Do Earnings Increase with Job Seniority?

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Abstract:

Cross-sectional wage regressions overstate the extent to which earnings increase with job seniority because they fail to take account of the sorting which occurs when high wage workers have lower rates of mobility. The main source of bias is a negative correlation between turnover probabilities and (unobserved) market valued individual characteristics which are transferable across firms. These results argue for the importance of theories which emphasize generally applicable individual differences and against those which focus on firm-specific attributes.

Article:

I. INTRODUCTION

Upwards sloping earnings profiles are a primary feature of most labor market theories. Wages rise with seniority in the human capital model (Becker, 1975), since job tenure is correlated with acquisitions of firm- specific skills, as well as in a number of more recently developed explanations. Lazear (1981) argues that rising wage profiles represent a bond which shirking workers risk losing. Similarly, deferred compensation reduces adverse selection in the testing models of Guasch and Weiss (1980, 1982) and might also insure against breach of implicit contracts in periods of above average demand.

These theories require wages to increase with job duration and draw support from the stylized results obtained from cross-sectional earnings regressions. However, such findings need not imply that compensation increases with seniority. Alternatively, *high paying jobs could last longer* than other employment. Sorting is then likely to result in cross-sectional wage regressions which overstate seniority earnings differentials.

There is good reason to expect an inverse relationship between wages and turnover rates. The matching models of Burdett (1978) and Jovanovic (1979) predict high pay and low mobility in jobs providing good "matches" and low earnings and higher turnover when the match is poor. Voluntary mobility may also be reduced in efficiency wages models, where firms pay high wages to raise effort and reduce turnover.[1] The relationship will be more pronounced if, in the presence of training or turnover costs, employers offer premiums to workers with low quit propensities. In each case, sorting will create a positive correlation between wages and seniority.

Recent empirical work confirms that tenure coefficients in cross-sectional wage regressions are seriously biased. Mincer and Jovanovic (1981) find that partial controls for previous mobility

***1.** See Yellen (1984) for a survey of this literature.*

reduce the estimated returns to tenure by up to 40%. Altonji and Shakotko (1987), using an instrumental variable technique measuring the difference between current and average tenure *within a given job*, show an almost 80% reduction in seniority wage premiums and Abraham and Farber (1987) find a similar decrease when the estimated *completed* duration of jobs is included in cross-sectional wage regressions.[2]

This paper investigates whether seniority premiums observed in cross-sectional data result from job-specific attributes or from a sorting process where superior workers simultaneously receive high wages and have low rates of turnover. It does so by examining whether involuntarily terminated employees are able to carry preseparation seniority differentials to their next job.[3] To the extent the premiums are transferred, they represent attributes valued by a number of employers. Conversely, if firm-specific attachments are of greater importance, previous job duration should be uncorrelated (or weakly correlated) with future wages.

Regardless of work history, most displaced individuals are required to search for new employment while jobless. Controlling for characteristics observable by potential employers, we therefore expect them to receive approximately the same distribution of new and wage and job offers. Unlike economically motivated quits, there is little reason for persons with good previous job matches to obtain better than average matches in their first post-layoff employment.[4]

II. SENIORITY AND WAGES

Standard cross-sectional earnings regressions take the form

$$W_{j,t} = Bs_{j,t} + e_{j,t}, \qquad (1)$$

where the subscripts *t* and *j* refer to the time period and job, *s* represents seniority (with all other observable characteristics suppressed to clarify exposition), and *e* is the regression error term. The disturbance term can be decomposed into an individual effect (*f*) which influences wages and is transferable across employers, a job-specific match effect (m_j), and a white noise error ($u_{j,t}$). Thus $e_{j,t} = f + m_j + u_{j,t}$.

The seniority coefficient estimated from (1) will be unbiased only if tenure is uncorrelated with both match and individual quality ($E(f|s) = E(m|s) = 0$). Conversely, if expected values of either variable increase with job duration, the seniority coefficient will be upwards biased. To see this let $\phi(\cdot)$ and $\omega(\cdot)$ indicate the expected value functions of *f* and *m*. In the simple case where $E(f|s) = \phi(s) = \phi s$ and $E(m|s) = \omega(s) = \omega s$, $\widehat{ß} = ß + \phi + \omega$. Since ϕ and ω may vary across firms, the bias in the regression coefficient captures an average of the aggregated firm effects. Employers

__2.__ Lang (1987) criticizes this research.

__3.__ Kletzer (1989) has recently used information on permanently terminated workers from the Displaced Workers Supplements to the Current Population Survey to investigate some of the questions focused upon in this paper.

__4.__ A weak relationship may persist to the extent that longer tenure workers receive longer prenotification of layoffs or greater reemployment assistance.

frequently terminating workers may have different expectation functions from those with low rates of turnover, thus the cross-sectional (pre-displacement) seniority premiums of laidoff workers may differ from those of job stayers.

Next consider individuals in their first employment following a permanent layoff. Using the subscripts *r* and *p* to denote the pre- and post-separation job, respectively, and suppressing the time subscript, post displacement wages are

$$w_p = Bs_p + f + m_P + u_P.$$

Since $E(f\,|s_r) = \phi(s_r)$, provides an unbiased estimate of 4, in regressions of

$$w_p = Bs_p + Cs_r + e_P, \tag{2}$$

if e_P is uncorrelated with s_r. This condition is fulfilled if m_P and u_p are orthogonal to s_r, in which case $\hat{C}$ indicates the extent to which market valued transferable characteristics are associated with previous tenure. If seniority differentials resulted exclusively from firm-specific acquisitions, $\hat{C}$ would equal zero.

u_p is a random disturbance and is therefore uncorrelated with prior seniority. Postseparation match quality will be positively correlated with previous tenure following *voluntary* turnover, however, since workers quit jobs (for economic reasons) only when superior offers are received. To reduce this correlation, I focus on involuntarily displaced employees, whose reemployment (and match quality) will be more random.

Even if matches occur randomly following permanent layoffs, some workers procure good matches (and high wages) and so will have relatively low subsequent turnover rates. Persons with highly valued transferable attributes may also be relatively immobile and obtain substantial seniority on the new job. As a result, both *m* and *f* will be positively correlated with s_p causing an upwards bias in $\hat{B}$.

Unbiased estimates of $\hat{B}$ can be obtained by replacing current seniority with an instrumental variable which is correlated with tenure but orthogonal to match and individual quality. I use an instrument (s_p*) which measures the deviation between current job duration (s_p) and average *within-job* seniority (rip).[5] This instrumental variable was developed and first used by Altonji and Shakotko (1987). Because it measures changes in seniority within a given job, both match and individual quality are held constant by definition. s_p* is therefore uncorrelated with the regression error term and $\hat{B}$ and $\hat{C}$, respectively, show the true returns to seniority and to transferable market valued individual characteristics which are correlated with previous tenure.

5. *Thus,* $s_p^* = s_p - \bar{s}_p$. *For example, if the post-layoff job continues for five survey periods, average within job tenure is 2.5 years and the value of the instrument in years one through five respectively is — 2, —1, 0, 1, and 2. Similar instruments can be constructed for other forms of the tenure variable. For example*$(s_p^2)^* = s_p^2 - (\overline{s_p^2})$.

Care is needed in generalizing the results described below because the seniority profiles of involuntary job leavers may differ from those of immobile workers. For example, layoffs may be concentrated among individuals with relatively poor matches and small investments in specific human capital, leading to relatively flat pre- separation wage profiles. This concern is partially mitigated by evidence that the seniority differentials of displaced workers are actually *steeper* than for their job stayer counterparts.

III. DATA AND ANALYSIS

This paper analyzes data on male heads of households from the 1969-1980 waves of the Panel Study of Income Dynamics (PSID).[6] Pooled data are used for individuals employed on their first jobs subsequent to involuntary (permanent) layoffs occurring between 1969 and 1975. Persons over the age of 55 or under 21, at the time of displacement, are excluded from the sample, as are those remaining out of the labor force for all of the 2 years following the calendar year of the termination. This sample includes 1373 individuals and 5559 person-year observations. For comparison purposes, a sample containing both laid-off and non-displaced workers was also constructed from pooled 1969-1975 data.

The dependent variable is the natural log of real weekly wages. Most regressors are specified in fairly standard ways. Data for actual work experience became available in 1974. For earlier years, experience was calculated as 1974 experience plus the difference between the survey year and 1974 (and bounded to be non-negative). Because seniority is available only as a categorical variable prior to 1976, dummy variables were created for 5 tenure groups—less than 1, 1 to 3, 3 to 9, 10 to 19, and 20 or more years seniority. Layoffs are classified as permanent if the worker fails to return to the original employer within 2 calendar years. The standard two stage "Heckman correction" for reemployment selection bias was used to provide consistent estimates in the post-separation wage regressions.[7]

Selected sample means are presented in columns (1) and (2) of table 1. Displaced males are younger, less experienced, less educated and have lower tenure than their full sample counterparts. They are more likely to be nonwhite or unmarried, more frequently work in blue collar jobs (72.1 vs. 54.4%) and receive lower weekly wages ($141 vs. $185) than the full sample: Post-separation earnings are approximately 16% higher than pre-layoff wages. This does not imply that involuntary terminations raise wages, however, since the compensation of other workers may be growing at a faster rate.[8]

Columns (3) and (4) show the results of period zero (preseparation) earnings regressions for the full sample and subsample of workers displaced in year one. The coefficients *TEN2* through

6. *Women are excluded because of the small number in the survey and because the process generating tenure premiums is likely to be substantially different than for men. Future work focusing on females is needed.*

7. *See Heckman (1979) for details. Wage observations are missing, due to sustained unemployment, for 8.8% (535 out of 6095) of the sample. The regressions were also estimated without correcting for selection bias. This did not lead to important changes in the coefficients of interest.*

8. *Recent work examining the wage consequences of mobility (i.e. Mincer 1986, Ruhm 1987a) finds greater gains following quits than layoffs and, controlling for the reason for turnover, for low than high tenure workers. Ruhm also uncovers important gender differences.*

TENS in column (3) indicate seniority differentials of job stayers. Wage premiums, over new job entrants, range from 10.7%, for workers with 1 to 3 years tenure, to a maximum of 20.7% after 10 to 19 years seniority.

LA Y1 through *LAYS* show how the preseparation wages of displaced workers compare to those of job stayers with equivalent seniority. The negative coefficients indicate lower pay. While the shortfalls are statistically significant and range between 10% and 13% for persons with less than a decade on the job, no significant differences exist for individuals with 10 or more years tenure. This suggests that firms initially use layoffs to eliminate less desirable workers, with senior employees being terminated more randomly. It also implies that period zero tenure profiles will be steeper for involuntary job leavers than for other workers. This can be seen by comparing the coefficients *TEN2* through *TENS* in column (4) with their counterparts in column (3). The maximum seniority premium is 20.7% for job stayers but almost 27% in the layoff subsample.

Columns (5) and (6) present estimates of the post-layoff wage regressions. *TEN2* through *TENS* indicate the correlation between *pre-displacement* seniority and *post-displacement* earnings and show the extent to which preseparation tenure premiums result from low turnover rates among persons with transferable market valued characteristics. *DUR* and *DURSQ* indicate how wages change with seniority in the immediately following employment. The coefficients in column (5) apply to instrumental variables measuring *within- job* deviations in seniority and are of primary interest. For comparison purposes, column (6) presents estimates of OLS regressions using actual post-separation tenure. If the post- separation seniority variable is of reasonably good quality and differences in unobserved job match or individual quality remain, after controlling for predisplacement seniority, we expect the OLS tenure coefficients to exceed the IV estimates.

Seniority premiums obtained prior to displacement continue onto the first subsequent job. Those with 1 to 19 years tenure actually earn a larger differential following the layoff than before it. For example, the 10 to 19 year group earns 27.4% more than to new job entrants and 16.8% greater than the 1 to 3 year category, prior to mobility, but 44.5% and 28.4% premiums, respectively, on the first post-displacement job. Seniority differentials transfer less completely for the longest tenure category —they earn 26% more than the less than 1 year group before the layoff but only 14% greater after it.

These results suggest that virtually the entire cross- sectional return to seniority (for workers with less than 20 years on the job) is explained by low turnover among individuals with market valued transferable characteristics. This interpretation receives further support from the small returns to post-layoff seniority in the instrumental variable regressions described in column (5). According to these estimates, earnings growth during the first 5 years on the post-displacement job averages only 2.5%. By contrast, the premium is a much larger 25.7% when actual postseparation tenure, rather than the within job instrument, is used (see column (6)). This indicates important unobserved differences in job or individual quality, even when prior seniority is controlled for.

Wage regressions were also estimated for subsamples stratified by ethnic status (white vs. nonwhite), industry (manufacturing vs. non-manufacturing), and occupation (blue collar vs. professional, managerial, and technical). Space does not permit a full discussion of the results, however, two findings are worth noting.[9] First, wage profiles are far steeper for professional,

9. *A table showing full regression results is available from the author.*

TABLE 1.—VARIABLE MEANS AND WAGE REGRESSIONS

	Sample Means		W_0		W_t	
	All Workers	Layoffs	All Workers	Layoffs	IV	OLS
EXP	18.5 yrs	15.1 yrs	0.033	0.038	0.041	0.037
			(25.70)	(6.60)	(8.20)	(7.46)
EXPSQ			−6.7E − 4	−8.9E − 4	−0.001	−9.6E − 4
			(−20.22)	(−5.46)	(−7.64)	(−7.17)
EDUC	11.2	10.6	−6.0E − 3	−0.035	−0.067	−0.053
			(−1.80)	(−2.15)	(−5.84)	(−4.74)
EDUCSQ			3.3E − 3	3.8E − 3	0.006	0.006
			(20.35)	(4.50)	(9.40)	(8.48)
NONWHITE	30.0%	40.3%	−0.291	−0.365	−0.419	−3.93
			(−37.09)	(−11.00)	(−13.28)	(−12.64)
MAR	90.1	83.2	0.286	0.288	0.295	0.286
			(25.69)	(6.99)	(11.41)	(11.18)
TEN2	24.7	31.6	0.102	0.087	0.118	0.124
			(9.56)	(2.44)	(6.83)	(7.23)
TEN3	26.4	17.0	0.170	0.164	0.290	0.273
			(15.95)	(3.78)	(7.51)	(7.14)
TEN4	20.0	7.6	0.188	0.242	0.368	0.359
			(16.29)	(4.01)	(10.77)	(10.58)
TEN5	12.4	4.0	0.147	0.237	0.131	0.154
			(11.05)	(2.94)	(2.98)	(3.54)
LAY1			−0.105			
			(−4.64)			
LAY2			−0.135			
			(5.51)			
LAY3			−0.128			
			(−3.93)			
LAY4			−0.047			
			(−0.97)			
LAY5			−0.036			
			(−0.54)			
DUR					0.019	0.047
					(1.42)	(4.28)
DURSQ					−2.8E−3	1.3E−3
					(−2.23)	(−1.09)
MILLS					0.929	0.791
					(3.20)	(2.76)
N			22,556	1373	5559	
RSQ			0.389	0.294	0.332	0.341
	(1)	(2)	(3)	(4)	(5)	(6)

Source: PSID 1969–1980. t-statistics are in parentheses. Also included in regressions are an intercept and dummy variables for city size and the survey year.

Note: *Variable Definitions*: *Continuous variables*—W_0 = log of real weekly wages in current year (full sample) or year prior to displacement; W_t = log of real weekly wages in period t, where layoff occurred in period one; *EXP* = work experience; *EXPSQ* = *EXP*EXP*; *EDUC* = education; *EDUCSQ* = *EDUC*EDUC*; *DUR* = actual or instrumental variable for seniority on post-displacement job; *DURSQ* = actual or instrumental variable for square of seniority on post-displacement job; *MILLS* = inverse Mill's ratio from probit reemployment equation. *Dummy variables* (equal one if)—*NONWHITE* = nonwhite; *MAR* = married; Period zero (preseparation) tenure: *TEN1* = < 1 year; *TEN2* = 1–3 years; *TEN3* = 3–9 years; *TEN4* = 10–19 years: *TEN5* = ≥ 20 years; *LAYX* = *LAYOFF*TENX* (for X = 1, 2, 3, 4, and 5), where *LAYOFF* equals one if a permanent layoff occurs in period 1.

managerial, and technical employees than for other workers. Where the estimated return to 5 years seniority is only around 2% for the entire sample and is zero or negative for blue collar occupations, it is approximately 33% for professionals. This result is consistent with recent research by Davis (1987) indicating much faster within-job earnings increases for managers than for other occupations. Second, there is some evidence of ethnic group differences in earnings profiles. The coefficients on the instrumented tenure variables are negative and statistically insignificant for nonwhites, whereas the estimated wage gain for staying 5 years on the first post-displacement job exceeds 9% for whites.

IV. CONCLUSION

Cross-sectional estimates of within-job earnings growth are strongly biased by a positive correlation between the regression error term and job seniority. This occurs mainly because individuals with high tenure possess market valued attributes which are transferable across firms. Differences in match quality appear much less important.

Using an instrumental variable measuring deviations between current and average *within job* tenure, there is virtually no reward for seniority on the first employment held after a permanent layoff. Conversely, the preseparation seniority differentials of persons leaving jobs of up to 20 years duration are retained (and frequently increased) upon reemployment. Individuals with still longer tenure appear more vulnerable to loss of seniority premiums following displacement.

For male head of households, the estimated return to five years seniority falls from 22.5% to 2.5% when match and individual quality are controlled for. The reductions for whites are slightly smaller—from 29.8% to 9.4%. Therefore, only 10% to 25% of cross-sectional tenure premiums represent actual within-job earnings increases. Other recent research yields similar results. Altonji and Shakotko (1987) find that the actual return to tenure, for white males, is only 9% to 30% as large as that indicated by cross-sectional estimates. Abraham and Farber (1987) argue that the corrected estimate (also for white males) is between one-fifth and one-third as large as in cross-sectional data.

These results question the importance of theories which emphasize match quality or firm-specific endowments. On the other hand, models which emphasize transferable (unobserved) worker traits probably deserve greater attention. It is not clear which specific characteristics are valued across firms. Related work (Ruhm, 1987b) suggests that heterogeneity in quit propensities may be of considerable importance. These findings should not be taken to imply that worker-firm attachments are never consequential. For example, managerial and professional workers gain considerable rewards for remaining with a single employer.

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