# THE EFFECT OF UNIONS ON WAGE INEQUALITY IN THE U.S. LABOR MARKET

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This study uses Current Population Survey micro data for 1973–74 and 1993 to evaluate the effect of changing union membership on trends in male and female wage inequality. Unionization rates of men fell between the two sample periods, with bigger declines among lower skill groups. These trends account for 15–20% of the rise in male wage inequality. Union membership rates of low-wage women also declined, while unionization increased among higher-wage women. On balance, shifting unionization accounts for very little of the rise in female wage inequality. Economy-wide trends in unionization mask a sharp divergence between the private sector, where unionism was declining, and the public sector, where it was rising. Comparisons across sectors suggest that unionization substantially slowed the growth in wage inequality in the public sector.

The fraction of trade union members in the U.S. labor market has fallen dramatically in recent decades (see, for example, Farber 1990; Riddell 1992), while the level of wage inequality has risen (for example, Katz and Autor 1999; Blackburn, Bloom, and Freeman 1990; Bound and Johnson 1992). Since unions historically exerted an equalizing effect on the distribution of wages (Freeman 1980; Freeman and Medoff 1984), many analysts believe that the fall in unionization has contributed to the rise in wage inequality. Indeed, a number of recent studies—including Freeman (1993) and DiNardo, Fortin, and Lemieux (1996)—estimate that the fall in union membership can account for up to one-quarter of the rise in male wage dispersion over the 1980s.[1]

This paper presents new estimates of the effect of changing unionization on wage inequality for male and female workers over

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A data appendix with additional results, and copies of the computer programs used to generate the results presented in the paper, are available from the author at the Department of Economics, University of California–Berkeley, 549 Evans Hall #3880, Berkeley, CA 94720-3880.

[1]The effect of unions on female wage inequality is relatively under-studied. DiNardo, Fortin, and Lemieux (1996) estimated that changing patterns of unionization explain relatively little of the recent rise in female wage dispersion. DiNardo and Lemieux (1997) studied the relative effect of unions on male wage inequality in the United States and Canada.

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the period from 1973 to 1993. The methodology extends the traditional two-sector framework developed by Freeman (1980) for measuring the equalizing effect of trade unionism in two ways. First, explicit attention is paid to the fact that unionization rates vary across the wage distribution, and that union membership has fallen disproportionately for lower-wage workers. This trend has reduced the equalizing effect of unionism in the economy. Second, the method accounts for differences in the relative wage effect of unions on different skill groups. A long-standing hypothesis in the literature is that unions raise wages more for lower-skilled workers (Lewis 1986). Indeed, conventional (ordinary least squares) estimates of the union wage gap for low-skilled workers are large and positive, while estimates for highly skilled men are small or even negative. Taken at face value, this pattern implies a substantial equalizing effect of unions. Evidence presented in Card (1996), however, suggests that unionized workers with low observed skill characteristics tend to have higher unobserved skills than their nonunion counterparts, contributing to their apparent wage advantage. Conversely, union members with higher observed skills tend to have below-average unobserved characteristics, explaining the negative union wage gap for highly educated and experienced workers. Estimates of the equalizing effect of unions that ignore these differential selectivity biases (for example, the re-weighting method used by DiNardo, Fortin, and Lemieux) may therefore overstate the role of unions in compressing wage differences across the skill distribution.

Although private sector union membership rates have declined sharply over the past 30 years, union densities have actually risen in the public sector (Freeman 1988). In light of this divergence, it is interesting to examine the effect of changing unionization on the growth of wage dispersion *within* the public and private sectors. An evaluation of the effect of unions on inequality in the public sector is particularly compelling because many observers believe that rises in public sector unionism have occurred for exogenous reasons, associated with changes in legal barriers to unionization, rather than for potentially endogenous reasons, such as shifts in demand that may have also contributed to rising wage inequality (Freeman 1986). Moreover, an analysis of union wage effects in the public and private sectors presents an opportunity to ask whether unions act differently in the two sectors—in particular, whether unions exert a greater equalizing effect across skill groups in a noncompetitive versus a competitive environment.

## Methods

To illustrate the potential effect of unions on wage inequality, it is useful to begin by assuming that workers can be classified into homogeneous skill groups—for example, categories based on detailed levels of education and age.[2] Let $w_i^n(c)$ represent the log wage that individual $i$ in skill category $c$ would earn in the nonunion sector, and let $w_i^u(c)$ represent the log wage for the same individual if he or she worked in a unionized job. Assume that

$$w_i^n(c) = w^n(c) + \varepsilon_i^n,$$

$$w_i^u(c) = w^u(c) + \varepsilon_i^u,$$

where $w^n(c)$ and $w^u(c)$ are the mean nonunion and union wages for individuals in skill group $c$, respectively, and that the residual components $\varepsilon_i^n$ and $\varepsilon_i^u$ satisfy the conditions

$$E[\varepsilon_i^n] = E[\varepsilon_i^u] = E[\varepsilon_i^n|\text{nonunion}] = E[\varepsilon_i^u|\text{union}] = 0.$$

One interpretation of these assumptions is that all workers in the same skill group are viewed as exchangeable (that is, equally productive) by potential employers. The observed union-nonunion gap in mean wages for workers in skill group $c$ is

$$\Delta_w(c) = w^u(c) - w^n(c).$$

[2]The following presentation borrows from Lemieux (1992); see Card (1992) for a parallel development.

Under the assumption that the conditional expectations of $\varepsilon_i^n$ and $\varepsilon_i^u$ are both zero, this is also the expected wage gain that a nonunion worker in skill group $c$ would receive if she could find a unionized job, or alternatively the expected wage loss that a union worker would suffer if he moved to the nonunion sector.

In addition to affecting the mean level of wages, unions can potentially influence the distribution of wages within skill categories. Let

$$\mathrm{Var}[\varepsilon_i^n | c] = v^n(c)$$

and

$$\mathrm{Var}[\varepsilon_i^u | c] = v^u(c)$$

denote the variances of log wage outcomes for individuals in skill group $c$ in the nonunion and union sectors, respectively. The union-nonunion variance gap for skill group $c$ will be denoted by

$$\Delta_v(c) = v^u(c) - v^n(c).$$

Finally, let $u(c)$ denote the fraction of workers in group $c$ whose wages are set by union contracts. In principle, $u(c)$ may be different from the fraction of trade union members. However, given the limitations of the available data (discussed below), such differences will be ignored.

Under the preceding assumptions, the mean log wage of all workers in group $c$ is

$$(1) \qquad w(c) = w^n(c) + u(c)\Delta_w(c).$$

The second term on the right-hand side of this expression is the average wage *gain* for workers in skill group $c$ associated with the presence of unionism (Lewis 1986), and is simply the product of the union coverage rate and the union wage gap. The variance of log wage outcomes for workers in skill group $c$ is[3]

$$(2) \qquad v(c) = v^n(c) + u(c)\Delta_v(c) + u(c)(1 - u(c))\Delta_w(c)^2.$$

This equation shows that unions exert a "within-sector" effect associated with any change in the dispersion of wage outcomes relative to those in the nonunion sector (the second term in equation 2), and a "between-sector" effect associated with the potential widening of mean wage outcomes between the union and nonunion sectors (the third term in equation 2).

Using equation (2), the variance of wage outcomes across all skill groups can be written as

$$\begin{aligned}(3) \qquad v &= \mathrm{Var}[w(c)] + \mathrm{E}[v(c)] \\ &= \mathrm{Var}[w^n(c) + u(c)\Delta_w(c)] + \mathrm{E}[v^n(c) \\ &\quad + u(c)\Delta_v(c) + u(c)(1 - u(c))\Delta_w(c)^2] \\ &= \mathrm{Var}[w^n(c)] + \mathrm{Var}[u(c)\Delta_w(c)] \\ &\quad + 2\mathrm{Cov}[w^n(c), u(c)\Delta_w(c)] + \mathrm{E}[v^n(c)] \\ &\quad + \mathrm{E}[u(c)\Delta_v(c)] + \mathrm{E}[u(c)(1 - u(c))\Delta_w(c)^2],\end{aligned}$$

where expectations (denoted by E[]), variances (denoted by Var[]), and covariances (denoted by Cov[]) are taken over the skill categories. In contrast, if all workers were paid according to the existing wage structure in the nonunion sector, the variance of wage outcomes would be

$$(4) \qquad v^n = \mathrm{Var}[w^n(c)] + \mathrm{E}[v^n(c)].$$

Thus, the effect of unions on the variance of wage outcomes, relative to the situation that would be observed if all workers were paid according to the existing wage structure in the nonunion sector,[4] is

$$\begin{aligned}(5) \qquad v - v^n &= \mathrm{Var}[u(c)\Delta_w(c)] \\ &\quad + 2\mathrm{Cov}[w^n(c), u(c)\Delta_w(c)] \\ &\quad + \mathrm{E}[u(c)\Delta_v(c)] + \mathrm{E}[u(c)(1 - u(c))\Delta_w(c)^2].\end{aligned}$$

A helpful aid to understanding this equation is to compare it to the simplified basis case in which union coverage rates, wage gaps, and variance gaps are all constant across skill groups (that is, $u(c) = u$; $\Delta_w(c) = \Delta_w$; $\Delta_v(c) = \Delta_v$). In this case the first two

[3]This equation follows from the standard decomposition of a variance into within-sector and between-sector components.

[4]Of course, in the absence of unions, the wage structure in the nonunion sector might change, as Lewis (1986) and others have often pointed out.

terms in equation (5) are zero and the effect of unions on the variance of wages reduces to the simple two-sector formula

$$(5') \qquad v - v^n = u\Delta_v + u(1-u)\Delta_w^2.$$

Relative to this benchmark case, variation in either the union coverage rate $u(c)$ or the union wage effect $\Delta_w(c)$ across skill groups introduces two additional factors into the overall wage dispersion effect. The first is a positive variance component that arises if the union wage gain $u(c)\Delta_w(c)$ varies across groups. The second is a covariance term that may be positive or negative, depending on whether the union wage gain is larger or smaller for higher- or lower-wage workers. If the unionization rate is higher for less-skilled workers, or if the union wage gap is higher for such workers, then the covariance will be negative, enhancing the equalizing effect of unions on wage dispersion.[5]

### Unobserved Heterogeneity

The preceding formulas have to be modified slightly if the union and nonunion workers in a given skill category have different productivity levels and would earn different wages even in the absence of unions. Such a phenomenon will arise if workers have productivity characteristics that are known to employers but not fully captured in the observed skill categories, and if the mean level of these unobserved skills is different between union and nonunion workers in a given skill group. As before, assume that workers are classified into skill categories on the basis of *observed* characteristics, and suppose that

$$(6a) \qquad w_i^n(c) = w^n(c) + a_i + \varepsilon_i^n,$$

$$(6b) \qquad w_i^u(c) = w^u(c) + a_i + \varepsilon_i^u,$$

where $a_i$ represents an unobserved skill component, and $E[\varepsilon_i^n|\text{nonunion}] = E[\varepsilon_i^u|\text{union}] = 0$. Note that $a_i$ is assumed to shift wages by the same amount in the union and nonunion sectors. Let

$$\theta(c) = E[a_i|\text{union}, c] - E[a_i|\text{nonunion}, c]$$

represent the difference in the mean of the unobserved skill component between union and nonunion workers in group $c$. The mean wage gap between union and nonunion workers in skill group $c$ then includes the true union wage premium and the difference attributable to unobserved heterogeneity:

$$E[w_i^u(c)|\text{union}] - E[w_i^n(c)|\text{nonunion}] = \Delta_w(c) + \theta(c).$$

Taking account of unobserved productivity differences between union and nonunion workers, the difference in the variance of wages in the presence of unions and in the counterfactual situation in which all workers are paid according to the nonunion wage structure is

$$(7) \qquad v - v^n = \mathrm{Var}[u(c)\Delta_w(c)] + 2\mathrm{Cov}[w^n(c), u(c)\Delta_w(c)] + E[u(c)\Delta_v(c)] + E[u(c)(1-u(c))\{(\theta(c)+\Delta_w(c))^2 - \theta(c)^2\}].$$

Only the last term of this equation, which reflects the gap in mean wages between union and nonunion workers with the same observed skills in the presence and absence of unions, differs from equation (5).[6]

The possibility that there are unobserved skill differences between union and nonunion workers with the same observable characteristics introduces a difficult empirical problem: how do we distinguish the

[5]If the unionization rate varies across skill groups but the union wage gap and union variance gap are constant (that is, $\Delta_w(c)=\Delta_w$ and $\Delta_v(c)=\Delta_v$), then $v - v^n = u\Delta_v + u(1-u)\Delta_w^2 + 2\Delta_w\mathrm{Cov}[w^n(c), u(c)]$, where $u$ is the average rate of unionization.

[6]The relatively simple form of equation (7) depends crucially on the assumption that unobserved skills are equally valuable in the union and nonunion sectors. If, for example,

$$w_i^u(c) = w^u(c) + k_c a_i + \varepsilon_i^u,$$

where $k_c < 1$, then the formula has to be modified to account for the fact that unionization affects the rewards to unobserved skills. Lemieux (1998) presented a model with this property.

true union wage effect $\Delta_w(c)$ for workers in a given skill group from the heterogeneity component $\theta(c)$? In the absence of unobserved heterogeneity, $\Delta_w(c)$ can be estimated by the mean wage difference between union and nonunion workers in skill group $c$. More generally, however, the observed difference in mean wages between union and nonunion workers reflects the sum of the true union wage effect and the mean difference in unobserved skills.[7]

A natural solution to this problem is to use longitudinal data on union status changers to evaluate the wage gains of union joiners and the wage losses of union leavers. Assuming that unobserved skills are rewarded equally in the union and nonunion sectors, the change in wages "differences out" the unobserved heterogeneity component, leaving only the change in the true union wage premium. In Card (1996) I considered wage changes for a longitudinal sample stratified into five observable skill groups on the basis of predicted wages in the nonunion sector. Lemieux (1992) considered wage changes in a Canadian data set for three similarly defined skill groups. The empirical results in these papers point to two important conclusions. First, in both the United States and Canada the "true" (that is, longitudinally based) union wage effect is higher for less-skilled workers. For example, the results in Card (1996) suggest that the union wage effect ranges from about 30% for men in the bottom quintile of the observed skill distribution to about 10% for men in the top quintile.[8] Second, in both countries union workers with lower observed skills tend to have higher *unobserved* skills than their nonunion counterparts, whereas union workers with higher observed skills tend to have lower unobserved skills than their nonunion counterparts. In terms of the notation introduced above, $\theta(c) > 0$ for lower skill groups and $\theta(c) < 0$ for higher skill groups.

This pattern suggests that the selection process controlling workers' union status differs between more- and less-skilled workers. In Card (1996) I hypothesized that the differences arise because workers must pass two "hurdles" to be observed in a union job. First, they have to find a unionized employer who will hire them. Second, they must prefer a union job to any nonunion alternatives. (See Abowd and Farber [1982] and Farber [1983] for similar models.) For workers with lower observed skills, the first of these hurdles is more likely to bind, because unionized employers typically have a queue of applicants and will reject workers with low education or limited experience unless they possess other skills that are not observable in a typical micro data set. For workers with higher observed skills the second hurdle is more likely to bind if unionized employers offer a "flatter" pay structure that is less attractive to older and better-educated workers. In this case, a union job will be more attractive to workers whose unobserved skills are below average. Thus, union workers with lower observed skills will be positively selected while those with higher observed skills will be negatively selected. In the analysis below I use results from Card (1996) to make a rough adjustment to the observed union wage gaps for workers in different deciles of the predicted wage distribution to account for these different selection biases.

## Data

This paper uses Current Population Survey (CPS) data on wages from the May 1973 and 1974 surveys, and from the 12 monthly surveys in 1993. The May 1973 sample is the first CPS that contains both union status information and wage data for individuals' current jobs. This sample is therefore

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[7]A similar problem arises in the estimation of $\Delta_v(c)$, if Var[$a_i$|union] ≠ Var[$a_i$|nonunion]. In particular, if the distribution of unobserved heterogeneity is not the same in the union and nonunion sectors, conditional on skill group, then there is a distinction between the observed gap in wage dispersion between union and nonunion workers and the gap attributable to the effect of unions. A full consideration of this possibility is beyond the scope of this paper.

[8]Lemieux's results for Canadian men and women are comparable, although the variation in the union wage effect across skill groups for women is smaller than for Canadian (or U.S.) men.

*Table 1.* Union Membership Rates for Men and Women: 1973–74 versus 1993.

| | Men | | | Women | | |
|---|---|---|---|---|---|---|
| *Group* | *1973–74* | *1993* | *Ratio 1993/1973–74* | *1973–74* | *1993* | *Ratio 1993/1973* |
| 1. *All* | 30.8 | 18.7 | 0.61 | 14.1 | 13.3 | 0.94 |
| 2. By Education: | | | | | | |
| < High School | 35.1 | 14.3 | 0.41 | 17.4 | 9.8 | 0.56 |
| High School | 39.3 | 24.5 | 0.62 | 13.5 | 11.8 | 0.87 |
| Some College | 22.7 | 19.5 | 0.86 | 9.0 | 10.5 | 1.17 |
| College or More | 10.7 | 12.4 | 1.16 | 14.8 | 20.4 | 1.38 |
| 3. By Age: | | | | | | |
| 16–30 | 24.9 | 10.4 | 0.42 | 11.5 | 7.3 | 0.63 |
| 31–45 | 32.6 | 20.7 | 0.63 | 15.5 | 15.0 | 0.97 |
| 46–65 | 37.3 | 26.5 | 0.71 | 16.7 | 18.0 | 1.08 |
| 4. By Race: | | | | | | |
| White | 30.2 | 18.3 | 0.61 | 13.5 | 12.4 | 0.92 |
| Black | 37.5 | 23.4 | 0.62 | 18.6 | 19.2 | 1.03 |
| Other | 28.3 | 14.8 | 0.52 | 18.8 | 14.6 | 0.78 |
| 5. By Region: | | | | | | |
| Northeast | 36.8 | 25.4 | 0.69 | 21.2 | 19.1 | 0.90 |
| Midwest | 38.2 | 23.9 | 0.63 | 17.2 | 14.9 | 0.87 |
| South | 19.5 | 10.9 | 0.56 | 6.8 | 7.4 | 1.09 |
| West | 31.3 | 18.7 | 0.60 | 13.8 | 15.4 | 1.12 |
| 6. By Sector: | | | | | | |
| Private | 31.1 | 14.9 | 0.48 | 13.0 | 7.1 | 0.55 |
| Public | 28.9 | 39.3 | 1.36 | 18.0 | 37.3 | 2.07 |
| 7. *No. Obs.* | 43,189 | 86,270 | — | 30,500 | 82,624 | — |

*Notes:* Based on samples derived from the May 1973/74 CPS and 1993 merged outgoing rotation group files. Samples include individuals age 16–65 who are not self-employed, and whose reported or constructed hourly wage is between $2.01 and $90.00 per hour in 1989 dollars. Samples are weighted by CPS sample weights.

the earliest benchmark to compare against later levels of unionization and wage inequality. In view of the relatively small sample size of the monthly CPS, I elected to pool the May 1973 and May 1974 data.[9] The 1993 CPS is the last survey prior to the introduction of a new computer-assisted survey instrument that substantially changed the nature of the earnings questions. I therefore use this sample to measure recent patterns of unionism and wage inequality.

Table 1 presents a descriptive overview of the changes in union membership between the early 1970s and the early 1990s. The samples underlying this table (and all subsequent tables in this paper) include employed individuals between the ages of 16 and 65 who reported an hourly or weekly wage for their main job.[10] Union status is measured by the individual's response to the question, "On this job (the main job) is the respondent a member of a labor union or an employee association similar to a union?" Recent CPS surveys have also collected union coverage information for non-

[9]The wage data from the two surveys are deflated to a common basis using the CPI.

[10]Self-employed workers are excluded. About 20% of individuals refuse to provide information on their earnings to the CPS. The 1993 sample includes allocated wages for these individuals, while non-respondents are dropped from the 1973 and 1974 samples.

*Table 2.* Characteristics of Union and Nonunion Workers in 1973–74 and 1993.

| | Men | | Women | |
|---|---|---|---|---|
| *Variable* | *Nonunion* | *Union* | *Nonunion* | *Union* |
| 1973–74 | | | | |
| 1. Education (years) | 12.3 | 11.2 | 12.1 | 11.7 |
| 2. Experience (years) | 16.9 | 21.5 | 17.6 | 20.8 |
| 3. Nonwhite (percent) | 9.1 | 11.7 | 11.8 | 16.4 |
| 4. Married (percent) | 71.1 | 81.4 | 59.5 | 63.1 |
| 5. Public Sector (percent) | 16.9 | 15.5 | 21.4 | 28.6 |
| 6. Mean Log Wage | 1.323 | 1.519 | 0.947 | 1.177 |
| 7. Unadjusted Union Wage Gap | — | 0.196 | — | 0.230 |
| 8. Adjusted Union Wage Gap | — | 0.178 | — | 0.220 |
| 9. Std. Dev. Log Wages | 0.553 | 0.354 | 0.442 | 0.383 |
| 10. Residual Std. Dev. Log Wage | 0.416 | 0.324 | 0.372 | 0.328 |
| 1993 | | | | |
| 1. Education (years) | 13.1 | 12.8 | 13.1 | 13.9 |
| 2. Experience (years) | 16.8 | 22.1 | 17.5 | 21.2 |
| 3. Nonwhite (percent) | 13.6 | 15.8 | 14.9 | 21.5 |
| 4. Married (percent) | 58.6 | 70.8 | 53.8 | 60.1 |
| 5. Public Sector (percent) | 11.4 | 32.2 | 14.7 | 57.3 |
| 6. Mean Log Wage | 2.359 | 2.613 | 2.153 | 2.466 |
| 7. Unadjusted Union Wage Gap | — | 0.254 | — | 0.313 |
| 8. Adjusted Union Wage Gap | — | 0.168 | — | 0.166 |
| 9. Std. Dev. Log Wages | 0.590 | 0.415 | 0.515 | 0.456 |
| 10. Residual Std. Dev. Log Wage | 0.446 | 0.363 | 0.423 | 0.379 |

*Notes:* See note to Table 1 for sample description. Education categories in 1993 CPS are re-coded to earlier basis. The adjusted union wage gap is the union coefficient from a regression model that also includes education, a cubic in potential experience, and indicators for nonwhite race, Hispanic ethnicity, marital status, and three regions. The residual standard deviation of log wages is the residual standard error from a similar regression fit separately to the union and nonunion samples.

members of unions; however, this information was not collected in the 1973 or 1974 surveys. For comparability over time I therefore use union membership status in both 1973–74 and 1993.[11]

The first row of Table 1 shows the well-known decline in union membership among male workers between 1973 and 1993, along with the fairly stable rate of union membership among women. Comparisons of membership patterns for different subgroups reveal that within the male and female labor forces some groups lost union membership while others gained. Younger and less-educated men and women experienced the largest drops in union membership, whereas union rates among college-educated men and women rose significantly. Men in different race groups and regions had fairly similar relative declines in union membership, whereas the patterns by race and region for women were more variable.

Row 6 of Table 1 illustrates what is probably the most important fact about union membership in the U.S. labor market over recent decades: the dramatic decline in unionism in the private sector (for both men and women) and the fairly rapid rise in public sector unionization. These figures indicate that the relative stability in union membership of women actually masked a shift in unionization from the private sector to the public sector. In 1973–

[11]In 1993, 2.1% of male nonunion members and 2.5% of female nonunion members reported that their wages were set by union contracts.

74, 29% of female union members worked in the public sector. By 1993 this ratio had risen to 57%. For men there was a similar shift—from 16% to 32%.

Table 2 presents comparisons of the characteristics of union and nonunion workers in the two sample periods. Unionized men were typically older, less educated, and more likely to be married than their nonunion counterparts. Interestingly, the mean gap in education narrowed over the two decades (from 0.9 years in 1973–74 to 0.3 years in 1993). This is consistent with the data in Table 1 showing that union densities fell most rapidly for less-educated workers. In 1973–74 unionized women were also older and less educated than nonunion women, but by 1993 the education differential had reversed, again consistent with the rapid rise in union membership of more-educated women.

The sixth through eighth rows of the upper and lower panels in Table 2 report mean log wages of union and nonunion workers in the two sample periods (row 6), the unadjusted differences in mean log wages between the sectors (row 7), and the adjusted wage gaps between union and nonunion workers (row 8), estimated from ordinary least squares (OLS) regression models for log hourly wages that include a union membership dummy and a standard set of control variables.[12] In 1973–74, the unadjusted gaps in mean log wages between union and nonunion men and women were very similar to the adjusted wage gaps (compare rows 7 and 8). In 1993, however, the unadjusted wage gaps were higher than the corresponding adjusted gaps (especially for women), implying that union workers had higher average skill characteristics than nonunion workers.

The ninth and tenth rows of Table 2 present measures of the dispersion in wages within the union and nonunion sectors. The entry in row 9 is just the standard deviation of log wages within each sector, while the entry in row 10 is the residual standard deviation after adjusting for the effects of a standard set of covariates (allowing separate coefficients in the union and nonunion sectors). Note that the union-nonunion difference in the residual standard deviation of earnings is smaller than the difference in the standard deviation in wages, particularly for men. This is mainly attributable to the compressed distribution of observable skill characteristics in the union sector. Comparisons of either measure of wage dispersion between sectors and over time illustrate three important facts. First, wages were less dispersed in the union sector, even after adjusting for differences in observable skills. Second, wages of women (in either union or nonunion jobs) had lower dispersion than wages of men, and the union-nonunion difference in dispersion was smaller for women than men. Third, wage inequality of male and female workers in both the union and nonunion sectors rose substantially between 1973–74 and 1993.

## Effects of Unions on Wage Inequality

### Naive Estimates

As a starting point for evaluating the contribution of changing unionism to the rise in inequality of wages, it is useful to begin with the simple two-sector framework developed by Freeman (1980).[13] Recall that if the union density $u(c)$ and the union relative wage effect $\Delta_w(c)$ are constant across skill groups, then the effect of unions on the variance of wages (relative to what would be observed if all workers were

[12]These are years of education, a cubic in potential experience, indicators for nonwhite race, hispanic ethnicity (available in 1993 only), and marital status, and a set of three indicators for region of residence.

[13]Freeman (1980) did not apply this framework to the overall labor force, but rather used it to study wage inequality within the manufacturing sector, assuming that unions raise the relative wages of blue-collar workers and lower their dispersion but have no effect on white-collar wages.

*Table 3.* Naive Estimates of the Contribution of Unions to Rising Wage Inequality, 1973–74 to 1993.

| *Description* | *Men* | *Women* |
|---|---|---|
| *1973–74* | | |
| 1. Variance of Log Wages | 0.258 | 0.195 |
| 2. Union Rate ($U$) | 0.308 | 0.141 |
| 3. Union Wage Gap ($\Delta w$) | 0.196 | 0.230 |
| 4. Union Variance Gap ($\Delta v$) | –0.180 | –0.049 |
| 5. Between-Sector Effect | 0.008 | 0.006 |
| 6. Within-Sector Effect | –0.056 | –0.007 |
| 7. Total Effect | –0.047 | 0.000 |
| *1993* | | |
| 1. Variance of Log Wages | 0.325 | 0.269 |
| 2. Union Rate ($U$) | 0.187 | 0.133 |
| 3. Union Wage Gap ($\Delta w$) | 0.254 | 0.313 |
| 4. Union Variance Gap ($\Delta v$) | –0.176 | –0.057 |
| 5. Between-Sector Effect | 0.010 | 0.011 |
| 6. Within-Sector Effect | –0.033 | –0.008 |
| 7. Total Effect | –0.023 | 0.004 |
| *Changes from 1973–74 to 1993* | | |
| Change in Variance of Wages | 0.067 | 0.074 |
| Change in Total Effect of Unions | 0.024 | 0.004 |
| Share Attributable to Unions | 0.365 | 0.056 |

*Note:* See text for formulas and Tables 1 and 2 for underlying data.

paid according to the existing nonunion wage structure) is

$$(5') \qquad v - v^n = u\Delta_v + u(1-u)\Delta_w^2.$$

A comparison of the size of this differential over time provides a first-pass estimate of the changing effect of unionism on wage inequality. Table 3 illustrates the application of this formula to data for men and women in 1973–74 and 1993, using the summary statistics from Table 2. Note that if the union density is constant across skill groups, and the union wage and variance effects are constant across skill groups, then it is appropriate to use the *unadjusted* union wage gap and unadjusted union variance gap in equation (5').[14] In fact, under these assumptions the adjusted union wage gap should equal the unadjusted gap, since the union membership rate is orthogonal to individual characteristics. As noted, this was roughly true in 1973–74, but not in 1993.

The results in Table 3 show that ignoring differences in union coverage rates and union effects across groups, the decline in unionism between 1973–74 and 1993 would have been expected to cause the variance of male wages to rise by 0.024 and the variance of female wages to rise by 0.004. Virtually all of the difference for men is attributable to the change in average union density (–0.121 = 0.308 – 0.187) multiplied by the union variance gap ($\Delta v \approx -0.18$). For women, the union variance gap is smaller than for men, and the decline in union density is negligible, so the net contribution of unionism to widening inequality is trivial. As shown at the bottom of Table 3, between 1973–74 and 1993 the variance of wages rose by 0.067 for men and 0.074 for women. Thus, a naive calculation suggests that falling unionism can explain about 36% of the rise in male wage inequality, but none of the rise in female inequality.

## Allowing for Differences across Skill Groups

As pointed out in the first section ("Methods"), there are several reasons to suspect that the naive calculations in Table 3 overstate the role of unions in the growth of wage inequality. Using the framework of equation (5) or (7), it is possible to refine these estimates to allow for differences in union coverage rates and union wage effects by skill group. A necessary first step, however, is to define skill groups. In this study I divided workers into observable groups based on their predicted wages in

[14]To see this, assume that the expected union wage in skill category $c$ is $w^u(c) = w^n(c)+\Delta_w$, and the variance of union wages in group $c$ is $v^u(c) = v^n(c)+\Delta_v$. Assuming that the union rate is constant across all skill groups, the mean union wage is $E[w^u(c)] = E[w^n(c)] + \Delta_w$, and the variance of union wages is $E[v^u(c)] = E[v^n(c)] + \Delta_v$.

*Table 4.* Distribution of Union Membership and Union Effects across Skill Deciles.

| | 1973–74 | | | | 1993 | | | |
|---|---|---|---|---|---|---|---|---|
| | | | *Raw Union Gaps:* | | | | *Raw Union Gaps:* | |
| *Decile* | *Percent Union* | *Decile Share of Union (%)* | *Wage* | *Variance* | *Percent Union* | *Decile Share of Union (%)* | *Wage* | *Variance* |
| *A. Men* | | | | | | | | |
| 1 | 12.1 | 3.9 | 0.42 | 0.02 | 5.3 | 2.9 | 0.29 | 0.05 |
| 2 | 26.2 | 8.5 | 0.35 | –0.01 | 11.0 | 5.9 | 0.29 | 0.02 |
| 3 | 35.2 | 11.4 | 0.31 | –0.03 | 16.3 | 8.7 | 0.29 | –0.02 |
| 4 | 39.5 | 12.8 | 0.25 | –0.08 | 21.1 | 11.3 | 0.29 | –0.05 |
| 5 | 40.9 | 13.3 | 0.23 | –0.06 | 24.8 | 13.3 | 0.26 | –0.07 |
| 6 | 39.9 | 13.0 | 0.18 | –0.07 | 26.7 | 14.3 | 0.20 | –0.11 |
| 7 | 36.3 | 11.8 | 0.11 | –0.10 | 27.6 | 14.8 | 0.17 | –0.11 |
| 8 | 38.1 | 12.4 | 0.08 | –0.09 | 22.2 | 11.9 | 0.10 | –0.11 |
| 9 | 27.8 | 9.1 | 0.00 | –0.09 | 15.9 | 8.5 | –0.01 | –0.11 |
| 10 | 11.6 | 3.8 | –0.10 | –0.10 | 15.5 | 8.3 | –0.05 | –0.09 |
| *B. Women* | | | | | | | | |
| 1 | 9.9 | 7.0 | 0.38 | –0.01 | 5.0 | 3.7 | 0.26 | 0.03 |
| 2 | 15.1 | 10.7 | 0.27 | –0.02 | 7.6 | 5.7 | 0.25 | 0.00 |
| 3 | 15.9 | 11.2 | 0.23 | –0.03 | 10.5 | 7.9 | 0.25 | –0.01 |
| 4 | 14.6 | 10.4 | 0.23 | 0.01 | 10.8 | 8.2 | 0.23 | –0.05 |
| 5 | 13.9 | 9.8 | 0.21 | –0.04 | 13.2 | 9.9 | 0.19 | –0.04 |
| 6 | 14.1 | 10.0 | 0.21 | –0.06 | 14.7 | 11.1 | 0.18 | –0.04 |
| 7 | 14.4 | 10.2 | 0.19 | –0.03 | 11.6 | 8.7 | 0.18 | –0.05 |
| 8 | 15.7 | 11.1 | 0.13 | –0.05 | 13.5 | 10.2 | 0.14 | –0.07 |
| 9 | 13.2 | 9.4 | 0.17 | –0.03 | 17.9 | 13.5 | 0.12 | –0.07 |
| 10 | 14.4 | 10.2 | 0.23 | –0.06 | 27.8 | 20.9 | 0.11 | –0.10 |

*Notes:* Skill deciles are based on the predicted wage in the nonunion sector. The decile share of union represents the percentage of all union workers in the skill decile. The wage gap is difference in mean log wages between union and non-union workers in the skill decile. The variance gap is the difference in variance of log wages between union and nonunion workers in the skill decile. See Table 1 for the sample definition; see text for a description of wage prediction models.

the nonunion sector (conditional on education, age, and race). In particular, I fit a set of wage prediction models to data for nonunion workers by gender and sample period, and then used the resulting coefficient estimates to assign all workers in a gender/year group to 10 equal-sized predicted wage deciles.[15]

Table 4 shows unionization rates, unadjusted union wage gaps, and unadjusted union variance gaps across skill deciles for men and women in 1973–74 and 1993. A key feature of the table is the pattern of union membership rates across skill groups. In 1973–74, union membership rates of men followed an "inverted-U" pattern, with the highest membership rates for workers in the middle of the skill distribution. The pattern of 1993 membership rates was similar, but with lower membership levels for all but the top skill decile. Among women, union rates were fairly constant across skill groups in 1973–74, but were rising across skill groups in 1993, with the highest membership rate in the top group.

[15]The prediction equation includes education, indicators for nonwhites and Hispanics (in 1993 only), a third-order polynomial in experience, interactions of indicators for three main levels of education with linear and quadratic experience, and interactions of the ethnicity dummies with education and linear and quadratic experience.

*Table 5.* Estimates of the Contribution of Unions to Rising Wage Inequality, 1973–74 to 1993.

| *Description* | *1973–74* | *1993* | *Change* |
|---|---|---|---|
| *A. Male Workers* | | | |
| Variance in Log Wages | 0.258 | 0.325 | 0.067 |
| Effect of Unions Using Naive Calculation (Equation 5') | –0.047 | –0.023 | 0.024 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | –0.027 | –0.015 | 0.012 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.019 | –0.011 | 0.008 |
| *B. Female Workers* | | | |
| Variance in Log Wages | 0.195 | 0.269 | 0.074 |
| Effect of Unions Using Naive Calculation (Equation 5') | 0.000 | 0.004 | 0.004 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | 0.000 | –0.002 | –0.002 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.002 | –0.004 | –0.002 |

*Notes:* See text for methods. Raw union wage differentials are actual differences in mean log wages between union and nonunion workers in each skill decile. Adjusted union wage differentials assume that the true union wage effect declines linearly from 0.30 for the lowest skill decile to 0.075 for the highest skill decile.

A second interesting feature is the pattern of the union wage gaps across skill groups. For men in 1973–74, these ranged from 40% for the lowest skill group to –10% for the highest skill group. Taken at face value, these estimates suggest that unions exerted a substantial "flattening" effect on the male wage structure in the early 1970s. This effect seems to have moderated slightly over the next two decades. In particular, the union wage gap for the bottom skill group was lower in 1993 than in 1973–74 (29% versus 42%). For women, the union wage gaps at the bottom of the skill distribution are comparable to those for men, but the decline in the unadjusted wage gaps across the skill distribution is less pronounced. Thus, it appears that unions may exert a more modest flattening effect on the female wage structure than on the male structure.

An important caveat to the interpretation of the wage gaps in Table 4 is the potential role of unobserved heterogeneity. Recall that the unadjusted union wage gap for any skill group is actually a combination of the true union wage effect and a selection effect equal to the difference in the unobserved skills of union versus nonunion workers in the group. If unionized workers in lower skill groups are positively selected and those in higher skill groups are negatively selected, then the flattening effect of unions is overstated by the unadjusted union wage gaps in Table 4. In Card (1996), I used longitudinal CPS data from 1986 and 1987 to estimate unadjusted and adjusted union wage gaps for five skill groups, based on predicted wages in the nonunion sector. As in Table 4, the unadjusted wage gaps decline sharply across the skill groups, from a high of 36% in the bottom skill quintile to a low of –13% in the top skill decile. The *adjusted* union wage gaps (based on wage changes for those who change union status) also decline across the skill distribution, but are lower for the least skilled group (28%) and higher for the most skilled group (11%).[16] Thus, the unadjusted union wage gap for low-skilled men overstates the true union wage effect for these workers, while the unadjusted gap for high-skilled workers actually understates the true union wage effect.

Table 5 presents a series of calculations that use the data in Table 4, together with

[16]A similar pattern arises in Lemieux's (1992) study of the Canadian labor market, based on wage changes for men in three skill groups.

the formulas given by equations (5) or (7), to re-estimate the contribution of changing unionism to rising wage inequality. (For reference, the table also reproduces the naive calculations from Table 3.) The estimates based on equation (5) ignore any unobserved skill differences between union and nonunion workers in the same skill decile, and use the unadjusted union wage gaps in Table 4 as estimates of the true union wage effects. The estimates based on equation (7) use adjusted wage gaps for each skill group derived from my 1986 paper. In the absence of longitudinal estimates for different time periods, or for women, I use a single set of estimates of the "true" union wage effects for each skill group that range from 30% for the lowest skill decile to 8% for the highest skill group.[17]

For women, the estimates of the effect of unionization are qualitatively and quantitatively similar, regardless of the method. In all cases, unions are estimated to have a negligible effect on cross-sectional wage inequality, or on changes in inequality. For men, the results from equation (5) or (7) are qualitatively similar to the results of the naive calculation (based on equation 5′), but the magnitude of the union effect is reduced. The main source of the difference between the estimate based on (5) versus the estimate from the naive two-sector model is that the average "within sector" effect of unions on the variance of wages (that is, the average of the $\Delta_v(c)$ terms across skill groups) is substantially smaller than the gross difference in the variance of wages between the union and nonunion sectors.[18] This difference arises because male union members tend to be drawn from the middle of the skill distribution—consequently, the union-nonunion gap in the overall variance of wages overstates the gap within any skill group. The main source of the difference between calculations based on the selection-adjusted wage gaps versus the unadjusted wage gaps is that the covariance term ($\mathrm{Cov}[w^n(c), u(c)\Delta_w(c)]$) in equation (7) is smaller in magnitude (less negative) when the wage gaps are adjusted for selection biases. The unadjusted gaps overstate both the positive effect of unions on low-wage workers and the negative effect on high-wage workers.

In principle, it is also possible to implement equation (7) using longitudinally based estimates of the union variance effect ($\Delta_v$) rather than the simple differences in the variances of wages between the union and nonunion sectors shown in Table 4. This would be appropriate if unobserved skills are rewarded equally in the union and nonunion sectors (as is assumed in equations 6a and 6b), but the variance of unobserved skill is different in the two sectors. Card (1992) and Lemieux (1992) both presented estimates of the effect of unions on the variance of wages based on the wage outcomes of union status changers. The longitudinal variance gap estimates presented in Card (1992) are relatively noisy, and on average only slightly smaller in absolute value than the corresponding cross-sectional estimates. Lemieux's estimates are also noisy but tend to be smaller (in absolute value) than the cross-sectional estimates. If the cross-sectional variance gaps in Table 4 are viewed as bounding the likely effect of unions on wage dispersion, then the estimates in Table 5 should be interpreted as *upper bound* estimates of the contribution of changing unionization to rising wage inequality. Taken as whole, then, it appears that the effect of unions on widening wage inequality may be relatively modest.

[17]These estimates were obtained by fitting a linear model to the adjusted union wage gaps for the five skill groups used in Card (1996), and then interpolating to a set of 10 skill groups. The adjusted union gap for skill decile $j$ is $0.30 - 0.0244 \times (j - 1)$.

[18]From Table 4, the average value of $\Delta_v(c)$ is about $-0.06$, while from row 4 of Table 3, $\Delta_v = -0.18$.

### Unionization and Inequality in the Public and Private Sectors

In light of the diverging rates of union membership in the public and private sectors, it is interesting to ask how much chang-

*Table 6.* Union Membership Rates in the Public Sector for Men and Women, 1973–74 versus 1993.

| | *Men* | | | *Women* | | |
|---|---|---|---|---|---|---|
| *Group* | *1973–74* | *1993* | *Ratio 1993/1973–74* | *1973–74* | *1993* | *Ratio 1993/1973* |
| *1. All* | 28.9 | 39.3 | 1.36 | 18.0 | 37.3 | 2.07 |
| *2. By Education:* | | | | | | |
| < High School | 30.2 | 29.0 | 0.96 | 17.6 | 24.1 | 1.37 |
| High School | 40.5 | 46.0 | 1.14 | 17.0 | 32.9 | 1.94 |
| Some College | 25.8 | 42.7 | 1.66 | 14.2 | 29.6 | 2.08 |
| College or More | 19.5 | 34.2 | 1.75 | 20.6 | 46.1 | 2.24 |
| *3. By Age:* | | | | | | |
| 16–30 | 24.9 | 28.8 | 1.16 | 16.4 | 27.0 | 1.65 |
| 31–45 | 29.7 | 42.9 | 1.44 | 18.6 | 38.5 | 2.07 |
| 56–65 | 31.7 | 40.8 | 1.29 | 19.2 | 41.5 | 2.16 |
| *4. By Race:* | | | | | | |
| White | 27.8 | 39.7 | 1.43 | 16.9 | 37.5 | 2.22 |
| Black | 37.0 | 39.8 | 1.08 | 23.9 | 36.9 | 1.54 |
| Other | 35.2 | 29.5 | 0.84 | 24.2 | 35.1 | 1.45 |
| *5. By Region:* | | | | | | |
| Northeast | 48.2 | 65.3 | 1.35 | 35.1 | 61.5 | 1.75 |
| Midwest | 31.9 | 44.1 | 1.38 | 20.9 | 41.7 | 2.00 |
| South | 14.7 | 21.0 | 1.43 | 6.0 | 19.9 | 3.32 |
| West | 28.0 | 41.2 | 1.47 | 17.5 | 42.7 | 2.44 |
| *6. By Industry:* | | | | | | |
| Education | 24.1 | 40.2 | 1.67 | 17.9 | 44.9 | 2.51 |
| Health/Hospital | 23.8 | 31.2 | 1.31 | 20.1 | 28.6 | 1.42 |
| Public Admin. | 32.8 | 36.7 | 1.12 | 17.2 | 28.1 | 1.63 |
| Other | 28.9 | 43.1 | 1.49 | 17.6 | 31.7 | 1.80 |
| *7. By Level of Government:* | | | | | | |
| Federal | — | 14.9 | — | — | 26.4 | — |
| State | — | 39.3 | — | — | 30.6 | — |
| Local | — | 39.3 | — | — | 42.8 | — |
| *8. No. Obs.* | 7,081 | 13,583 | — | 6,814 | 17,117 | — |

*Notes:* Based on samples derived from the May 1973/74 CPS and 1993 merged outgoing rotation group files. Samples include individuals age 16–65 who are not self-employed, and whose reported or constructed hourly wage is between $2.01 and $90.00 per hour in 1989 dollars.

ing unionism affected the inequality of wages within and between the two sectors. Tables 6 and 7 present some simple comparisons of unionization rates across different subgroups of the two sectors, while Figures 1 and 2 show unionization rates in the two sectors by predicted skill group in 1973–74 and 1993.[19]

An examination of the data for the public sector in Table 6 and Figure 1 suggests that public sector union rates rose for almost all groups after the early 1970s, with relatively larger gains for workers in the top two predicted skill deciles. Much of this rise is attributable to the rise in unionization among teachers: as shown in Table 6,

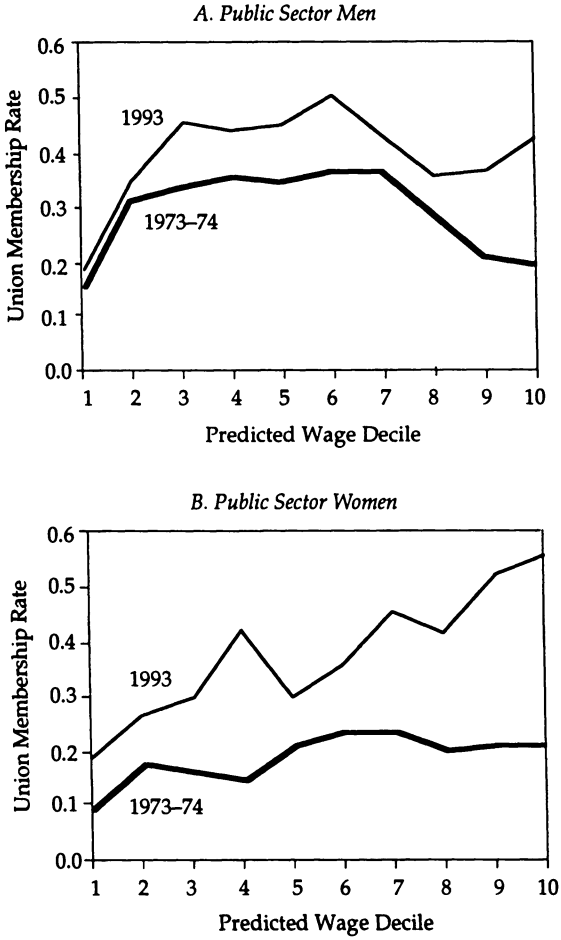
[19]The predicted skill groups for the two sectors are based on sector-specific wage equations, fit to the nonunion workers in each sector. The estimated coefficients from the two wage models were used to assign a predicted nonunion wage for all workers in each sector, and the samples of public and private sector workers were then divided into 10 equal-sized groups.

the union membership rate of men in the education sector rose by 67% between the early 1970s and the early 1990s, while the rate for women rose by 150%. Since teachers and related workers make up such a large share of public sector employment (30% of men and around 50% of women), the rise of teacher unions has been a key determinant of the growth of public sector unionism, accounting for 40% of the rise in union membership among public sector men between 1973 and 1993, and 70% of the rise for public sector women.

The institutional factors controlling the process of unionization in the public sector vary by state: some states prohibit collective bargaining for certain groups of state or local employees, while others have adopted more or less "pro-union" legislation (see, for example, Freeman [1986], and the papers in the volume edited by Freeman and Ichniowski [1988]). This variation is reflected in Table 6 by the widely different levels of public sector unionization across regions. Nevertheless, the rates of growth between 1973 and 1993 are fairly similar across regions, especially for men. It is also interesting to compare unionization rates between the federal, state, and local levels. Unfortunately, information on the level of government was not collected in the 1973 or 1974 CPS surveys, so this comparison is not possible for the base period, but the data for 1993 show generally higher union rates at the local level, and fairly comparable densities across regions at the federal and state levels.

In contrast to the pattern of increasing union membership in the public sector, the data in Table 7 and Figure 2 show uniformly decreasing private sector union rates. On average, union rates fell by about 50%, with larger declines for younger and less-educated workers, but with fairly similar declines across regions and major industries.[20] The similarity of the trends in union membership for men in construction, manufacturing, transportation, communications, and retail trade is notable because these industries experienced very different employment trends over the sample periods. As noted by Farber (1990), the fact that unionization rates declined at comparable rates across industries that experienced very different sectoral growth rates makes it difficult to find support for a theory of union decline linked to sector-specific demand conditions. On the other hand, explanations linked to the institutional or legal environment might be bet-

*Figure 1.* Union Membership Rates in the Public Sector by Skill Group.




[20]The union membership trend for women in construction is very imprecisely estimated because of the small number of women in this industry.

*Table 7.* Union Membership Rates in the Private Sector for Men and Women, 1973–74 versus 1993.

| | *Men* | | | *Women* | | |
|---|---|---|---|---|---|---|
| *Group* | *1973–74* | *1993* | *Ratio 1993/1973–74* | *1973–74* | *1993* | *Ratio 1993/1973* |
| 1. *All* | 31.1 | 14.9 | 0.48 | 13.0 | 7.1 | 0.55 |
| 2. *By Education:* | | | | | | |
| < High School | 35.7 | 13.3 | 0.37 | 17.4 | 8.4 | 0.48 |
| High School | 39.1 | 21.7 | 0.55 | 12.8 | 8.2 | 0.64 |
| Some College | 22.0 | 15.0 | 0.68 | 7.5 | 6.3 | 0.84 |
| College or More | 5.9 | 5.2 | 0.88 | 5.2 | 5.5 | 1.06 |
| 3. *By Age:* | | | | | | |
| 16–30 | 24.9 | 8.5 | 0.34 | 10.3 | 4.6 | 0.45 |
| 31–45 | 33.2 | 16.4 | 0.49 | 14.4 | 8.0 | 0.56 |
| 56–65 | 38.7 | 22.6 | 0.58 | 15.9 | 9.4 | 0.59 |
| 4. *By Race:* | | | | | | |
| White | 30.6 | 14.6 | 0.48 | 12.5 | 6.3 | 0.50 |
| Black | 37.6 | 19.0 | 0.51 | 16.4 | 12.5 | 0.76 |
| Other | 26.7 | 11.8 | 0.44 | 17.0 | 9.4 | 0.55 |
| 5. *By Region:* | | | | | | |
| Northeast | 34.7 | 18.3 | 0.53 | 17.8 | 9.8 | 0.55 |
| Midwest | 39.2 | 20.8 | 0.53 | 16.2 | 8.7 | 0.54 |
| South | 20.5 | 9.0 | 0.44 | 7.0 | 3.8 | 0.54 |
| West | 32.1 | 14.3 | 0.45 | 12.6 | 7.8 | 0.62 |
| 6. *By Industry:* | | | | | | |
| Construction | 40.9 | 21.6 | 0.53 | 2.0 | 3.3 | 1.65 |
| Durable Mfg. | 45.3 | 22.9 | 0.51 | 28.6 | 14.9 | 0.52 |
| Nondurable Mfg. | 37.7 | 21.2 | 0.56 | 27.4 | 11.6 | 0.42 |
| Transportation | 59.8 | 31.7 | 0.53 | 20.7 | 21.2 | 1.02 |
| Communication | 54.3 | 30.6 | 0.56 | 53.9 | 36.2 | 0.67 |
| Public Utilities | 45.8 | 36.6 | 0.80 | 19.7 | 20.8 | 1.06 |
| Retail Trade | 13.3 | 7.2 | 0.54 | 9.5 | 5.5 | 0.58 |
| 7. *No. Obs.* | 36,108 | 72,687 | — | 23,686 | 65,507 | — |

*Notes:* See notes to Table 6.

ter able to explain the uniformity of trends across private sector industries.

How do unions affect the structure of wages in the public versus private sectors? Figures 3 and 4 plot unadjusted union wage gaps by skill decile for men and women in the two sectors. A comparison of the wage gaps by skill level in the public and private sectors suggests that unions exerted a surprisingly similar effect on the wage structures in the two sectors. In particular, the union wage gaps are large and positive for the least skilled men in both sectors, and decline rather quickly across skill groups, with *negative* wage gaps for the most highly skilled men in both sectors. Nevertheless, the average union wage gap is smaller for men in the public sector than in the private sector (Lewis 1988), primarily because the wage gaps for workers in the middle of the skill distribution are lower in the public sector.[21] The unadjusted union wage gaps

[21]Conventional union wage gaps by sector (estimated from simple models fit by sector and gender) are presented in Appendix Table A1, and show wage gaps of about 10% for public sector men in both 1973–74 and 1993, versus 19% for private sector men in both years. One caveat to comparison of union-nonunion wage gaps in the public and private sectors is the possibility that "spillovers" from the unionized sector to the nonunionized sector may be more important in the public sector—see Belman, Heywood, and Lund (1997), for example. If this is the case, then the presence of public sector unionism may have a

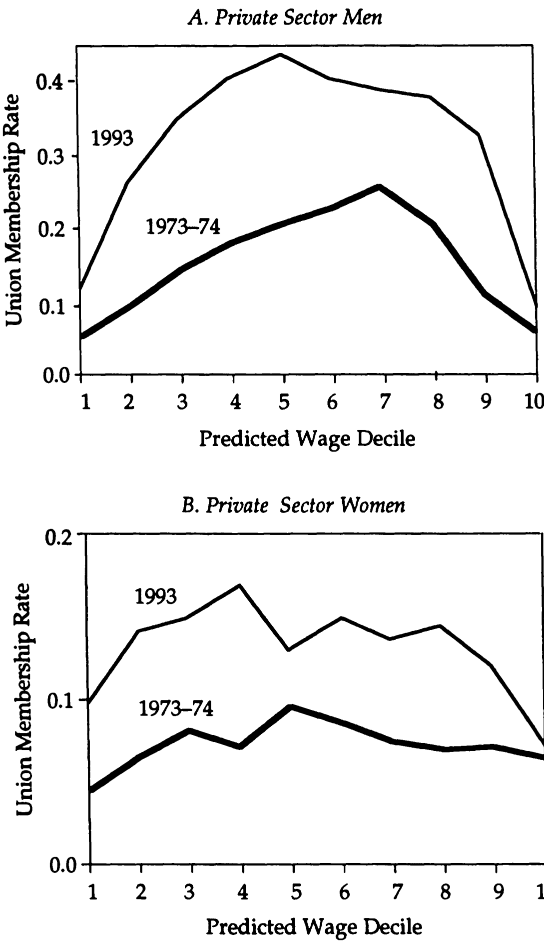
for women in the public and private sectors are even more similar, and indeed the averages of the unadjusted wage gaps across all 10 skill deciles are comparable in the two sectors.[22]

Another interesting feature of the wage gaps in Figures 3 and 4 is the similarity of the patterns in 1973–74 and 1993. Despite the rapid growth of public sector unionism, the effects of unions on wages in the public sector seem to have changed relatively little over the 1970s and 1980s. By the same token, despite dramatic declines in private sector unionization, union wage effects for different skill groups remained fairly constant, with only a modest decline in the union wage advantage for the least skilled men in the private sector. In the absence of longitudinally based estimates of the "true" union wage effects for the two years, these changes must be interpreted cautiously, however, since the processes of selection into the union sector may have also changed, leading to shifts in the magnitude of selection biases in the observed wage gaps.

Table 8 uses data by predicted skill decile for men and women in the public and private sectors to estimate the effects of unions on wage inequality in the two sectors in 1973–74 and 1993. As in Table 5, I have computed the effects of unions using two alternative sets of union wage gaps: the observed gaps (shown in Figures 3 and 4) and adjusted gaps based on the estimates in my 1996 paper.[23]

The results for private sector men and women in Table 8 are fairly close to the results for all workers in Table 5: changes in unionism can account for 15–20% of the rise in wage inequality among private sector men, and virtually none of the rise in inequality for private sector women. The results for public sector workers suggest a more important role for unions. Changes in public sector unionism apparently "held back" rising wage inequality to a significant degree. For men, the estimates suggest that the variance of wages would have risen an additional 30–40% in the absence of unions, while for women the variance of

*Figure* 2. Union Membership Rates in the Private Sector by Skill Group.




relatively bigger effect than is estimated using the counter-factual of the current nonunion wage structure.

[22]As shown in Appendix Table A1, conventional union wage gaps for women based on models fit to the public and private sectors are only slightly smaller in the public sector than in the private sector.

[23]In principle, one might prefer to modify the adjusted gaps for the public and private sectors. I experimented with several alternatives and found that they gave results similar to the ones presented in Table 8.

*Figure 3.* Union Wage Gaps in the Public Sector by Skill Group.

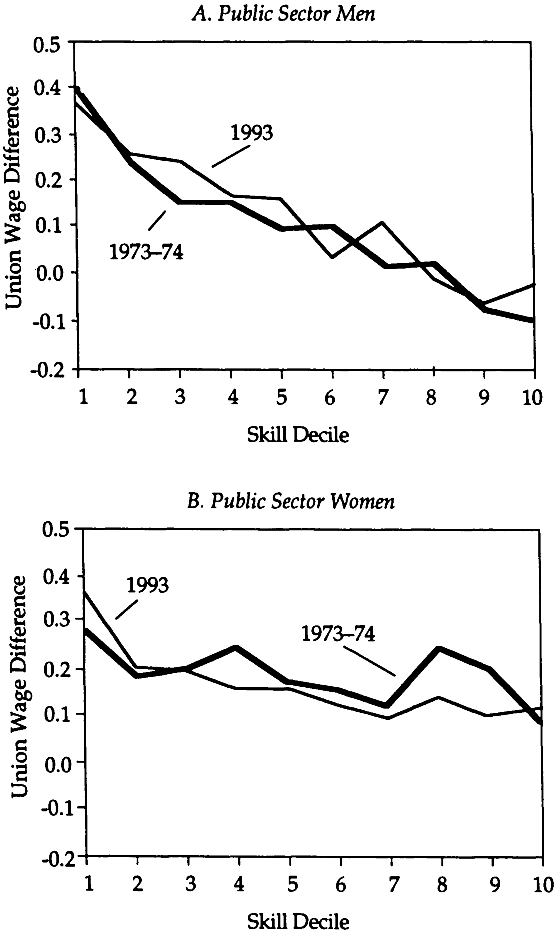
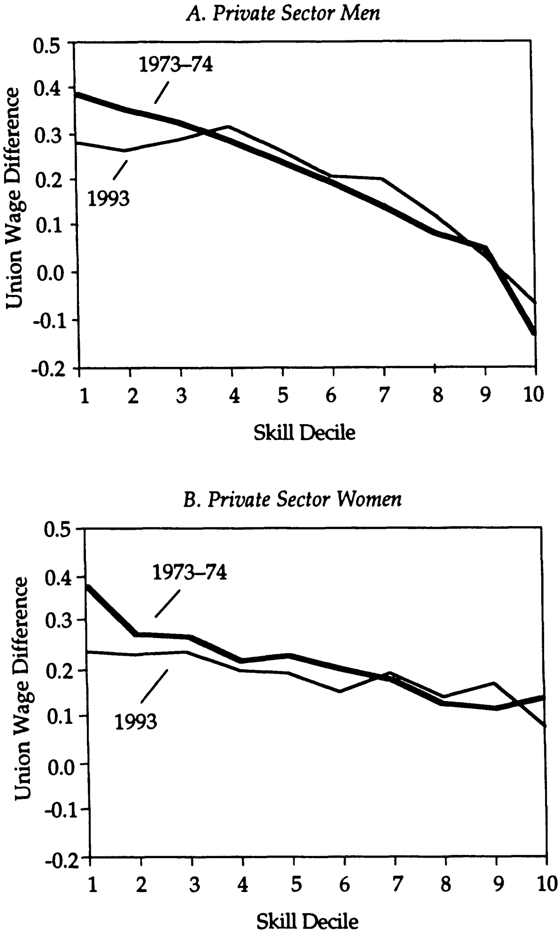



*Figure 4.* Union Wage Gaps in the Private Sector by Skill Group.




wages would have risen an additional 40%. Comparing the changes in wage inequality in the public and private sectors, differential trends in union membership can potentially account for 50–80% of the slower rise in wage inequality for men in the public sector, and 20–30% of the slower rise in wage inequality for women in the public sector.

Another important aspect of the differential trends in unionization in the public and private sectors is the potential effect on public-private wage gaps. For men, union membership rose 10.4 percentage points in the public sector and fell 16.2% in the private sector, implying a 26.6 percentage point divergence. Assuming a mean union wage gap of about 15%, this divergence would have caused mean public sector wages to rise by about 4 percentage points relative to private sector wages. A similar calculation for women shows a 25.2 percentage point divergence in union coverage, also implying a roughly 4 percentage point widening of the mean public-private wage gap. On average, the public-private wage gap for men rose slightly (about 3 percentage points) between 1973 and 1993, whereas it

*Table 8.* Estimates of the Contribution of Unions to Rising Wage Inequality: Public and Private Sectors, 1973–74 to 1993.

| *Description* | *1973–74* | *1993* | *Change* |
|---|---|---|---|
| A. *Public Sector Male Workers* | | | |
| Variance in Log Wages | 0.233 | 0.266 | 0.033 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | –0.029 | –0.043 | –0.014 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.024 | –0.033 | –0.009 |
| B. *Public Sector Female Workers* | | | |
| Variance in Log Wages | 0.204 | 0.237 | 0.033 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | 0.001 | –0.013 | –0.014 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.003 | –0.016 | –0.013 |
| C. *Private Sector Male Workers* | | | |
| Variance in Log Wages | 0.260 | 0.328 | 0.068 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | –0.024 | –0.009 | 0.015 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.018 | –0.008 | 0.010 |
| D. *Private Sector Female Workers* | | | |
| Variance in Log Wages | 0.173 | 0.264 | 0.091 |
| Effect of Unions Using Raw Union Wage Differentials (Equation 5) | –0.002 | 0.000 | 0.002 |
| Effect of Unions Using Adjusted Differentials (Equation 7) | –0.002 | –0.001 | 0.001 |

*Notes:* See notes to Table 5.

fell about 8% for women.[24] Thus, differential shifts in unionization can potentially explain most of the movement of the mean public-private wage gap for men over the two decades, but none of the shift for women. Indeed, in the absence of changing relative union patterns, the public sector wage gap for women would have fallen even faster.

## Conclusions

The primary objective of this paper has been to reassess the connection between declining unionization and widening wage inequality using data for men and women from the early 1970s and early 1990s. The evidence points to three main findings on this issue. First, since the fraction of women belonging to unions was relatively stable over the two decades under examination, shifts in unionization explain almost none of the rise in overall wage inequality among female workers. Second, the decline in union membership among men explains a modest share—15–20%—of the rise in overall male wage inequality. Third, within the public sector, rising unionism was a significant force in forestalling rising wage inequality for both male and female workers. For men, the differences in trends in union membership between the public and private sectors can explain 50–80% of the slower growth of wage inequality in the public sector than in the private sector. For women a similar calculation shows that differences in unionism can explain 20–30% of the difference in the growth of wage inequality between the sectors.

A secondary goal of the paper has been to develop a deeper understanding of union membership patterns and union wage effects in the labor market as a whole and in the public and private sectors. As late as

[24]In 1973/74, public sector men earned about 1% lower wages than private sector men, controlling for education, experience, and race, whereas public sector women earned about 14% more than private sector women, controlling for the same factors. In 1993, comparable public-private wage gaps were 2% for men and 6% for women.

1974, union membership in the U.S. economy was concentrated among men with average or slightly below-average education working in the private sector. In 1993, the highest union membership rates occurred for highly educated women in the public sector. Despite this dramatic shift, an important characteristic of unions—their tendency to raise wages more for workers with lower measured skills—persisted. Indeed, there was remarkable stability in the structure of union-nonunion wage gaps across different skill groups over the sample periods. A comparison of union relative wage structures in the public and private sectors over time suggests that unions exerted about the same effect on different skill and gender groups in the two sectors, and that despite the dramatic shifts in union membership, the structure of union relative wage effects was about the same in the mid-1990s as in the mid-1970s. How and why unions were able to maintain such a stable effect on the structure of wages among their membership is an interesting question for further research.

**Appendix Table A1**
**Conventional Union Wage Gaps by Sector**

| | *Men* | | | *Women* | | |
|---|---|---|---|---|---|---|
| *Year* | *All* | *Public* | *Private* | *All* | *Public* | *Private* |
| 1973–74 | 0.178 | 0.095 | 0.194 | 0.220 | 0.182 | 0.213 |
| | (0.004) | (0.011) | (0.005) | (0.006) | (0.020) | (0.007) |
| 1993 | 0.168 | 0.096 | 0.194 | 0.166 | 0.142 | 0.183 |
| | (0.004) | (0.008) | (0.005) | (0.004) | (0.007) | (0.007) |

*Notes:* Standard errors in parentheses. Entries are estimated union coefficients from weighted OLS regression models that include years of education, a cubic in experience, dummies for marital status, nonwhite race, Hispanic ethnicity (in 1993 only), and three region dummies. The estimation uses CPS sampling weights.

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