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Richard Edwards
University of Massachusetts, Amherst, MA, redwards1@unl.edu

Paul Swaim
University of Massachusetts, Amherst, MA

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Union-Nonunion Earnings Differentials and the Decline of Private-Sector Unionism

By RICHARD EDWARDS AND PAUL SWAIM*

Recent years have been difficult ones for the American labor movement. Especially during the past half-decade, the economic and political environment for unions has become increasingly hostile—dominated by a growing anti-union sentiment in management; the adverse effects of industrial restructuring, import competition, deregulation, and high unemployment; and the tightening constraints of a labor law and NLRB enforcement mechanism that have become markedly less supportive of unions. These and other external changes, as well as some continuing internal weaknesses, have thrust unions into a period of declining union membership, eroding bargaining strength, repeated contract concessions, and what at least some observers have perceived as the beginning of a "new era" in industrial relations (Thomas Kochan and Michael Piore, 1985; Edwards and Michael Podgursky, 1986).

In the specific area of wage setting, recent and continuing concession bargaining by unions has attracted the most attention. Union retreats in steel, autos, and transportation have been highly publicized, but several recent studies have suggested that the decline in union bargaining strength has extended as well to construction, retail foodstores, and other industries not immediately affected by such pressures as import competition or deregulation. A large and unprecedented fraction of settlements now involve wage freezes or reductions. And, if current union members face bleak settlements, the proliferation of two-tier wage provisions may presage even more severe cuts for future employees (Charles Craypo, 1981; Daniel Mitchell, 1985).

Despite these developments, there remains the question of whether bargaining under these conditions has in fact resulted in a substantial compression of union-nonunion earnings differentials. The union-nonunion wage effect has typically been estimated by cross-section studies using micro data on union and nonunion workers (Richard Freeman and James Medoff, 1981, Table 1). But until recently it has not been possible to replicate these studies for the 1980's, since collection of union membership data in the *Current Population Survey* (*CPS*) was temporarily suspended.

Drawing upon recently released *CPS* data for 1984, we find that the union relative wage gap in the private sector remains substantial and does not seem to have narrowed since 1979. Indeed, in contrast to what perhaps has been commonly assumed, union and nonunion wages may actually have drifted farther apart. While the causes of this outcome remain to be studied, we conclude that the most striking aspect of recent experience is not a rapid compression of union earnings premiums, but rather their general persistence. In effect, labor's changing circumstances seem to be reflected in a quantity rather than a price adjustment: the union wage differential has not changed, and instead there has been a rapid substitution of nonunion for union workers.

I. Model and Results

In order to determine the effect of unionization on earnings we estimated the following reduced-form earnings equation:

$$\ln(W_i) = B_1X_i + B_2U_i + e_i. \tag{1}$$

The dependent variable is the natural log of hourly earnings for the ith worker. The term X_i is a set of independent variables including demographic variables, education, region, oc-

*Professor and Assistant Professor of Economics, respectively, University of Massachusetts, Amherst, MA 01003. This paper is part of a joint project undertaken with Michael Podgursky. We thank Marc Kitchel for excellent research assistance.

cupation, and industry dummies; these variables capture variation in human capital investments and other worker and job characteristics influencing earnings. The U is a dummy variable taking the value one if the worker is covered by a collective bargaining agreement at his or her current job. Thus, the estimated coefficient of this variable measures the natural log of the ratio of union to nonunion earnings, which is approximately the proportionate union-nonunion earnings gap (the exact proportionate differential is $\exp(B_2)-1$). We use a union coverage variable rather than a union membership dummy since nonunion workers covered by collective bargaining agreements (some 2–3 percent of the labor force) by law receive identical benefits as union members. Hence, a union membership dummy may yield a biased estimate of the union wage effect. Finally, we assume that unmeasured variation in remunerative worker and job characteristics is uncorrelated with our independent variables and is imbedded in the mean-zero independent and identically distributed residual e_i.

Equation (1) was estimated on a large sample of nonfarm private wage and salary workers from the May and June, 1984 *CPS*. In order to capture industry interaction effects, we stratified the sample by broad industry groups and estimated equation (1) within each group. For comparison we estimated the same models on a smaller sample of workers from the May 1979 *CPS Pension Survey*, which includes matched June earnings records (Wesley Mellow, 1983). In both samples we excluded managers, professional workers, and private household workers in order to focus on a somewhat more homogeneous and organized segment of the labor force. F-tests allowed us to reject the hypothesis that the earnings equations were identical across these broad industry groups. Race, sex, and Hispanic ethnicity union interactions were insignificant in the stratified regressions in both years.

Ordinary least squares estimates of the union coverage coefficient (B_2) and related statistics are reported in Table 1. The coefficient of union coverage took its expected positive coefficient and was highly significant in all sectors in both 1979 and 1984. The

TABLE 1—ESTIMATED UNION-NONUNION EARNINGS DIFFERENTIALS AND UNIONIZATION RATES: MAY–JUNE 1979, AND MAY–JUNE 1984[a]
(Dependent Variable = Natural Log of Hourly Earnings)

	May–June, 1979		May–June, 1984		
Industry	$\bar{U}$	$\hat{B}$	$\bar{U}$	$\hat{B}$	$\hat{B}$
Mining, Forestry and Fisheries	.470	.222[c] (.049)	.204	.218[c] (.062)	−.004 (.079)
Construction	.375	.369[c] (.027)	.274	.436[c] (.024)	.067 (.036)
Manufacturing	.440	.130[c] (.011)	.320	.150[c] (.011)	.020 (.016)
Trans., Comm., and Public Utilities	.615	.229[c] (.026)	.462	.300[c] (.023)	.069[b] (.035)
Trade and Services	.112	.202[c] (.015)	.088	.217[c] (.014)	.015 (.021)

Sources: May 1979 *CPS Pension Survey* and 1984 Earnings File. Microdata tapes available from the Bureau of Labor Statistics.

[a]Means and *OLS* estimates of the coefficient of union contract coverage dummy variable were obtained by estimating equation (1) within each of the industrial groups shown above. Within each industry group, managers, professionals, and private household workers were excluded. In addition to union coverage, the regression included controls for education, race, Hispanic ethnicity, sex, years of labor market experience, household head, veteran, SMSA residence, region, part-time work, and 7 occupation dummy variables. Industry dummies varied with the sample, ranging from 1 in Mining, Forestry, and Fisheries, to 16 in manufacturing. A complete set of regression coefficients and related statistics are in a separate appendix available from the authors.

[b]Significant at a .05 level of confidence.

[c]Significant at a .01 level of confidence.

union coefficient differed significantly across sectors, and in 1984 it was lowest in Manufacturing and highest in Construction and Transportation, Communication, and Public Utilities. This cross-section pattern may reflect a more elastic demand for union labor in Manufacturing since offshore production, plant relocation, or imports do not provide as ready substitutes for union labor in the latter industries as they do in Manufacturing.

Our focus, however, is not on cross-section patterns, but on changes over time—and here the results are rather surprising. The unionization rate in our sample declined from 27.8 to 19.0 percent over these five years, reflecting the sharp medium-term decline reported by the BLS and continuing the longer-term trend noted by many researchers

(Larry Adams, 1985; Freeman and Medoff, 1984, ch. 15). Sharp declines also occurred within each of our broad industry groups, ranging from approximately 3 percentage points in services to over 25 percentage points in Mining, Forestry, and Fisheries.

In spite of very sharp declines in the unionization rate, the union-nonunion earnings differential remained relatively stable or widened within each of these broad industrial sectors. The last column of the table shows that, with the exception of Mining, Forestry, and Fisheries, the estimated union-nonunion differential widened in every industry group examined. The widening is statistically significant (and then only at a 5 percent level of confidence) only in Transportation, Communication, and Public Utilities.

The estimates in Table 1 constrain the union wage effect to be the same across industries within these rather broad industrial groups. We also examined possible industry-union interaction effects within several of these broad industrial groups. In Transportation, Communications, and Public Utilities, we tested whether the union-nonunion wage gap in recently deregulated industries exhibited a different trend. The only significant interaction was trucking services, where the union-nonunion gap dropped from .328 to .308, a statistically insignificant decline. Union-industry interactions in the earnings equation for manufacturing showed a widening gap in nine industries and narrowing gap in nine others. In no industry, however, was the change significant at a 5 percent level of confidence. In four cases the change was significant at a 10 percent level of confidence: three with a widening gap (Food and Tobacco, Basic and Fabricated Metal, and Electrical Machinery and Equipment); and one with a compressed gap (Rubber and Plastic Products). The 1979 employment-weighted sum of the changes was +1.2 percentage points, approximately equal to the change reported in Table 1.

II. Discussion

These findings are surprising given the considerable attention that has focused on concession bargaining by unions. Before attempting to explain them, we note that our findings are consistent with more aggregated wage-trend data published by the Bureau of Labor Statistics. The Wage and Salary Employment Cost Index indicates that between 1979 and 1984 union wages rose 5.1 percent more than nonunion wages in manufacturing and 5.6 percent more in nonmanufacturing industries (U.S. Department of Labor, 1985, p. 38). Nonetheless, we view the results in Table 1 as preliminary since the matched May–June 1979 sample may not be fully comparable with the 1984 sample, which simply pools May and June *CPS* records (difficulties in matching records could have resulted in nonrandom attrition from the sample). We are currently constructing data files for 1978 and 1979 that avoid this potential problem.

Perhaps the most straightforward interpretation of these findings would focus on price-quantity adjustments. It could be argued that, faced with increasingly elastic demand for their members' services, most unions have evidently chosen to maintain established wage levels; the result has been dramatic reductions in employment. This interpretation, of course, is at odds with journalistic accounts emphasizing a perceived willingness of many union negotiators to trade substantial wage and benefit concessions for enhanced job security. It would similarly confound Colin Lawrence and Robert Lawrence's provocative 1985 analysis of recent wage developments in manufacturing, since their analysis relies heavily on the notion that union wage demands vary inversely with the elasticity of demand. But, in this interpretation, our results would indicate that unions assign a lower priority to maintaining membership levels than is commonly believed.

There are, however, several reasons to be cautious about this wage inflexibility interpretation. First, in the Wage and Salary Employment Cost Index mentioned above, nonunion earnings growth exceeded union earnings growth in manufacturing industries in 1983 and 1984, and in nonmanufacturing industries in 1984. Thus, it may be that concession bargaining has only very recently become widespread enough to be reflected in broad wage averages. In the same vein, as was already mentioned, two-tier wage scales have sometimes been implemented that partially shield existing employees from the full brunt of the compensation cuts that have been negotiated. If these provisions remain in effect, the impact on earnings will continue to grow as the current work force is progressively replaced by new workers.

Second, while our list of control variables is extensive and includes general labor market experience, we were not able to control for a possible rise in the average seniority of unionized workers, since data on employer-specific job tenure was not collected in the 1984 *CPS* surveys. Seniority plays an important role in layoff and recall in most union contracts, so it is highly likely that workers who have weathered the shakeout in the union sector have greater seniority relative to an average nonunion worker in 1984 as compared to 1979. Thus, the coefficients for union coverage in the 1984 earnings regressions could be biased upward relative to the 1979 coefficients and the extent of union wage flexibility correspondingly understated.

The May 1979 *CPS*, which included information on employer-specific seniority, showed, not surprisingly, a strong correlation between seniority and age. Indirect evidence on the magnitude of effect of changes in seniority on estimated union wage effects can thus be gleaned from an examination of union-nonunion age gaps. In 1979, union workers were already older than nonunion workers, yet these average gaps widened appreciably between 1979 and 1984 in all sectors except trade and service. If we assume that the same five-year changes apply to comparisons of average seniority levels between union and nonunion workers, then the product of these estimated increases in average seniority gaps with estimates of the marginal earnings effect of more seniority provides an indication of the extent of upward bias in the union coverage coefficients. The results of these calculations (described in a statistical appendix available upon request), indicate that plausible shifts in average seniority levels were appreciable, but probably too small to have obscured a substantial compression of union sector wages.

The apparent predominance of quantity adjustments over price adjustments revealed in Table 1 may also reflect a second compositional effect that further obscures the extent of union wage flexibility. Just as higher seniority members were probably more successful in negotiating the rapid decline in union membership within each bargaining unit, differential survival probabilities across bargaining units may also have played a role in maintaining the union wage differentials inherited from the 1970's. One source of the rapid decline in union membership has been the loss of locals representing workers at establishments that were either closed or reorganized to operate nonunion. This attrition may have been most pronounced in establishments where unions already failed to negotiate compensation levels commensurate to those achieved by other union locals. A "survival of the fittest" effect could then have resulted in stable or rising average union-nonunion earnings differentials even though many surviving unions made substantial concessions: the "give-backs" negotiated by surviving unions being offset by the disproportionate elimination of the weakest union contracts from the universe of current contracts.

Although the *CPS* data provide no information with which to test the hypothesis just offered, it is consistent with the higher compensation levels typical of larger employers and with what anecdotal evidence suggests is a more rapid de-unionization of small firms. In some manufacturing industries, for example, where the large, core employers remain unionized, the deunionization of the periphery has been encouraged by the expansion of outsourcing to small, nonunion firms. Similarly, in trucking and construction, large unionized core employers have themselves acquired small nonunion peripheral firms in order to expand into new markets—a phenomenon termed "double-breasting" by unions. Both of these practices particularly disadvantage small, unionized firms that find themselves in direct competition with these nonunion producers. The higher mortality rates of small establishments (David Birch, 1979) combined with the low and falling success rate of unions in representation elections (Freeman and Medoff, 1984, ch. 15) also suggest that union penetration has eroded most rapidly in smaller establishments. Thus, the union-nonunion gap may have widened not because unions continued to obtain relatively higher wage settlements than their nonunion counterparts, but because union control of their jurisdiction is slipping away most rapidly where wages are lowest.

III. Conclusion

Our examination of the changes in unionization rates and earnings differentials between 1979 and 1984 indicates that nonunion workers have very rapidly been substituted for union workers in the face of a stable or slightly widened union-nonunion earnings differential. The failure of the earnings differential to decline is surprising, given the widely reported and apparently substantial erosion of union bargaining power.

Several different interpretations of these results are possible. One could interpret them as suggesting that in fact union bargaining power has not declined. This seems unlikely and inconsistent with the sharp fall in unionization rates and much other evidence. Alternatively, one could argue that our results indicate a lack of wage flexibility in the union sector: while unions have been able to maintain the relative price of their labor, in so doing they have placed the burden of adjustment to their deteriorating circumstances on the quantity response. Such an interpretation might be used to suggest that if unions had been willing to reduce the union wage premium, the resulting quantity adjustment would have been less.

A third explanation would place more stress on the institutional relationships involved to suggest that our results may be masking a somewhat more complicated story. Union jobs may have been lost in marginal and peripheral enterprises due to import penetration, deregulation, and the relative ease with which union workers could be replaced by nonunion workers, resulting in the subsequent retrenchment of the union sector to a pool of older, high-tenure workers in a shrinking core of large establishments. If true, these compositional shifts have apparently more than offset the very poor (by historical standards) wage settlements that a number of unions have been forced to accept. Under these circumstances, there may have been no reasonable concessions that unions could have made that would have stemmed the loss of union jobs.

The last interpretation, while consistent with fragmented and anecdotal data, is obviously speculative and requires confirmation from further survey data. Moreover, the second and third explanations are not mutually exclusive and may be valid in differing degrees for different industries. Finally, the impact of continued concession bargaining on broad measures of union wage effects may become much more pronounced as these compositional effects run their course.

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