The Wage Premium for Job Seniority During the 1980s and Early 1990s

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In this article I use Panel Study of Income Dynamics (PSID) data to study whether the rewards for job tenure have fallen since the early 1980s. An upward-sloping seniority-wage profile is generally thought to be an important dimension of compensation during a career. However, recent interest in the incidence of layoffs and changes in the structure of firms gives rise to the possibility that employment contracts are less often being structured to reward tenure or that such contracts are more difficult to honor. Using a two-stage estimator to attempt to control unmeasured individual and job match effects on wages, I find some evidence that the wage premium paid to senior workers has declined moderately. However, I find that these results are mildly sensitive to alternative methods of handling the relatively noisy PSID tenure data.

UNTIL RECENTLY, A RISING SENIORITY-WAGE profile was a readily accepted empirical finding and conventionally assumed to be a primary determinant of compensation in the U.S. labor market. Beginning in the mid-1980s, however, a number of researchers suggested that the positive relationship between wages and tenure at one firm is simply a statistical artifact brought about by higher wages paid at all points in a career to those workers who eventually attain high levels of seniority. It is the purpose of this article to determine whether skepticism about the magnitude of the slope of the seniority-wage profile ought rightly to have a second dimension. Not only may the seniority-wage profile at any point in time

be flatter than previously thought, but there also is reason to believe that the premium workers earn for additional years of job tenure may be falling over time.

Because the relationship between seniority and wages is integral to our understanding of compensation and employment, determining whether the wage premium for job tenure is falling over time is important. Sizable tenure-wage premia are theorized to be due to returns on joint investments in training by workers and firms or to benefits workers receive for engaging in long-term contracts with their employers. Falling tenure premia would imply that wage growth over a career will depend less on these arrangements between workers and firms and more on workers' own efforts to continually develop marketable productive abilities.

Using Panel Study of Income Dynamics (PSID) data for 18- to 60-year-old men, I estimate the degree to which the tenure wage profile may have changed between 1981 and 1992. I use variation in within-job tenure to help identify the effect of tenure on wages. Unlike direct measures of tenure across jobs, deviations from within-job tenure means are uncorrelated with unmeasurable fixed individual and job-match effects on wages. I base this method on work by Altonji and Shakotko (1987) and find some evidence of a decline in the value of seniority during the 1980s and early 1990s.

Background

Because of their perceived empirical regularity, a number of important theories of employer-employee contracting have been developed to explain why we might observe senior workers earning more than otherwise comparable junior workers. Becker (1962) and Mincer (1974) attribute the above-market earnings for senior workers to returns these workers earn on investments they and their employers made in the development of firm-specific skills early in their tenure. These investments mean that senior workers earn more because they are more productive.

Medoff and Abraham (1980) provide evidence that it is not always the case that senior workers are more productive, even when they earn more. Accordingly, a second class of models has shown that the earnings differential between junior and senior workers may have nothing to do with productivity differences. Rather, some firms may structure their earnings payments to reward seniority so as to provide incentives for junior workers (Lazear, 1981) or to boost morale.

However, a variety of evidence suggests that employment relationships may be changing in a way that undermines seniority-rewarding contracts.

First, on different levels, researchers have identified an increase in the use of short-term contracting, raising the possibility that firms and workers are becoming less willing to engage in the type of long-term contracts that could pay seniority premia. For example, Belous (1989) employs Current Population Survey (CPS) data to show that the proportion of workers in short-term employment contracts is rising. Abraham (1990) provided survey evidence that human resource directors increasingly prefer such short-term arrangements.

Second, recent evidence suggests that the stability of the employment relationship has declined, at least for certain groups of workers. Swinnerton and Wial (1995), using the CPS data, and Rose (1995) and Boisjoly, Duncan, and Smeeding (1994), using the Panel Study of Income Dynamics data, present evidence suggesting that jobs have become less stable overall. Similarly, Monks and Pizer (1996) and Bernhardt et al. (1997) find a general increase in job turnover between the early 1970s and early 1990s using National Longitudinal Surveys (NLS) data. Together these findings are consistent with a decline in job stability that predated the recession of the early 1990s, which generated much popular interest in the topic.

Furthermore, Farber (1993), using the CPS Dislocated Workers Surveys, finds that involuntary job separations were more common during the last recession than during the recession of the early 1980s for college-educated workers, the group whose jobs traditionally have been most stable. Farber (1997a) also finds that the overall rate of job loss has increased in the 1990s. Finally, Farber (1997b) has found recently that the fraction of workers reporting job tenure of 10 or 20 years fell during the 1990s, suggesting that increases in job loss and instability may be affecting the prevalence of long-term jobs.

Using the CPS data, Diebold, Neumark, and Polsky (1997) have presented evidence that questions the extent to which declines in job stability have been pervasive. However, as Marcotte (1995) reviews, while important aspects of the findings of Diebold, Neumark, and Polsky are at odds with those of Swinnerton and Wial (1995) and researchers who have used the PSID and NLS, all researchers have found evidence consistent with declining job stability for certain groups of workers. In particular, there is broad agreement that black workers generally and workers without any college education have seen declines in job stability.

Certainly no consensus has been reached on the importance or extent of any general decline in job stability. Nonetheless the popular accounts of mass layoffs and declining commitment between firms and workers that gave rise to the nascent literature on job stability trends, as well as the

empirical results presented in that literature, have important bearing on our understanding of seniority premia. The popular press has devoted substantial ink to descriptions of the labor market as one with little job security and a greater need for workers to devote themselves to their skills or occupations, not their employers.[1] Just as these stories raised testable hypotheses about job security trends, they also raise empirical questions about patterns of compensation within firms if long-term commitments from firms or workers are eroding. Similarly, the emerging pattern of diminished job stability for certain groups of workers raises questions about whether the premium for job seniority for these groups has declined.

Employment contracts marked by significant seniority premia necessarily involve long-term commitments on the part of both workers and firms. Whether seniority premia are due to returns on investments in human capital or to deferred compensation schemes, both firms and workers must agree to and honor relatively long-term contracts that permit training investments to be recouped or enable workers to collect deferred earnings. A decline in the stability of the employment relationship, either for certain groups of workers or more generally, raises the possibility that firms and workers are less often engaging in or honoring contracts that pay supramarket earnings to senior workers.

This article examines the question of whether the past decade and a half have witnessed an erosion in the value of job seniority. It follows on substantial interest in the empirical study of tenure returns during the 1980s. This interest focused on the possibility that observed seniority premia are due to unobserved attributes of senior workers or their jobs rather than to their seniority itself.

Beginning in the mid-1980s, a series of challenges arose to the idea that observed premia paid to more senior workers were due to their seniority. These challenges, by Abraham and Farber (1987), Altonji and Shakotko (1987), and Topel (1986), suggested that workers with more seniority were more able workers or in better job matches. As a result, these workers earn more at every stage of their careers. Because of their characteristics and/or good job matches, these highly paid workers stay in a firm longer and are disproportionately represented among the most senior workers. To the researcher, then, what appears to be a positive return to tenure could be payment for unobserved individual ability or a worker's

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[1] The *New York Times* series, "The Downsizing of America" (March 3–9, 1996), and the *Time* magazine cover story, "Whatever Happened to the Great American Job" (November 22, 1993), are perhaps the best known of a very large number of the former stories. *Fortune* magazine's, "Getting Past Insecurity" (April 17, 1995), is one of the many examples of the latter.

relatively high productivity in a job for which he or she is well suited. More recently, however, Topel (1991) has provided evidence that returns to job seniority may not be insignificant.

In this article I build on these empirical studies. I do so not to help identify the size of the seniority premium per se but to investigate whether there has been an erosion of the reward for job tenure over time. Below I will briefly describe the methods and data I use to examine this question.

Methods and Data

Basic model. As a first step in describing the methods used to study changes in tenure rewards, I will illustrate the nature of the problem of identifying the degree to which higher wages paid to senior workers are due to their job tenure. Consider the following wage equation:

$$\ln w_{ijt} = X_{ijt}\beta_0 + T_{ijt}\beta_1 + T_{ijt}^2\beta_2 + OLDJOB_{ijt}\beta_3 + t\beta_4 + v_{ijt} \qquad (1)$$

where

$$v_{ijt} = u_i + q_{ij} + e_{ijt}$$

T_{ijt}, T_{ijt}^2, and $OLDJOB_{ijt}$ each measure some aspect of a worker's current job tenure. T_{ijt} and T_{ijt}^2 measure worker i's tenure in job j at time t in quadratic form. $OLDJOB_{ijt}$ is a dummy variable equaling 1 if worker i has been in job j for more than 1 year at time t. I include this to allow for the possibility that very junior workers may agree to a brief period of low wages, consistent with an initial probationary or training period. X_{ijt} is a vector of worker and job characteristics thought to influence wages. The measure of time t is intended to pick up any independent change in real wages over time. β_0 is a conformable vector of coefficients that relates X_{ijt} to $\ln w_{ijt}$, the log of real wages, and β_1, β_2, β_3, and β_4 are scalars mapping the various aspects of job tenure and time to $\ln w_{ijt}$.

The problem on which the literature has focused is that v_{ijt} is not orthogonal to measures of job tenure. The error term v_{ijt} contains the individual fixed effect u_i and the job-match component q_{ij}, as well as a transitory component e_{ijt} assumed to be orthogonal to all regressors and error components. As long as more able workers are more likely to be retained in a job from year to year and/or workers with better job matches are more likely to keep their jobs, then the covariance between the various measures of job tenure and v_{ijt} will

be positive. As a result, if Eq. (1) is estimated via ordinary least squares (OLS), estimates of the effects of job tenure on wages will be biased upward.

Instrumental variables (IV) approach. A number of methods have been employed to generate a consistent estimate of the effects of job tenure on wages. A straightforward approach is to find instruments for the measure of tenure. Appropriate instruments would vary with tenure but be uncorrelated with associated individual and job-match components of the residual. If such instruments can be found, here represented by T_{ijt}^{IV}, $T_{ijt}^{2\,\text{IV}}$, and $OLDJOB_{ijt}^{\text{IV}}$, we can estimate the following equation:

$$\ln w_{ijt} = X_{ijt}\beta_0 + T_{ijt}^{\text{IV}}\beta_1 + T_{ijt}^{2\,\text{IV}}\beta_2 + OLDJOB_{ijt}^{\text{IV}}\beta_3 + t\beta_4 + v_{ijt} \qquad (2)$$

Suitable instruments would be orthogonal to v_{ijt}, and the IV estimates of β_1, β_2, and β_3 would be consistent estimators of the effects of the various aspects of job tenure on wages.

It is this method, or variants of it, that has been used to challenge the conventional notion that wages rise with seniority. This is also the method that will be employed here to examine if consistent estimates of tenure returns have been falling over time.

As always, the difficulty in employing the instrumental variables method is finding appropriate instruments that are both correlated with the explanatory variables of interest and uncorrelated with nonrandom unobservables. Fortunately, the job is made less difficult because of the previous interest in examining the importance of worker heterogeneity and job-match quality in generating apparent seniority premia. A straightforward approach was employed by Altonji and Shakotko (1987), who used deviations from observed within-job means as their principal instrument for measures of job tenure:

$$\widetilde{T}_{ijt} = T_{ijt} - \overline{T}_{ij}$$
$$\widetilde{T}_{ijt}^2 = T_{ijt}^2 - \overline{T}_{ij}^2 \qquad (3)$$
$$\widetilde{OLDJOB}_{ijt} = OLDJOB_{ijt} - \overline{OLDJOB}_{ij}$$

where $\overline{T}_{ij}$ is the mean tenure level for worker i in job j over the period in which the job is observed, $\overline{T}_{ij}^2$ is the mean of the square of job tenure for individual i in job j, and $\overline{OLDJOB}_{ij}$ is the mean value of the dummy variable indicating whether individual i had held job j for more than 1 year.

The IV procedure is a 2SLS estimator. Here, X_{ijt} and $\widetilde{T}_{ijt}$, $\widetilde{T}_{ijt}^2$ and $\widetilde{OLDJOB}_{ijt}$ are used as instruments for T_{ijt}, T_{ijt}^2, and $OLDJOB_{ijt}$. As

Altonji and Shakotko make clear, these deviations from within-job means serve as useful instruments for the job tenure measures here because each sums to zero for any worker regardless of the actual level of tenure during the period of observation.[2] As a result, they are uncorrelated with the fixed individual and job-match components of the error term v_{ijt}. Further, because both T_{ijt}, T_{ijt}^2, and $OLDJOB_{ijt}$ and the deviations from their within-job means increase as worker i remains in job j, the instruments employed by Altonji and Shakotko are correlated with the aspects of job tenure of interest here.[3]

I use this instrumental variables method as the basis of my test of whether the value of job tenure has fallen during the 1980s and early 1990s. To do this, I first estimate β_1^{IV}, β_2^{IV}, and β_3^{IV} for the entire 1981–1992 period. This provides a baseline estimate of the parameters of the seniority-wage profile during the entire period.

What is of interest here, however, is whether the seniority-wage profile is changing over time, not time-invariant estimates of that profile. This issue can be addressed by generalizing the relationship described in Eq. (2) to explicitly account for the possibility that the value of job tenure may vary over time:

$$\ln w_{ijt} = X_{ijt}\beta_0 + T_{ijt}^{IV}\beta_1 + T_{ijt}^{2\,IV}\beta_2 + OLDJOB_{ijt}^{IV}\beta_3 +$$
$$t\beta_4 + \left(t * T_{ijt}^{IV}\right)\beta_5 + \left(t * T_{ijt}^{2\,IV}\right)\beta_6 + \left(t * OLDJOB_{ijt}^{IV}\right)\beta_7 + v_{ijt} \tag{4}$$

In this parameterization, interactions with the job-tenure instruments and the measure of time t are included on the right-hand side. These interaction terms allow the effect of job tenure to vary over time. Testing whether there have been substantial changes in the effects of job tenure on wages during the period considered here thus is simplified to a test of significance on the coefficients of these interaction terms.

PSID sample. I implement this test using PSID data, which allow for observations on workers' wages and relevant productive and job

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[2] Consider as an example observing two workers during a 3-year period with 10, 11, and 12 years and 1, 2, and 3 years of tenure on their jobs, respectively. For each worker, $\widetilde{T}_{ijt}$ will be -1, 0, and 1 in the first, second, and third years of the period during which we observe them. As a result, $\widetilde{T}_{ijt}$ takes on similar values for high-tenure and low-tenure workers and is uncorrelated with the fixed individual and job-match components of the error term v_{ijt}.

[3] It should be noted that the instrumental variables method employed here may not purge all correlation between the error term and all regressors. For example, there may be correlation between experience and job-match effects if workers sort into better jobs as their time in the labor market increases.

characteristics. Since 1968, the PSID has collected data on over 5000 U.S. families. The PSID data are particularly appealing here because they are the data employed in many of the previously cited empirical studies of tenure returns. There is an important difference in the use of the PSID data here, though. In previous studies, tenure returns were estimated using PSID waves beginning in 1969. I use PSID waves beginning in 1981, discarding earlier data. I made this restriction because of changes in the wording of the tenure question and in the survey guidelines determining which sample members were asked about their tenure prior to 1981.[4] Since 1981, a consistent question about job tenure has been asked of the same set of PSID respondents.

For the purpose of this article, I restrict the sample used here to 18- to 60-year-old male workers who were household heads and were not self-employed.[5] I further restrict the sample to include only those who reported positive tenure and wages and who reported working fewer than 4000 hours in any calendar year. Finally, because the PSID oversamples low-income households, all estimates presented here are based on the application of sample weights. Descriptive statistics of the resulting sample are presented in Table 1.

Data issues and consistency. Use of the PSID data to study job tenure is complicated by problems with consistency over time. In an important paper on the subject, Brown and Light (1992) illustrate that in many instances, changes in PSID respondents' reported job tenure between interviews do not accord with changes in calendar time. These changes in reported job tenure over time take a variety of forms. Some are consistent with job change, and some are not. For example, reported tenure may be 50 months in year t and 2 months in year $t + 1$. Or reported tenure may be 12 months in year t and 36 months in year $t + 1$. In the first case, this

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[4] First, prior to 1976, respondents were asked about the length of time in their current positions. Beginning in 1976, respondents were asked about the length of time with their current employer. Respondents may have interpreted the word *position* in the earlier question to mean a particular job with their current employer. So we could observe workers recently promoted within their firm providing very low values of "tenure." When the Current Population Survey made a similar switch in the wording of their tenure question between 1981 and 1983, the mean level of reported tenure increased by more than 1 full year. This would have been a remarkable shift in mean job tenure over the course of a 2- year period and likely reflects differences in the way respondents answer the two questions. Second, in 1978, only sample members under the age of 45 were asked about their tenure with their current employer. In 1979 and 1980, the job-tenure question was omitted from the PSID survey.

[5] The sample was restricted to household heads because all relevant information is available only for them. Women were excluded because married women usually were not considered household heads. So selecting on household heads resulted in a female sample that was young and unmarried and not likely to be representative of the experiences of all working women.

TABLE 1

DESCRIPTIVE STATISTICS FOR THE
PSID SAMPLE

Variable	
Median hourly wage	12.21
Median age (years)	35
Median tenure (years)	5.5
Proportion nonwhite	0.118
	0.002
Proportion HS dropout	0.168
	0.002
Proportion HS degree	0.229
	0.002
Proportion 1–3 years college	0.337
	0.003
Proportion college degree	0.266
	0.003
Proportion in argiculture	0.015
	0.001
Proportion in const./mining	0.087
	0.002
Proportion in manufacturing	0.301
	0.003
Proportion in trade	0.169
	0.002
Proportion in transp./util.	0.102
	0.002
Proportion in fire	0.037
	0.001
Proportion in bus. services	0.042
	0.001
Proportion in pers. services	0.022
	0.001
Proportion in prof. services	0.136
	0.002
Proportion in pub. admin.	0.089
	0.002
	N = 30,281

Note: Wages in 1992 dollars; standard deviations in italics
below.

change in reported tenure between interview years may be consistent with
a change of employers between interviews, or it may be reporting error. In
the second case, the change in reported tenure may be the result of round-
ing error, reporting error, or perhaps the possibility that the respondent

changed employers and returned to work for an employer with which he had previously accrued tenure.[6]

Because of inconsistencies in tenure responses over time, it can become difficult to determine when one job ends and another job begins and consequently to determine within-job tenure for the purposes of constructing instruments. Brown and Light devote considerable attention to these problems of inconsistency and of partitioning tenure responses between jobs. They point out that there are a variety of ways to partition tenure responses into different jobs and to deal with these inconsistencies. Each of the methods they examine is intuitively reasonable. While Brown and Light illustrate that some partitioning methods are preferred, their principal conclusion is that researchers using these data need to examine whether key empirical findings are sensitive to the choice of partitioning method. Because of this, I make use of a number of different methods to partition jobs and attempt to ensure consistency of the tenure data.

In order to employ the instrumental variables method, I first adopt the partitioning method that Brown and Light point to as the best among a number of alternatives.[7] This partitioning method assumes that a new job has begun whenever reported tenure is less than elapsed time since the last interview (less 1 month to limit rounding error). Like Brown and Light, I refer to this partitioning method as T. I view the results based on the T partition as the preferred IV estimates.

However, as a check on the sensitivity of the estimates obtained from the preferred partitioning, I also employ a different method to partition tenure responses between different jobs. Instead of partitioning based on absolute *levels* of reported job tenure, a new job can be inferred whenever *changes* in reported tenure between two different years are sufficiently different from changes in calendar time that it appears that an individual held two different jobs in those years. Thus, whenever tenure *increases* by more than a threshold number of months in excess of the elapsed calendar time between interviews, or whenever tenure *decreases* by more than that threshold number of months below the change in elapsed calendar time, I assume that there has been a job change of some sort under this alternative

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[6] It should be noted that time between interviews varies moderately from individual to individual. This fact has been accounted for in the analysis presented here but is implicitly ignored for this simple example.

[7] Because employer codes are available on National Longitudinal Survey data, Brown and Light check the accuracy of different partitioning methods for those data by comparing the resulting jobs to changes in employer codes. Doing so reveals that the partition method used as a baseline (their T partition) here is superior to all other alternatives, with a lower sum of type I and type II errors than any other partitioning method. Without employer codes, checking the accuracy of partitioning methods with PSID cannot be done with certainty. With this caveat, Brown and Light suggest that the T partitioning method is an "acceptable method for assigning observations to jobs" (p. 255).

partitioning. For my purposes, I employ three different thresholds values. First, I assume a new job is observed whenever the absolute value of the change in reported tenure between 2 years is greater than 6 months more than the change in calendar time between interviews. Next, I change the threshold values to 12 months and then to 18 months. Because of its relative symmetry, I refer to this partitioning method as S.[8] Below I report estimates based on both the T and S partitioning methods.

As a second step to determine the sensitivity of the IV estimates based on the preferred T partitioning method to how the relatively noisy PSID tenure data are handled, I consider alternative methods to control for inconsistency in tenure responses. First, I temporarily drop from consideration all information from sample individuals who report tenure changes that are inconsistent with calendar time. In a second, less restrictive consistency check, I eliminate from consideration tenure information for an individual only during years in which some threshold of inconsistency appears to have been crossed but keep information for other years. And last, I make no restrictions at all to limit tenure responses that are inconsistent with calendar time.

For these purposes, I define as inconsistent those tenure responses in one year (t) that exceed tenure in the previous year ($t - 1$) by at least 13 months *more* than the change in calendar time between interviews. The first consistency check would then drop all data for all years for individuals who report such inconsistent tenure information. The second consistency check would not use data from year t in estimating values for the instruments discussed earlier. However, later years' consistent data would be included (and assumed to be part of another job).[9]

Results

In Table 2 I present the results of the OLS and IV estimated wage equations for the entire 1981–1992 period, with no interaction terms included.

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[8] Brown and Light refer to this partitioning method by the number of months allowable between differences in change in reported tenure and change in calendar time, such as 0, 6, or 12.

[9] Note that the consistency checks are not applied to the S partitioning methods. The S partition infers that job change occurs whenever changes in job tenure between interviews do not match up with changes in time. The consistency checks employed along with the T partition infer that large disparities of reported tenure change and time are more likely due to reporting error, and a change into a new job is inferred only when tenure in any one year is very low in absolute terms. This reflects an implicit difference in the assumptions behind the S partition and the consistency checks employed with the T partition. The S partition treats large differences between time and reported changes in job tenure as consistent with job change. The consistency checks treat such differences as noise and are more cautious about inferring that a job change has occurred.

TABLE 2
### OLS and IV Estimated Wage Equations: Full Period, with No Trends

Variable	OLS	IV
Intercept	1.783*	1.749*
	(0.027)	(0.030)
Potential experience	0.0296*	0.0329*
	(0.002)	(0.003)
Potential experience squared	−0.0006*	−0.0007*
	(0.00005)	(0.00006)
Nonwhite	−0.0637*	−0.0606*
	(0.014)	(0.015)
H.S. dropout	−0.0968*	−0.0957*
	(0.014)	(0.016)
Some college	0.4638*	0.4784*
	(0.024)	(0.029)
College degree	0.5923*	0.5910*
	(0.012)	(0.013)
Residence in Midwest	0.0650*	0.0641*
	(0.011)	(0.012)
Residence in Northwest	0.0952*	0.0922*
	(0.012)	(0.012)
Residence in West	0.1529*	0.1547*
	(0.012)	(0.013)
Agriculture ind.	−0.2339*	−0.2292*
	(0.044)	(0.047)
Bus. services ind.	0.0685*	0.0713*
	(0.023)	(0.024)
Const./mining ind.	0.1802*	0.1760*
	(0.018)	(0.019)
Fire ind.	0.1258*	0.1285*
	(0.023)	(0.025)
Manufacturing ind.	0.1658*	0.1670*
	(0.013)	(0.015)
Pers. services ind.	−0.2153*	−0.2103*
	(0.032)	(0.034)
Prof. services ind.	−0.1155*	−0.1057*
	(0.015)	(0.017)
Public admin. ind.	0.0692*	0.0818*
	(0.017)	(0.019)
Transp./util. ind.	0.2346*	0.2370*
	(0.016)	(0.018)
Tenure	0.0136*	0.0088*
	(0.0008)	(0.002)
Tenure squared	−0.0001*	−0.00001*
	(0.00001)	(0.00003)
OLDJOB	0.0561*	0.0994*
	(0.014)	(0.017)
Time	−0.0061*	−0.0059*
	(0.001)	(0.001)
R-squared	0.4472	0.3813

Note: Tenure is measured in 6-month intervals. *Significant at 5 percent level.

These results suggest that fixed individual and job-match effects account for a portion of the observed effect of tenure on wages. The OLS estimates suggest that wages increase significantly with job tenure. In fact, the coefficients on each of the job-tenure parameters are significant and of the expected sign. The IV estimates, on the other hand, suggest a somewhat flatter seniority wage profile, consistent with the notion that at least part of the observed increase in wages with tenure is attributable to unobserved individual and/or job match characteristics.

In Figure 1 the different seniority-wage profiles implied by the OLS and IV estimates are illustrated more clearly. Each is upward sloping and mildly concave.[10] Notice that while each seniority-wage profile rises significantly, the IV-estimated seniority-wage profile is flatter, its rise more mild than the OLS-estimated profile.

These somewhat different seniority-wage profiles serve as the point of departure for the remaining analysis. Below I present the results of the various analyses to consider whether these patterns observed for the entire

FIGURE 1
OLS and IV Estimated Seniority-Wage Profiles: 1981–1992 Period

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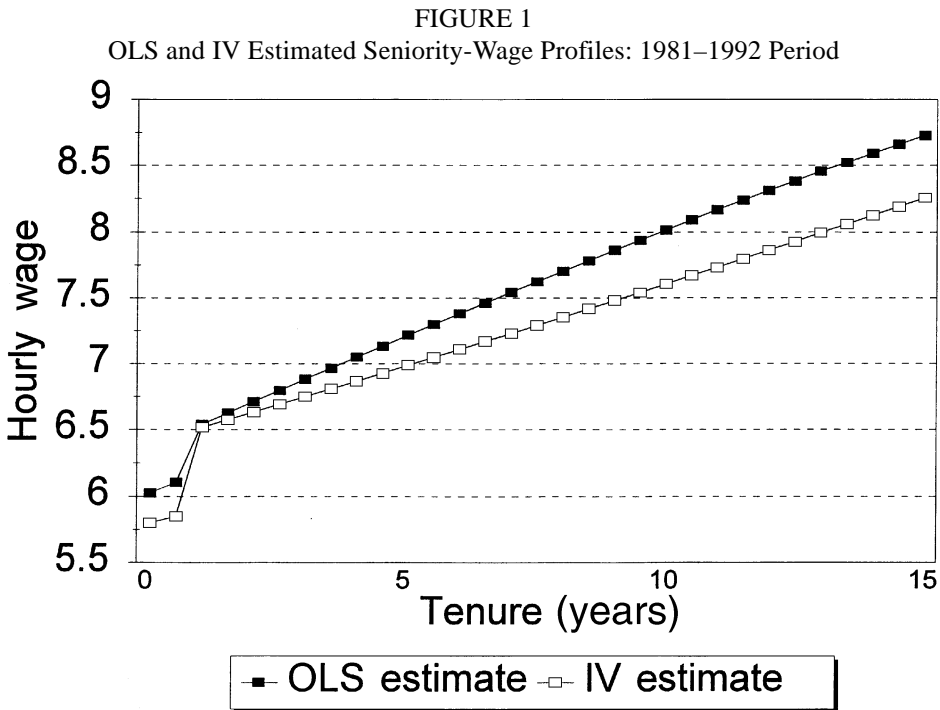
[10] For the purposes of this figure only, I consider the seniority-wage profile of a high school graduate white worker as he enters the labor market.

1980s to early 1990s period may in fact be masking a downward trend in the reward for job seniority within the period.

In Table 3 I present the results of the OLS and IV estimations of the basic model described in Eq. (4). This model employs the first check to limit inconsistent reporting of tenure, and the IV estimation is based on the T partitioning. Like the results presented in Table 2, the OLS estimates of the effects of an additional year of job tenure on wages are positive and significant, and workers in their first year with an employer earn significantly less than workers on the job more than 1 year. Moreover, the OLS results suggest that there have been no significant changes in the relationship between tenure and wages over time: None of the interactions between time and the tenure parameters are significant.

However, the IV results presented in Table 3 suggest a different pattern than that identified by the OLS results. The estimated coefficients on tenure, tenure squared, and *OLDJOB* suggest that at the beginning of the period, wages rose significantly with job tenure. However, the coefficients on the time and tenure interaction terms suggest that the relationship between tenure and wages has itself changed over time. The negative coefficient on the interaction of time and tenure suggests that the direct marginal effect of tenure on wages fell over the period. At the same time, the relationship between tenure and wages appears to have lost some of its concave shape: The coefficients interacting time with tenure-squared and *OLDJOB* suggest that each of these parameters of the tenure-wage relationship lost some of their power over the period.

Figure 2 illustrates the net effect of these changes on the relationship between tenure and wages. The figure presents the seniority-wage profiles implied by the IV coefficients from Table 3 at the beginning and end of the 1981–1992 period. The seniority-wage profile estimated to prevail by the end of the period rises more gradually than the profile estimated for the beginning of the period, consistent with an interesting decline in seniority returns during the period.

While these results provide some evidence that the relationship between tenure and wages may have been changing during this period, they should appropriately be viewed tentatively. As discussed earlier, responses to interview questions about individuals' tenure with their employers are subject to well-known error. If we are to conclude that the wage premium for job tenure did indeed decline during the 1980s and early 1990s, we need to consider the extent to which the results presented in Table 3 are sensitive to alternative methods for drawing inference about employer change from the tenure data and different methods for limiting inconsistency in the responses to questions about tenure.
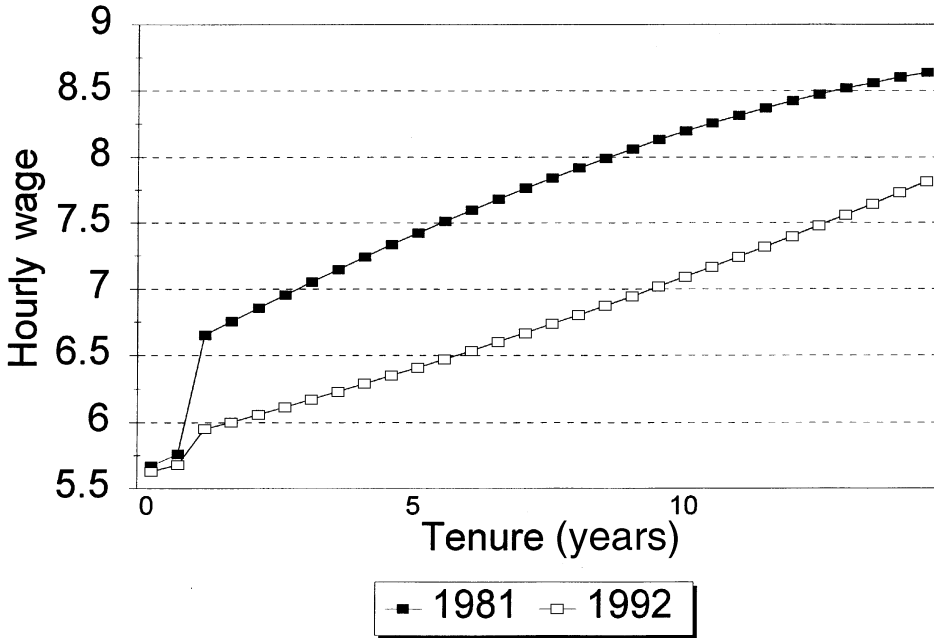
TABLE 3

OLS- AND IV-ESTIMATED WAGE EQUATIONS: FULL PERIOD, WITH TRENDS

Variable	OLS	IV
Intercept	1.773**	1.719**
	(0.031)	(0.041)
Potential experience	0.0298**	0.0302**
	(0.002)	(0.003)
Pot. exp. squared	−0.0006**	−0.0006**
	(0.00005)	(0.00006)
Nonwhite	−0.0649**	−0.0660**
	(0.014)	(0.015)
H.S. dropout	−0.0949**	−0.0887**
	(0.014)	(0.016)
Some college	0.4664**	0.4727**
	(0.024)	(0.029)
College degree	0.5936**	0.5979**
	(0.012)	(0.013)
Residence in Midwest	0.0644**	0.0638**
	(0.011)	(0.012)
Residence in Northwest	0.0947**	0.0953**
	(0.012)	(0.012)
Residence in West	0.1531**	0.1568**
	(0.012)	(0.013)
Agriculture ind.	−0.2324**	−0.2228**
	(0.044)	(0.047)
Bus. services ind.	0.0698*	0.0731**
	(0.023)	(0.024)
Const./mining ind.	0.1804**	0.1754**
	(0.018)	(0.019)
Fire ind.	0.1256*	0.1277*
	(0.023)	(0.025)
Manufacturing ind.	0.1657**	0.1688*
	(0.013)	(0.015)
Pers. services ind.	−0.2154**	−0.2080**
	(0.032)	(0.034)
Prof. services ind.	−0.1164**	−0.1153**
	(0.015)	(0.017)
Public admin. ind.	0.0697**	0.0721**
	(0.017)	(0.019)
Transp./util. ind.	0.2351**	0.2367**
	(0.016)	(0.018)
Tenure	0.0133**	0.0177**
	(0.002)	(0.003)
Tenure squared	−0.0001**	−0.0002**
	(0.00002)	(0.00005)
OLDJOB	0.0803**	0.1371**
	(0.027)	(0.033)
Time	−0.0048*	6E-07
	(0.003)	(0.005)
Tenure × time	−3E-07	−0.0009**
	(0.0002)	(0.0004)
Tenure squared × time	3E-067	2.6E-05**
	(0.00004)	(0.000006)
OLDJOB × time	−0.0038	−0.0090**
	(0.004)	(0.005)
R-squared	0.4476	0.3853

Note: Tenure is measured in 6-month intervals. *Significant at 10 percent level. **Significant at 5 percent level.

FIGURE 2
IV-Estimated Seniority-Wage Profiles, 1981 and 1992

In Table 4 I present the OLS- and IV-estimated effects of job tenure on wages based on the different methods of limiting inconsistency and partitioning jobs described earlier. The first three columns present the OLS-estimated effects of the principal dimensions of job tenure on wages interacted with the time trend. The coefficients reported in each of the columns are based on different consistency checks. The last six columns present IV estimates based on different partitioning methods.[11] All these results come from estimation of the same model used to derive the results presented in Table 3. I report the coefficients on the tenure and time variables here to conserve space and focus on the key variables.

Using either or none of the checks to ensure consistency, the OLS results suggest no significant decline in the value of job tenure during the period. Quite differently, IV estimates suggest important changes in the relationship between tenure and wages that are largely consistent with a decline in real tenure returns. While these results seem to be mildly sensitive to the manner in which the tenure data are handled, this pattern is generally consistent across the IV results.

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[11] Note that tenure responses are not partitioned into different jobs for the purposes of OLS estimation.

Under the preferred T partitioning method, whether or not consistency is imposed, the IV estimates provide evidence consistent with a decline in tenure returns. The results based on the S partitioning method are somewhat sensitive to the threshold chosen for partitioning tenure responses between jobs, with a significant decline in the marginal effect of tenure on wages observed for only two of the three threshold values. However, taken as a group, these results suggest that the IV-estimated decline in seniority premia presented in Table 3 also is observed when alternative methods of handling the PSID tenure data are employed.

This decline in the IV-estimated wage advantage associated with tenure, compared with the OLS-estimated trend, is itself interesting. One explanation consistent with the growing difference between the two estimates is the possibility that the value of unobserved ability or job-match quality grew during the period. This occurrence would not be unexpected if indeed there has been a decline in job stability. If firms are less willing or able to offer long-term employment to their workers, they would likely be more selective in deciding which workers to offer long-term contracts or in which cases to honor such contracts. As a result, the senior workers observed in later years in the period may be even more heavily comprised of those with high levels of ability or in good job matches, and the OLS and IV estimates would increasingly differ as the correlation between unmeasured components of the OLS error term and measures of job tenure increased.

Reconsideration of the instrument used here

Because a consistent estimate of tenure returns is a matter of some debate, a brief comment is needed here about the use of the instrumental variables method proposed by Altonji and Shakotko for the purpose of evaluating changes in tenure returns over time. Topel (1991) has presented estimates of seniority returns that challenge what has been the developing consensus that such returns are quite small. Moreover, Topel makes the case that Altonji and Shakotko's methods produce an inconsistent estimate of the returns to job tenure. In light of his review, I will discuss how the points raised in Topel's critique impede inference for the purposes of this article.

First, Topel points out the problems that can occur due to measurement error in recorded job tenure. His main point refers to biases that arise due to the use of tenure recorded in interval form. This point does not apply to the current results, since they are based only on the waves of the PSID for which job tenure is recorded not in intervals but in actual months. Topel's

TABLE 4

ESTIMATED EFFECTS OF JOB TENURE ON WAGES UNDER ALTERNATIVE METHODS FOR PARTITIONING JOBS AND CHECKING DATA CONSISTENCY

| | OLS Estimates | | | IV Estimates | | | | | |
| | | | | T Partition | | | S Partition | | |
	First Consistency Check	Second Consistency Check	No Consistency Checks	First Consistency Check	Second Consistency Check	No Consistency Checks	6-Month Threshold	12-Month Threshold	18-Month Threshhold
Tenure	0.0133**	0.0134**	0.0122**	0.0177**	0.0138**	0.0122**	0.0133**	0.0133**	0.0143**
Tenure sqaured	−0.0001**	−0.0001**	−0.0001**	−0.0002**	−0.0002**	−0.0002**	−0.0001**	−0.0002**	−0.0002**
OLDJOB	0.0803**	0.0752**	0.0722**	0.1371**	0.1485**	0.1478**	0.1391**	0.1851**	0.1929**
Tenure × time	−3E-07	−3E-05	7E-05	−0.0009**	−0.0005*	−0.0009**	−0.0005*	−0.0004	−0.007**
Tenure squared × time	3E-06	4E-06**	2E-06**	2.6E-05**	1.75E-05**	2.7E-05**	1.4E-05**	2E-05**	2E-05**
OLDJOB × time	−0.0038	−0.0017	−0.0027	−0.009*	−0.0053*	−0.0038	0.0019	−0.004	−0.003
Time	−0.0048	−0.0059**	−0.0058**	−6E-07	−0.0032	−0.0034	−0.008**	−0.007**	−0.007**
R-squared	0.4476	0.4439	0.4300	0.3853	0.3765	0.3714	0.3697	0.3753	0.3797

Notes: *Significant at 10 percent level. **Significant at 5 percent level.

point applies to more general consistency problems. I attempt to circumvent these problems by checking the sensitivity of the results presented here to alternative methods for checking data consistency.

Topel also points out that Altonji and Shakotko's instrument relies on the assumption that the job effect on wages (q_{ij}) is fixed. If the job-match effect is time-variant, the Altonji and Shakotko instrument will yield inconsistent estimates of the impact of job tenure on wages. To see the problem Topel points out, reconsider Eq. (2). T_{ijt}^{IV}, $T_{ijt}^{2\,\text{IV}}$, and $OLDJOB_{ijt}^{\text{IV}}$ will yield consistent estimates of the effects of job tenure on wages only so long as they are uncorrelated with the individual and job-match elements of the residual, u_i and q_{ij}. If the job-match effect varies over time, then estimates of the impact of tenure on wages will be biased upward or downward depending on whether the match evolves or devolves.

However, even in this case, the Altonji and Shakotko instrument is useful for the present article. What I am exploring is the possibility that tenure premia themselves are falling over time. For a time-variant job effect to impede inference for my purposes, the nature of that time variation in the job effect would itself have to vary over time. Specifically, if the observed decline in job tenure premia were due to *changes* in the time variation of the job-match effect on wages, the significance of the job-match component of wages would have to be falling during the period. I know of no reason to believe that this has occurred. However, the observed pattern in tenure returns may be overstated if indeed such a change in the job-match effect on wages has occurred.

Implications and conclusions

The empirical results presented here provide the first evidence that seniority-wage premia may have become less important during the 1980s and early 1990s. While these results are novel, they are consistent with a widespread popular perception that workers and firms are in some ways less committed to one another than in the past. Such perceptions are fueled by media attention to large layoffs accompanying firms' downsizing strategies, as well as stories suggesting that successful professional careers are requiring a commitment to a profession (or a knowledge base) and not to a company. If either of these popular stories is in anyway accurate, the economic bases for rewarding seniority would be weakened.

To the extent that the results presented in this article represent any trend toward lower seniority premia, a number of important implications would result. Wage growth over a career is the result of rewards workers receive for their experience in the labor market, as well as their tenure within a

firm. Lower tenure premia would mean that wage growth throughout a career would depend more heavily on the returns earned on general labor market experience. Such returns are the result of continuing productivity enhancement throughout the career. As a consequence, workers who are ill-prepared for a career that involves continued general skills development rather than specific, employer-provided training may face careers marked by stagnant wage growth.

Unfortunately, workers' abilities to make investments in their own productivity are likely to be uneven. Liquidity constraints for the poorest workers and bad previous experiences with schooling for the lowest-educated workers have played a part in making continuing education largely an avenue for relatively well-off and educated workers to enhance their skills (Martin, 1993; Watts and Boss, 1987). Altonji and Spletzer (1991) and Lillard and Tan (1986) also provide evidence that education and wages are correlated with other forms of private investment in skill development. If wage growth depends more on such investments, the wage differences between more and less educated workers could continue to grow.

While the results of this article are suggestive of important changes in the employment relationship, they should appropriately be viewed with some reservation for at least two reasons. First, survey data on job tenure are notoriously noisy. While the conclusions drawn here hold for the preferred method of handling the PSID tenure data, they are somewhat sensitive to alternative methods. Second, the recent literature on the economic value of job tenure has not settled on the best way to control for the biases caused by unmeasurable individual and job-match effects. More confidence about the extent to which the present results accurately reflect a trend in the U.S. labor market will require similar analyses with different data and different methods. Clearly, however, continued attention to such questions will be important as we attempt to understand whether and how the employment relationship may be changing.

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