There is a large literature that estimates turnover probabilities (essentially an inverse measure of job duration) as a function of wage rates.[14] Another way of thinking about the approach taken here is that the earnings function in equation (1) is the reduced form for one equation in a two equation earnings/duration model. It is precisely this reduced form that is of interest here, but the coefficient estimates are biased by the omitted quality measure ($\mu$). Thus, we include job duration in the earnings function precisely because it is correlated with the person, job, and/or match quality component of the error that most would argue is the source of any "simultaneity bias".[15] Despite the ambiguity regarding causation, job duration provides exactly the information required of it in this study.

## III. 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. Clearly, a data set suitable for this task must follow individual workers over time so that one can observe how long the jobs they hold ultimately last. The data set should also have information on the individual workers' characteristics and their

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14. Parsons (1977) and Freeman and Medoff (1984) present surveys of some of this literature. Mincer and Jovanovic (1981) present a joint analysis of earnings and turnover.

15. Measurement of these unobservable factors is likely to be what is required to identify the underlying structural relationships between earnings and job duration.

wages/earnings. The Michigan Panel Study of Income Dynamics (PSID) satisfies these requirements and is used in the empirical analysis. Given the completed duration of every job in the sample, the analysis would proceed by including this measure in a "standard" earnings function as a control. 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 the expected completed job duration conditional on the job lasting at least as long as the last observed seniority level. In the estimation of the earnings function, 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 to 60 who participated in the Michigan Panel Study of Income Dynamics (PSID).[16] We used only observations from the random national sample portion of the PSID (the so-called Survey Research Center or SRC

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16. Unfortunately, the design of the PSID precludes meaningful examination of females for the purposes of this study. This is because complete information is available only for household heads, and, where households contain both male and female adults, the male is assumed to be the head by default.

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 where the individual reported that he was unionized.[17] 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, one comprised of nonunion professional, technical and managerial employees and the second comprised of nonunion blue collar employees; results for a nonunion sample covering all occupations are included for comparison purposes. 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 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.[18]

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17. Over the course of the PSID different definitions of unionization were used. In some years unionization refers to union membership, and in other years unionization refers to working on a job covered by a collective bargaining agreement. Where both measures were available, collective bargaining coverage was used.

18. Joseph Altonji kindly provided us with an extract containing these 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 (1985). 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 (SEO). We also smoothed the tenure variable in two ways in instances where a given individual was assigned the midpoint value of a tenure interval. First, 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. Second, if all observations for the individual on a given job were in the same tenure interval and thus would have been counted as having the same tenure in all years on the job, 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.

There are 1048 jobs held by 748 individuals represented in the nonunion professional, technical and managerial sample, 1623 jobs held by 958 individuals represented in the nonunion blue collar sample, and 2857 jobs held by 1581 individuals represented in the full nonunion sample. Our concern at this point is with ascertaining how long each of these jobs ultimately lasted.[19]

Various characteristics of the jobs in each of the three samples are reported in Table 1. Variables that can change over time in an unpredictable fashion (e.g., 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 professional/technical/managerial 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.[20] There were 91 cases in which an individual reported moving from blue collar status to professional, technical or managerial status and 87 cases in which an individual reported moving from professional, technical or managerial status to blue collar status while a job

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19. The two occupational subgroups for which separate models are estimated do not cover all nonunion jobs. Approximately 15 percent of the nonunion jobs 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. We chose to exclude clerical and sales jobs rather than to include them with professional, technical and managerial jobs in a broader white collar grouping.

20. For example, if an individual 1) was observed on a single job for ten years running; 2) reported being a blue collar worker for the first five years and a professional employee for the next five years; and 3) reported having 13 years seniority in the last observed year, then the job would appear in both the blue collar and the professional/technical/managerial subsamples with a last observed seniority of 13 years in both. The job would appear once in the overall nonunion jobs sample, and it would be classified on the basis of the first observed occupation (blue collar).

Table 1:
Selected Characteristics of Jobs Samples for Occupational Subgroups[a]

| | Prof,Tech, Managerial, Nonunion | | Blue Collar Nonunion | | All Nonunion | |
|---|---|---|---|---|---|---|
| | Complete | Censored | Complete | Censored | Complete | Censored |
| Proportion with years of tenure at last date job observed in range: | | | | | | |
| $T \leq 1$ | .486 | .146 | .626 | .278 | .602 | .242 |
| $1 < T \leq 3$ | .248 | .222 | .218 | .204 | .221 | .221 |
| $3 < T \leq 10$ | .203 | .298 | .123 | .291 | .141 | .283 |
| $T > 10$ | .0620 | .335 | .0334 | .227 | .0365 | .254 |
| Mean [standard deviation] of: | | | | | | |
| Years of tenure at last date job observed | 2.97 [4.07] | 9.19 [8.93] | 1.99 [3.31] | 6.79 [7.95] | 2.15 [3.38] | 7.35 [8.26] |
| Years of pre-job experience | 10.04 [7.82] | 9.50 [8.27] | 9.85 [8.94] | 11.85 [9.92] | 9.84 [8.50] | 11.0 [9.44] |
| (Years pre-job experience)$^2$ | 161.9 [235.3] | 158.8 [267.4] | 176.9 [322.5] | 239.0 [366.0] | 169.1 [291.1] | 210.9 [334.8] |
| Years of education | 14.5 [2.08] | 14.7 [2.10] | 11.4 [2.31] | 11.3 [2.49] | 12.5 [2.65] | 12.9 [2.89] |
| Proportion: | | | | | | |
| Nonwhite | .0521 | .0310 | .136 | .134 | .107 | .0856 |
| Married | .856 | .871 | .828 | .880 | .833 | .869 |
| Disabled | .0571 | .0543 | .0751 | .115 | .0759 | .0938 |
| Prof,tech | .481 | .547 | -- | -- | .134 | .214 |
| Managerial | .519 | .453 | -- | -- | .123 | .148 |
| Foreman, craft | -- | -- | .415 | .455 | .239 | .222 |
| Oper,labor | -- | -- | .585 | .545 | .342 | .277 |
| Number of observations | 403 | 645 | 839 | 784 | 1396 | 1461 |

[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. The industry and region characteristics of the samples are summarized in Table 1 of Appendix B.

was in progress.[21]

In all three samples, we observe the end of a substantial fraction of the jobs represented in the sample. In the professional, technical and managerial sample, we observe the actual completed duration for 403 of 1048 jobs; in the blue collar sample, for 839 of 1623 jobs; and in the full nonunion sample, for 1396 of 2857 jobs. Not surprisingly, a large proportion of the completed jobs are relatively short: 73 percent of the completed jobs in the professional, technical and managerial sample, 84 percent of the completed jobs in the blue collar sample and 82 percent of the completed jobs in the full nonunion sample lasted no more than three years. However, in all three samples, there are a sizeable number of completed jobs lasting 3 to 10 years and over 10 years. Longer jobs are more common among the still-in-progress jobs: 34 percent of the incomplete jobs in the professional, technical and managerial sample, 23 percent of the incomplete jobs in the blue collar sample and 25 percent of the incomplete jobs in the full nonunion sample had lasted more than 10 years as of the last date they were observed.

With regard to other characteristics, the completed jobs and the incomplete jobs generally look similar. In the blue collar sample, completed jobs tend to have occured slightly earlier in the worklife (9.9 years of pre-job experience for completed jobs versus 11.9 years of pre-job experience for incomplete jobs); the same is not true for the professional, technical and managerial sample. In the professional, technical and managerial sample, professionals account for a smaller share of the completed jobs than of the incomplete jobs (48.1 percent versus 54.7 percent); in the blue collar sample, operatives and laborers account for a larger fraction of the completed jobs

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21. 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.

than of the incomplete jobs ( 58.5 percent versus 54.5 percent). In the full nonunion sample, completed jobs appear to have occurred earlier in the worklife; professionals are under-represented and operatives and laborers over-represented among the completed jobs.

B. Specification of the Job Duration Model

In order to use completed tenure as an earnings equation control variable in the fashion described in Section II, we need a method of determining the expected completed duration of the incomplete jobs. We specify and estimate a parametric model of completed job duration, and then use the estimated parameters to predict the expected completed length of jobs still in progress as of the last date we observe them.

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

(16) $\Pr(D \geq T) = \exp[-\lambda T^{\alpha}]$

where $\alpha$ is a positive parameter. The proportional hazard assumption implies that

(17) $\lambda = e^{-Z\gamma}$,

where Z is a vector of observable individual characteristics hypothesized to affect job duration and $\gamma$ is a vector of parameters. The separation hazard associated with this distribution is

(18) $H(t) = \lambda\alpha t^{(\alpha-1)}$.

Clearly, where $\alpha=1$ this distribution collapses to the exponential as a special case. In this case the hazard is constant over time and equal to $\lambda$. Where $\alpha>1$, this distribution has an associated hazard rate which is increasing as the spell continues. In other words, the probability that a spell ends at a point in time given that it has lasted to that point is increasing.

Analogously, where $\alpha<1$, this distribution has an associated hazard rate which is decreasing as the spell continues. In other words, the probability that a spell ends at a point in time given that it has lasted to that point is decreasing.

If the parameters of a Weibull duration model are estimated, there is some ambiguity in the interpretation of the estimate of $\alpha$. The obvious interpretation is that the estimated value of $\alpha$ indicates "true" duration dependence. In other words, it indicates how the instantaneous probability of a spell ending moves over time. An alternative interpretation is that the estimate of $\alpha$ is biased downward by unmeasured heterogeneity in match quality so that the estimate (particularly if less than one, as is often the case) must be interpreted with care. Consider the simple case where there is, in fact, no true duration dependence so that the hazard rate associated with a given spell is some constant $\lambda$. Heterogeneity in the hazard rates across jobs could arise because individuals have different temperaments (and thus different separation propensities), because jobs differ in their relative attractiveness (and thus in their turnover rates) and/or because there is variation in the quality of the individual/job match (and thus in the probability of the match being broken off). If there is unmeasured heterogeneity in the hazard across spells, the spells that last a long time are more likely to be the spells with low hazards. These spells are less likely to end than the randomly selected spell so it appears that spells are progressively less likely to end as time goes by. However, this apparent duration dependence is spurious and simply an artifact of a sample of surviving spells that is becoming increasingly dominated by spells with low hazard rates.

For the purposes of this study, we are not interested in distinguishing between true and spurious duration dependence. We are simply interested in estimating a parametric model of completed job duration that is flexible enough to make allowance for both true duration dependence and unmeasured heterogeneity in hazards. While we could specify a particular distribution for the unobserved heterogeneity in the hazard along with the Weibull form of duration dependence, it is well known that there are problems with robustness of the estimates of heterogeneity and duration dependence with regard to changes in the assumed distributions in models of this sort.[22] We take the approach of estimating a simple Weibull model of completed job duration without any explicit representation for unmeasured heterogeneity. However, the estimate of the parameter $\alpha$ is interpreted here as representing some (unspecified) combination of true duration dependence and unmeasured heterogeneity. The resulting predicted expected completed job durations, which are the important end products of this analysis, are directly affected both by this parameter and by the last observed value of seniority. Thus, both duration dependence and unmeasured heterogeneity are accounted for, and these predicted values should be considered to be the product of a rather flexible specification for completed job duration. The important point is that we are interested in accurately predicting completed job duration, not in isolating the degrees of true duration dependence and unmeasured heterogeneity. The Weibull model should be adequate for this purpose.

As in the case of exponential durations, the contribution to the likelihood function made by a completed job is the probability-density that the job lasted <u>exactly</u> $S_f$ years given that the job lasted at least $S_0$ years.[23]

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

23. 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

Given a Weibull distribution for duration, this is

$$(19)\quad \Pr(D=S_f|D>S_0) = \lambda\alpha S_f^{\alpha-1}\exp[-\lambda(S_f^{\alpha}-S_0^{\alpha})].$$

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

$$(20)\quad \Pr(D\geq S_f|D>S_0) = \exp[-\lambda(S_f^{\alpha}-S_0^{\alpha})].$$

The log-likelihood function is formed from these probabilities as

$$(21)\ \ln(L) = \sum_j(C_j\ln\Pr(D_j\geq S_{fj}|D_j>S_{0j}) + (1-C_j)\ln\Pr(D_j=S_{fj}|D_j>S_{0j}))$$

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).

Note that this specification of the likelihood function assumes that unmeasured factors affecting completed job durations are independent across spells. However, within each of the three samples, there are multiple observations on job durations for some individuals. A more complete specification would account explicitly for the possibility that some individuals are inherently less likely to change jobs than others for reasons that are unmeasured. This would induce a correlation across jobs in the completed job durations for a given individual after controlling for observed differences in the hazard. Given the highly nonlinear nature of the model, an appropriate tractable procedure for accounting for this correlation is not obvious. At this point the analysis proceeds under the assumption that unmeasured factors affecting completed job durations are independent.

C. Estimation of the Job Duration Model

Column 1 and column 2 of Table 2 contain estimates of the Weibull job duration model estimated over the subsamples of 1048 professional, technical and managerial jobs and 1623 blue collar jobs, respectively. These estimates were derived by maximizing the likelihood function defined above with respect to the parameters $\gamma$ and $\alpha$.[24] In interpreting the estimates of the determinants of the baseline hazard ($\lambda$), recall that the hazard rate was specified such that $\lambda = e^{-Z\gamma}$. Thus, an increase in a variable with a positive coefficient reduces $\lambda$ and increases the expected duration of the job.

The two sets of estimates exhibit some interesting differences.[25] The marginal effect of pre-job experience on job duration for professional, technical and managerial workers is never statistically significant at the .05 level or better. In contrast, among blue collar workers, having more pre-job experience has a significant positive association with completed job duration.

Recall that the explanation for a positive relationship between experience and earnings based on sorting of workers into successively better jobs also had the implication that completed job duration would be positively related to pre-job experience. For blue collar workers there is evidence consistent with this sorting. However, this evidence is also consistent with the view that blue collar workers mature over time into "naturally" more stable workers. There is no evidence for sorting or maturation for professional, technical, and managerial workers.

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24. The algorithm described by Berndt, Hall, Hall, and Hausman (1974) was used to find the maximum.

25. Based on a Wald test, 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 $\chi^2$ with 22 degrees of freedom, is 77.9. The critical value of this distribution at the one percent level of significance is 40.29. The independence assumption of this test is not strictly satisfied, since the two samples contain some jobs in common

Table 2:
Selected Coefficients from Final Tenure Models[a]

| | (1) Prof,Tech Managerial Nonunion | (2) Blue Collar Nonunion | (3) All Nonunion |
|---|---|---|---|
| Y (Inverse Baseline Hazard, $\lambda = e^{-zY}$) | | | |
| Years of experience | -.0238 | .0548 | .0278 |
| | (.0197) | (.0096) | (.0080) |
| (Years of experience)$^2$ | .00106 | -.00095 | .00028 |
| | (.00023) | (.00068) | (.00026) |
| Years of education | .0740 | .0272 | .0386 |
| | (.0241) | (.0130) | (.0107) |
| Nonwhite (yes = 1) | -.407 | -.00137 | -.0996 |
| | (.226) | (.08381) | (.0747) |
| Married (yes = 1) | .298 | .477 | .416 |
| | (.134) | (.070) | (.059) |
| Disabled (yes = 1) | .0577 | .215 | -.0295 |
| | (.2164) | (.107) | (.0809) |
| Manager (yes = 1) | -.119 | --- | -.237 |
| | (.111) | | (.101) |
| Clerical, sales (yes = 1) | --- | --- | -.297 |
| | | | (.096) |
| Foreman, craftworker (yes = 1) | --- | .109 | -.645 |
| | | (.059) | (.089) |
| Operative, laborer (yes = 1) | --- | --- | -.726 |
| | | | (.092) |
| "Duration" Parameter | | | |
| $\alpha$ | .380 | .400 | .401 |
| | (.027) | (.016) | (.013) |
| Log-Likelihood | -927.0 | -1206.4 | -2397.4 |
| Sample size | 1048 | 1623 | 2857 |

[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 column (1) and column (3) models and operatives/laborers are the omitted occupational group in the column (2) model. The models also include industry and region controls: these coefficient estimates are reported in Table 2 of Appendix B. The numbers in parentheses are asymptotic standard errors.

The estimates also suggest that education has a stronger positive relationship with job duration in white collar occupations than in blue collar occupations. In both samples, being married raises expected job duration.[26] Disabled blue collar workers actually appear to have somewhat longer expected job durations than non-disabled blue collar workers. Nonwhites have shorter job durations than whites in the professional/technical/managerial subsample, though the relevant coefficient is only marginally significant. There is no significant difference in job durations by race among blue collar workers. As might have been expected given the lower percentage of professional/technical workers in the completed job subsample than in the incomplete job subsample, the column (1) estimates are weakly consistent with the notion that, within the white collar group, professional and technical workers' jobs last longer than managers' jobs. Within the blue collar group, operatives and laborers have shorter jobs than do foremen and craftsworkers. The estimates of $\alpha$ are significantly and substantially less than one for both white and blue collar jobs. We interpret this as implying some combination of negative duration dependence and unmeasured heterogeneity.

Overall, the most interesting difference between the professional, technical, and managerial job duration model and the blue collar job duration model is the difference in the effects of prior experience on job duration. We return to this finding below.

The third column of Table 2 contains the estimates of the Weibull final tenure model estimated for the full sample of 2857 nonunion jobs. These results are intermediate between those for the professional, technical and managerial sample and those for the blue collar sample. The estimated

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26. If we accept the notion that married workers are more "mature", then this result is consistent with the view that more "mature" workers are naturally more stable.

coefficients imply that jobs last longer in cases where the person taking the job has more previous work experience, has more education, is married and is not disabled. Jobs held by whites do not last significantly longer than jobs held by nonwhites. The omitted occupational category in the model is the professional/technical category. Being a professional or technical worker increases expected job duration compared to that associated with any of the other occupational categories. The occupational category with the shortest expected job duration is the operative/laborer category. The estimate of $\alpha$ is strongly significantly less than one, which we again interpret as reflecting some mixture of heterogeneity and negative duration dependence.

D. 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 ($S_f$ years). Note that the job duration model we have estimated is based on data for the pre-retirement 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:

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

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

$$\begin{aligned} &\Pr(D>S_f)=\exp[-\lambda S_f^{\alpha}], \\ (23)\quad &\Pr(D\geq S_{65})=\exp[-\lambda S_{65}^{\alpha}], \text{ and} \\ &\lambda=e^{-Z\gamma}. \end{aligned}$$

With an appropriate change of variables, the integral on the right side of the conditional expectation in equation (22) can be expressed as incomplete gamma functions. This is

$$(24)\quad \int_{S_f}^{S_{65}} \lambda\alpha t^{\alpha} e^{-\lambda t^{\alpha}}\, dt = \lambda^{(1/\alpha)}\left[\int_0^{y_{65}} e^{-y} y^{(1/\alpha)} dy - \int_0^{y_f} e^{-y} y^{(1/\alpha)} dy\right]$$

where $y_f=\lambda S_f^{\alpha}$ and $y_{65}=\lambda S_{65}^{\alpha}$. Both of these incomplete gamma functions can be evaluated numerically.[27] Estimates of the conditional expectation in equation (22) computed using the appropriate set of parameters from Table 2 produce the job duration measure used for incomplete jobs in the earnings function estimation in the next section.

We also use estimates of the square of completed job duration in the earnings function estimation. For incomplete jobs, this is estimated similarly to completed job duration itself as

$$(25)\quad \hat{E}(D^2|D>S_f) = \frac{1}{\Pr(D>S_f)}\int_{S_f}^{S_{65}} \lambda\alpha t^{(1+\alpha)} e^{-\lambda t^{\alpha}}\, dt + \frac{\Pr(D\geq S_{65})}{\Pr(D\geq S_f)} * S_{65}^2$$

A similar change of variables to that which produced equation (24) yields easily approximated incomplete gamma functions which are used, along with the appropriate set of parameters from table 2, to compute our estimate of the square of completed job duration for incomplete jobs.

As noted earlier, actual job duration was observed for 403 of the 1048 jobs represented in the professional, technical and managerial sample, for 836

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27. The approximations used were taken from Abromovitz and Steigun (1972).

of the 1623 jobs in the blue collar sample, and for 1396 of the 2857 jobs in the all occupations sample. Many of the professional, technical and managerial jobs were quite long; of the 1048 completed and incomplete jobs represented in the sample, 19 percent were predicted to have completed durations of 1 year or less, 10 percent to have completed durations of 1 to 3 years, 18 percent to have completed durations of 3 to 10 years and 53 percent to have completed durations of more than ten years.[28] More of the blue collar jobs were relatively short; of the 1623 completed and incomplete jobs represented in the sample, 33 percent were predicted to have completed durations of 1 year or less, 15 percent to have completed durations of 1 to 3 years, 26 percent to have completed durations of 3 to 10 years, and 26 percent to have completed durations of more than ten years.[29] The distribution of predicted completed job durations for the all occupations sample is intermediate between that for the two subsamples.[30]

It is not easy to assess intuitively whether the parameter estimates in this completed tenure model are sensible. One check on the model is to explore whether its implications regarding the distribution of completed job durations are consistent with other available evidence. Hall (1982) has recently reported estimates of various statistics relating to the distribution of the lengths of completed jobs constructed using a radically different approach than we have used here. We have computed statistics which can be

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28. Of the 403 completed professional, technical and managerial jobs in the sample, 82 (20.3%) had completed durations of three to ten years and 25 (6.2%) had completed durations of more than ten years.

29. Of the 839 completed blue collar jobs in the sample, 103 (12.3%) had completed durations of three to ten years and 28 (3.3%) had completed durations of more than ten years.

30. The distribution for the full sample is not simply the weighted average of the distributions of the two subsamples for three reasons: 1) Some jobs appear in both of the occupational subsamples; 2) A few clerical and sales jobs appear in neither subsample; and 3) The predicted durations for the three samples are based on different Weibull model estimates.

compared to some of those he reports. For the purposes of this comparison, all estimates of expected job duration were computed using the model of job duration contained in the third column of Table 2.

Hall (1982) uses the CPS tenure data to draw conclusions concerning the distribution of completed job duration. In essence, his approach is to infer the probability that a job that has already lasted N years will last another i years by comparing the proportion of employed persons aged A plus i years in jobs that have lasted N plus i years to the proportion of employed persons aged A in jobs that have lasted N years. Among other statistics, he reports the probabilities that jobs in various tenure brackets will end up lasting at least 20 years for workers aged 40 to 45. His estimates are reported in the first column of Table 3.

It is relatively easy to compute similar statistics based on the job duration model in column (3) of Table 2. For comparability with Hall's results, we perform calculations for an individual aged 42 with the sample mean amount of education who we assume started school at age 6 and has worked continuously since leaving school. We assume sample mean values of all other characteristics except pre-job experience used in making our predictions; pre-job experience is set so that age upon attaining the mean amount of schooling plus seniority plus pre-job experience equals 42. Important differences between the two sets of estimates include the fact that our sample includes only nonunion male heads of households, whereas Hall's estimates reflect the experiences of all 40 to 45 years olds, both union and nonunion and both female and male. Our estimates of the statistics reported by Hall appear in the last column of Table 4.

The two sets of estimates are quite similar overall. The major difference between them is that we generally find somewhat higher probabilities that short jobs will end up lasting over 20 years. The fact

Table 3:
Aggregate Final Tenure Distributions:[a]
Estimated Using Alternative Approaches

| Probability that a job in indicated tenure interval will end up lasting 20 or more years: | Estimates reported by Hall (1982) | Probability that a job with indicated current tenure will end up lasting 20 or more years: | Estimates based on Table 2: |
|---|---|---|---|
| 0 - 0.5 years | .046 | 0.25 years | .116 |
| 0.5 - 1 year | .078 | 0.75 years | .147 |
| 1 - 2 years | .113 | 1.50 years | .181 |
| 2 - 3 years | .157 | 2.50 years | .218 |
| 3 - 5 years | .204 | 4.00 years | .268 |
| 5 - 10 years | .355 | 7.50 years | .384 |
| 10 - 15 years | .590 | 12.50 years | .571 |
| 15 - 20 years | .980 | 17.50 years | .826 |
| 20 + years | 1.000 | 20.00 years | 1.000 |

[a]The Hall (1982) figures apply to all 40 to 45 year olds. Our estimates are computed for a 42 year old white male with mean characteristics for the full nonunion jobs sample using the parameter estimates contained in the third column of table 2.

that Hall's estimates include women while ours do not could explain this discrepancy, though the fact that Hall's estimates include union jobs while ours do not seems likely to work in the opposite direction. Overall, we find it reassuring that Hall's nonparametric procedure and our parametric procedure produce similar results and take this as evidence that our estimates of completed job duration are reasonable.

## IV. Earnings Function Estimates

Having derived an estimate of completed tenure which can be used as a control variable in earnings functions, we turn now to the empirical investigation of the central questions motivating this research. First, do those on long jobs receive higher earnings than those on short jobs? Second, if so, what fraction of the positive association between seniority and earnings in the cross section reflects the fact that those with longer service tend to be in better paying, longer jobs? Third, is there any evidence that the positive cross-sectional association between labor market experience and earnings in significant measure reflects the fact that more experienced workers have found better jobs and/or better matches?

Consider a standard earnings function of the form:

$$(26)\quad \ln(W) = X\beta + \varepsilon$$

where ln(W) is the logarithm of real average hourly earnings, X is a vector of individual characteristics, $\beta$ is a vector of parameters, and $\varepsilon$ represents unmeasured factors affecting earnings. This earnings function is estimated using each of three samples of individual-year observations from the three samples of jobs from the PSID discussed in Section III. As previously noted, in presenting our empirical findings we have chosen to focus primarily on the results for the two broad occupational subgroups: nonunion professional, technical and managerial employees and nonunion blue collar employees

Results for the full nonunion sample are also presented for comparison purposes. Recall that the samples consist only of nonunion male heads of households. There are 3603 individual-year observations on workers in the 1048 professional, technical, and managerial jobs; 3926 individual-year observations on workers in the 1623 blue collar jobs; and 8688 individual-year observations on workers in the 2857 jobs on the overall sample.[31]

The first column in each of tables 4a, 4b and 4c contain, for each of the samples, the means and standard deviations of the central variables used in the analysis. Given that we wish to identify the incremental return to seniority over and above the return to general labor market experience, the earnings functions we estimate include total labor market experience rather than pre-job experience.[32] The professional, technical and managerial sample has slightly more total labor market experience (18.14 years versus 17.62 years) and slightly longer current tenure (8.85 years versus 6.46 years) than the blue collar sample. As explained in Section III, expected job duration and the expected value of its square were computed based on the Weibull model for the relevant sample of jobs where the end of the job was not observed; for the remaining jobs, actual completed job duration and its square are used. Expected job durations are somewhat longer in the professional, technical and managerial sample than in the blue collar sample (20.78 years versus 13.94 years). The characteristics of the full nonunion sample are intermediate

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31. Persons who were retired, permanently disabled, self-employed, employed by the government or residents of Alaska or Hawaii were also excluded. The overall nonunion sample includes clerical and sales workers.

32. If both seniority and the experience term were entered linearly, the two specifications would yield identical fits. If the squares of seniority and/or experience appear in the earnings function, the two specifications will not yield identical fits. However, we have estimated all of the Section IV models 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.

between those of the two occupational subsamples.[33]

As with the sample of jobs, the individual-year sample has multiple observations for some individuals. If there are unmeasured individual factors affecting earnings then one might want to estimate some sort of errors components model that accounts for correlations across observations in $\varepsilon$. While this is relatively straightforward to implement in a linear model, the estimates presented in this section are standard ordinary least squares (OLS) estimates that do not account for such correlations. The reason for this is fundamentally related to the ultimate goal of using a measure of completed job duration to account for heterogeneous worker, job, and/or match quality and yield better estimates of the returns to seniority and experience. If we estimate a standard fixed effect earnings model, which is equivalent to including a separate constant term for each individual in the sample in the earnings function, we can learn nothing about the relative returns to seniority and experience from those for whom only one job is observed. Unless a job change occurs, seniority and experience both increase by one each year, so that their effects are not separately identifiable. Moreover, the completed job duration measure will vary within the observations on a particular individual only for those individuals observed on more than one job. Observing multiple jobs for an individual will be more likely where jobs are relatively short. It is likely that the returns to seniority and experience among the group of workers who change jobs during the period we observe them will differ from the returns to seniority and experience in the sample as a whole. In addition, the relationship of the key job duration

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33. As was true with the jobs samples, the all occupations means are not simply the weighted averages of the means for the two occupational subgroups, since some jobs appear in both of the occupational subsamples, a few clerical and sales jobs appear in neither, and the predicted job durations for the three samples are based on different Weibull models.

measure with earnings cannot be adequately estimated through a sample of short jobs. On balance, it seems preferable to leave the fixed effects unaccounted for rather than to have the entire analysis rely on jobs with short completed durations.

Given that a key variable, completed job duration, is proxied by our estimate of the expectation of completed job duration, the question of potential measurement error in this variable ought to be addressed. There are actually two sources of error in the predicted completed job duration measure. The first 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 will be consistent. The second source of error is that the expected value of completed job duration is used in place of the actual realization of completed job duration. This is not a problem because, by definition, the expected value of completed job duration is uncorrelated with the deviation between the expected and the true value. Since the measurement error is exactly the deviation between the expected value and the true value, there is no correlation induced between the included regressor and the error term in the regression. Thus, there is no bias from this source in our estimated earnings function coefficients.

While these 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. The standard errors presented are corrected for the effects of these errors and for general heteroskedasticity of the form analyzed by White (1980).[34]

Table 4a contains ordinary least squares (OLS) regression results for the nonunion professional, technical and managerial sample. Column (1) is a standard earnings equation; it suggests that professional, technical and managerial employees enjoy sizeable returns to both labor market experience and seniority with a particular employer. Column (2) differs from the standard earnings function in that it also includes expected job duration and the expectation of the square of job duration.[35]

The first thing to note about the coefficient estimates in column (2) is that there is a very strong positive association between job length and earnings. A man on a job that ends up lasting 10 years rather than 5 earns 7.6 percent more in each year on that job; a man on a job that ends up lasting 20 years rather than 5 earns 18.2 percent more in each year on that job. The estimates in column (3) allow the effects of completed job duration and its square to vary with seniority. The substance of the results do not change and the evidence is persuasive that workers in long jobs earn more uniformly over the life of the job.[36]

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34. See Newey (1984) for a detailed discussion of the sort of standard error calculations used here.

35. We include the expectation of the square of job duration because we do not wish to be unduly restrictive about the functional form of the underlying relationships.

36. The positive relationship between job duration and earnings, which exists for both of the occupational subsamples and the full nonunion sample, provides strong evidence against the Harris and Holmstrom (1982) interpretation of the return to experience. In their model, the worker and all firms are initially uncertain about the worker's productivity. The initial wage equals the expected value of the worker's productivity minus an insurance premium and the employer guarantees not to reduce the initial wage. Workers revealed to have high productivity 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, the simple Harris and Holmstrom model predicts a negative correlation between job duration and earnings; our results indicate that this correlation is strongly positive.

Table 4a:
Selected Coefficients from ln (average hourly earnings) Models
Managerial and Professional Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of experience | 18.14 [10.11] | .0289 (.0029) | .0283 (.0026) | .0255 (.0030) |
| (Years of experience)$^2$ | 431.20 [411.46] | -.00047 (.00007) | -.00047 (.00006) | -.00041 (.00007) |
| Years of current seniority | 8.85 [8.32] | .0220 (.0026) | .00543 (.00173) | .00407 (.00247) |
| (Years of current seniority)$^2$ | 147.47 [227.79] | -.00044 (.00009) | --- | --- |
| E(completed job duration) | 20.78 [12.12] | --- | .0201 (.0025) | --- |
| {E(completed job duration)}$^2$ | 628.88 [501.82] | --- | -.00036 (.00006) | --- |
| E(job duration)*[=1 if seniority ≤1] | 1.44 [4.47] | --- | --- | .0201 (.0078) |
| {E(job duration)}$^2$*[=1 if seniority ≤1] | 31.96 [119.1] | --- | --- | -.00039 (.00027) |
| E(job duration)*[=1 if 1 < seniority ≤ 3] | 2.29 [6.08] | --- | --- | .0297 (.0057) |
| {E(job duration)}$^2$*[=1 if 1 < seniority ≤ 3] | 54.47 [173.9] | --- | --- | -.00067 (.00017) |
| E(job duration)*[=1 if 3 < seniority ≤ 10] | 6.03 [10.46] | --- | --- | .0260 (.0039) |
| {E(job duration)}$^2$*[=1 if 3 < seniority ≤ 10] | 165.7 [324.9] | --- | --- | -.00051 (.00010) |
| E(job duration)*[=1 if 10 < seniority] | 376.7 [569.1] | --- | --- | .0193 (.0030) |
| {E(job duration)}$^2$*[=1 if 10 < seniority] | | --- | --- | -.00030 (.00007) |
| $R^2$ | --- | .3681 | .3817 | .3832 |

[a]All models also include controls for education, race, marital status, disability, occupation, industry, region, and year. These coefficient estimates are reported in Table 3a of Appendix B. E(completed job duration) is computed using the estimates in column (1) of table 2. The numbers in parentheses are standard errors. The standard errors in columns (2) and (3) are asymptotic and corrected for the fact that E(completed job duration) is predicted. Sample size=260[illegible]

A second noteworthy feature of the results in column (2) is that there is a marked fall in the estimated return to seniority compared to the standard model. In the model of column (1), an additional year of seniority at the mean was associated with 1.4 percent higher earnings; in the column (2) model which includes completed job duration as a control, an additional year of seniority is associated with only 0.5 percent higher earnings. In the column (1) model, 10 years of seniority contribute 19.2 percent to earnings and 20 years of seniority contribute 30.2 percent; in the column (2) model, 10 years of seniority contribute 5.6 percent to earnings and 20 years of seniority contribute 11.5 percent to earnings. Thus, roughly two thirds of the estimated return to seniority in the cross section appears to reflect bias associated with higher earnings on longer jobs, rather than true returns to seniority *per se*.[37]

A third question of interest is whether standard estimates of the return to experience include a component reflecting returns to search. The fact that the estimated return to experience does not fall when completed job duration and its square are introduced into the model suggests that, for professional, technical and managerial employees, the answer to this question is "no"; the positive association between experience and earnings appears to reflect true returns to experience rather than sorting of more experienced workers into better jobs.

Table 4b reports a similar set of earnings equations estimated using OLS for the sample of blue collar nonunion employees. Column (1) is again the standard cross section earnings equation. In this model, there appear to be

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37. With both seniority and seniority squared in the column (2) model, the coefficient of neither was significantly different from zero. Results based on the estimates of this model are virtually identical to those we report

Table 4b:
Selected Coefficients from ln (average hourly earnings) Models
Blue Collar Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of experience | 17.62 [11.19] | .0146 (.0025) | .0103 (.0025) | .0104 (.0025) |
| (Years of experience)$^2$ | 435.59 [475.31] | -.00030 (.00006) | -.00022 (.00005) | -.00023 (.00006) |
| Years of current seniority | 6.46 [7.45] | .0263 (.0026) | .00389 (.00200) | .00109 (.00282) |
| (Years of current seniority)$^2$ | 97.16 [193.90] | -.00051 (.00010) | --- | --- |
| E(completed job duration) | 13.94 [11.62] | --- | .0182 (.0020) | --- |
| {E(completed job duration)}$^2$ | 362.05 [438.37] | --- | -.00026 (.00006) | --- |
| E(job duration)*[=1 if seniority ≤1] | 1.14 [3.14] | --- | --- | .0298 (.0088) |
| {E(job duration)}$^2$*[=1 if seniority ≤1] | 16.85 [63.40] | --- | --- | -.00084 (.00041) |
| E(job duration)*[=1 if 1 < seniority ≤ 3] | 1.58 [4.12] | --- | --- | .0423 (.0056) |
| {E(job duration)}$^2$*[=1 if 1 < seniority ≤ 3] | 25.61 [86.82] | --- | --- | -.00119 (.00023) |
| E(job duration)*[=1 if 3 < seniority ≤ 10] | 4.66 [8.24] | --- | --- | .0337 (.0040) |
| {E(job duration)}$^2$*[=1 if 3 < seniority ≤ 10] | 102.7 [208.5] | --- | --- | -.00079 (.00014) |
| E(job duration)*[=1 if 10 < seniority] | 6.57 [12.89] | --- | --- | .0154 (.0030) |
| {E(job duration)}$^2$*[=1 if 10 < seniority] | 216.9 [457.9] | --- | --- | -.000092 (.000069) |
| $R^2$ | --- | .3810 | .3934 | .3996 |

[a]All models also include controls for education, race, marital status, disability, occupation, industry, and region. These coefficient estimates are reported in Table 3b of Appendix B. E(completed job duration) is computed using the estimates in column (2) of table 2. The numbers in parentheses are standard errors. The standard errors in columns (2) and (3) are asymptotic and corrected for the fact that E(completed job duration) is predicted. Sample size=3926.

substantial returns to both labor market experience and tenure. If anything, it appears that seniority is worth more to blue collar workers than to white collar workers (2.0 percent per year versus 1.4 percent per year at the respective sample means. However, blue collar workers appear to have lower returns to labor market experience (0.4 percent per year versus 1.2 percent per year at the respective sample means).

The estimates in column (2) of table 4b, which include expected job duration and the expectation of its square, imply that, like professional, technical and managerial employees, blue collar workers in long jobs earn substantially more than blue collar workers in short jobs. A man in a job that ends up lasting 10 years earns 7.5 percent more in each year of that job than a man in a job that ends up lasting 5 years; a man in a job that ends up lasting 20 years earns 19.4 percent more than a man in a job that ends up lasting 5 years. The estimates in column (3) allow the effects of completed job duration and its square to vary with seniority. As with the professional, technical and managerial workers, the substance of the results do not change and the evidence is persuasive that workers in long jobs earn more uniformly over the life of the job.[38]

For the blue collar sample, as with the professional, technical, and managerial sample, the positive association between job duration and earnings appears to account for most of the apparent return to seniority in the standard cross section earnings equation; the estimated return to seniority at the mean drops from 2.0 percent in the standard column (1) model to under 0.4 percent per year in the column (2) model and loses statistical significance. The point estimates of the coefficients in the column (1) model imply that 10

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38. Once again, this is strong evidence against the Harris and Holmstrom (1982) interpretation of the return to experience. See note 36.

years of seniority add 24 percent to earnings and 20 years of seniority add 38 percent; those in the column (2) models, that 10 years of seniority contribute 4 percent to earnings and 20 years of seniority contribute 8 percent.[39]

With regard to the return to general labor market experience, the results for blue collar workers are somewhat different than for the professional, technical, and managerial workers. For those white collar workers, there was no evidence of a positive search component in the return to experience estimated using the standard (column 1) specification. In contrast, the estimated return to experience for blue collar workers falls by approximately 20 to 30 percent when our measures of completed job duration are introduced into the regression (column (2)). Essentially, a piece of the already-low estimated return to labor market experience among blue collar workers appears to reflect an underlying correlation between experience and job and/or match quality, rather than returns to experience <u>per se</u>. The estimated return to an additional year of experience at the mean drops from 0.4 percent in column (1) to 0.3 percent in column (2). In the column (1) model, 10 years of experience contribute 12 percent to earnings and 20 years of experience contribute 19 percent; in the column (2) model, 10 years of experience add 8 percent to earnings and 20 years of experience add 13 percent.

Finally, Table 4c presents results like those just discussed for the full nonunion sample. With one exception, the full sample results are intermediate between those for professional, technical and managerial employees and those for blue collar workers. The standard earnings function

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39. Again, in a model like that in column (2) that included both seniority and its square, neither had a coefficient that was significantly different from zero. Results based on the estimates of this model are very similar to those we report.

Table 4c:
Selected Coefficients from ln (average hourly earnings) Models
Full Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of experience | 17.61 [10.68] | .0227 (.0018) | .0210 (.0017) | .0193 (.0018) |
| (Years of experience)$^2$ | 424.13 [443.47] | -.00038 (.00004) | -.00037 (.00004) | -.00034 (.00004) |
| Years of current seniority | 7.38 [7.89] | .0272 (.0018) | .00690 (.00125) | .00396 (.00179) |
| (Years of current seniority)$^2$ | 116.77 [208.68] | -.00059 (.00006) | --- | --- |
| E(completed job duration) | 16.59 [12.02] | --- | .0213 (.0015) | --- |
| {E(completed job duration)}$^2$ | 459.84 [466.28] | --- | -.00040 (.00004) | --- |
| E(job duration)*[=1 if seniority ≤1] | 1.28 [3.73] | --- | --- | .0223 (.0054) |
| {E(job duration)}$^2$*[=1 if seniority ≤1] | 22.8 [86.3] | --- | --- | -.00055 (.00021) |
| E(job duration)*[=1 if 1 < seniority ≤ 3] | 1.95 [5.05] | --- | --- | .0315 (.0037) |
| {E(job duration)}$^2$*[=1 if 1 < seniority ≤ 3] | 38.4 [124.2] | --- | --- | -.00075 (.00013) |
| E(job duration)*[=1 if 3 < seniority ≤ 10] | 5.11 [8.99] | --- | --- | .0276 (.0026) |
| {E(job duration)}$^2$*[=1 if 3 < seniority ≤ 10] | 123.1 [246.5] | --- | --- | -.00058 (.00008) |
| E(job duration)*[=1 if 10 < seniority] | 8.25 [14.14] | --- | --- | .0215 (.0020) |
| {E(job duration)}$^2$*[=1 if 10 < seniority] | 275.6 [502.2] | --- | --- | -.00034 (.00005) |
| $R^2$ | --- | .4790 | .4889 | .4903 |

[a]All models also include controls for education, race, marital status, disability, occupation, industry, and region. These coefficient estimates are reported in Table 3c of Appendix B. E(completed job duration) is computed using the estimates in column (3) of table 2. The numbers in parentheses are standard errors. The standard errors in columns (2) and (3) are asymptotic and corrected for the fact that E(completed job duration) is predicted. Sample size=8688.

estimates of column (1) indicate the existence of substantial returns to both labor market experience and seniority. In column (2), both expected job duration and the expectation of its square are significant, and the estimated association between job duration and earnings looks similar to that for the occupational subgroups. The estimates in column (3) allow the effects of completed job duration and its square to vary with seniority. As with the occupational subgroups, the evidence is persuasive that workers in long jobs earn more uniformly over the life of the job.

The noteworthy difference between the full sample results and the occupational subgroup results is that the estimated return to seniority remains slightly larger in the full nonunion sample model than in either of the occupational subsample models (0.7 percent per year and highly significant, versus 0.5 percent per year in the professional, technical and managerial sample and an insignificant 0.4 percent per year in the blue collar sample). This pattern of results may imply that there is a return to seniority with a particular employer for some workers that takes the form of promotion from a blue collar position to a white collar position.[40] However, even in the estimates for the full nonunion sample, no more than a third of the apparent return to seniority estimated using the standard cross section earnings function appears to reflect a true return to seniority *per se*.

The estimated return to labor market experience in the full sample column (2) model is between that for the two occupational subsamples.

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40. As noted earlier, fewer than 178 of the 2857 jobs in the full nonunion sample involve a move from blue collar status to professional, technical or managerial status or vice versa. However, if there were substantial returns to seniority for even a small number of workers taking the form of cross-occupational category promotions, this could conceivably yield returns to seniority for the full sample that exceed the returns to seniority for the occupational subsamples by the 0.2 to 0.3 percentage points we find.

## V. 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 return to seniority in excess of the return to general labor market experience for the average worker. For both the nonunion professional, technical and managerial sample and the nonunion blue collar sample, the point estimate of the seniority coefficient in the model including a control for job duration suggests that the true return to seniority is on the order of 0.5 percent per year, rather than the 1.4 to 2.0 percent per year suggested by the standard cross section model. The seniority coefficient in the model for the full nonunion sample that includes a control for job duration is slightly larger and more precisely estimated than those in the models for the occupational subgroups, but even that estimate suggests true returns to seniority that are no more than a third as large as implied by the coefficient estimates for the same sample from a standard cross section model.[41]

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 return to seniority of 0.5 percent per year could translate into a

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41. Our findings regarding the returns to seniority are consistent with those obtained by Altonji and Shakotko (1985) using a much different instrumental variables approach.

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.[42] However, earnings deferral under implicit contracts appears to be a much less important factor in both white collar and blue collar labor markets than has generally been believed.

A second significant result to emerge from our analysis is the finding that labor market experience plays a different role for professional, technical and managerial employees than for blue collar employees. There is no evidence that nonunion professional, technical and managerial employees sort themselves into better jobs over the course of their working lives; for this group, there is no association between pre-job experience and job duration, and introducing job duration as a control variable into the earnings function does not affect the relatively large estimated return to labor market experience. In contrast, among nonunion blue collar employees, greater pre-job experience is associated with significantly longer job durations and introducing job duration as a control variable into the earnings function lowers the already-low estimated return to labor market experience by 20 to 30 percent. One interpretation of the results for blue collar workers is that young blue collar workers pass through a period in which they hold a series of short, relatively unrewarding jobs, and that, after some time, either because they find a job that is good enough to stay with (i.e., a job that is unlikely to be dominated by another) and/or because they mature into more stable and more productive workers, they settle into a longer lasting, more remunerative

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42. Freeman and Medoff (1984) 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.

position.[43]

The results regarding experience and seniority were derived by including a measure of completed job duration in an earnings function. One benefit of this approach is that we can examine the relationship of earnings with completed job duration directly. In all cases it was found that workers who are on jobs that will end up lasting a long time earn significantly more in every year of the job.[44] Indeed, this relationship largely explains the positive cross sectional association between seniority and earnings.

We would like to be able to say more about the structural foundations of the positive relationship between earnings and job duration. One key to untangling the simultaneous nature of this structural relationship is the measurement of the relative importance of individual heterogeneity versus job/match heterogeneity as causal factors in determining earnings and job duration. If such a measurement found that individual heterogeneity accounts for a significant part of the positive association between earnings and job duration, the positive relationship may be thought of as a return to stability, per se, that stable workers could receive from many employers. If, on the other hand, it were found that job and/or match heterogeneity accounted for a significant part of the positive association between earnings and job duration, there may be evidence for efficiency wage behaviour. For example, if there are employers for whom turnover is particularly costly, monitoring worker shirking is particularly expensive, or worker loyalty is particularly valuable, these employers may agree to pay their workers more throughout the

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43. Osterman (1980) characterizes the typical labor market behaviour of non-college-educated men as consisting of a "moratorium phase" followed by a "settling down" phase of the sort just described.

44. This relationship is also implicit in the large body of literature on the negative relationship between separation rates and earnings. See Parsons (1977) and Freeman and Medoff (1984) for surveys of this literature.

job. Workers in such jobs will be unlikely to quit. Alternatively, the positive association between earnings and job duration would reflect the desire of employers to preserve idiosyncratically good matches.

In conclusion, we can offer no tangible evidence regarding why earnings deferral under implicit contracts does not seem to be an important overall factor in labor markets. It may be that workers are unable or unwilling to accept earnings deferral. Minimum wage laws might prevent some workers from posting the bond required under an earnings deferral contract. Probably more important, risk aversion and uncertainty about whether the firm will honor its implicit commitments seem likely to make many workers unwilling to enter into such arrangements. The nature of long term employment relationships are certainly deserving of further study.

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APPENDIX A

Bias in Seniority and Experience Coefficients

in Cross-section Earnings Equations

This appendix derives the theoretical results concerning bias in the coefficients on seniority and experience in cross sectional earnings functions reported in Section II of the paper. A simple base model is presented first; the consequences of relaxing two of this model's assumptions are then explored.

Base Model

Suppose that the earnings of a particular worker on a particular job at a point in time can be written:

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

where

$W$ = hourly earnings,

$S$ = is current seniority (tenure),

$EXP$ = is pre-job experience,

$\mu$ = 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 with the same observed characteristics on a randomly selected job,

$\eta$ = is a person/job/time-period specific error term,

$i$ = index of individuals,

$j$ = index of jobs,

$t$ = index of years within jobs, and

$\beta$'s = parameters of interest.

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, the coefficient

$\beta_2$ captures the returns to experience *per se*, not any increases in earnings that occur because more experience workers have sorted themselves into better jobs and/or matches. The $\mu_{ij}$ term captures the net influence of unobserved person quality, unobserved job quality and unobserved match quality on hourly earnings. The error $\mu_{ij}$ is assumed to be fixed over the course of a job. The error $\eta_{ijt}$ is assumed to be orthogonal to S, EXP and $\mu$.

In practice, earnings functions are generally estimated using cross section data and $\mu_{ij}$ is not observable. When $\mu_{ij}$ is omitted, equation (A.1) can be rewritten as

$$\text{(A.2)} \quad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \nu_{ijt}.$$

The error term ($\nu_{ijt}$) is defined in terms of the underlying stochastic strucure as

$$\text{(A.3)} \quad \nu_{ijt} = \mu_{ij} + \eta_{ijt}.$$

Suppose that the completed duration of jobs is positively correlated with $\mu_{ij}$ so that the following relationship holds:

$$\text{(A.4)} \quad D_{ij} = \gamma\mu_{ij} + \varepsilon_{ij},$$

where $D_{ij}$ is the completed length of the current job, $\mu_{ij}$ is as defined above, $\gamma$ is a parameter which summarizes the relationship between D and $\mu$, and $\varepsilon_{ij}$ captures the variation in completed job duration that cannot be linked to the earnings advantage associated with worker, job and/or match quality. Suppose further that workers with more pre-job experience are likely to have found better jobs and/or better matches. One simple approximation to such a relationship is

$$\text{(A.5)} \quad \mu_{ij} = \alpha EXP_{ij} + \phi_{ij},$$

where $\mu$ and EXP are as defined above, $\alpha$ is a parameter that summarizes the relationship between $\mu$ and EXP and $\phi_{ij}$ is the variation in $\mu$ not systematically related to EXP. We assume for the present that pre-job experience is not correlated with completed job duration except insofar as more experienced workers are more likely to have found a good job and/or a good match, so that $E(EXP_{ij}*\varepsilon_{ij})$ is zero. Substituting equation (A.5) into equation (A.4) yields

(A.6) $D_{ij} = \alpha\gamma EXP_{ij} + \gamma\phi_{ij} + \varepsilon.$

It is straightforward to show given these relationships and the stated assumptions that

$$(A.7)\quad \begin{aligned} &cov(EXP_{ij},\mu_{ij}) = \alpha*var(EXP_{ij}),\\ &var(D_{ij}) = (\alpha\gamma)^2*var(EXP_{ij}) + \gamma^2*var(\phi_{ij}) + var(\varepsilon_{ij}),\\ &cov(D_{ij},EXP_{ij}) = \alpha\gamma*var(EXP_{ij}),\\ &cov(D_{ij},\phi_{ij}) = \gamma*var(\phi_{ij}),\\ &cov(D_{ij},\varepsilon_{ij}) = var(\varepsilon_{ij}), \text{ and}\\ &cov(D_{ij},\mu_{ij}) = \alpha^2\gamma*var(EXP_{ij}) + \gamma*var(\phi_{ij}). \end{aligned}$$

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. More formally,

(A.8) $E(S_{ijt}) = 1/2*D_{ij}$

and

(A.9) $S_{ijt} = 1/2*D_{ij} + \xi_{ijt}$ ,

where $\xi_{ijt}$ is a random variable with zero mean. The distribution of $\xi_{ijt}$ will vary depending upon the completed length of the job. However, its mean is always zero, as is $cov(\mu_{ij},\xi_{ijt})$, $cov(D_{ij},\xi_{ijt})$, and $cov(EXP_{ij},\xi_{ijt})$. It is straightforward to show that

$$(A.10)\quad \begin{aligned} &cov(S_{ijt},D_{ij}) = 1/2*var(D_{ij}),\\ &cov(S_{ijt},EXP_{ij}) = 1/2*cov(D_{ij},EXP_{ij}),\\ &cov(S_{ijt},\phi_{ij}) = 1/2*cov(D_{ij},\phi_{ij}),\\ &cov(S_{ijt},\varepsilon_{ij}) = 1/2*cov(D_{ij},\varepsilon_{ij}),\\ &cov(S_{ijt},\mu_{ij}) = 1/2*cov(D_{ij},\mu_{ij}), \text{ and}\\ &var(S_{ijt}) = 1/4*var(D_{ij}) + var(\xi_{ijt}). \end{aligned}$$

The standard formula for the bias in ordinary least squares regression coefficients is:

$$(A.11)\quad E(\hat{\beta}) - \beta = (X'X)^{-1}(X'\tau),$$

where $\hat{\beta}$ is the estimated coefficient vector, $\beta$ is the true coefficient vector, X is the vector of explanatory variables, and $\tau$ is the equation error. Applying the standard bias formula to the seniority coefficient in equation (A.2) estimated from cross sectional data yields

$$(A.12)\quad E(\hat{\beta}_1)-\beta_1 = (1/DET_1)*[var(EXP_{ij})*E(S_{ijt}*\mu_{ij})- cov(EXP_{ij},S_{ijt})*E(EXP_{ij}*\mu_{ij})]$$

where $DET_1$ is the determinant of the (X'X) matrix from equation (A.2) and the moments are as defined above. This simplifies to:

$$(A.13)\quad E(\hat{\beta}_1)-\beta_1=(1/2)*(1/DET_1)*\gamma*var(EXP_{ij})var(\phi_{ij})$$

which is positive provided $\gamma$ (the coefficient summarizing the relationship between D and $\mu$) is positive.

The expected difference between the pre-job experience coefficient in equation (A.2) estimated from cross sectional data and $\beta_2$ is

$$(A.14)\quad E(\hat{\beta}_2)-\beta_2=(1/DET_1)* [-cov(EXP_{ij},S_{ijt})*E(S_{ijt}*\mu_{ij}) + var(S_{ijt})*E(EXP_{ij}*\mu_{ij})]$$

This can be rewritten as:

(A.15) $E(\hat{\beta}_2)-\beta_2=(1/DET_1)*$

$$[\alpha*var(EXP_{ij})*\{var(S_{ijt}) - 1/4*var(D_{ij}) + 1/4*var(\varepsilon_{ij})\}],$$

which is also positive so long as $\alpha$ (the coefficient summarizing the relationship between $EXP_{ij}$ and $\mu_{ij}$) is positive.

Augmenting equation (A.2) by adding $D_{ij}$ as an explanatory variable in a cross section yields the following specification:

(A.16) $\ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \beta_3 D_{ij} + \omega_{ijt}$,

where $\omega$ is the estimating equation error. Using the relationships in equations (A.3) and (A.4), this error is

(A.17) $\omega_{ijt} = -(1/\gamma)*\varepsilon_{ij} + \eta_{ijt}$ .

How does introducing $D_{ij}$ into the earnings equation affect the estimated tenure and experience coefficients? The bias in the tenure coefficient in the augmented equation (A.16) model estimated using a cross section is:

(A.18) $E(\hat{\beta}_1)-\beta_1=(1/DET_2)*$

$$[\{var(EXP_{ij})*var(D_{ij})-cov^2(EXP_{ij},D_{ij})\}*E(-S_{ijt}*\varepsilon_{ij}/\gamma) -$$
$$\{cov(EXP_{ij},S_{ijt})*var(D_{ij})-cov(EXP_{ij},D_{ij})*cov(S_{ijt},D_{ij})\}$$
$$*E(-EXP_{ij}*\varepsilon_{ij}/\gamma) +$$
$$\{cov(EXP_{ij},S_{ijt})*cov(EXP_{ij},D_{ij})-$$
$$var(EXP_{ij})*cov(S_{ijt},D_{ij})\}$$
$$*E(-D_{ij}*\varepsilon_{ij}/\gamma)]$$

where $DET_2$ is the determinant of the (X'X) matrix from equation (A.16). Noting that $E(-EXP_{ij}*\varepsilon_{ij})=0$ (by assumption) and that $cov(S_{ijt},Z)=1/2*cov(D_{ij},Z)$ for any variable Z (from equation A.10), this bias can readily be shown to equal zero.

The expected difference between the experience coefficient in the augmented equation (A.15) model estimated using a cross section and $\beta_2$ is:

$$\text{(A.19)}\quad E(\hat{\beta}_2)-\beta_2=(1/DET_2)*$$

$$[-\{var(D_{ij})*cov(EXP_{ij},S_{ijt})-$$

$$cov(S_{ijt},D_{ij})*cov(EXP_{ij},D_{ij})*E(-S_{ijt}*\varepsilon_{ij}/Y)\ +$$

$$\{var(\ _{ijt})*var(D_{ij})-cov^2(S_{ijt},D_{ij})\}*E(-EXP_{ij}*\varepsilon_{ij}/Y)$$

$$-\{var(S_{ijt})*cov(EXP_{ij},D_{ij})-$$

$$cov(EXP_{ij},S_{ijt})*cov(S_{ijt},D_{ij})\}$$

$$*E(-D_{ij}*\varepsilon_{ij}/Y)].$$

Again noting that $E(-EXP_{ij}*\varepsilon_{ij})=0$ and that $cov(S_{ijt},Z)=1/2*cov(D_{ij},Z)$ for any variable Z, this bias expression simplifies to

$$\text{(A.20)}\quad E(\hat{\beta}_2)-\beta_2=(1/DET_2)*\{var(S_{ijt})-1/4*var(D_{ij})*$$

$$\alpha*var(EXP_{ij})*var(\varepsilon_{ij}).$$

This is still positive so long as $\alpha$ is positive, though straightforward comparison of equation (A.20) against equation (A.15) reveals that introducing $D_{ij}$ into the earnings equation reduces the bias in the experience coefficient. Note that the bias in $\beta_2$ falls to zero as $var(\varepsilon_{ij})$ falls to zero, that is, as $D_{ij}$ becomes a better proxy for $\mu_{ij}$.

Relaxing the cov(EXP,$\varepsilon$) = 0 Assumption

The preceding has assumed that pre-job experience affects completed job duration only through its association with job and/or match quality. If more experienced (older) workers are simply more stable than less experienced (younger) workers, then experience may have a direct positive relationship with completed job duration that is not reflected in higher earnings once experience is controlled for. An obvious generalization of the base model is thus to allow for the possibility that $cov(EXP_{ij},\varepsilon_{ij})$ might be positive. This affects several of the moments specified in equation (A.7) and equation (A.10). The following replace the corresponding expressions in equation (A.7):

$$\mathrm{var}(D_{ij}) = (\alpha\gamma)^2*\mathrm{var}(EXP_{ij}) + \gamma^2*\mathrm{var}(\phi_{ij}) +$$

$$\mathrm{var}(\varepsilon_{ij}) + 2\alpha\gamma*\mathrm{cov}(EXP_{ij},\varepsilon_{ij})$$

$$\text{(A.21)}\quad \mathrm{cov}(D_{ij},EXP_{ij}) = \alpha\gamma*\mathrm{var}(EXP_{ij}) + \mathrm{cov}(EXP_{ij},\varepsilon_{ij})$$

$$\mathrm{cov}(D_{ij},\varepsilon_{ij}) = \mathrm{var}(\varepsilon_{ij}) + \alpha\gamma*\mathrm{cov}(EXP_{ij},\varepsilon_{ij})$$

$$\mathrm{cov}(D_{ij},\mu_{ij}) = \alpha^2\gamma*\mathrm{var}(EXP_{ij}) + \gamma*\mathrm{var}(\phi_{ij}) + \alpha*\mathrm{cov}(EXP_{ij},\varepsilon_{ij})$$

It remains true that $\mathrm{cov}(S_{ijt},Z) = 1/2*\mathrm{cov}(D_{ij},Z)$ for all $Z$, so that the above replacements also affect $\mathrm{cov}(S_{ijt},D_{ij})$, $\mathrm{cov}(S_{ijt},EXP_{ij})$, $\mathrm{cov}(S_{ijt},\varepsilon_{ij})$ and $\mathrm{cov}(S_{ijt},\mu_{ij})$.

Now consider the standard cross section earnings equation:

$$\text{(A.22)}\quad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \xi_{ijt}.$$

The bias in the estimated seniority coefficient is again given by equation (A.12); although some of the moments in equation (A.12) take on different values when the assumption that $\mathrm{cov}(EXP_{ij},\varepsilon_{ij})=0$ is dropped, the bias in the estimated seniority coefficient still reduces to the expression in equation (A.13).

The expected difference between the estimated experience coefficient and $\beta_2$ is again given by equation (A.14). In the present case, this reduces to:

$$\text{(A.23)}\quad E(\beta_2)-\beta_2=(1/DET_1)*$$

$$[\alpha*\mathrm{var}(EXP_{ij})*\{\mathrm{var}(S_{ijt}) - 1/4*\mathrm{var}(D_{ij})\} +$$

$$(1/4\gamma)*\{\mathrm{cov}(EXP_{ij},D_{ij})*\mathrm{cov}(D_{ij},\varepsilon_{ij}) -$$

$$\mathrm{var}(D_{ij})*\mathrm{cov}(EXP_{ij},\varepsilon_{ij})\}$$

When $\mathrm{cov}(EXP_{ij},\varepsilon_{ij}) = 0$, this reduces to the expression in equation (A.15). Assuming $\alpha$ to be positive, a sufficient condition for this expression to be positive is that

$$\text{(A.24)}\quad \mathrm{cov}(D_{ij},\varepsilon_{ij})/\mathrm{var}(D_{ij}) > \mathrm{cov}(EXP_{ij},\varepsilon_{ij})/\mathrm{cov}(EXP_{ij},D_{ij})$$

Now consider the effect of introducing completed job duration into the earnings equation:

$$\text{(A.25)}\quad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \beta_3 D_{ij} + \omega_{ijt}$$

As in the base case, the resulting bias in the tenure coefficient is given by equation (A.18). Again, even though some of the moments in equation (A.18) take on different values when the assumption that $cov(EXP_{ij}, \varepsilon_{ij}) = 0$ is dropped, it can be shown that this expression equals zero. That is, introducing completed job duration into the earnings function eliminates the bias in the tenure coefficient.

The expected difference between the experience coefficient in an earnings function that also includes completed job duration as a control and $\beta_2$ is given by equation (A.19). This reduces to:

$$(A.26)\ E(\beta_2)-\beta_2=(1/\gamma*DET_2)*\{var(S_{ijt}) - (1/4)*var(D_{ij})\}* \{cov(EXP_{ij},D_{ij})*cov(D_{ij},\varepsilon_{ij})-var(D_{ij})*cov(EXP_{ij},\varepsilon_{ij})\}$$

This expression is positive so long as the condition stated in equation (A.24) above holds. However, if

$$(A.27)\ cov(_{Dij},\varepsilon_{ij})/var(D_{ij}) \leq cov(EXP_{ij},\varepsilon_{ij})/cov(EXP_{ij},D_{ij}),$$

then the experience coefficient in the earnings function that includes the measure of completed job duration will be downward biased. If the condition holds with equality, then including the measure of completed job duration eliminates the bias in $\beta_2$. It is straightforward to show that, so long as both $\alpha$ and $\gamma$ are positive, the experience coefficient in equation (A.23) (the standard cross section earnings function) is larger than that in equation (A.27) (the earnings function that includes completed job duration). Thus, under the condition stated in equation (A.24), introducing completed job duration into the earnings equation reduces but does not eliminate upward bias in the experience coefficient.

Overall, then, generalizing the base model to allow for a positive correlation between $EXP_{ij}$ and $\varepsilon_{ij}$ does not change the main conclusions of the analysis. The standard cross section estimate of the tenure coefficient $\beta_1$ is upward biased as before, and introducing $D_{ij}$ into the equation eliminates that bias. Under reasonable conditions (as given in equation (A.24)), the standard cross section estimate of the experience coefficient $\beta_2$ is also biased upwards as an estimate of the returns to experience _per se_ and introducing $D_{ij}$ into the earnings equation reduces but does not eliminate that bias.

Relaxing the Single Factor Assumption

A second potential generalization of the base model would be to relax the assumption that individual-, job-, and match-specific earnings components all have the same incremental association with job duration. Consider a model in which individual and job/match factors enter separately. In this more general specification, an earnings increment attributable to unobserved individual characteristics may bear a different relationship to job duration than one attributable to job/match characteristics. Equation (A.4) is rewritten as

$$\text{(A.28)} \quad D_{ij} = \gamma_1\mu_i + \gamma_2\theta_{ij} + \varepsilon_{ij},$$

where $\mu_i$ is the earnings increment attributable to unobserved individual differences, $\theta_{ij}$ is the earnings increment attributable to unobserved job/match characteristics, $\varepsilon_{ij}$ captures the variation in completed job duration that cannot be linked to the earnings advantage associated with either worker or job/match quality, i indexes individuals and j indexes jobs. The parameters $\gamma_1$ and $\gamma_2$ capture the relationships of individual and job/match characteristics, respectively, with job duration; both $\gamma_1$ and $\gamma_2$ are assumed to be positive. Assume further that $\mu_i$ and $\theta_{ij}$ are independent. Experience would affect job/match specific attributes but not (fixed) individual attributes, so equation (A.5) would be rewritten in terms of $\theta_{ij}$:

$$\text{(A.29)} \quad \theta_{ij} = \alpha EXP_{ij} + \phi_{ij}.$$

Substituting equation (A.29) into equation (A.28) yields

$$\text{(A.30)} \quad D_{ij} = \gamma_1\mu_1 + \gamma_2\alpha^*EXP_{ij} + \gamma_2\phi_{ij} + \varepsilon_{ij}.$$

Given these relationships, it follows that:

$$\mathrm{cov}(EXP_{ij},\mu_i) = 0,$$

$$\mathrm{cov}(EXP_{ij},\theta_{ij}) = \alpha*\mathrm{var}(EXP_{ij}),$$

$$\mathrm{var}(D_{ij}) = (\gamma_1)^2*\mathrm{var}(\mu_i) + (\alpha\gamma_2)^2*\mathrm{var}(EXP_{ij}) + (\gamma_2)^2*\mathrm{var}(\phi_{ij}) + \mathrm{var}(\varepsilon_{ij}),$$

$$\text{(A.31)}\quad \mathrm{cov}(D_{ij},EXP_{ij}) = \alpha\gamma_2*\mathrm{var}(EXP_{ij}),$$

$$\mathrm{cov}(D_{ij},\mu_i) = \gamma_1*\mathrm{var}(\mu_i),$$

$$\mathrm{cov}(D_{ij},\theta_{ij}) = \alpha^2\gamma_2*\mathrm{var}(EXP_{ij}) + \gamma_2*\mathrm{var}(\phi_{ij}),$$

$$\mathrm{cov}(D_{ij},\phi_{ij}) = \gamma_2*\mathrm{var}(\phi_{ij}), \text{ and}$$

$$\mathrm{cov}(D_{ij},\varepsilon_{ij}) = \mathrm{var}(\varepsilon_{ij}).$$

The relationship between $S_{ijt}$ and $D_{ij}$ is as before, so that $\mathrm{cov}(S_{ijt},Z) = 1/2*\mathrm{cov}(D_{ij},Z)$ for all Z.

Now consider the standard cross-section earnings equation:

$$\text{(A.32)}\quad \ln W_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \nu_{ijt}.$$

The error term $(\nu_{ijt})$ in this specification equals:

$$\text{(A.33)}\quad \nu_{ijt} = \mu_i + \theta_{ij} + \eta_{ijt},$$

with the error $\eta_{ijt}$ assumed to be orthogonal to S, EXP, $\mu$ and $\theta$. The bias in the estimated seniority coefficient equals

$$\text{(A.34)}\quad E(\beta_1)-\beta_1=(1/DET_1)*\{\mathrm{var}(EXP_{ij})*E(S_{ijt}*(\mu_i+\theta_{ij})) - \mathrm{cov}(EXP_{ij},S_{ijt})*E(EXP_{ij}*(\mu_i+\theta_{ij}))\}$$

which reduces to

$$\text{(A.35)}\quad E(\beta_1)-\beta_1=(1/DET_1)*1/2*\mathrm{var}(EXP_{ij})*[\gamma_1*\mathrm{var}(\mu_i)+\gamma_2*\mathrm{var}(\phi_{ij})].$$

So long as both $\gamma_1$ and $\gamma_2$ are positive, the seniority coefficient in the standard cross section model is biased upwards. The expected difference between the estimated experience coefficient and $\beta_2$ equals

$$\text{(A.36)}\quad E(\beta_2)-\beta_2=(1/DET_1)*\{-\mathrm{cov}(EXP_{ij},S_{ijt})*E(S_{ijt}*(\mu_i+\theta_{ij}))+ \mathrm{var}(S_{ijt})*E(EXP_{ij}*(\mu_i+\theta_{ij}))\}$$

which reduces to

(A.37) $E(\beta_2)-\beta_2=(1/DET_1)*\{\alpha*var(EXP_{ij})*[var(S_{ijt})-1/4*var(D_{ij})+$

$1/4*var(\varepsilon_{ij})+1/4*\ {}_{1}*(\gamma_1-\gamma_2)*var(\mu_i)]$

Assuming $\alpha$ to be positive, a sufficient condition for this expression to be positive is that

(A.38) $var(\varepsilon_{ij}) + \ {}_{1}(\gamma_1-\gamma_2)*var(\mu_i) > 0.$

This condition obviously holds whenever $\gamma_1 > \gamma_2$.

Now consider the effect of introducing completed job duration, $D_{ij}$, into the earnings function:

(A.39) $lnW_{ijt} = \beta_1 S_{ijt} + \beta_2 EXP_{ij} + \beta_3 D_{ij} + \omega_{ijt}.$

The error term in equation (A.39) equals

(A.40) $\omega_{ijt} = (1-\beta_3\gamma_1)*\mu_i + (1-\beta_3\gamma_2)*\theta_{ij} - \beta_3\varepsilon_{ij} + \eta_{ijt}.$

The bias in the estimated equation (A.39) seniority coefficient equals

(A.41) $E(\hat{\beta}_1)-\beta_1=(1/DET_2)*$

$[\{var(EXP_{ij})*var(D_{ij})-cov^2(EXP_{ij},D_{ij})\}*$

$E(S_{ijt}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})]) -$

$\{cov(EXP_{ij},S_{ijt})*var(D_{ij})-cov(EXP_{ij},D_{ij})*cov(S_{ijt},D_{ij})\}$

$*E(EXP_{ij}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})]) +$

$\{cov(EXP_{ij},S_{ijt})*cov(EXP_{ij},D_{ij})-$

$var(EXP_{ij})*cov(S_{ijt},D_{ij})\}$

$*E(D_{ij}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})])$

which can be shown to equal zero. The expected difference between the estimated equation (A.39) experience coefficient and $\beta_2$ equals

$$\text{(A.42)}\quad E(\hat{\beta}_2)-\beta_2=(1/DET_2)*[-\{var(D_{ij})*cov(EXP_{ij},S_{ijt})-cov(S_{ijt},D_{ij})*cov(EXP_{ij},D_{ij})\}*E(S_{ijt}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})])+\{var(EXP_{ij})*var(D_{ij})-cov^2(S_{ijt},D_{ij})\}*E(EXP_{ij}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})])-\{var(S_{ijt})*cov(EXP_{ij},D_{ij})-cov(EXP_{ij},S_{ijt})*cov(S_{ijt},D_{ij})\}*E(D_{ij}*[(1-\beta_3\gamma_1)*\mu_i+(1-\beta_3\gamma_2)*\theta_{ij}-\beta_3\varepsilon_{ij})])$$

After some manipulation, this reduces to

$$\text{(A.43)}\quad E(\hat{\beta}_2)-\beta_2=(1/DET_2)*\alpha*var(EXP_{ij})*[var(S_{ijt})-1/4*var(D_{ij})]*[var(\varepsilon_{ij})+\gamma_1(\gamma_1-\gamma_2)*var(\mu_i)]$$

which is positive so long as the condition specified in equation (A.38) holds. It is straightforward to show that, so long as both $\gamma_1$ and $\gamma_2$ are positive, the experience coefficient in equation (A.32), the standard cross section earnings function, is larger than the experience coefficient in equation (A.39), the earnings function that includes completed job duration as a control.

In sum, generalizing the base model to allow for separate individual and job/match effects does not affect the conclusions concerning bias in the seniority coefficient; the standard cross section estimate of $\hat{\beta}_1$ is still biased upward, and this bias is eliminated by including a measure of completed job duration in the regression. With regard to the estimated effect of experience, as long as the condition given in equation (A.38) is satisfied, the basic results concerning the experience coefficient follow: the experience coefficient in the standard cross section earnings function is upward biased as an estimate of the returns to experience _per se_ and introducing completed job duration into the model reduces but does not eliminate this upward bias.

Appendix B
Table 1:
Regional and Industrial Distribution of Jobs Samples Subgroups[a]

| | Prof.Tech, Managerial, Nonunion | | Blue Collar Nonunion | | All Nonunion | |
|---|---|---|---|---|---|---|
| | Complete | Censored | Complete | Censored | Complete | Censored |
| Proportion by Region: | | | | | | |
| Pacific | .129 | .132 | .128 | .102 | .125 | .122 |
| Mountain | .0670 | .0558 | .0453 | .0536 | .0494 | .0527 |
| West North Central | .0893 | .102 | .0906 | .115 | .0903 | .103 |
| East North Central | .201 | .216 | .157 | .148 | .171 | .180 |
| West South Central | .0844 | .0853 | .157 | .139 | .130 | .108 |
| East South Central | .0323 | .0481 | .0775 | .0944 | .0666 | .0746 |
| South Atlantic | .156 | .126 | .191 | .189 | .185 | .162 |
| Middle Atlantic | .169 | .157 | .120 | .114 | .140 | .138 |
| New England | .0720 | .0775 | .0334 | .0446 | .0423 | .0595 |
| Proportion by Industry: | | | | | | |
| Ag, For, Fish | .0124 | .0109 | .0584 | .0561 | .0380 | .0329 |
| Mining | .00248 | .0140 | .0107 | .0204 | .00716 | .0151 |
| Dur Goods Man | .139 | .211 | .222 | .236 | .181 | .203 |
| NonDur Goods Man | .0397 | .0775 | .106 | .111 | .0788 | .0903 |
| Construction | .0496 | .0481 | .181 | .162 | .122 | .105 |
| Trans, Comm, Util | .0645 | .0713 | .0822 | .116 | .0731 | .0931 |
| Trade | .238 | .163 | .169 | .120 | .221 | .172 |
| Fin, Ins, RE | .0918 | .0651 | .0155 | .00765 | .0638 | .0527 |
| Services | .362 | .340 | .155 | .170 | .214 | .235 |
| Number of observations | 403 | 645 | 839 | 784 | 1396 | 1461 |

[a]The other characteristics of the samples are summarized in Table 1.

Appendix B

Table 2:

Additional Coefficients from Final Tenure Models[a]

| | (1) Prof,Tech Managerial Nonunion | (2) Blue Collar Nonunion | (3) All Nonunion |
|---|---|---|---|
| Y (Inverse Baseline Hazard, $\lambda = e^{-ZY}$) | | | |
| Construction (yes = 1) | -.0513 | -.297 | -.277 |
| | (.4401) | (.121) | (.116) |
| Durable manufacturin (yes = 1) | .192 | -.122 | -.106 |
| | (.385) | (.119) | (.112) |
| Nondurable manufacturing (yes = 1) | .499 | -.150 | -.00623 |
| | (.439) | (.137) | (.12966) |
| Transportation, communications, utilities (yes = 1) | -.191 | -.0187 | -.0786 |
| | (.408) | (.1395) | (.1272) |
| Trade (yes = 1) | -.420 | -.353 | -.414 |
| | (.379) | (.122) | (.112) |
| Finance, insurance, real estate (yes = 1) | -.329 | -1.372 | -.420 |
| | (.400) | (.204) | (.144) |
| Services (yes = 1) | -.175 | -.176 | -.279 |
| | (.369) | (.124) | (.112) |
| Mountain (yes = 1) | -.647 | .408 | .150 |
| | (.211) | (.133) | (.106) |
| West north central (yes = 1) | .124 | .632 | .351 |
| | (.195) | (.119) | (.096) |
| East north central (yes = 1) | .0689 | .397 | .216 |
| | (.1647) | (.093) | (.078) |
| West south central (yes = 1) | -.0760 | .305 | .119 |
| | (.2045) | (.092) | (.082) |
| East south central (yes = 1) | .588 | .539 | .358 |
| | (.344) | (.119) | (.102) |
| South atlantic (yes = 1) | -.00783 | .483 | .231 |
| | (.17182) | (.096) | (.080) |
| Middle atlantic (yes = 1) | .00822 | .401 | .168 |
| | (.17452) | (.105) | (.085) |
| New England (yes = 1) | .0380 | .461 | .242 |
| | (.2190) | (.164) | (.124) |

[a]The experience, education, race, marital status, disability status, and occupational coefficients from these models are reported in Table 2 in the

Appendix B

Table 3a:

Additional Coefficients from ln (average hourly earnings) Models

Managerial and Professional Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of education | 14.90 [ 2.05] | .0775 (.0034) | .0736 (.0036) | .0743 (.0034) |
| Nonwhite (yes = 1) | .025 [.157] | -.113 (.029) | -.0946 (.0403) | -.0955 (.0404) |
| Married (yes = 1) | .900 [.300] | .0640 (.0206) | .0519 (.0216) | .0517 (.0217) |
| Disabled (yes = 1) | .044 [.206] | -.0791 (.0299) | -.0723 (.0285) | -.0727 (.0286) |
| Manager (yes = 1) | .463 [.499] | .0778 (.0145) | .0794 (.0145) | .0787 (.0144) |
| Construction (yes = 1) | .036 [.187] | .142 (.055) | .153 (.0635) | .156 (.0632) |
| Durable manufacturing (yes = 1) | .231 [.421] | .143 (.046) | .154 (.0520) | .154 (.0517) |
| Nondurable manufacturing (yes = 1) | .081 [.273] | .162 (.049) | .160 (.0546) | .167 (.0540) |
| Transportation, communication, utilities (yes=1) | .061 [.238] | .0113 (.0509) | .0227 (.0581) | .0245 (.0576) |
| Trade (yes = 1) | .152 [.359] | -.0650 (.0479) | -.0367 (.0554) | -.0384 (.0550) |
| Finance, insurance, real estate (yes = 1) | .070 [.256] | .116 (.050) | .137 (.0579) | .138 (.0574) |
| Services (yes = 1) | .350 [.477] | -.0383 (.0457) | -.0278 (.0522) | -.0284 (.0519) |
| Mountain (yes = 1) | .040 [.197] | -.146 (.035) | -.122 (.0322) | -.122 (.0321) |
| West north central (yes = 1) | .103 [.304] | -.159 (.026) | -.161 (.0267) | -.161 (.0265) |
| East north central (yes = 1) | .241 [.428] | -.0265 (.0219) | -.0308 (.0234) | -.0320 (.0233) |
| West south central (yes = 1) | .078 [.269] | -.230 (.028) | -.231 (.0300) | -.230 (.0299) |
| East south central (yes = 1) | .052 [.223] | -.164 (.032) | -.173 (.0327) | -.168 (.0328) |

Appendix B - Table 3a: (cont.)

| | | | | |
|---|---|---|---|---|
| South atlantic (yes = 1) | .124 [.329] | -.0676 (.0249) | -.0716 (.0287) | -.0690 (.0286) |
| Middle atlantic (yes = 1) | .180 [.384] | .0213 (.0229) | .0217 (.0252) | .0214 (.0251) |
| New England (yes = 1) | .063 [.243] | -.0823 (.0302) | -.0787 (.0310) | -.0753 (.0310) |
| 1969 (yes = 1) | .059 [.235] | .0667 (.0359) | .0735 (.0365) | .0702 (.0365) |
| 1970 (yes = 1) | .058 [.234] | .0891 (.0360) | .0973 (.0351) | .0923 (.0351) |
| 1971 (yes = 1) | .061 [.238] | .0410 (.0357) | .0538 (.0343) | .0498 (.0342) |
| 1972 (yes = 1) | .067 [.251] | .0538 (.0349) | .0655 (.0352) | .0606 (.0353) |
| 1973 (yes = 1) | .072 [.258] | .0517 (.0345) | .0646 (.0339) | .0611 (.0340) |
| 1974 (yes = 1) | .074 [.261] | .0377 (.0343) | .0500 (.0345) | .0471 (.0345) |
| 1975 (yes = 1) | .085 [.278] | .0215 (.0334) | .0302 (.0319) | .0272 (.0319) |
| 1976 (yes = 1) | .083 [.276] | .0154 (.0336) | .0250 (.0329) | .0217 (.0330) |
| 1977 (yes = 1) | .092 [.289] | .0374 (.0329) | .0479 (.0321) | .0463 (.0322) |
| 1978 (yes = 1) | .093 [.290] | .0752 (.0329) | .0797 (.0316) | .0772 (.0316) |
| 1979 (yes = 1) | .097 [.297] | .0845 (.0326) | .0860 (.0323) | .0823 (.0324) |
| 1980 (yes = 1) | .105 [.306] | .0309 (.0322) | .0342 (.0325) | .0290 (.0326) |
| Constant | 1.000 [.000] | .0707 (.0797) | .0189 (.0847) | .0206 (.0856) |

[a] The experience, seniority and completed job duration coefficients for these models are reported in Table 4a of the text.

Appendix B

Table 3b:

Additional Coefficients from ln (average hourly earnings) Models

Blue Collar Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of education | 11.02 [ 2.55] | .0449 (.0028) | .0436 (.0029) | .0427 (.0029) |
| Nonwhite (yes = 1) | .131 [.337] | -.119 (.019) | -.116 (.0203) | -.113 (.0202) |
| Married (yes = 1) | .891 [.311] | .132 (.020) | .109 (.0214) | .106 (.0214) |
| Disabled (yes = 1) | .091 [.287] | -.0854 (.0212) | -.0855 (.0221) | -.0820 (.0221) |
| Foreman or craft worker (yes = 1) | .506 [.500] | .203 (.013) | .196 (.0135) | .201 (.0135) |
| Construction (yes = 1) | .146 [.353] | .165 (.028) | .182 (.0321) | .178 (.0319) |
| Durable manufacturing (yes = 1) | .263 [.440] | .279 (.026) | .286 (.0300) | .285 (.0297) |
| Nondurable manufa turing (yes = 1) | .108 [.311] | .243 (.029) | .251 (.0335) | .250 (.0333) |
| Transportation, communication, utilities (yes=1) | .103 [.304] | .307 (.030) | .310 (.0366) | .313 (.0363) |
| Trade (yes = 1) | .120 [.325] | .101 (.029) | .121 (.0331) | .120 (.0329) |
| Finance, insurance, real estate (yes = 1) | .009 [.093] | -.0257 (.0681) | .00882 (.0566) | .0254 (.0549) |
| Services (yes = ) | .176 [.381] | .132 (.027) | .139 (.0316) | .136 (.0313) |
| Mountain (yes = 1) | .051 [.221] | -.042 (.033) | -.0511 (.0294) | -.0462 (.0294) |
| West north cent al (yes = 1) | .094 [.292] | -.195 (.028) | -.214 (.0281) | -.199 (.0280) |
| East no th central (yes = 1) | .155 [.362] | -.0522 (.0248) | -.0718 (.0222) | -.0635 (.0221) |
| West south central (yes = 1) | .141 [.348] | -.195 (.025) | -.210 (.0244) | -.202 (.0241) |
| East south central (yes = 1) | .100 [.299] | -.235 (.028) | -.262 (.0278) | -.247 (.0278) |

Appendix B - Table 3b: (cont.)

| | | | | |
|---|---|---|---|---|
| South atlantic (yes = 1) | .219 | -.196 | -.217 | -.203 |
| | [.414] | (.024) | (.0229) | (.0230) |
| Middle atlantic (yes = 1) | .101 | -.117 | -.128 | -.124 |
| | [.301] | (.027) | (.0246) | (.0246) |
| New England (yes = 1) | .040 | -.0352 | -.0458 | -.0303 |
| | [.197] | (.0356) | (.0302) | (.0300) |
| 1969 (yes = 1) | .051 | .0325 | .0304 | .0280 |
| | [.219] | (.0373) | (.0363) | (.0361) |
| 1970 (yes = 1) | .058 | .0598 | .0599 | .0558 |
| | [.234] | (.0361) | (.0357) | (.0358) |
| 1971 (yes = 1) | .056 | .0397 | .0413 | .0355 |
| | [.230] | (.0364) | (.0359) | (.0356) |
| 1972 (yes = 1) | .063 | .0479 | .0534 | .0467 |
| | [.243] | (.0355) | (.0365) | (.0363) |
| 1973 (yes = 1) | .076 | .105 | .114 | .108 |
| | [.265] | (.034) | (.0348) | (.0347) |
| 1974 (yes = 1) | .071 | .0867 | .0925 | .0869 |
| | [.257] | (.0347) | (.0351) | (.0350) |
| 1975 (yes = 1) | .077 | .0259 | .0338 | .0268 |
| | [.267] | (.0342) | (.0351) | (.0350) |
| 1976 (yes = 1) | .086 | .0837 | .0950 | .0849 |
| | [.280] | (.0336) | (.0358) | (.0358) |
| 1977 (yes = 1) | .093 | .112 | .123 | .116 |
| | [.291] | (.033) | (.0345) | (.0344) |
| 1978 (yes = 1) | .102 | .132 | .146 | .138 |
| | [.302] | (.033) | (.0348) | (.0347) |
| 1979 (yes = 1) | .111 | .102 | .118 | .110 |
| | [.314] | (.032) | (.0350) | (.0350) |
| 1980 (yes = 1) | .105 | .0942 | .107 | .101 |
| | [.307] | (.0325) | (.0353) | (.0352) |
| Constant | 1.000 | .197 | .208 | .188 |
| | [.000] | (.054) | (.0588) | (.0559) |

[a] The experience, seniority and completed job duration coefficients for these models are reported in Table 4b of the text.

Appendix B

Table 3c:

Additional Coefficients from ln (average hourly earnings) Models All Occupations Nonunion Sample[a]

| | Mean [s.d.] | (1) | (2) | (3) |
|---|---|---|---|---|
| Years of education | 12.96 [ 2.92] | .0629 (.0020) | .0610 (.0020) | .0612 (.0020) |
| Nonwhite (yes = 1) | .077 [.267] | -.131 (.017) | -.128 (.0173) | -.127 (.0172) |
| Married (yes = 1) | .893 [.309] | .101 (.014) | .0834 (.0144) | .0847 (.0144) |
| Disabled (yes = 1) | .069 [.253] | -.0705 (.0165) | -.0627 (.0177) | -.0610 (.0177) |
| Manager (yes = 1) | .192 [.394] | .0619 (.0137) | .0673 (.0135) | .0645 (.0134) |
| Clerical (yes = 1) | .133 [.340] | -.0680 (.0159) | -.0679 (.0162) | -.0712 (.0162) |
| Foreman, craft worker (yes = 1) | .229 [.420] | -.0729 (.0147) | -.0604 (.0134) | -.0652 (.0134) |
| Operative, laborer (yes = 1) | .223 [.416] | -.260 (.016) | -.240 (.0162) | -.246 (.0162) |
| Construction (yes = 1) | .084 [.278] | .142 (.025) | .160 (.0281) | .159 (.0281) |
| Durable manufacturing (yes = 1) | .230 [.421] | .212 (.023) | .223 (.0259) | .222 (.0258) |
| Nondurable manufa turing (yes = 1) | .095 [.293] | .197 (.025) | .200 (.0280) | .202 (.0279) |
| Transportation, communication, utilities (yes=1) | .084 [.278] | .195 (.025) | .201 (.0298) | .202 (.0297) |
| Trade (yes = 1) | .167 [.373] | .0589 (.0236) | .0803 (.0275) | .0794 (.0274) |
| Finance, insurance, real estate (yes = 1) | .058 [.234] | .147 (.028) | .168 (.0318) | .167 (.0317) |
| Services (yes = ) | .239 [.427] | .0464 (.0229) | .0549 (.0265) | .0538 (.0264) |
| Mountain (yes = 1) | .046 [.210] | -.0730 (.0228) | -.0731 (.0208) | -.0718 (.0208) |
| West north cent al (yes = 1) | .099 [.298] | -.163 (.018) | -.169 (.0183) | -.165 (.0184) |

Appendix B - Table 3c: (cont.)

| | | | | |
|---|---|---|---|---|
| West south central (yes = 1) | .110 [.313] | -.174 (.018) | -.180 (.0178) | -.179 (.0177) |
| East south central (yes = 1) | .075 [.263] | -.186 (.020) | -.199 (.0198) | -.195 (.0198) |
| South atlantic (yes = 1) | .173 [.378] | -.130 (.016) | -.138 (.0164) | -.134 (.0164) |
| Middle atlantic (yes = 1) | .143 [.350] | -.00947 (.01653) | -.0130 (.0166) | -.0122 (.0166) |
| New England (yes = 1) | .049 [.216] | -.0431 (.0223) | -.0488 (.0213) | -.0434 (.0213) |
| 1969 (yes = 1) | .054 [.227] | .0492 (.0250) | .0495 (.0241) | .0496 (.0241) |
| 1970 (yes = 1) | .058 [.234] | .0674 (.0246) | .0694 (.0237) | .0671 (.0237) |
| 1971 (yes = 1) | .059 [.236] | .0285 (.0245) | .0306 (.0234) | .0281 (.0234) |
| 1972 (yes = 1) | .066 [.249] | .0489 (.0239) | .0536 (.0236) | .0505 (.0236) |
| 1973 (yes = 1) | .074 [.263] | .0732 (.0233) | .0806 (.0229) | .0781 (.0229) |
| 1974 (yes = 1) | .073 [.260] | .0604 (.0234) | .0654 (.0230) | .0637 (.0231) |
| 1975 (yes = 1) | .080 [.272] | .0262 (.0230) | .0283 (.0221) | .0268 (.0222) |
| 1976 (yes = 1) | .085 [.278] | .0530 (.0228) | .0565 (.0228) | .0537 (.0228) |
| 1977 (yes = 1) | .092 [.289] | .0690 (.0225) | .0741 (.0221) | .0731 (.0221) |
| 1978 (yes = 1) | .097 [.297] | .108 (.022) | .111 (.0222) | .110 (.0223) |
| 1979 (yes = 1) | .104 [.305] | .0971 (.0220) | .0995 (.0224) | .0975 (.0224) |
| 1980 (yes = 1) | .104 [.305] | .0652 (.0220) | .0668 (.0227) | .0651 (.0227) |
| Constant | 1.000 [.000] | .226 (.046) | .185 (.0470) | .189 (.0474) |

[a]The experience, seniority and completed job duration coefficients in these models are reported in Table 4c of the text.