Canadian International Labour Network

Labour Market Outcomes:

A Cross-National Study

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McMaster University

DEPARTMENT OF ECONOMICS

Vulnerable Seniors:

Unions, Tenure and Wages Following Permanent Job Loss

Peter Kuhn[1]

Arthur Sweetman[2]

A well known finding in the literature on displaced workers is the apparent "portability" of tenure across firms: controlling for experience and other observable characteristics, workers with high levels of predisplacement tenure earn higher postdisplacement wages (e.g. Kletzer 1989). Using four data sets on displaced workers, we show that this finding is reversed for workers losing unionized jobs. Our finding cannot be explained by firm- or industry-specific human capital accumulation, deferred-pay policies, standard matching models, or by a correlation between tenure and re-entry rates into unionized jobs. We argue instead that it can reflect only two possible processes: negative selection of senior union workers, or a negative causal effect of unionism on workers' alternative skills. An important implication of our findings is that, despite a much flatter predisplacement tenure-wage profile, displaced union workers' wage losses increase with tenure at a comparable or higher rate to that of nonunion workers.

[1]McMaster University
[2]University of Victoria

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1. Introduction

The growing literature on displaced workers' wage outcomes has yielded two well-known and robust results. One is that, controlling for experience and other observables, displaced workers' wage losses increase with tenure, or seniority, on the lost job; this finding has often been interpreted as evidence of firm-specific training. The other is that, controlling for experience and observables, the *level* of postdisplacement wages is positively correlated with predisplacement tenure (Kletzer 1989). In other words, some of the wage-enhancing aspects of predisplacement tenure seem to be "portable" into postdisplacement jobs.[1] This phenomenon has been interpreted as evidence of upward ability bias in tenure-wage equations, though Neal (1995) has argued that it represents industry-specific human capital accumulation as well.

One limitation of existing research on displaced workers is the fact that very little is known about the consequences of displacement for unionized workers.[2] This is unfortunate for at least two reasons. First, given a union wage premium, unionized workers may experience particularly severe consequences of displacement, which are worthy of attention in their own right. Second, displaced-worker data provide an important new source of information about the processes of union wage-setting and selection into unions that have long been a subject of interest among labor economists (e.g. Lewis 1986, Card 1996). Because displaced workers change jobs for reasons that are arguably orthogonal to their unobserved ability, an examination of their wage outcomes may thus provide some new insights into these selection processes.

In this paper we use four different data sets —one American and three Canadian— with information on displaced workers' union status to analyse the postdisplacement wages of workers losing union and nonunion jobs. In all the data sets, unionized workers experience much greater wage losses than other displaced workers, largely because of the loss of union coverage associated with displacement. We also find, however, that the two "stylized facts" about displacement noted above do not extend in a straightforward way to workers losing unionized jobs. Most significantly, the second is reversed: In each of the four data sets examined here, the level of alternative, or postdisplacement wages, *declines* with tenure on the lost job for unionized workers.

What explains this decline in alternative wages? Because —by definition— neither firm-specific capital, deferred-compensation policies, nor job match quality are portable across firms, it is very difficult to attribute this correlation to any of these processes. In the paper we also show, empirically, that these declines in alternative wages cannot be wholly explained by correlations of postdisplacement union status or industry-switching with predisplacement tenure. We argue, instead, that our results are consistent with only two possible scenarios (or a combination of them). In the first, senior union workers are negatively selected relative to junior union workers. This is consistent with a model in which workers' general ability is revealed over time, but ability is less rewarded in the union than nonunion sector.[3] In this world, union workers receiving positive informational shocks about their productivity either quit, or are promoted out of, unionized jobs; only those with negative shocks accumulate high tenures in that sector.

The second scenario incorporates a possibility that has been largely ignored in any of the literatures on tenure, displacement, and unions: the depreciation of "alternative" skills, defined as skills that are not useful in the current job but are useful elsewhere.[4] In this scenario, spending time in a unionized work environment causes an atrophy of skills which are useful in the alternative labor market. This could, of course, happen in all jobs, but it should be more important in unionized jobs to the extent they are less multiskilled (Ichniowski 1990), and —at least *ex ante*— are more secure (Freeman 1980).

Overall, regardless of which of the two above causal mechanisms is at work, our findings indicate very strongly that cross-sectional union wage-tenure profiles dramatically understate the amount of job rents earned by senior union workers, as well as the rate of increase of such rents with tenure. Thus the results provide renewed support for theories, including the exit-voice (Freeman 1980), and discriminating-monopoly (Kuhn 1988) models, that predict correlations of job rents with tenure for union workers that are at least comparable to those of nonunion workers. Our findings also highlight the need to think about selection into, and out of, the union sector in a dynamic, life-cycle context, rather than the simple static frameworks —based on the Roy model— commonly used. Further theoretical and empirical work in this area might help distinguish the two scenarios outlined above.

Section 2 of the paper describes each of the four data sets used in our analysis. Section 3 presents our basic results, while Section 4 examines whether they can be easily explained by firm wage effects, differential re-entry of senior versus junior displaced workers into the union sector, or differential rates of industry switching among displaced workers. Section 5 discusses some of the implications of the findings and concludes.

2. Data

In our empirical work we use one data source from the United States —the pooled 1994 and 1996 Displaced Worker Surveys (USDWS)— and three Canadian ones: the Canadian Displaced Worker Survey (CDWS), the Canadian Out-of-Employment Panel (COEP), and an

Ontario Ministry of Labor (OML) survey. Unlike the earlier Displaced Worker Surveys conducted in the United States, which have been used in most studies of worker displacement to date, the more recent ones contain information on the worker's union status. Aside from providing more observations, the Canadian data have the advantage of providing a higher fraction of union observations from an industrial relations context very similar to that in the United States, and of being constructed using three very different sampling methods. Given some of the difficulties associated with recall of past displacements in the retrospective structure of the USDWS (e.g. Evans and Leighton 1995), this provides an added check on the robustness of our results. All four data sets contain measures of union status prior to displacement, and all but the CDWS also have information on postdisplacement union status.

The USDWS, conducted biannually as a supplement to the Current Population Survey (CPS), has been widely used, and is documented elsewhere (e.g. Farber 1997); we do not describe it further here. The CDWS was conducted only once, as a supplement to the January 1986 Labor Force Survey.[5] The target population was all individuals who permanently lost a job due to a shortage of work in the calendar years 1981-1985. The COEP is a survey of individuals separating from jobs for any reason in one of two window periods, January 31-March 13, and April 25-June 5, 1993. Approximately 10,000 individuals were interviewed three times, between 23 and about 60 weeks following their job separation.[6] To be included in the "displaced worker" subsample used in this paper, a worker had to be part of a mass layoff, or laid off on an individual basis because of a shortage of work and not recalled. The OML is a 1982 survey of workers involved in total and partial plant closings in Canada's most populous province in the preceding three years. It sampled 21 firms from a census of mass layoffs (50 or more workers in a 4 week period; in practice 18 of the 21 were complete plant closures) and is discussed in Jones and Kuhn (1995). While the OML's sample size is small, it allows workers to be matched to their predisplacement employer and permits a focus on workers displaced in mass layoffs. Both the CDWS and COEP record the starting wage on the postdisplacement job, but the OML and USDWS surveys obtain the wage at the survey date.[7]

The sample for analysis from each of the surveys comprises men who lost full time employment, who were between 20 and 64 when displaced, and who were not self-employed. Observations were dropped due to missing information, and those with tenure greater than 20 years were removed since the cell sizes were too small to provide useful inference.[8] It should be noted that both displaced worker surveys, and the COEP sample, include workers who were displaced on an individual basis in addition to those who were part of a mass layoff. Unfortunately, while a sample composed only of mass layoff victims, such as the OML data, increases our confidence in the orthogonality of worker characteristics and the event of job loss, the subsamples in the other data sets are too small to permit separate inference.[9]

Union membership, in all three Canadian surveys, is defined as being covered by collective bargaining. The USDWS asks respondents only if they were *members* of a union or an employee association similar to a union in their predisplacement job, whereas the CPS, which provides the postdisplacement information, asks first about both membership and coverage under a collective contract. However, the CPS restricts the postdisplacement questions to respondents in rotation groups 4 and 8: only 244 individuals from our sample, 30 of whom were unionized prior to displacement. Thirty of these workers were union members in their postdisplacement jobs, and three of the remaining 214 workers indicated that they were covered by collective bargaining. Assuming a similar union coverage to membership ratio for the predisplacement jobs, this suggests, very roughly, that just under 10% of the displaced workers who would be classified in the union sector by the coverage definition used in the Canadian surveys are classified as being in the nonunion one in the USDWS.

A summary of the predisplacement sample characteristics by union status is given in Table 1. Some of the observed differences in worker characteristics and outcomes undoubtedly results from variations in survey methodology and business conditions across the locations and time periods covered, as well as cross-national differences in economic institutions. The robustness of our results to these differences, however, increases our confidence in them. Substantial variation in the tenure distributions exist across the four samples. On average, the COEP has the shortest tenure, the CDWS has the next longest, the USDWS has higher tenure again, and the OML is comprised of a large fraction of very senior workers. This is probably because the COEP is a flow sample, while the USDWS and CDWS are retrospective and select on the longest tenured job from which a worker was displaced, and the OML focuses on mass layoffs which implicitly eliminate small firms from the sample. Relatedly, the CDWS sample is, on average, younger; the COEP union sample's similar average age is similar to that of the OML and USDWS despite its lower average tenure.

The education variables are not strictly comparable across the four data sets; in the USDWS the high school group includes incomplete postsecondary, and the CDWS high school indicator subsumes the “some high school” and “high school” groups. Similarly, the “postsecondary” indicator in the OML sample includes those who have incomplete postsecondary education (“some postsec.”), and complete community, 2 year, college (“postsec.

cert".) and university. The OML survey contains an additional variable that measures other formal vocational or technical training. Once these differences are understood, the average education levels of the CDWS and the COEP look quite similar to each other and the OML nonunion sample. The OML union sample, however, has a much lower average education level. In contrast, the USDWS has a much higher fraction of university graduates in both sectors, and fewer workers with only elementary level education. The American and Canadian displaced workers surveys report similar fractions, about 35%, of workers being displaced by a plant closing or moving, though the USDWS has a somewhat lower number for the union sample; this compares to about 20% in the COEP. A large part of this difference may be again attributable to dissimilarities in the sampling schemes. Years since displacement is presented for the USDWS, CDWS and OML samples which, in contrast to the COEP, are all retrospective. A large fraction of the OML sample were displaced in the 1981 recession.

One striking difference among the Canadian data sets is that the OML nonunion sample is just under half the size of the union one rather than about twice the size as in the CDWS and the COEP. This likely reflects the predominance of manufacturing firms in the OML data, as well as its sampling strategy, which, as previously mentioned, excludes small establishments. As seen in the CDWS, the only survey to measure establishment size, union workers tend to be displaced from larger firms than nonunion workers. Over 50% of nonunion displaced workers in the CDWS sample are from establishments with less than 20 employees, whereas about 50% of the union sample are displaced from establishments with over 500 employees. Controlling for firm size in at least one data set increases our confidence in the robustness of our results.

Monetary values are measured as weekly earnings for the USDWS and CDWS, and as hourly wages for the COEP and the OML. Wage losses of workers displaced from union jobs are larger than those of workers displaced from nonunion jobs in all but the OML sample. As we shall see, this difference —which is reversed in the regressions that control for tenure and personal characteristics below— results from the fact that, again given the OML's sampling structure, a large fraction of the nonunion workers are managerial and supervisory personnel in establishments that are more than 90 percent unionized. Average wage losses in the USDWS are about 16% and 12% for the predisplacement union and nonunion samples respectively. Union losses in the CDWS are similar, on average, at about 14%, and they are about 7% in the COEP. But the average nonunion losses in the CDWS and COEP, in contrast to the USDWS, are close to negligible. This results, in part, from the higher tenure of nonunion workers displaced in the USDWS sample. Overall, in the USDWS, CDWS and COEP, unionized workers earn more initially, and while they lose more upon displacement, they still have a higher postdisplacement average wage. In contrast, in the OML sample the nonunion workers have a higher average predisplacement wage and the union workers experience a slightly smaller average wage loss.

3. Results

Table 2 presents unadjusted pre- and postdisplacement wages, and wage losses, by (predisplacement) tenure category for the pooled 1994-1996 USDWS. The top half of the table presents simple means of log wages and log wage losses. The bottom half presents coefficients on three dummy variables for the same tenure categories (less than one year is the left-out group) from log wage (or wage loss) regressions that also control for the predisplacement variables presented in Table 1: age (specified as a quadratic), education, marital status, whether the individual was displaced individually or *en masse*, the year of the survey, 3 region indicators, 9 predisplacement industry dummy variables and, in the postdisplacement wage and wage loss regressions only, years since the displacement, and weeks unemployed. The specification is intended to be similar to that of Neal (1995) and Kletzer (1989).

Columns (3), (6) and (9) of Table 2 do not condition on union status; they are provided to allow comparisons with previous work (e.g. Addison and Portugal 1989; Kletzer 1989; Ruhm 1990). The results are very similar to those obtained from earlier years of the USDWS: Predisplacement wages are strongly, and positively, correlated with predisplacement tenure, whether observed characteristics are held constant or not. However, because some of the apparent effects of predisplacement tenure are "portable" into postdisplacement jobs, displacement-induced wage losses increase considerably less rapidly with tenure than predisplacement wages. This increasing pattern of postdisplacement wages with predisplacement tenure has been interpreted, as mentioned, as positive sorting on ability by the above authors. In other words, predisplacement tenure is not believed to affect the postdisplacement wage causally; instead it is interpreted as proxying for an element of unobserved ability that is transferable across jobs. Neal (1995), using USDWS data, stratifies his sample on industry switchers and stayers, and finds that while the coefficients on predisplacement tenure in postdisplacement earnings regressions are statistically significant in both, the return is much larger, and very similar to those in the predisplacement wage equations, for the stayers. He argues that this is evidence of industry-specific human capital accumulation. None of the above authors differentiate between unionized and nonunionized workers, although Neal presents some regressions using a "mostly nonunion" sample.

In contrast to the traditional results, we observe important heterogeneity across the union and nonunion sectors; interestingly, the postdisplacement wage-tenure profile changes dramatically for union workers when the sample is disaggregated. Mirroring previous studies of cross-sectional union wage patterns (e.g. using United States data: Borjas 1979; Farber 1983; Pearce 1990; and using Canadian data: Green 1990, Doiron and Riddell 1994) column 1 shows a very flat wage-tenure profile among unionized workers, (indeed the profile is negatively sloped when personal characteristics are held constant), while the nonunion profile in column 2 remains upward sloping, reflecting the dominance of nonunionized workers in the United States sample. Columns 4 and 5 present what are probably the paper's most novel and surprising results: the opposite correlation of postdisplacement wages with predisplacement tenure for union versus nonunion workers. Whether or not personal characteristics are held constant, senior nonunion workers clearly have better labor market alternatives than junior nonunion workers, while the opposite is true among union workers. However, the combination of the differences in the pre- and postdisplacement tenure profiles gives tenure-wage loss profiles that are more similar across the two sectors. The point estimates, when we control for observables, suggest that the union tenure-wage loss profile initially rises more steeply, but the highest tenure category's losses are slightly larger for the nonunion group.

Are the patterns observed in the 1994 and 1996 USDWS simply a quirk of the sampling structure, time period, and economic institutions affecting those particular surveys? To address this possibility, Tables 3 and 4 conduct similar exercises to those in Table 2 in the three Canadian displaced-worker data sets at our disposal. Table 3 presents unadjusted log wages, and Table 4 presents tenure coefficients from log wage regressions that control for very similar sets of characteristics as were used in the Table 2 regressions. Considering first unadjusted log wages, it is clear that the trends in predisplacement wages are similar in all three data sets, showing a wage gap of 30 to 40 log points between the most junior and senior nonunion workers, and little or no gap between junior and senior union workers. Not surprisingly, the union profile is everywhere higher than the nonunion one in the CDWS and COEP, but is below the nonunion profile in the OML data, where union and nonunion workers are largely drawn from within the same firms. Columns 3 and 4 again show opposite correlations between predisplacement tenure and postdisplacement wages for workers displaced from union versus nonunion jobs: While senior nonunion workers do better than junior nonunion workers in the alternative labor market, senior union workers do absolutely *worse* in the outside labor market than their junior colleagues.

The difference between pre- and postdisplacement wages, (i.e. the wage loss associated with displacement), is examined in columns 5 and 6 of Table 3. Primarily because of the deterioration of union workers' alternative wages with tenure, wage losses are large and increasing with tenure in the union sector, whereas they are smaller and more slowly increasing in the nonunion one in all three Canadian data sets. These small wage losses for senior nonunion workers contrast somewhat with the highest tenure category in the USDWS; a difference that might warrant further investigation that pays close attention to institutional differences between the countries.

Turning to the regression-adjusted estimates in Table 4, columns 1 and 2 present tenure coefficients from standard cross-section regressions of predisplacement wages on predisplacement characteristics (as defined earlier and in Table 1), and the results conform with both the USDWS and with much earlier work. In particular, the tenure coefficients are small in the union sector, tracing out an essentially flat profile, whereas the nonunion regressions show a sharply increasing wage profile with increasing tenure. Turning to postdisplacement wages in columns 3 and 4, the results mirror both the unadjusted means and our findings from the USDWS: in each data set, the postdisplacement wage is increasing in predisplacement tenure for the nonunion sample, and is decreasing in tenure for the union sample. The effect is especially large in the COEP which is the only flow sample without the recall problems of the other surveys.

Columns 5 and 6 of Table 4 give estimates of the relation between tenure and the cost of job loss, controlling for observable characteristics. It is clear that the costs are increasing with tenure in the union sector, and are much flatter in the nonunion sector, especially in the COEP and OML data. This differs somewhat from the USDWS results in Table 2 where, relative to the nonunion sector, the union coefficient point estimates were larger for the 1 to 4, and 5 to 9, year tenure groups, but a bit lower for the those with 10-20 years of tenure. Overall, however, in all four data sets, these estimates of unionized workers' job rents show an opposite pattern to that suggested by the well-known, very flat, cross-sectional union tenure-wage profile (which does not incorporate the correlation of alternative wages with tenure). Job rents, measured as the wage losses experienced by permanently displaced workers, increase at least as rapidly, and in the Canadian data more rapidly, with tenure in the union than in the nonunion sector. The main reason is the “vulnerability” of senior union workers, whose labor market alternatives are strictly worse than those of their more junior colleagues.

4. Explanations for the Deterioration in Union Workers’ Alternative Wages with Tenure

What explains the very robust negative correlation between tenure and union workers' alternative wages documented in the previous section? To see why this correlation might be of some interest, note first that, unlike the correlation between tenure and predisplacement wages, it cannot be attributed to firm-specific capital accumlation and/or deferred compensation systems; these should affect predisplacement, but not postdisplacement wages. Also, note that standard search and matching models (e.g. Burdett 1978; Topel 1991) are likewise incapable of generating this correlation because these models generally treat displacement –in contrast to voluntary job mobility– as a re-initialization of the matching process with a random new draw from the overall wage offer distribution.

What possibilities remain? One is that, in contrast to the standard assumption, postdisplacement job match quality is not orthogonal to predisplacement match quality. For example, workers in high-wage firms before displacement might be more likely to get help from those firms in finding new jobs in other high-wage firms, including affiliates of the original company. Of course, for this to affect our estimates of the correlation between predisplacement tenure and postdisplacement wages, predisplacement match quality needs to be correlated with predisplacement tenure, a possibility which is by no means guaranteed in standard matching models (for a treatment of these models and their implications for estimating the returns to tenure, see Crossley 1998). In any case, to the extent that match quality is captured by firm wage effects, its effects on our results can easily be measured in one of our data sets, which identifies workers' predisplacement firms. This possibility is explored in what follows.

Two other, relatively simple, explanations for our main result can also be tested with our data. One is differential access by displaced workers *back into* union-sector jobs: senior union workers may simply have greater difficulty (than junior union workers) in finding new, unionized jobs. Finally, our result may have something to do with industry-specific skills. As Neal (1995) has pointed out, industry tenure is almost certainly positively correlated with firm tenure, so that a positive effect of predisplacement tenure on postdisplacement wages might just reflect the accumulation of industry-specific skills, which truly are portable across firms. To explain the *negative* correlation between predisplacement tenure and postdisplacement wages we see among unionized workers, the above effect would have to be outweighed by a greater tendency of senior than junior workers to switch industries after displacement. We check for this possibility by stratifying the sample into industry changers and stayers below. To the extent that our main result remains after exploring the above possibilities, we are driven to an interpretation based either on negative selection of senior union workers on unobservables, or on a causal, negative effect of union tenure on workers' alternative labor market skills.

(a) Matching and Firm Wage Effects

As mentioned, there is a possibility that something about the kinds of predisplacement firms in which senior versus junior workers tend to be found might be explaining our main result. For example, "good" (high-wage) firms might have more senior workers, and might provide either more or less help to displaced workers in finding good, new matches. A simple way to address this issue is to add (predisplacement) firm dummy variables to the regressions, which is possible in one of our data sets (the OML). To that end, Table 5 presents tenure coefficients based on the OML data and similar to those in Table 4, but which include firm dummies. Interestingly, in both the union and nonunion predisplacement wage equations of columns 1 and

2, the highest tenure coefficients are substantially smaller than their counterparts in Table 4, suggesting that firm heterogeneity affects the estimated returns to predisplacement tenure in the predisplacement firm. Focussing however on the postdisplacement wage equation's tenure coefficients in columns 3 and 4, the results are very similar to those without firm dummies in Table 4. At least to the extent that match quality is captured by firm effects, the standard maintained assumption in the literature, that postdisplacement jobs represent random new draws of job match, is thus supported by these results. In our view, this makes it very difficult to attribute the apparent deterioration of union workers' outside options with tenure in our data to any story based on matching.

(b) Differential access to union-sector jobs

Another potentially important explanation for the decline in union workers' alternative wages with tenure is the rate of access back into unionized jobs after displacement. Could it be that workers' ability to become re-employed in a unionized job declines with tenure, and that this alone explains the pattern we observe? To explore this possibility we first examine displaced workers' reemployment rates into the union sector in Table 6, for the three data sets in which information on postdisplacement union status is available. These are the percentages of workers in each predisplacement sector-tenure cell that are reemployed into the union sector (recall that all the workers in our sample are re-employed somewhere). For all three data sets it is clear that union workers are much more likely to be reemployed in the union sector, but the tenure patterns in union re-employment differ somewhat across surveys. Considering first the overall rate of "de-unionization" among displaced workers, 58, 57 and 55% of the predisplacement union workers are reemployed in the union sector, compared to only 7, 12 and 22% of the predisplacement nonunion workers for the USDWS, COEP and OML samples respectively.[10]

Gross averages, however, mask an important trend in the evolution of the de-unionization rate by tenure. As seen in column 1 of Table 6, in the USDWS predisplacement union workers in the two most junior tenure categories are about twice as likely to be covered by collective bargaining following displacement as their more senior colleagues. Further, as seen in column 2, very few predisplacement nonunion workers become unionized. The profile for the COEP, in columns 3 and 4, is both smoother and steeper; junior workers are much more likely to be reemployed in the union sector than more senior displaced workers. For predisplacement union workers the probability of reemployment into a unionized job drops from 70%, for those with less than 1 year of tenure, to a low of 15% for those with more than 10. For the predisplacement nonunion workers, the probability of reemployment in the union sector drops from 14 to 3% for the same tenure groups. The deunionization trend with tenure is not nearly as strong for the union sector in the OML data as seen in column 3. Further, that for the nonunion side, column 4, is flat.

Overall, the above results suggest that differential access back into the union sector has the potential to explain much of our result: to a moderate degree in two data sets, and very strongly in a third, senior union workers appear to have greater difficulty than juniors in finding postdisplacement jobs in the union sector. Perhaps this, alone, explains their lower postdisplacement wages. To see whether this is the case, in Table 7 we replicate elements of the preceding analysis, restricting attention to workers who do not change union status: if our results persist for these subsamples, they cannot be explained by changes in union status alone. While the USDWS's sample with known postdisplacement union status is too small to profitably carry out this exercise, this issue can be addressed using the COEP and the OML.[11]

Unadjusted log postdisplacement wages, as well as tenure coefficients from log postdisplacement wage regressions that are restricted to the sample of those who maintain their union status across jobs in the COEP and OML are reported in Table 7. Consider first the COEP, comparing the unadjusted postdisplacement wages in the upper half of the table, the postdisplacement profiles in columns 1 and 2, while noisier because of the smaller sample size, are similar to those in columns 3 and 4 of Table 3, though the U-U profile in Table 7 is, unsurprisingly, above that in Table 3. Thus even in this sample, senior union workers still obtain lower postdisplacement wages than their more junior colleagues, whereas the converse is true for nonunion workers. The lower portion of the columns, which contain regression coefficients, tell the same story and the estimates are remarkably similar to those in Table 4. Notice that the union profile appears steeper once we controlled for observables. For the OML the results are likewise in accord with the earlier findings. The disaggregated results in Table 7 therefore suggest that changes in union status are not the main explanation of the divergent patterns of postdisplacement wages with respect to tenure we observe in the union and nonunion sectors.

(c) Differential Propensities to Switch Industries.

An additional possible explanation of our results, motivated by Neal (1995) and Parent (1995), is that across tenure categories in the union and nonunion sectors, workers may have systematically different propensities to switch industries. Tenure coefficient profiles in the postdisplacement wage equations may, therefore, be influenced by the portability of industry specific human capital. This issue can be addressed using the USDWS and the COEP.[12] Industry changing in the USDWS is defined as being coded in a different "detailed industry" (with 52 categories) as described in the U.S. Department of Commerce, Bureau of the Census (1996) documentation. In the COEP each respondent is directly asked whether he changed industries. Table 8 presents predisplacement tenure coefficient estimates from postdisplacement wage regressions similar to those in Neal (1995) which stratify the sample on industry switching, but we also stratify the sample on predisplacement union status. The previously observed differences across the union and nonunion sectors remain in both data sets. For both industry switchers and stayers, predisplacement union workers have negative coefficient estimates, while the nonunion ones have (with one exception) positive ones. Industry switching does not appear to be driving the observed differences in the correlation between predisplacement tenure and postdisplacement wages across the two sectors. This is perhaps not unexpected, for a reason noted earlier: To explain the negative effect in the union sector, there would need to be an increase with tenure in the tendency to switch industries that is strong enough to outweigh the positive direct effects of higher industry tenure on wages.

4. Discussion.

Using displaced worker data from two countries, this paper demonstrates a significant and robust difference in the correlation between alternative market wages and predisplacement tenure between the union and nonunion sectors. This correlation is positive for workers displaced from nonunion jobs, but negative for workers displaced from unionized jobs. These correlations are of interest because, as has been pointed out by several authors, they are a simple source of quite subtle information about the role of general worker ability in the joint process generating tenure, wages, and (in the current context) union status. Further, because of data limitations affecting earlier work, the negative correlation for union workers has, to our knowledge, not been documented before.

What, then, does this negative correlation mean? In the paper we argue that it cannot, by definition, be explained by specific human capital accumulation or firm-specific deferred-pay policies. We also argue, and demonstrate empirically, that it cannot easily be attributed even to fairly complicated variants of search and matching models, of the type set out in Burdett (1978), or Topel (1991). Nor can it be explained by differential rates of access back into the union sector, or industry-specific human capital. In our reckoning two candidate explanations for the negative correlation remain. Either senior union workers are negatively selected (relative to junior union workers), or the accumulation of long tenures in the union sector has an adverse causal effect on workers' alternative labor market skills. The former, "selection" hypothesis suggests a scenario in which union firms initially select a highly able workforce (as is shown by the high overall level of union workers' wages in our data, even after displacement). However, because ability is less rewarded in the union than nonunion sector, the workers with the most advantageous shocks to their ability tend disproportionately to leave the union sector, and not to accumulate high tenures there. In contrast, the "atrophy" hypothesis suggests that the relative security and narrow job definitions usually associated with unions might cause a greater depreciation of alternative skills with tenure.

Interestingly, both the above scenarios are consistent with another pattern that emerges quite strongly in our data: unionized firms seem particularly unwilling to hire displaced senior union workers. Clearly, we view the disentangling of these two scenarios as an important area for further work, but argue that making such subtle distinctions will be much more likely using data, unlike that here, with more than one wage observation per worker-job match. Our results also strongly suggest that any such future work on unions and selectivity, rather than relying on static Roy models, would profit from a more dynamic perspective in which new information on productivity is revealed over time, inducing mobility out of and into a union sector where ability is differentially rewarded from the nonunion sector.

Overall, however, whichever of the two causal mechanisms identified above is at work, our findings indicate very strongly that, in a cross-section of union workers, the more senior workers have strictly inferior labor market alternatives than the juniors. As a result, cross-sectional union wage-tenure profiles dramatically understate the amount of job rents earned by senior union workers, as well as the rate of increase of such rents with tenure. Thus the results provide renewed support for theories, including the exit-voice (Freeman 1980), and discriminating-monopoly (Kuhn 1988; Kuhn and Robert 1989) models, that predict —for either political or efficiency reasons— correlations of job rents with tenure for union workers of a comparable or greater magnitude to those of nonunion workers. These theories, of course, also have implications for our understanding of union employment effects, and the nature of the union wage premium.[13]

Our finding also has implications for the measurement of "the" return to tenure and the nature of specific human capital. A series of papers have addressed this issue with a focus on biases that arise from worker and job heterogeneity and the endogenous accumulation of tenure (e.g. Abraham and Farber 1987; Altonji and Shakotko 1987; Topel 1991; Altonji and Williams

1997). While only Abraham and Farber (1988) distinguish between the union and nonunion sectors, the results presented here imply that care must be taken to distinguish between them since wages evolve quite differently with tenure in each and likely reflect quite different underlying processes.

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Table 1 - Descriptive Statistics by Predisplacement Union Status

	USDWS		**CDWS**		**COEP**		**OML**	
	Pre-U	Pre-NU	Pre-U	Pre-NU	Pre-U	Pre-NU	Pre-U	Pre-NU
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Tenure:								
<1[1]	.21	.27	.17	.31	.63	.56	.14	.26
1-4[1]	.35	.39	.49	.53	.19	.26	.16	.16
5-9	.19	.20	.18	.11	.09	.10	.28	.25
10-20	.24	.14	.15	.05	.09	.08	.42	.33
Education:								
elementary	.01	.03	.13	.12	.05	.03	.23	.07
some high sch.	.08	.08	--	--	.32	.23	.39	.30
high school[2]	.68	.53	.60	.55	.30	.34	.24	.25
some postsec.	--	--	.08	.11	.06	.15	--	--
postsec. cert[3]	.10	.10	.16	.16	.20	.16	.12	.38
university	.10	.19	.02	.05	.04	.09	--	--
post-grad	.02	.08	--	--	--	--	--	--
oth tech/voc	--	--	--	--	--	--	.26	.42
single	.15	.23	.17	.31	.18	.35	.24	.21
plant closure	.26	.36	.34	.35	.20	.19	--	--
Blue collar	--	--	--	--	--	--	.88	.50
Establishment Size:								
<20			.34	.56				
21-99			.15	.21				
100-499			.20	.11				
≥ 500			.51	.12				
1996 U.S. DWS	.51	.50						
Weeks Unempl.	7.3	6.0						
	(.942)	(.338)						
Years Since Displacement								
1	.32	.37	.17	.15			.14	.40
2	.33	.32	.29	.25			.78	.57
3	.35	.31	.26	.22			.08	.03
4			.16	.22				
5			.11	.15				
Age (years)	38.1	36.8	32.1	29.2	38.4	33.8	37.9	38.0
	(.553)	(.236)	(.475)	(.341)	(.605)	(.355)	(.611)	(.803)
Predisp wage	732.67	630.22	445.94	310.30	15.64	11.12	8.61	9.95
	(19.4)	(9.69)	(6.93)	(4.59)	(.347)	(.202)	(.105)	(.258)
Postdisp wage	609.81	554.90	381.23	302.33	14.47	11.10	7.94	9.01
	(19.5)	(8.78)	(8.03)	(10.49)	(.334)	(.200)	(.130)	(.234)
N	271	1808	389	830	285	637	340	165

Note: Age is recorded as the mid-point of a range for the CDWS and the OML survey. The CDWS and OML wages are in 1981, the COEP wage in 1993, and the USDWS in 1996 constant dollars. Standard errors in parentheses.

[1]Tenure categories in the OML are 0-2 and 3-4 years respectively.
[2]Includes some high school in CDWS.
[3]Includes some postsecondary and university in OML.

Table 2 - Unadjusted and Regression-Adjusted log Wages (Std Errs) by Predisplacement Tenure Category and Union Status: US Displaced Worker Survey, 1994 and 1996. (ln weekly wage)

	Pre-Wage			**Post-Wage**			**ln(Pre-Wage/Post-Wage)**		
	Pre-U	Pre-NU	All	Pre-U	Pre-NU	All	Pre-U	Pre-NU	All
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Tenure (years):	*Unadjusted log wages (weekly)*								
<1	6.54	5.98	6.03	6.42	5.92	5.98	.11	.05	.06
	(.028)	(.014)	(.014)	(.034)	(.015)	(.014)	(.032)	(.014)	(.012)
1-4	6.39	6.22	6.24	6.21	6.13	6.14	.18	.08	.10
	(.030)	(.014)	(.013)	(.036)	(.015)	(.013)	(.030)	(.013)	(.012)
5-9	6.52	6.39	6.40	6.21	6.23	6.23	.30	.15	.17
	(.031)	(.014)	(.013)	(.044)	(.016)	(.015)	(.036)	(.014)	(.013)
10-20	6.53	6.61	6.59	6.19	6.23	6.22	.33	.37	.36
	(.020)	(.014)	(.012)	(.035)	(.017)	(.015)	(.036)	(.016)	(.015)
	Tenure Coefficients from log wage regressions								
1-4	-0.07	0.15	.12	-0.14	0.14	.10	0.09	0.00	.02
	(0.083)	(0.037)	(.035)	(0.097)	(0.038)	(.036)	(0.096)	(0.038)	(.035)
5-9	-0.01	0.22	.20	-0.16	0.14	.10	0.17	0.07	.09
	(0.092)	(0.044)	(.041)	(0.116)	(0.048)	(.045)	(0.107)	(0.049)	(.044)
10-20	-0.05	0.38	.32	-0.22	0.12	.06	0.18	0.26	.26
	(0.072)	(0.047)	(.042)	(0.107)	(0.057)	(.051)	(0.112)	(0.059)	(.052)
N	271	1808	2079	271	1808	2079	271	1808	2079
R^2	.27	.39	.37	.23	.25	.23	.10	.09	.08

Note: Standard errors are in parentheses; those for the regressions are heteroskedasticity consistent. Additional control variables (coefficients not shown) are as shown in Table 1. Also, included are 3 region dummy variables, and 9 industry dummy variables.

Table 3: Unadjusted log Wages (Std Errs) by Predisplacement Tenure Category: CDWS, COEP and OML

	Pre-Wage		Post-Wage		ln(Pre-Wage/Post-Wage)	
	Pre-U	Pre-NU	Pre-U	Pre-NU	Pre-U	Pre-NU
	(1)	(2)	(3)	(4)	(5)	(6)
CDWS (ln weekly wage)						
<1	6.07	5.48	5.97	5.50	.10	-.02
	(.038)	(.024)	(.049)	(.023)	(.044)	(.022)
1-4	6.04	5.71	5.83	5.67	.21	.04
	(.025)	(.019)	(.033)	(.019)	(.024)	(.017)
5-9	6.01	5.79	5.77	5.73	.25	.05
	(.042)	(.039)	(.059)	(.040)	(.051)	(.035)
10-20	6.08	5.83	5.85	5.74	.23	.09
	(.036)	(.062)	(.050)	(.071)	(.039)	(.059)
COEP (ln hourly wage)						
<1	2.71	2.26	2.68	2.30	.02	-.03
	(.031)	(.022)	(.029)	(.020)	(.028)	(.021)
1-4	2.62	2.31	2.53	2.33	.09	-.02
	(.053)	(.033)	(.058)	(.031)	(.050)	(.026)
5-9	2.49	2.38	2.38	2.32	.12	.06
	(.073)	(.049)	(.070)	(.046)	(.064)	(.040)
10-20	2.75	2.63	2.34	2.53	.41	.09
	(.067)	(.063)	(.077)	(.069)	(.069)	(.073)
OML (ln hourly wage)						
0-2	2.04	2.07	2.08	2.06	-.04	.01
	(.029)	(.053)	(.034)	(.060)	(.035)	(.040)
3-4	2.11	2.19	2.06	2.05	.05	.13
	(.032)	(.047)	(.040)	(.083)	(.037)	(.058)
5-9	2.09	2.20	2.03	2.13	.06	.08
	(.025)	(.049)	(.032)	(.039)	(.028)	(.034)
10-20	2.18	2.44	1.99	2.26	.20	.17
	(.020)	(.036)	(.029)	(.038)	(.033)	(.035)

Note: Standard errors are in parentheses.

Table 4 - Tenure Coefficients from Wage Regressions by Predisplacement Union Status Using Predisplacement Characteristics

	Pre-Wage		Post-Wage		ln(Pre-Wage/Post-Wage)	
	Pre-U	Pre-NU	Pre-U	Pre-NU	Pre-U	Pre-NU
	(1)	(2)	(3)	(4)	(5)	(6)
CDWS						
1-4 yrs	0.01	0.12	-0.09	0.09	0.09	0.03
	(.043)	(.027)	(.059)	(.029)	(.052)	(.027)
5-9 yrs	-0.01	0.19	-0.17	0.15	0.16	0.04
	(.049)	(.044)	(.076)	(.045)	(.068)	(.042)
10-20 yrs	-0.01	0.22	-0.15	0.18	0.14	0.05
	(.053)	(.062)	(.076)	(.080)	(.065)	(.072)
N	389	830	389	830	389	830
R^2	.31	.35	.23	.21	.11	.07
COEP						
1-4 yrs	- 0.06	0.03	-0.17	0.03	0.10	0.01
	(.065)	(.035)	(.060)	(.035)	(.063)	(.035)
5-9 yrs	- 0.02	0.15	-0.20	0.05	0.18	0.10
	(.088)	(.055)	(.080)	(.049)	(.091)	(.049)
10-20 yrs	0.10	0.28	-0.26	0.22	0.37	0.06
	(.082)	(.057)	(.096)	(.067)	(.086)	(.072)
N	285	637	285	637	285	637
R^2	.38	.40	.42	.30	.24	.08
OML						
3-4 yrs	0.05	0.12	-0.04	0.02	0.09	0.10
	(.045)	(.048)	(.031)	(.071)	(.032)	(.053)
5-9 yrs	0.03	0.04	-0.11	0.00	0.14	0.04
	(.036)	(.070)	(.042)	(.076)	(.038)	(.059)
10-20 yrs	0.14	0.19	-0.15	0.08	0.27	0.10
	(.055)	(.070)	(.078)	(.075)	(.076)	(.064)
N	340	165	340	165	340	165
R^2	.24	.52	.08	.31	.11	.11

Note: Standard errors, in parentheses, for the COEP and CDWS regressions are heteroskedasticity consistent; those in the OML regressions are heteroscedasticity consistent and adjusted to account for the clustered nature of the sample. Another OML regression (not shown) with additional controls: ln(local population) and the local unemployment rate, gave almost identical results. Additional control variables (coefficients not shown) are as shown in Table 1. Also, included in the CDWS and COEP regressions are 9 provincial dummy variables, and the CDWS and COEP have 11 and 12 industry dummy variables respectively.

Table 5 - Predisplacement Tenure Coefficients from Wage Regressions Using Predisplacement Characteristics and Firm Dummy Variables - OML

	Pre-Wage		Post-Wage		ln(Pre-Wage/Post-Wage)	
	Pre-U	Pre-NU	Pre-U	Pre-NU	Pre-U	Pre-NU
	(1)	(2)	(3)	(4)	(5)	(6)
3-4 yrs	0.05	0.11	-0.04	0.06	0.09	0.05
	(.035)	(.042)	(.039)	(.068)	(.028)	(.047)
5-9 yrs	0.05	0.03	-0.06	-0.01	0.11	0.04
	(.031)	(.064)	(.035)	(.054)	(.026)	(.066)
10-20 yrs	0.05	0.10	-0.16	0.05	0.22	0.05
	(.027)	(.057)	(.070)	(.040)	(.072)	(.060)
R^2	.56	.67	.25	.46	.22	.23
N	340	165	340	165	340	165

Note: The standard errors are heteroscedasticity consistent and adjusted to account for the clustered nature of the sample. Not shown are the coefficients for: age, age^2, education, marital status, and a blue collar indicator as well as firm dummies. Another regression (not shown) with additional controls: ln(local population) and the local unemployment rate, gave almost identical results.

Table 6 - Postdisplacement Unionization Rates by Predisplacement Tenure and Union Status

	USDWS		COEP		OML	
	Pre-U	Pre-NU	Pre-U (%)	Pre-NU (%)	Pre-U (%)	Pre-NU (%)
Tenure	(1)	(2)	(3)	(4)	(5)	(6)
<1[1]	75	6	70	14	59	23
1-4[1]	82	8	46	10	65	23
5-9	38	7	33	9	55	19
10-20	40	0	15	3	49	24
Total	58	7	57	12	55	22

Note: In the USDWS postdisplacement union status is asked only of those in rotation groups 4 and 8. Sample sizes are thus only 33 and 211 for the predisplacement union (membership) and nonunion samples used to calculate these postdiplacement union (coverage) rates by tenure. Both the COEP and OML surveys report union coverage for the entire sample.

[1]Tenure categories in the OML are 0-2 and 3-4 years respectively.

Table 7 - Unadjusted and Regression-Adjusted Postdisplacement Wages Conditioning on Union-Status Transition Paths: COEP

	COEP		OML	
	U-U	NU-NU	U-U	NU-NU
	(1)	(2)	(3)	(4)
Unadjusted log Wages				
Tenure:				
<1 yrs	2.78	2.26	2.15	2.09
	(.031)	(.023)	(.043)	(.077)
1-4 yrs	2.56	2.31	2.16	2.04
	(.074)	(.033)	(.038)	(.104)
5-9 yrs	2.65	2.28	2.13	2.11
	(.067)	(.046)	(.034)	(.041)
10-20 yrs	2.43	2.57	2.15	2.29
	(.063)	(.075)	(.028)	(.047)
Tenure Coefficients From log Wage Regressions				
1-4 yrs	-0.15	0.02	-0.03	0.07
	(.070)	(.037)	(.034)	(.095)
5-9 yrs	-0.19	-0.03	-0.09	-0.05
	(.108)	(.054)	(.029)	(.070)
10-20 yrs	-0.34	0.20	-0.05	0.06
	(.163)	(.060)	(.032)	(.066)
R^2	.38	.26	.16	.34
N	161	557	186	128

Note: Standard errors are in parentheses; those for the regressions are heteroskedasticity consistent. Not shown are the coefficients for: age, age^2, 5 education dummies, marital status, a dummy variable to indicate whether the plant closed or moved, and 9 provinces.

Table 8: Postdisplacement Tenure-Wage Coefficients that control for changes in Industry, USDWS and COEP

	Pre-U		Pre-NU	
	Same Ind.	Change Ind.	Same Ind.	Change Ind.
	(1)	(2)	(3)	(4)
USDWS				
3-4 yrs	-0.16	-0.08	0.02	0.17
	(.098)	(.139)	(.068)	(.044)
5-9 yrs	-0.15	-0.04	0.05	0.17
	(.139)	(.165)	(.091)	(.055)
10-20 yrs	-0.05	-0.16	0.08	0.12
	(.106)	(.147)	(.100)	(.068)
R^2	.41	.26	.33	.25
N	108	163	411	1397
COEP				
3-4 yrs	-0.14	-0.16	0.01	0.02
	(.078)	(.118)	(.048)	(.057)
5-9 yrs	-0.37	-0.20	-0.07	0.04
	(.092)	(.112)	(.069)	(.070)
10-20 yrs	-0.23	-0.57	0.23	0.20
	(.129)	(.141)	(.114)	(.094)
R^2	.57	.55	.34	.30
N	180	105	297	340

Note: Heteroskedasticity consistent standard errors are in parentheses. Additional control variables (coefficients not shown) are as listed in Table 1. Also included in the USDWS regressions are 3 region, and 9 industry, dummy variables; and the CDWS regressions contain 9 provincial, and 11 industry, dummy variables.

Notes

1. To our knowledge, this was first pointed out by Kletzer (1989). Standard search and matching models (e.g. Burdett 1978; Topel 1991) cannot generate this correlation because these models generally treat displacement –in contrast to voluntary job mobility– as a re-initialization of the matching process with a random new draw from the overall wage offer distribution.

2. Before 1994, the United States' Displaced Worker Survey did not contain measures of individual union status; thus most authors do not distinguish union and nonunion workers. Some authors (e.g. Neal 1995) attempted to generate results that are representative of the nonunion population by restricting attention to industries with very low unionization rates. Very little is known, however, about the union population.

3. The lesser tendency of unionized firms to reward both observed and unobserved ability has been well documented in a number of contexts, including Freeman (1982) and Lemieux (1998). Note also that our results do *not* imply that union workers, as a group, are negatively selected relative to nonunion workers; except at the highest tenure levels the postdisplacement wages of workers displaced from union jobs exceed those of workers displaced from nonunion jobs. Our findings are more suggestive of a world in which unionized firms initially hire a positively selected group of workers, of whom the most able eventually leave.

4. Perhaps surprisingly, it is not widely appreciated that the standard categorization of skills into specific (useful only in the current firm) and general (useful in both the current and alternative firms) is not exhaustive. We call the missing category “alternative skills” (useful only in alternative firms). Examples are obvious and widespread: for example any kind of literacy, numeracy or interpersonal skills that are not used in one’s current job but might very well be useful elsewhere.

5. For a comparison of the Canadian and American displaced worker surveys see Zagorsky (1995).

6. For more details on the design and characteristics of the COEP see Browning, Jones and Kuhn (1994).

7. The CDWS also has a measure of current wages, but it is not useful in this study since it is only asked of those who are still in the first job found following displacement and the sample size is very small.

8. For the CDWS at least one of the two wages is missing for 32% and 38% of the respective union and nonunion workers who satisfy the above criteria. The randomness of this missing wage information may of course be a concern, as it is in all analyses of wage changes in displaced worker surveys. In an attempt to understand what influence this might have on our study, predisplacement wage regressions replicating those in Table 4 were run for all those observations for whom we have appropriate wage data. The coefficients in each were very similar.

9. Inasmuch as nonunion firms may selectively lay off individuals of lower ability, while union firms may be more constrained by seniority systems, displaced nonunion workers might represent a more negatively selected sample and, therefore, experience lower postdisplacement wages than would an average nonunion worker were he displaced. As a result, our estimates may understate the level of postdisplacement wages for the average nonunion worker.

10. Recall that the USDWS asks only a subset of the respondents about their postdisplacement union status, and that union status is measured as membership in the predisplacement job, but as coverage in the postdisplacement job.

11. Recall that the U-U path in the USDWS contains only 30 observations. The NU-NU path

has 211, but the results are essentially identical to those presented in Table 2.

12. Although the CDWS captures *current* industry, this does not accord with the wage measure which is for the *first* job following displacement. Given the resulting small size of the sample where the first job is also the current one, we do not pursue this issue with that data set.

13. Our finding that alternative wages vary substantially *among* union workers of different tenure levels suggests, for example, that empirical tests of "monopoly" versus "efficient contracts" models of union wage-setting (e.g. Brown and Ashenfelter 1986; MaCurdy and Pencavel 1986) may need to be rethought. These tests are based on the response of union employment to a single (and often imprecisely-measured) alternative wage for *all* union workers.

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