

Unions and Wage Inequality

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I. Introduction

How unions affect the distribution of income is a subject that has long intrigued social scientists. The publication of *What Do Unions Do?* and the related papers by Freeman (1980, 1982, 1984) represented a watershed in the evolution of economists' views on this question. Until the 1970s the dominant view was that unions tended to increase wage inequality (Johnson, 1975). Using micro data on individual workers in the union and nonunion sectors, Freeman (1980) presented results that challenged this view. He showed that the inequality-reducing effects of unions were quantitatively larger than the inequality-increasing effects. The equalizing effect of unions became a key chapter in *What Do Unions Do?* and an important component of the authors' overall assessment of the social and economic consequences of unions.

Recently the relationship between unions and inequality has attracted renewed interest as analysts have struggled to explain increases in wage inequality in many industrialized countries. The fact that two of the countries with the largest declines in unionization – the U.S. and the U.K. – also experienced the biggest increases in wage inequality raises the question of whether these two phenomena are linked. If so, how much of the growth in earnings inequality can be attributed to the fall in union coverage?

We make several contributions to this issue. We begin by presenting a simple framework for measuring the effect of unions on wage inequality, based on the potential outcomes framework that is now widely used in program evaluation. Our framework emphasizes three key aspects of collective bargaining: How does the probability of union coverage vary for workers who would earn more or less in the nonunion sector? How much do unions raise average wages for workers in different skill groups? How do unions affect the dispersion of wages within narrow skill groups? Next, we trace the evolution of economists' views on the impacts of unions on the wage distribution. This section places the contributions of Freeman (1980, 1982, 1984) and Freeman and Medoff (1984) in historical context. Third, we present new evidence on the relationship between unions and wage inequality for three countries -- Canada, the U.K., and the U.S. -- during the past three decades. Finally, we assess whether the position put forward in

What Do Unions Do? regarding unions and wage inequality has held up to the scrutiny of subsequent research, including the new evidence reported herein.

Our analysis of unions and wage inequality in the U.S., the U.K., and Canada is motivated by several factors. One is to better understand trends in income inequality. Several previous studies have concluded that falling unionization contributed to the steep increase in wage inequality in the U.S. and the U.K. during the 1980s. Wage inequality did not rise as quickly in these countries in the 1990s. This raises the question of whether the evolution of union coverage and union wage impacts can account for some of the changing trend in wage inequality. More generally, differences across these countries in the timing of changes in unionization and in wage inequality provide an opportunity for further assessing the contribution of institutional change to trends in income inequality.

Our empirical analysis is also motivated by the fact that in these three countries the institutional arrangements governing unionization and collective bargaining provide an environment that is suitable for estimating how unions affect wage inequality. As with other aspects of the economy, collective bargaining institutions in these countries are broadly similar. In particular, negotiations are conducted at the enterprise level, and there is no general mechanism to extend union wage floors beyond the organized sector. The fraction of workers covered by collective agreements in the three countries is also relatively modest – currently under one-third of wage and salary workers. Thus it is possible to compare the structure of wages for workers whose wages are set by union contracts, and those wages are not, and potentially infer the effect of unions on overall wage inequality. A similar task is far more difficult in other countries (including the major European countries and Australia) because there is no clear distinction between the union and nonunion sectors. Collective bargaining in these countries is conducted at the industry or sectoral level, and the provisions are formally or informally extended to most of the labor force. Moreover, in many countries, unions exert considerable influence on political decisions (such as minimum wages) that directly affect labor market outcomes.

We also seek to assess whether there are common patterns in the impact of unions on the wage structure in countries with economies and industrial relations systems that are broadly similar. Of particular interest are patterns in union coverage and union wage impacts by gender and skill. To do so, we use micro data samples to compare the incidence and average wage effect of unions by skill level on male and female workers in the three countries, and measure recent trends in union coverage by skill level. Despite some differences in the institutional systems that govern the determination of union status in the three countries, we find remarkable similarity in the overall patterns of union coverage and in the degree to which unions affect wages of different skill groups. Within narrowly defined skill groups, wage inequality is always lower for union workers than nonunion workers. For male workers, union coverage

tends to be concentrated at the middle of the skill distribution, and union wages tend to be “flattened” relative to nonunion wages. As a result, unions have an equalizing effect on the dispersion of male wages across skill groups in the three countries, complementing the effect on within-group inequality. For female workers, however, union coverage is concentrated near the top of the skill distribution, and there is no tendency for unions to flatten skill differentials across groups. Thus, unions tend to *raise* inequality between more and less skilled women in the three countries, offsetting their effect on within-group inequality.

As a final step, we use data from the past 25 years to compute the changing effect of unionization on wage inequality. During the 1980s and 1990s, unionization rates fell in all three countries, with the most rapid decline in the U.K. and the slowest fall in Canada. These trends contributed to rising male wage inequality, particularly in Britain. Indeed, we estimate that the precipitous fall in U.K. unionization can explain up to two thirds of the difference in the trend in male wage inequality between Britain and the U.S.

II. Unions and Wage Inequality

Conceptual Framework. A useful framework for studying the effect of unions on wage inequality is the potential outcomes model now widely used in program evaluation (Angrist and Krueger, 1999). Assume that each worker faces two potential wages: a log wage in the union sector, W_i^U , and a log wage in the nonunion sector W_i^N . Ignoring dual job holders, a given individual is either in one sector or the other at any point in time, so one of these outcomes is observed and the other is not. Letting U_i denote an indicator for union status, the observed wage of individual i is

$$W_i = U_i W_i^U + (1 - U_i) W_i^N.$$

Let W^U and W^N represent the means of the potential wage outcomes in the two sectors, and let V^U and V^N represent the corresponding variances. Finally, let W and V represent the mean and variance of observed wages in the economy as a whole. In this setting, a natural measure of the effect of unions on wage inequality is $V - V^N$: the difference between the observed variance of wages and the variance that would prevail if everyone was paid his or her nonunion potential wage.

There are two problems with this measure. The first is purely practical: how do we estimate V^N ? The second is conceptual. Arguably, any given individual in the union sector has a well-defined potential wage in the nonunion sector. But if the union sector disappeared, the equilibrium set of wage offers in the nonunion sector could change.¹ Thus V^N is a function of the size of the union sector, $V^N(u)$, where $0 \leq u \leq 1$ indexes the fraction of workers in the union sector. In the absence of any unionization, the variance of observed wages would be $V^N(0)$. Thus, the effect of unionization on wage inequality, taking account of

the general equilibrium impact of the presence of the union sector, is:

$$V - V^N(0) .$$

Despite its theoretical appeal, it is difficult to imagine developing a credible estimate of $V^N(0)$. Under strong assumptions, however, it may be possible to estimate $V^N(U)$, where U is the current fraction of unionized workers. The advantage of this measure is that potential nonunion wage outcomes *under the current level of unionism* are at least partially observed (for all current nonunion workers). Since

$$V - V^N(U) = V - V^N(0) + \{ V^N(0) - V^N(U) \} ,$$

the difference $V - V^N(U)$ overstates or understates the “true” effect of unions by a term reflecting how much the variance of nonunion wage outcomes would change if the union sector was eliminated. While acknowledging this potential bias, in the rest of this analysis we focus on comparisons between V , the observed variance of wages, and $V^N(U)$, the variance that would prevail if everyone were paid according to the *current* nonunion wage structure.

Estimating the Variance of Potential Nonunion Wages. In order to estimate V^N we have to make an assumption about how current union workers would be paid if they worked in the nonunion sector. One starting point is the assumption that union status is “as good as randomly assigned,” conditional on observed skill characteristics. In this case, the counterfactual variance V^N can be estimated as the variance of wages for a suitably reweighted sample of nonunion workers. In this section we show how the resulting calculations are related to three key factors: the variation in the union coverage rate by wage level in the absence of unions, the size of the union wage effect for different skill groups, and the union-nonunion difference in the variance of wages within skill categories. We then show how the assumption that union status is independent of unobserved productivity factors can be relaxed.

Let $W_i^N(c)$ represent the log wage that individual i in skill group c would earn in the nonunion sector, and let $W_i^U(c)$ denote the log wage for the same individual if employed in a union job. Assume that

$$W_i^N(c) = W^N(c) + e_i^N \quad \text{and}$$

$$W_i^U(c) = W^U(c) + e_i^U ,$$

where $W^N(c)$ and $W^U(c)$ are the mean nonunion and union log wages for individuals in skill group c , respectively, and the random terms e_i^N and e_i^U are *independent* of actual union status (conditional on the observed skill level c). Let $V^U(c)$ and $V^N(c)$ denote the variances of potential wage outcomes for individuals in skill group c in the union and nonunion sectors, respectively. The union-nonunion gap in average wages for workers in skill group c is

$$\Delta_w(c) = W^U(c) - W^N(c) ,$$

while the corresponding variance gap is

$$\Delta_v(c) = V^U(c) - V^N(c) .$$

Under the independence assumption, $W^N(c)$ and $V^N(c)$ provide unbiased estimates of the mean and variance of nonunion wage outcomes for all workers in skill group c , not just those who are actually working in the nonunion sector. The variance of wages in the nonunion sector will not necessarily equal V^N , however, if the distribution of nonunion workers across skill groups differs from the distribution of the overall work force. A simple way to estimate V^N is to reweight individual observations from the nonunion work force to account for this difference. Letting $U(c)$ denote the fraction of workers in skill group c in union jobs, the appropriate weight for nonunion workers in group c is $1/(1-U(c))$.

While reweighting provides a convenient way to calculate V^N , it is nevertheless instructive to develop an analytical expression for $V - V^N$ under the conditional independence assumption. Analogous expressions were first derived by Freeman (1980) and used extensively in Freeman and Medoff (1984). To begin, note that if a homogeneous group of workers in some skill group c is split into a union and a nonunion sector, and if unions alter the mean and variance of wages by $\Delta_w(c)$ and $\Delta_v(c)$, respectively, then the overall mean and variance of wage outcomes for workers in skill group c will be

$$W(c) = W^N(c) + U(c) \Delta_w(c) . \quad (1)$$

$$V(c) = V^N(c) + U(c)\Delta_v(c) + U(c)(1-U(c))\Delta_w(c)^2. \quad (2)$$

The first of these equations says that the average wage of workers in skill group c will be raised relative to the counterfactual nonunion average wage by the product of the unionization rate $U(c)$ and the union wage gap $\Delta_w(c)$. The second expression shows that the presence of unions exerts two potentially offsetting effects on the dispersion of wages, relative to the counterfactual $V^N(c)$. First is a “within-sector” effect that arises if wages are more or less disperse under collective bargaining than in its absence. This is just the product of the extent of unionization and the union effect on the variance of wages. Second is a positive “between-sector” effect, reflecting the wedge between the average wage of otherwise identical union and nonunion workers.

If there are many skill groups in the economy, the variance of log wages across all workers is the sum of the variance of mean wages across groups and the average variance within groups:

$$V = \text{Var}[W(c)] + E[V(c)] .$$

where expectations (denoted by $E[]$), and variances (denoted by $\text{Var}[]$) are taken across the skill categories. Using equations (1) and (2), this expression be rewritten as:

$$\begin{aligned} V &= \text{Var}[W^N(c) + U(c)\Delta_w(c)] + E[V^N(c) + U(c)\Delta_v(c) + U(c)(1-U(c))\Delta_w(c)^2] \\ &= \text{Var}[W^N(c)] + \text{Var}[U(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), U(c)\Delta_w(c)] \\ &\quad + E[V^N(c)] + E[U(c)\Delta_v(c)] + E[U(c)(1-U(c))\Delta_w(c)^2] , \end{aligned} \quad (3)$$

where $\text{Cov}[,]$ denotes the covariance across skill groups. In contrast to equation (3), if all workers were paid according to the wage structure in the nonunion sector, the variance of wage outcomes would be

$$V^N = \text{Var}[W^N(c)] + E[V^N(c)].$$

The effect of unions on the variance of wage outcomes, relative to what would be observed if all workers were paid according to the current wage structure in the nonunion sector, is therefore

$$V - V^N = \text{Var}[U(c)\Delta_v(c)] + 2\text{Cov}[W^N(c), U(c)\Delta_w(c)] + E[U(c)\Delta_v(c)] + E[U(c)(1-U(c))\Delta_w(c)^2]. \quad (4)$$

Substituting observed values for $W^N(c)$, $U(c)$, $\Delta_v(c)$, and $\Delta_w(c)$ into equation (4) leads to an expression that is numerically equal to the difference between the observed variance of wages V and the “reweighting” estimate of V^N , derived by reweighting each nonunion worker by $1/(1-U(c))$.

To understand the implications of equation (4) it is helpful to begin by considering a case where the union coverage rate, the union wage effect $\Delta_w(c)$, and the union variance gap are all constant across skill groups. In this case the first two terms in equation (4) are 0, and the effect of unions reduces to the simple “two sector” equation introduced by Freeman (1980):

$$V - V^N = U\Delta_v + U(1-U)\Delta_w^2, \quad (4')$$

where U represents the overall unionization rate, Δ_v represents the difference in the variance of log wages between union and nonunion workers, and Δ_w represents the difference in mean wages between union and nonunion workers.

When union coverage rates or union effects on the level or dispersion of wages vary by skill group, the effect of unions also depends on how the union wage gain $U(c)\Delta_w(c)$ varies with the level of wages in the absence of unions (the covariance term in equation (4)), and on by how much unions raise the variation in mean wages across different groups (the $\text{Var}[U(c)\Delta_v(c)]$ term in equation (4)). In particular, if union coverage is higher for less-skilled workers, or if the union wage impact is higher for such workers, then the covariance term will be negative, enhancing the equalizing effect of unions on wage dispersion.

Allowing for Unobserved Heterogeneity. The assumption that union status is “as good as random” conditional on observed skills is convenient but arguably too strong. In this section we show how the presence of unobserved productivity differences between union and nonunion workers biases the calculation that ignores these differences. As before, assume that workers are classified into skill categories on the basis of observed characteristics, and suppose that potential nonunion and union wages are given by:

$$W_i^N(c) = W^N(c) + a_i + e_i^N, \quad (5a)$$

$$W_i^U(c) = W^U(c) + a_i + e_i^U, \quad (5b)$$

where a_i represents an unobserved skill component that is equally rewarded in the union and nonunion sectors. Continue to assume that e_i^N and e_i^U are independent of union status, and let

$$\theta(c) = E[a_i | U_i=1, c] - E[a_i | U_i=0, c]$$

represent the difference in the mean of the unobserved skill component between union and nonunion workers in group c . If union workers in group c have higher unobserved skills than their nonunion counterparts, for example, then $\theta(c) > 0$.

The *observed* wage gap between union and nonunion workers in group c includes the true union wage premium and the difference attributable to unobserved skills:

$$D_w(c) = \Delta_w(c) + \theta(c).$$

Similarly, assuming that e_i^N and e_i^U are independent of a_i , the *observed* difference in the variance of wages between union and nonunion workers in group c is:

$$D_v(c) = \Delta_v(c) + \text{Var}[a_i | U_i=1, c] - \text{Var}[a_i | U_i=0, c],$$

which is a combination of the true union effect on within-group inequality and any difference in the variance of the unobserved productivity effects between union and nonunion workers. If union workers tend to have a narrower distribution of unobserved skills, for example, the observed variance gap $D_v(c)$ will be biased downward relative to the “true” union effect $\Delta_v(c)$.

Assuming that potential wage outcomes are generated by equations (5a, 5b), it can be shown that 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

$$\begin{aligned} V - V^N &= \text{Var}[U(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), U(c)\Delta_w(c)] \\ &\quad + E[U(c)\Delta_v(c)] + E[U(c)(1-U(c))\{(\theta(c)+\Delta_w(c))^2 - \theta(c)^2\}]. \end{aligned} \quad (6)$$

Only the last term of this equation differs from equation (4), the expression that applies when $\theta(c)=0$ for all groups.² In the presence of unobserved heterogeneity, however, $\Delta_w(c)$ and $\Delta_v(c)$ can no longer be estimated consistently from the observed differences in the means and variances of union and nonunion workers in skill group c . By the same token, it is no longer possible to use a reweighting procedure based on the fraction of union members in different observed skill groups to estimate V^N .

It is instructive to compare the estimated effect of unions under the “as good as random” assumption to the true effect, when potential wages are generated by equations (5a,5b). The estimated effect is given by equation (4), using the observed within-skill group union differences $D_w(c)$ and $D_v(c)$ as estimates of $\Delta_w(c)$ and $\Delta_v(c)$. The true effect is given by equation (6). The difference is

$$\text{Bias} = \text{Var}[U(c)D_w(c)] - \text{Var}[U(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), U(c)(D_w(c) - \Delta_w(c))]]$$

$$+ E[U(c)(D_w(c) - \Delta_w(c))] + E[U(c)(1-U(c)) \{ D_w(c)^2 - \Delta_w(c)^2 - 2 \theta(c) \Delta_w(c) \}] .$$

There are various competing factors here. For example, if $D_w(c)$ varies more across skill groups than $\Delta_w(c)$, the sum of the first two terms is likely to be positive. On the other hand, if $D_w(c)$ is more strongly negatively correlated with nonunion wages across skill groups than $\Delta_w(c)$ (as argued in Card, 1996), then the third (covariance) term will be negative, leading to an overstatement of the equalizing effect of unions. We return to this issue below.

III. *A Review of the Literature on Unions and Inequality*

Until recently, most economists believed that unions tended to raise inequality. For example, Friedman (1956) – appealing to Marshall’s laws of derived demand – argued that craft unions will be more successful in raising the wages of their members than industrial unions. Following this logic, Friedman (1962, p. 124) concluded:

If unions raise wage rates in a particular occupation or industry, they necessarily make the amount of employment available in the occupation or industry less than it otherwise would be – just as any higher price cuts down the amount purchased. The effect is an increased number of persons seeking other jobs, which forces down wages in other occupations. Since unions have generally been strongest among groups that would have been high-paid anyway, their effect has been to make high-paid workers higher paid at the expense of lower-paid workers. Unions have therefore not only harmed the public at large and workers as a whole by distorting the use of labor; they have also made the incomes of the working class more unequal by reducing the opportunities available to the most disadvantaged workers.

As this quote makes clear, Friedman posited two channels for the disequalizing effect of unions. One is the “between-sector” effect – the gap in wages between otherwise similar workers in the union and nonunion sectors. The other is a hypothesized positive correlation between the union wage gain and the level of wages in the absence of unions – that is, an assumption that the covariance term in equation (4) is positive. Even economists more sympathetic to unions than Friedman shared this view. For example, Rees (1962) suggested that “theory and evidence” both predict unions will have a bigger effect on high-skilled workers. Noting that union membership (in 1950) was concentrated among workers in the upper half of the earnings distribution, Rees concluded that the overall effect of unions was probably to increase inequality.

Not all scholars accepted this position. Following their detailed analysis of the evolution of the wage structure in several industries, Reynolds and Taft (1956, p. 194) concluded that: “Summing up these diverse consequences of collective bargaining, one can make a strong case that unionism has at any rate

not worsened the wage structure. We are inclined to be even more venturesome than this, and to say that its net effect has been beneficial.” Much of the reasoning behind this position was based on evidence of unions negotiating “standard rates” that resulted in greater uniformity of wages within and across establishments.

Evidence on these issues was scanty and inconclusive until the widespread availability of microdata in the 1970s. Stieber (1959) examined the effect of unions in the steel industry and concluded that during the 1947-1960 period collective bargaining did not flatten the wage distribution. In an interesting contribution, Ozanne (1962) tabulated data for McDormick Deering (a farm machine company) over the period 1858 to 1958. During this period many different unions unionized the same plant. He found no tendency for unions per se to reduce or increase intra-firm wage inequality. Skill differentials narrowed during some regimes and they widened during other periods. However, there was a general tendency for industrial unions to lower skill differentials and for craft unions to raise them. In his classic study of union relative wage effects, Lewis (1963) examined the correlation between estimates of the union wage differential and wage levels. He concluded that unionism increased the inequality of average wages across industries by 2 to 3 percentage points.

Some contrary evidence appeared in the late 1960s and early 1970s. Stafford (1968), Rosen (1970), and Johnson and Youmans (1971) found that unions compress the wage structure by raising wages of less skilled workers relative to their more skilled counterparts, while Ashenfelter (1972) found that unions contributed to the narrowing of the black-white wage gap. Nonetheless, in a survey written in the mid-1970s, Johnson (1975, p. 26) concluded that “union members generally possess characteristics which would place them in the middle of the income distribution, ... so that unionism probably has a slight disequalizing effect on the distribution of income.”

The direction of subsequent research was fundamentally altered by the methods and findings in Freeman’s (1980) important study, which first laid out the two-sector framework described in Section II.³ Freeman (1980) also used establishment-level data to measure the effect of unions on the wage gap between blue-collar and white-collar workers. Since few white-collar workers are unionized, this exercise extended the simple two-sector model to incorporate a “between group - within sector” effect analogous to the “between” and “within” effects in the basic two-sector model. The key finding in Freeman’s study – and a result that was largely unanticipated by earlier analysts – is that the “within-sector” effect of unions on wage inequality is large and negative, especially in manufacturing. Freeman attributed the compression of wages in the union sector to explicit union policies that seek to standardize wages within and across firms and establishments. Coupled with the tendency for unions to reduce the wage gap between blue-collar and more highly paid white-collar employees, the equalizing effects of unions more

than offset the “between-sector” disequalizing effect between union and nonunion workers. In nonmanufacturing industries, Freeman concluded that the net impact of unions was smaller, reflecting both a smaller “within-sector” effect and larger “between-sector” effect. Subsequent research has confirmed that wage differences between different demographic and skill groups are lower, and often much lower, in the union sector than in the nonunion sector.⁴ The residual variance of wages within demographic and skill groups is also generally lower in the union sector.

Analysis of longitudinal data by Freeman (1984) confirmed the finding of lower wage inequality in the union sector, even controlling for individual worker effects. In particular, Freeman documented that wage dispersion tends to fall when workers leave nonunion for union jobs and to rise when they move in the opposite direction. The effect of unions on wage dispersion estimated from longitudinal data is, however, smaller than comparable estimates using cross-sectional data. This lower estimate appears to be at least partly due to measurement error in union status.

Freeman (1993), using more recent longitudinal data from the 1987-88 CPS, confirmed that unionization reduces wage inequality. On the basis of his longitudinal estimates, he concluded that declining unionization accounted for about 20 percent of the increase in the standard deviation of male wages in the U.S. between 1978 and 1988. Using a more sophisticated econometric approach (see the discussion of Card, 1996, below), Card (1992) also concluded that the drop in unionization explained around 20 percent of the increase in wage inequality during the 1980s. Gosling and Machin (1995) reached a similar conclusion that the fall in unionization accounts for around 15 percent of the increase in male wage inequality among semi-skilled workers in Britain between 1980 and 1990.

Second Generation Studies. Beginning with Freeman (1980), the first generation of micro-based studies significantly altered views regarding the relationship between unionization and wage inequality. But these studies tell an incomplete story. On the one hand, they focus on male, private sector workers. On the other hand, they tended to ignore variation in the union coverage rate and the union wage effect across different types of workers.

A second generation of studies used variants of the framework underlying equation (4) to develop a more complete picture of the effect of unions. DiNardo and Lemieux (1997) implemented a reweighting technique to construct estimates of the sum of the terms in equation (4) for men in the U.S. and Canada in 1981 and 1988. They estimated that in 1981 the presence of unions reduced the variance of male wages by 6 percent in the U.S. and 10 percent in Canada. The corresponding estimates in 1988 are 3 percent in the U.S. and 13 percent in Canada. Thus, they estimated that changing unionization contributed to the rise in U.S. wage inequality in the 1980s, but worked in the opposite direction in

Canada. Their decompositions also showed that in both countries unions lower the variation in wages within and between groups, with a larger net effect within skill groups.

A related study by DiNardo et al. (1996) examined both men and women in the U.S. in 1979 and 1988. DiNardo et al. (hereinafter, DFL) use the reweighting technique applied by DiNardo and Lemieux. DFL focus on explaining the rise in wage inequality over the 1979-1988 period. For men, their methods suggested that shifts in unionization accounted for 10-15 percent of the overall rise in wage dispersion in the 1980s, with most of the effect concentrated in the middle and upper half of the wage distribution. For women, on the other hand, the estimated contribution of changing unionization was very small. DFL also estimated that falling unionization explains about one-half of the rise in the wage premium between men with a high school diploma and dropouts, and about a quarter of the rise in the college-high school wage gap for men.

The study by Bell and Pitt (1998) used DFL's method to analyse the impact of declining unionization on the growth in wage inequality in Britain. Depending on the data source used, they found that between 10 and 25 percent of the increase in male wage inequality can be explained by the fall in unionization. Machin (1997) reached similar conclusions.

Card (2001) examined the contribution of unions to wage inequality among U.S. men and women in 1973-1974 and in 1993. Card reported estimates based on the simple two-sector formula (equation (4')), and on a variant of equation (4) obtained by dividing workers into 10 equally-sized skill groups, based on predicted wages in the nonunion sector. Two key findings emerged from this analysis. First, the presence of unions was estimated to have reduced the variance of men's wages by about 12 percent in 1973-1974 and 5 percent in 1993. Overall, shifts in unionization can explain about 15-20 percent of the rise in male wage inequality in the 1973-1993 period. Second, although the within-group variance of wages is lower for women in the union sector than the nonunion sector (i.e. $\Delta_v(c)$ is on average negative), this equalizing effect is counteracted by a positive between-group effect, so overall unions had little net effect on wage inequality among U.S. women in 1973-74 or 1993.

Card (2001) also conducted separate analyses of the effects of unions on men and women in the public and private sectors in 1973-1974 and 1993. The trends in unionization were quite different in the two sectors, with *rises* in union membership in the public sector for both men and women, and declines in the private sector. Nevertheless, comparisons of the patterns of union wage gaps by skill group suggest that unions affect the wage structure very similarly in the two sectors, with a strong tendency to "flatten" wage differences across skill groups for men, and less tendency for flattening among women. Overall, Card's estimates implied that unions reduced the variance of men's wages in the public sector by 12 percent in 1973-74 and 16 percent in 1993. In the private sector, where union densities declined, the

union effect fell from 9 percent in 1973-1974 to 3 percent in 1993. An interesting implication of these estimates is that differential trends in unionization among men in the public and private sectors can potentially explain a large share (up to 80 percent) of the greater rise in wage inequality in the private sector. The estimated effects of unions on women's wage inequality are all close to zero, except in the public sector in 1993, when the effect is about -5 percent.

Gosling and Lemieux (2001) examined the effects of unions on the rise in wage inequality in the U.S. and the U.K. between 1983 and 1998, using the DFL reweighting method. Their estimates suggested that in both the U.S. and the U.K., unions have a much smaller equalizing effect on female wage inequality than male inequality. They estimated that shifts in union coverage among men in the U.K. can explain up to one-third of the rise in wage inequality there between 1983 and 1998, while in the U.S. the decline in unions can explain up to 40 percent of the rise in inequality. Consistent with findings in DFL and Card (2001) they concluded that changes in unionization had little net effect on female wage inequality in either country.

Studies That Correct for Unobserved Skill Differences. A potential problem with estimates of the equalizing effect of unions based on equations (4') or (4) is that union workers may be more or less productive than otherwise similar nonunion workers. In this case, comparisons of the mean and variance of wages for union and nonunion workers with the same observed skills confound the true "union effect" and unobserved differences in productivity. Traditionally, economists have argued that union workers are likely to have higher unobserved skills than their nonunion counterparts (Lewis, 1986). This prediction arises from the presumption that in a competitive environment, unionized employers will try to counteract the effect of above-market wage scales by hiring the most productive workers. If total productivity of worker i consists of an observed component p_i and another component a_i that is observed by labor market participants but unobserved by outside data analysts, and if an employer who if forced to pay a union wage W^U hires only those workers with $p_i + a_i > W^U$, then p_i and a_i will be *negatively* correlated among those who are hired. Workers with the lowest observed skills will only be hired if they have relatively high unobserved skills, whereas even those with below-average unobserved skills will be hired if their observed skills are high enough. This line of reasoning suggests that the "flattening" of the wage structure in the union sector arises from selectivity bias, rather than from the wage policy of unions per se.

If unions really flatten the wage structure, however, then there is another side to the story, since highly skilled workers gain less from a union job. A worker with observed productivity skills p_i and unobserved skills a_i can expect to earn $p_i + a_i$ in a competitive labor market. Such a worker will only take a union job paying W^U if $p_i + a_i < W^U$. In this case union members are negatively selected: Workers with

the highest observed skills will only accept a union job if their unobserved skills are low. This view also implies that the wage structure in the union sector will appear “flatter” than the nonunion wage structure. Combining the two sides of the market, one might expect union workers with low unobserved skills to be positively selected, since for these workers the demand side is the binding constraint, whereas unionized job holders with high unobserved skills are negatively selected, since for these workers the supply side constraint is the more serious constraint.

Some evidence of this “two-sided” view of the determination of union status was developed by Abowd and Farber (1982), who used information on workers who reported that they would prefer a union job, as well as on those who held union jobs, to separate the roles of employer and employee choice. They found that workers with higher experience were less likely to want a union job (consistent with the idea that wages for highly experienced workers were relatively low in the union sector), but were more likely to be hired for a union job, conditional on wanting one (consistent with the idea that employers try to choose the most productive workers).

Several recent studies attempt to assess the effect of unions on the wage structure, while recognizing that union workers may be more or less productive than otherwise similar nonunion workers. Lemieux (1993) and Card (1996) measure the wage outcomes of job changers who move between the union and nonunion sectors, distinguishing between workers in groups defined by observed productivity characteristics. A limitation of these studies is that they implicitly assume that the rewards for unobserved ability are similar in the union and nonunion sectors. Lemieux (1998) adopted a more general approach that allows the union sector to flatten the returns to unobserved ability relative to the nonunion sector.

Lemieux (1993) studied men and women in Canada in the late 1980s, and reported separate estimates of the effect of unions for three skill categories in the public and private sectors separately, and in the overall economy. For men, his results showed that unionized workers from the lowest skill group are positively selected (i.e., they have higher unobserved skills than do nonunion workers in the same group), whereas those in the upper skill groups are negatively selected. This result – which is consistent with a simple two-sided selection model – echoes a similar finding in Card (1996) for U.S. men in the late 1980s. An implication of this pattern is that the between-group “flattening effect” of unions apparent in the raw data is somewhat exaggerated, although there is still evidence that unions raise wages of low-skilled men more than those of high-skilled men. Lemieux also examined the changes in the variance of wages, and concluded that some of the apparent reduction in variance in the union sector may be due to selectivity, rather than to a within-sector flattening effect. Unfortunately, this inference is confounded by the potential selectivity of the group of union status changers, and the fact that the variability of wages

may be temporarily high just before and just after a job change. Overall, Lemieux concluded that the presence of unions lowers the variance of male wages in Canada in the late 1980s by about 15 percent. A similar calculation for U.S. men, based on Card (1996), shows a 7 percent effect. These effects are somewhat smaller than corresponding estimates that fail to correct for unobserved heterogeneity.

Lemieux's findings for women in Canada were much different than those for men. In particular, neither the cross-sectional nor longitudinal estimates of the union wage gaps showed a systematic flattening effect of unions. Coupled with the fact that union coverage is lower for less-skilled women, these results implied that unions raise the between-group variance of wages for women. This effect is larger than the modest negative effect on the within-group variance, so Lemieux's results imply that on net unions raised wage dispersion among Canadian women.

Lemieux (1998) presented an estimation method that accounts for the potential "flattening" effect of unions on the returns to individual skill characteristics that are constant over time but unobserved in conventional data sets. Using data on men who were forced to change jobs involuntarily, he concluded that unions tend to "flatten" the pay associated with observed and unobserved skills. Moreover, the variance of wages around the expected level of pay is lower in the union sector. As a result of these tendencies, Lemieux's results implied that unionization reduced the variance of wages among Canadian men by about 17 percent – not far off the estimate in his 1993 study.

IV. Estimating the Effect of Unions on Wage Inequality

Data Sources. We use a variety of micro data files to compare the effects of unions on wages in the U.S., the U.K., and Canada over the past thirty years. Our U.S. samples are the most straightforward, since the Current Population Survey (CPS) has been collecting data on wages and union status annually since 1973. We use the pooled May 1973 and May 1974 CPS samples as our first U.S. observation. For later years, we use the monthly earnings supplement files (the so-called "outgoing rotation group") for 1984, 1993, and 2001. The earnings and union status information all pertain to an individual's main job as of the survey week.

In 1993 the U.K.'s Labour Force Survey (UKLFS) began asking questions on union status and earnings that are comparable to the CPS questions. Strictly comparable data are unavailable for earlier years. The 1983 General Household Survey (GHS) is the only large scale micro data set that contains information on union status and wages in the U.K. prior to the 1990s. While this data source has several limitations, we elected to combine the 1983 GHS with the 1993 and 2001 UKLFS samples for our U.K. analysis.

The Canadian Labour Force Survey (CLFS) added questions on earnings and union status in

1997. To supplement these data, we combine two smaller surveys – the 1991 and 1995 Surveys on Work Arrangements – as a source of information for the early 1990s, and use the 1984 Survey of Union Membership as a source of information for the early 1980s. All three of these surveys were conducted as supplements to the regular CLFS.

In addition to the usual problems that arise in comparing survey responses over time and across countries, a significant issue for our analysis is the measurement of union status. The 1984 and later CPS files include questions on both union membership and union coverage. The 1973 and 1974 May CPS files, however, only ask about union membership. For comparability reasons, we therefore focus on union membership as our measure of union status in the U.S. Our U.K. data sets include data on union membership as well as responses to a question about whether there is a “union presence” at the individual’s place of employment. As noted in Bland (1999, Table 6), however, the latter question significantly overstates coverage under collective bargaining agreements. As in our U.S. analysis, we therefore use union membership as our measure of unionization in the U.K. With respect to Canada, consistent information on union membership cannot be recovered from the 1991 and 1995 SWA’s, so we use union coverage as our measure of unionization in Canada. We believe that this choice has little effect on the results, since only about two percent of Canadian employees are covered by collective agreements but are not union members.⁵

In the data appendix we explain in detail how we process the various data sets to arrive at our final estimation samples. Generally speaking, our samples include only wage and salary workers age 16 to 64 (15 to 64 in Canada) with non-allocated wages and earnings (except in 1984 and 2001 in Canada). We use hourly wages for workers who are paid by the hour and compute average hourly earnings for the other workers by dividing weekly earnings by weekly hours (or earnings for a longer time period divided by the corresponding measure of hours). We also exclude workers with very low or very high hourly wage values. Sample weights are used throughout except in the 1983 GHS for which sample weights are not available.

To implement the methods developed in Section II, we divide workers in each sample into skill groups, based on age and educational attainment. The number of skill groups used varies by country, reflecting differences in the sample sizes and the age and education codes reported in the raw data files. In the earlier Canadian data sets, age is only reported in 10-year categories (a total of 5 categories for workers age 15 to 64), and education can only be consistently coded into 5 categories. Thus we only use 25 skill groups for Canada. Given the small sample sizes available in the 1983 GHS and the 1993 UKLFS, we use the same number of skill groups for the U.K. (five age and five education groups). In our U.S. samples, we are able to use a much larger number of skill categories because of the larger sample

sizes and detailed age and education information in the CPS. We have re-analyzed the U.S. data using about the same number of skill groups as in Canada and the U.K., however, and found that this has little impact on our results.

Patterns of Union Coverage and Union Wage Effects. To set the stage for our analysis it is helpful to begin by looking at how union coverage and the size of the union wage gap vary by skill level. Figures 1-3 show the unionization rates of men and women in the U.S., the U.K. and Canada, by the level of real hourly wages. These graphs are constructed by calculating union membership/coverage rates for workers in narrow wage bins, and smoothing across bins.⁶ In all three countries, unionization rates of men tend to follow a hump-shaped pattern, peaking for workers near the middle or upper middle of the wage distribution. By comparison, unionization rates of women in the U.S. and Canada are about the same for highly paid workers as for those in the middle. This pattern is driven in part by relatively high rates of unionization for teachers, nurses, and other public sector workers, who are near the top of the female wage distribution. In the U.K. there is more of a fall-off in union membership among the highest paid women, especially in the more recent data. Comparisons of the unionization rates in different years reveal the rapid decline in union membership among U.S. and U.K. men. Declines are also evident for Canadian men and for women in all three countries.

The framework developed in Section II suggests that the effect of unions on wage inequality depends in part on how the union wage gap varies by skill. Figures 4-6 provide some simple evidence on this variation, using data from the early 1990s for the three countries. These figures plot mean wages for unionized workers in a given age-education group (i.e., $W^U(c)$ in the notation of Section II) against the corresponding mean for nonunion workers with the same skill level (i.e., $W^N(c)$) for 25 age-education groups in Canada and the U.K., and about 150 groups in the U.S.). In interpreting these figures, note that if union and nonunion workers in a given skill group have the same average wages, the points in these graphs will lie on the 45-degree line. On the other hand, if the union wage gap $D_w(c)$ is positive, the points will lie above the 45-degree line. Moreover, if $D_w(c)$ is larger for lower wage workers, the points will tend to be further above the 45 degree line for low-wage skill groups (on the left side of the graph) than for high-wage groups (on the right). This is in fact the case for U.S. men. The best-fitting line relating $W^U(c)$ to $W^N(c)$ is also shown in the figure, and lies above the 45 degree line but with a slope of less than 1.

Interestingly, the same pattern is true for men in Canada and the U.K., as shown in Figures 5a and 6a. For skill groups with low average wages (e.g., less educated and relatively young men) the mean union wage tends to be substantially higher than the mean nonunion wage, while for groups with high average wages (e.g., middle-age college or university graduates) the mean union wage is not too much

above the mean nonunion wage. Thus, in all three countries $D_w(c)$ is larger for low-wage men than high-wage men, implying that unions tend to “flatten” wage differentials across skill groups. As previously discussed, one caveat to this conclusion is that there may be unobserved skill differences between union and nonunion workers in different age-education groups that tend to exaggerate the apparent negative correlation between wages in the nonunion sector and the union wage gap. We address this issue further in the next section.

For women, the patterns of union wages relative to nonunion wages are also remarkably similar in the three countries. Unlike the patterns for men, however, the union wage gaps for women are roughly constant. Coupled with the tendency for unionization rates of women to rise across the wage distribution, the absence of a “flattening” effect of unions on female wages implies that covariance between the nonunion wage $W^N(c)$ and the union wage gain $U(c)\Delta_w(c)$ is either zero or positive, limiting the potential equalizing effect of unions on female inequality.

Although the data in Figures 4-6 pertain to the early 1990s, similar plots from other years show that the basic patterns have been very stable in all three countries over the past 20-30 years. In all our sample years, the union-nonunion wage gap for men tends to be highest for the least skilled workers, and to be relatively small (or even negative) for highly skilled men. The union gap for women, on the other hand, tends to be stable or only slightly declining with skill level. The consistency of these patterns over time and across the three countries is remarkable.

The Effect of Unions on Wage Inequality. With this background, we turn to our analysis of the effect of unions on wage inequality in the three countries. Tables 1, 2, and 3 summarize a variety of facts about unionization and the structure of wages for the U.S., Canada, and the U.K., respectively. Reading across the columns of Table 1, a comparison of the entries in the first row confirms the steep decline in U.S. unionization rates documented in many studies. As illustrated in Figures 1a and 1b, however, these aggregate figures hide a sharp difference between men and women. Between 1973 and 2001, the unionization rate of women declined only about 2 percentage points, from 14 to 12 percent, while for men it fell by 50 percent, from 31 to 15 percent. This sharp male-female difference has much to do with the gradual shift of unionization from the private to the public sector. For instance, Card (2001) showed that for both men and women, unionization rates declined by about 50 percent in the private sector between 1973 and 1993. During the same period, however, unionization rates increased sharply in the public sector. Women in general, and unionized women in particular, are much more concentrated in the public sector than their male counterparts. As a result, the rise in public sector unionism has largely offset the decline in private sector unionization among women.

The trends in unionization in Canada between 1984 and 2001 (Table 2) are similar to those in the

U.S. The male unionization rate declined by 14 percentage points, even more than the 9 percentage point decline in the U.S. over the same period. As in the U.S., the decline for women was more modest (4 percentage points). The drop in unionization in our Canadian samples is much steeper than the decline registered in membership tallies obtained from union reports, but is consistent with the trends reported by Riddell and Riddell (2001) based on similar micro-data sources.

Table 3 shows that unionization rates have also fallen sharply in the U.K. in the past two decades: by 27 percentage points for men and by 14 percentage points for women. As in the U.S. and Canada, the faster decline in male unionization is linked to the relative shift of unionization from the private to the public sector (Gosling and Lemieux, 2001). In the U.K., this shift was compounded by privatization of many nationalized industries, which transferred sizeable numbers of mainly male workers from the unionized public sector to the much less organized private sector (Gosling and Lemieux, 2001). Interestingly, the relatively faster decline of male unionization in the three countries meant that by 2001, male and female unionization rates were not too different in the U.S., Canada, or the U.K. This near equality marks a sharp departure from the historical pattern of greater unionization among men.

The next set of rows in Tables 1-3 show the evolution of mean wages of nonunion and union workers and the trend in the union wage gap. We also report an adjusted wage gap, calculated from a regression that includes dummies for each skill category. As in the case for the unionization rates, the estimated wage gaps show a remarkably similar pattern across the three countries. The unadjusted wage gaps tend to be larger for women than for men. The adjusted wage gaps are uniformly smaller than the unadjusted gaps, and in all three countries the divergence has increased over time, implying that union membership rates have fallen more for relatively unskilled workers (as is apparent in Figures 1-3).

Like the unadjusted union wage gap, the adjusted wage gap is typically larger for women than for men. Nevertheless, gender differences in the adjusted gaps are less pronounced than the corresponding differences in the unadjusted gaps, especially in more recent years. For example, the unadjusted wage gaps in the U.S. in 2001 were 0.233 for men and 0.305 for women, versus adjusted wage gaps of 0.156 and 0.149. This pattern is consistent with Figures 1 to 3, which show that unionized women are more highly concentrated in the upper end of the skill distribution than unionized men. As a result, controlling for the skill composition of the work force reduces the union wage gap far more for women than for men.

Another trend that is shared by all three countries is a gradual decline in the adjusted union wage gap, by 5 to 10 percentage points (depending on gender and country) between the early 1980s and 2001. Since the rate of unionization also declined sharply during this period, the implied effect of unions on average wages – the union wage gain $E[U(c)D_w(c)]$ – has declined dramatically over the last two decades. For example, the adjusted impact of unions on male wages in the U.K. went from 9.2 percentage points in

1983 (unionization rate of 0.57 times an adjusted gap of 0.162) to 1.7 percentage points in 2001 (0.307 times 0.045). In the U.S., the effect on average wages of men fell from 5.7 percentage points (unionization rate of 0.307 times an adjusted gap of 0.185) to 2.3 percent in 2001 (0.149 times 0.156).

The next rows in Tables 1-3 report measures of wage dispersion within the union and nonunion sectors. Once again, the results are remarkably consistent across countries. As first documented in Freeman (1980), the standard deviation of wages is always smaller in the union than in the nonunion sector. Moreover, the gap between the standard deviation in the union and nonunion sector is always larger for men than for women. These observations are confirmed by Figures 7 to 12, which show kernel density estimates of the densities of log hourly wages in the union and nonunion sectors, and for the two sectors pooled together, by gender and time period.⁷ For example, Figure 7 displays the wage distribution for U.S. males. In all four time periods, wages are more tightly distributed in the union than the nonunion sector. In particular, while the upper tails of the union and nonunion densities look qualitatively similar, the lower tail goes much further to the left in the nonunion sector. By contrast, the inter-sectoral differences in wage dispersion are much less striking for U.S. women (Figure 8). In 1984, for example, the union and nonunion distributions show different skewness, and average wages are higher in the union sector. However, whether wages are more narrowly distributed in the union or nonunion sector is unclear.

Inspection of Figures 7 and 8 (and the corresponding figures for Canada and the U.K.) suggests that the minimum wage is an important factor in explaining overall trends in wage inequality, particularly for nonunion female workers.⁸ An interesting conjecture is that unions appear to have a more limited effect on the dispersion of female wages in part because minimum wages limit the amount of dispersion in the lower tail of the female wage distribution.

The wage densities for Canadian men (Figure 9) and women (Figure 10) are qualitatively similar to those in the U.S. In particular, it is clear that male wages are more narrowly distributed in the union sector than the nonunion sector. Things are not as clear for women, in part because of the minimum wage, which has a surprisingly large visual impact in the nonunion sector, especially in 2001. Relative to the U.S. or Canada it is more difficult to see union wage compression effects for U.K. males (Figure 11) or females (Figure 12). Comparing the reported standard deviations of wages in the two sectors in Table 3, however, the union-nonunion gaps are nonetheless quite similar to those in Canada or the U.S.

The bottom rows of Tables 1-3 show the various components of our analysis of the effect of unions on wage inequality. For reference purposes, we first present a simplified analysis based on the two-sector model that ignores any differences across skill groups (equation (4')). Comparing the results across countries and over time, the results from this simplified analysis are remarkably consistent. For

men, the within-sector effect is substantially larger (in absolute value) than the between-sector effect, implying that unions reduce wage dispersion. Relative to the overall variance, the compression effect ranges from 31 percent in the U.K. in 1984, when the unionization rate was 57 percent, to 6 percent in the U.S. in 2001 (unionization rate of 15 percent). More generally, the compression effect of unions is highly correlated with the overall level of unionization.⁹

In contrast to the situation for men, the simplified analysis of equation (4') implies that unions have either no effect on female wage inequality, or a slightly disequalizing effect. This contrast is attributable to three complementary factors. First, the female unionization rate is lower, reducing the size of the within-sector effect. Second, the gap in overall wage dispersion between union and nonunion workers is much smaller for women than men. Third, the union wage gap is systematically larger for women than men, yielding a larger (more positive) between-sector effect $U(1-U)D_w^2$. Indeed, in the later years of our analysis, the between-sector effect dominates in all three countries. Consistent with findings reported in Card (2001) and Lemieux (1993), unions thus tend to increase the variance of wages among women.

The final rows in Tables 1-3 show the effect of unions on the variance of wages when we distinguish among skill groups. Recall from equation (4) that this analysis includes three components: an average within-sector effect, $E[U(c)D_v(c)]$, an average between-sector effect, $E[U(c)(1-U(c))D_w(c)^2]$, and the sum of two “between-skill-group” terms, $\text{Var}[U(c)D_w(c)] + 2\text{Cov}[W^N(c), U(c)D_w(c)]$, that reflect the rise in inequality between groups if the union wage gain varies by skill group and any tendency of unions to raise wages more or less for higher wage workers.

Starting with men, the introduction of controls for observable skill systematically reduces the magnitudes of both the within- and between-sector effects. It is easy to see why this happens in the case of the between-sector effect. As noted earlier, adjusting for characteristics reduces the union wage gap, and thus decreases the size of the between-group effect. In other words, part of the measured between-sector effect in the simple two-sector calculation is a spurious consequence of that fact that union workers are more skilled, on average, than nonunion workers. A similar reasoning can be used to understand why the within-group effect also declines when differences in observed skills are taken into account. Recall from Figures 1a-3a that unionized men are more concentrated in the middle of the wage distribution than nonunion men. Part of the lower dispersion of wages in the union sector is thus a spurious consequence of the fact that union workers are more homogenous.

Interestingly, adjusting for observed skill characteristics also reduces the magnitude of the between-sector effect for women but increases (or leaves unchanged in the U.S.) the magnitude of the within-group effect. The latter finding means that union women are no more homogenous (in terms of

their observable skills) than their nonunion counterparts, which is consistent with the evidence reported in Figures 1b, 2b, and 3b. Once worker characteristics are taken into account, the within-sector effect tends to dominate the between-sector effect for both men and women. Thus, the results from a simplified analysis which ignores measured skill differences tends to overstate male-female differences in the effect of unions on wage inequality.

The final components of the union effect are the two terms which reflect the effect of unions on the distribution of wages across skill groups. As highlighted in our discussion of Figures 4-6, the union wage effect $D_w(c)$ is systematically lower for high-wage men, inducing a negative covariance between $W^N(c)$ and $U(c)D_w(c)$. By contrast, the wage gap for women is not much lower for high-wage groups, and the higher unionization rate for those groups induces a positive covariance between $W^N(c)$ and $U(c)D_w(c)$.

The results in Tables 1-3 are broadly consistent with this prediction. As expected, unions tend to reduce wage dispersion across skill groups for men (except in recent years in Canada where the effect is essentially zero). Also as expected, unions tend to increase wage dispersion across skill groups for women in Canada and the U.K. In the U.S., however, unions have little effect on female wage dispersion across skill groups from 1973 and 1993 and actually reduce wage dispersion in 2001. A natural explanation for the difference between the U.S. on one hand, and Canada and the U.K., on the other, is that the union wage gap for U.S. women tends to decline slightly with higher nonunion wages (see Figure 4b). This lowers the covariance between $W^N(c)$ and $U(c)D_w(c)$ for U.S. women relative to the other two countries.

Once all three factors are taken into consideration, our calculations show that unions systematically reduce the variance of wages for men. By contrast, the effects for women tend to be small and slightly positive (i.e., unions raise inequality). This pattern of result is quite similar to what we found with the simpler model, though the magnitude of the effects tend to be smaller when we control for worker's characteristics.

Biases from Unobserved Heterogeneity. As noted earlier, a potential problem with estimates of the equalizing effect of unions based on equation (4) or (4') is that union workers may be more or less productive than otherwise similar nonunion workers. In this case, comparisons of the mean and variance of wages for union and nonunion workers with the same observed skills confound the true union effect and unobserved differences in productivity. Studies by Lemieux (1993), Card (1996), and Lemieux (1998) have attempted to use data on job changers to measure the extent to which union and nonunion workers in different skill groups have different unobserved productivity characteristics.¹⁰ The two Lemieux studies analyzed data for men and women in Canada, while Card examined data for men in the U.S. All three studies found that among North American men, unions tend to raise wages more for less

skilled workers, but that simple comparisons which ignore unobserved skill components tend to *overstate* the flattening effect. Lemieux's results for Canadian women, on the other hand, show little evidence of flattening, either in simple cross sectional comparisons or in more sophisticated longitudinal estimators. Lemieux (1998) also uses longitudinal data to examine the apparent effect of unions on the dispersion of wages controlling for observed and unobserved skill components. This analysis suggests that some of the apparent reduction in variance in the union sector may be due to selectivity, rather than to a within-sector effect. Unfortunately, this inference is confounded by the potential selectivity of the group of union-status changers, and the fact that the variability of wages may be temporarily high just before and just after a job change. Overall, Lemieux (1998) concluded incorporating unobserved heterogeneity effects leads to a small reduction in the apparent effect of unions on male wage inequality. A similar conclusion was reached in Card (1996).

Based on these findings, we conclude that the estimates of the equalizing effect of unions on male workers in the U.S., Canada, and the U.K. in Tables 1-3 are likely to slightly *overstate* the true equalizing effects. For women, the estimated effects in Tables 1-3 are very small anyway, and the existing longitudinal research suggests there is no important bias.

Unions and Differences in the Trends in Wage Inequality. To what extent can changes in the strength of unions explain the evolution of wage inequality over time and the differences in inequality across countries? In light of the results of Tables 1-3, we look at this question for men only since unions appear to have little effect on wage inequality for women.¹¹ Starting with the U.S., Table 1 shows that the variance of male wages increased from 0.258 to 0.340 (a rise of 0.082) between 1973/1974 and 2001. During the same period, the effect of unions on the variance of wages computed using the simplified model declined from -0.047 to -0.021 (a rise of 0.026). If this effect had remained constant over time, overall wage inequality would have grown by 31 percent less (0.026/0.082) than it actually did. The contribution of unions to the growth of inequality remains important though only about one-half as big (14 percent) when we use the more sophisticated estimates of the union effect that control for observable skills.

The results for the U.K. are qualitatively similar. Between 9 and 29 percent of the 0.087 growth in the variance of log wages between 1983 and 2001 can be accounted for by the decline in union compression effects. Furthermore, in both the U.S. and U.K. union wage compression effects remained relatively constant between 1993 and 2001. In particular, the effects from the analysis that controls for workers' characteristics are essentially unchanged in the period from 1993 to 2001. This is consistent with the slowdown in the growth of inequality in both countries in the 1990s, relative to the 1980s.

As in the U.S. and U.K., the union wage compression effect has been steadily declining for

Canadian men since 1984. Unlike the U.S. and U.K., however, overall inequality has remained very stable in Canada over time, so overall inequality would have actually declined if union wage impacts had remained at their 1984 levels. Several developments may have offset the pressures toward increased inequality associated with the decline in union strength. The real minimum wage in Canada rose from the mid-1980s to the late-1990s, in contrast to the situation in the U.S. where the real minimum wage was approximately constant over this period (Kuhn, 2000). In addition, there is some evidence that the much more rapid growth in educational attainment in Canada compared to the U.S. during the 1980s and 1990s reduced the tendency for widening earnings differentials between less educated and more educated workers (Murphy et al., 1998).

Turning to cross-country differences in wage inequality, first note that in 1983/1984 the variance of wages was lowest in the U.K. (0.216) followed by Canada (0.231) and the U.S. (0.289). By contrast, union wage compression effects (from the model that controls for skill differences) were highest in the U.K. (-0.050), followed by Canada (-0.037) and the U.S. (-0.017). The pattern of cross-country differences in wage inequality is thus consistent with the pattern of union wage compression effects. For instance, differences in union wage compression effects account for 45 percent of the U.K.-U.S. difference in the variance of wages in the early 1980s. By 2001, the U.S.-U.K. difference in the variance of wages had fallen to 0.037, while the U.S.-U.K. difference in the union compression effect had fallen to 0.027. This indicates that over 70 percent of the difference in wage inequality can now be explained by union wage compression effects. In 2001, however, union wage compression effects cannot account for the much lower variance of wages in Canada.

In summary, union wage compression effects help explain a reasonable fraction of the secular growth in male wage inequality and of cross-country differences in male wage inequality. One exception is the surprising lack of growth in male wage inequality in Canada relative to the other two countries. An assessment of the relative importance of the various influences on wage inequality among Canadian men is a worthwhile subject for future research.

V. Summary and Conclusions

The impact of unions on the structure of wages has recently attracted renewed interest as analysts have struggled to explain the rise in earnings inequality in several industrialized countries. Canada, the U.K., and the U.S. provide a potentially valuable set of countries for examining this question. All three countries now collect comparable data on wages and union status in their regular labor force surveys. Several features of the collective bargaining institutions of these countries make them suitable for studying the relationship between unions and wage inequality. Bargaining is highly decentralized; there

are no general mechanisms for extending collective bargaining provisions beyond the “organized” sector; and the fraction of the work force covered by collective bargaining is relatively modest. Thus it is possible to compare the structure of wages for workers covered by union contracts to those who are not covered, and potentially infer the effect of unions on overall wage inequality.

A number of previous studies, including Freeman (1980, 1982, 1984, 1993), Lemieux (1993), Card (1996, 2001), DiNardo et al. (1996), DiNardo and Lemieux (1997), Machin (1997), and Gosling and Lemieux (2001), have examined the relationship between unionization and wage inequality in these countries individually or in country pairs. Most of the previous work has focused on men and on the 1970s and 1980s. One contribution of this study is to provide a comprehensive analysis of the evolution of unionization and wage inequality for both men and women in all three countries over the past two to three decades. Following the approach developed in Lemieux (1993) and Card (1996), we also take into account variation in collective bargaining coverage and union wage impacts across workers with different levels of observable skills.

In his landmark paper, Freeman (1980) concluded that, overall, unions tend to reduce wage inequality among men because the inequality-increasing “between-sector” effect is smaller than the dispersion-reducing “within-sector” effect. Our analysis indicates that this finding is very robust across countries (U.S., U.K., and Canada) and time periods (from the early 1970s to 2001). Controlling for worker characteristics alters the magnitudes of the “within-sector” and “between-sector” effects, and introduces additional terms that reflect differences in union coverage and union wage effects across skill groups. For men in all three countries both the “within-” and “between-” sector effects decline when we control for the skill composition of the work force. Because union workers are more skilled, on average, than nonunion workers, adjusting for characteristics reduces the magnitude of the “between-sector” effect. The decline in the within-group effect reflects the fact that unionized men are more homogeneous than their nonunion counterparts.

We find remarkably similar patterns in union representation and union wage impacts across skill groups for men in all three countries. Union coverage tends to be concentrated in the middle of the skill distribution, and union wages tend to be compressed relative to nonunion wages. As a consequence, unions have an equalizing effect on the dispersion of wages across skill groups in the three countries, complementing the effect on “within-group” inequality.

Once all these factors are taken into consideration, our calculations imply that unions systematically reduce the variance of wages for men in all three countries, though the magnitudes of the effects are smaller when we control for the skill composition of the work force.

Interestingly, an equally robust finding that emerges from this paper is that unions *do not* reduce

wage inequality among women. In all three countries, this important male-female difference in the impact of unionism is due to a combination of three factors. First, unionized women are more concentrated in the upper end of the wage distribution than their male counterparts. Second, the union wage gap is larger for women than for men, resulting in a larger “between-sector” effect. Third, the union wage gap is larger for lesser than higher skilled men, while this is not the case for women.

Another important conclusion is that the impacts of unions on the wage structure in the U.S., Canada, and the U.K. have followed remarkably similar trends over the last two decades. In all three countries, the unionization rate and the union wage differential have declined substantially since the early 1980s. For men, this has resulted in a steady erosion of the equalizing effect of unions that explains a significant fraction of the growth in wage inequality in the U.S. and U.K. The decline of female unionization has been much smaller than that of men. As a consequence, unionization rates of men and women are nearly equal now in all three countries, marking a sharp departure from the historical pattern. However, the modest decline in union coverage among women had little impact on female wage inequality.

Interestingly, in both the U.S. and the U.K. our estimates of the effects of unions on wage inequality were virtually unchanged between 1993 and 2001. This is consistent with the slowdown in the growth of inequality in both countries during the 1990s, relative to the 1980s. However, in Canada there was little change in wage inequality during the 1980s and 1990s, despite a moderate drop in union coverage among men. The Canadian experience suggests that other factors offset the pressures toward widening inequality associated with the decline in unionization.

Although trends in union coverage and union wage effects are very similar in the three countries, there are substantial differences in the levels of unionization and wage inequality. The pattern of cross-country differences in wage inequality is consistent with the pattern of wage compression effects. Our calculations indicate that differences in union wage compression effects can account for almost one-half of the U.K. - U.S. differential in the variance of wages in the early 1990s, and over two-thirds of the differential in 2001.

In *What Do Unions Do?*, the impact of unions on the distribution of income is a leading example of how the “voice” aspect of unionism (reduced inequality among union workers – the “within-sector” effect) dominates its “monopoly” face (inequality between union and nonunion workers – the “between-sector” effect). Twenty years later, our study confirms that this key finding remains robust to the choice of country and time period. An important qualification, however, is the case of women where we show that these two aspects of the impact of unions on wage inequality more or less offset each other.

What Do Unions Do? recognizes that there is no consensus on the social benefits of the equalizing effects of unions on the distribution of income. It states that “For readers to whom greater inequality is a plus, what do unions do here is definitely good. For readers to whom greater equalization of income is undesirable, what unions do is definitely bad” (p. 247). On balance, however, *What Do Unions Do?* clearly sides with those who think that unless the equalizing effects of unions result in large costs due to allocative inefficiencies, what unions do here is socially good. Although we do not assess the impacts of unions on resource allocation and economic performance, we share the view that the consequences of unions for wage inequality are beneficial from a social point of view.

The last paragraph of *What Do Unions Do?* starts with a dire warning: “All told, if our research findings are correct, the ongoing decline in private sector unionism – a development unique to the United States among developed countries – deserves serious public attention as being socially undesirable.” (p. 251). By linking the decline in unionism to the dramatic increase in wage inequality in the United States since the 1970s, our research strongly confirms that the ongoing decline in private sector unionism indeed had socially undesirable consequences. In retrospect, this sentence of *What Do Unions Do?* only erred by stating that the decline in private sector unionism was a development *unique* to the U.S. In the twenty years following the publication of the book, unionism sharply declined in the United Kingdom and fell moderately in Canada. In the case of the United Kingdom, the decline in unionism has resulted in a steep growth in wage inequality among men. These recent developments show that the social consequences of the decline in unionism deserve even broader attention (north of the border and across the Atlantic) than at the time *What Do Unions Do?* was first published.

NOTES

¹ This possibility was emphasized by Lewis (1963). The presence of unionized employers may lead to higher wages in the nonunion sector (if nonunion employers raise wages to deter unionization efforts) or to lower wages (if unionization reduces employment in the union sector, increasing labor supply in the nonunion sector).

² Equation (6) is only correct if unobserved skills are rewarded equally in the union and nonunion sectors, although it may provide a good first approximation if the rewards for unobserved ability in the union sector are not too much lower than in the nonunion sector. Lemieux (1998) presented a model in which unobserved attributes are rewarded differently in the union and nonunion sectors.

³ A couple of studies in the late 1970s and early 1980s also pointed to the conclusion that unions lowered wage inequality. Hyclak (1979) analyzed the determinants of inequality in wage and salary income in urban labor markets and found that higher union coverage was associated with lower earnings inequality. Hyclak (1980) found a negative relationship between the state mean of union density and the percentage of families with low earnings. Hirsch (1982) performed a cross-sectional study at the industry level using a simultaneous equations model of earnings, earnings dispersion, and union coverage. He concluded that the equalizing effects of unions on earnings inequality are larger when allowance is made for the joint determination of union coverage and wage dispersion. Metcalf (1982) also looked at the dispersion of wages across industries in the U.K. (without controlling for the joint determination of earnings and union coverage) but concluded that union coverage widened the pay structure across industries.

Metcalf also showed, however, that the variation of weekly earnings was lower in the union sector and that unions narrowed the pay structure by occupation and race.

⁴ See Lewis (1986) for a review of U.S. studies and Meng (1990) and Lemieux (1993) for Canadian evidence.

⁵ In the 2001 CLFS, 2.4 percent of male workers and 1.9 of female workers were covered by collective bargaining but not members of a union. The two different measures of unionization lead to nearly identical estimates of the union wage premium in a conventional linear regression of wages on union status, education, and experience.

⁶ In the U.S., for example, we use bins for the log hourly wage of width 0.05. We use smaller bins for our U.K. and Canadian samples.

⁷ The densities are estimated using a bandwidth of 0.05. See DiNardo et al. (1996) for more detail.

⁸ This is similar to DiNardo et al. (1996) who showed that the minimum wage has a much larger impact on women than on men.

⁹ The derivative of the right hand side of equation (4') with respect to the unionization rate is

$\Delta_v + (1-2U)\Delta_w^2$. This is negative as long as Δ_v is large relative to Δ_w^2 .

¹⁰ Taking a more direct approach, Hirsch and Schumacher (1998) examined test-score data and found that union members with high measured skills had relatively low test scores.

¹¹ DiNardo et al. (1996), Card (2001), and Gosling and Lemieux (2001) all concluded that de-unionization explains very little of the increase in wage inequality among women in the U.S. or U.K.

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DATA APPENDIX

U.S. Data: Since 1979, the U.S. Census Bureau has been collecting data on weekly hours, weekly earnings, and hourly earnings (for workers paid by the hour) for all wage and salary workers in the “outgoing rotation group” (ORG) of the Current Population Survey (CPS). Beginning in 1983, the ORG supplement of the CPS also asks about the union status of workers (and union coverage). Similar variables are also available in the May supplement of the CPS between 1973 and 1978, though only union membership (and not coverage) is available for this period.

In both the May and ORG supplements of the CPS, workers paid by the hour are asked their hourly rate of pay. We use this variable, which is collected in a consistent fashion over time, as our measure of the hourly wage rate for these workers. The May and ORG supplements also provide information on usual weekly earnings for all workers. For workers not paid by the hour, we use average hourly earnings (weekly earnings divided by weekly hours) as our measure of the wage rate.

Note, however, that weekly earnings are not measured consistently over time. From 1973 to 1993, this variable was collected by asking individuals directly about their earnings on a weekly basis. From 1994 to 2001, individuals had the option of reporting their usual earnings on the base period of their choice (weekly, bi-weekly, monthly, or annually). Weekly earnings are then obtained by normalizing the earnings reported by workers to a weekly basis. The available evidence does not suggest, however, that this change in the way earnings are collected had a significant impact on the distribution of wages (see Card and DiNardo (2002) and Gosling and Lemieux (2001) for more detail).

Another potential problem is that weekly earnings are top-coded at different values for different years throughout the sample period. Before 1988, weekly earnings were top-coded at \$999. The top-code was later increased to \$1,923 in 1988 and \$2,884 in 1998. For an individual working 40 hours a week, the weekly earnings top code corresponds to an hourly wage ranging from of \$42.6 in 1984 (\$2001) to \$99.6 in 1973 (\$2001). To keep the wage samples relatively comparable over time, we trim observations with wages above \$63 (\$2001). We also trim observations with wages below \$2.5 (\$2001), which typically corresponds to about half of the minimum wage. The wage deflator used is the Consumer Price Index (CPI-U). All the U.S. wage statistics reported herein are also weighted using the CPS earnings weights.

Questions about educational achievement were changed substantially in the early 1990s. Until 1991, the CPS asked about the highest grade (or years of schooling) completed. Starting in 1992, the CPS moved to questions about the highest degree. We have recoded the post-1992 data in term of completed years of schooling to have a measure of schooling that is consistent over time. We then use years of schooling to compute the standard measure of years of potential experience (age-schooling-6). Only observations with potential experience larger or equal than zero are kept in the analysis samples.

Finally, in the 1979-2001 ORG supplements of the CPS, wages or earnings of workers who refuse to answer the wage/earnings questions were allocated using a “hot deck” procedure. We exclude observations with allocated wages and earnings for two reasons. First, wages and earnings were not allocated in the May 1973-1978 CPS. We thus need to exclude allocated observations from the 1984, 1993, and 2001 ORG supplement data to maintain a consistent sample over time. Second, union status is not one of the characteristics used to match observations with missing earnings to observations with non-missing earnings in the imputation procedure (hot deck) used by the U.S. Census Bureau. As a result, estimates of union wage effects obtained from a sample with allocation observations included can be severely biased downward (see Hirsch and Schumacher, 2003 for more details).

U.K. Data: As mentioned in the text, for the U.K. we use data from the 1983 GHS and the 1993 and 2001 UKLFS. For the sake of consistency, we exclude observations from Northern Ireland since this region was sampled in the UKLFS but not in the GHS. Real wages are obtained by deflating nominal wages with the Consumer Price Index (Retail Price Index). To limit the effect of outliers, we only keep observations with an hourly wage rate between 1.5 and 50 pounds (in 2001 pounds).

In general, we process the U.K. samples to make them as comparable as possible to the U.S. samples. In both the UKLFS and the GHS, we use observations for wage and salary workers with non-missing wages and earnings. We also use the sample weights whenever available (there are no sample weights in the GHS). Since education is not consistently measured over time, we recode education into five broad categories that are consistent over time: university graduates, higher-level vocational training and A-level qualifications, middle-level vocational training or O-level qualifications, lower-level vocational training, and no qualifications or diploma.

Canadian Data: As mentioned in the text, for Canada we use the 2001 Labour Force Survey (CLFS), the 1991 and 1995 Surveys on Work Arrangements (SWA), and the 1984 Survey of Union Membership (SUM). These data sets are all relatively comparable since both the SUM and the SWA were conducted as supplements to the Labour Force Survey. Relative to the U.S. and U.K. data however, there are some important limitations in the Canadian data. First, as mentioned in the text, it is not possible to distinguish union membership from union coverage in the SWA. For the sake of consistency over time, we thus use union coverage as our measure of unionization in Canada.

A second limitation is that in the 1984 SUM and the 2001 CLFS missing wages and earnings were allocated but no allocation flags are provided. We thus have to include observations with allocated wages and earnings in the analysis with generates an inconsistency relative to the SWA (where missing wages and earnings are not allocated) and the U.S. and U.K. data. This likely understates the effect of unions on wages in 1984 and 2001, though it is not possible to quantify the extent of the bias. Another limitation is that age is only provided in broad categories, unlike in the U.S. and U.K. data where age is reported in years. In particular, it is not possible to separate workers age 15 from those age 16. This explains why we use all wage and salary workers age 15 to 64 in Canada, compared to workers age 16 to 64 in the two other countries.

A further limitation is that hourly wages are top-coded at a relatively low level in the Canadian data. The top codes are \$45 in the 1984 SUM, \$50 in the 1991 SWA, \$40 in the 1995 SWA, and \$100 in the 2001 CLFS. For the sake of consistency, we trim observations with hourly wages above \$44 (in \$2001). Wages are deflated using the Canadian CPI for all items. We also trim observations with wages below \$2.5 in \$2001, which represents about half of the minimum wage.

One final limitation is that only five education categories are consistently available over time. These categories are: 0 to 8 years of school, high school (some or completed), some post-secondary education, post-education degree or diploma (less than university), and university degree. As in the CPS and the UKLFS, all statistics for Canada are computed using sample weights.

Table 1
Effect of Unions on Wage Structure of U.S. Workers, 1973-2001

	1973/1974		1984		1993		2001	
	male	female	male	female	male	female	male	female
Fraction Union Members	0.307	0.141	0.236	0.141	0.185	0.132	0.149	0.121
<u>Mean Log Wages (2001\$):</u>								
Nonunion Workers	2.646	2.270	2.573	2.276	2.535	2.337	2.667	2.457
Union Workers	2.841	2.499	2.866	2.605	2.838	2.686	2.899	2.761
Union Gap (unadjusted)	0.196	0.230	0.293	0.329	0.304	0.349	0.233	0.305
Union Gap (adjusted)	0.185	0.220	0.208	0.228	0.210	0.210	0.156	0.149
<u>Standard Deviation Log Wages:</u>								
Nonunion Workers	0.553	0.442	0.563	0.467	0.594	0.515	0.601	0.538
Union Workers	0.354	0.383	0.363	0.408	0.399	0.444	0.417	0.460
Union Gap	-0.198	-0.059	-0.199	-0.058	-0.194	-0.071	-0.184	-0.077
<u>Variance Decomposition:</u>								
Overall Variance	0.258	0.195	0.289	0.223	0.331	0.270	0.340	0.289
<i>Two sector model</i>								
Within-sector effect	-0.055	-0.007	-0.044	-0.007	-0.036	-0.009	-0.028	-0.009
Between-sector effect	0.008	0.006	0.015	0.013	0.014	0.014	0.007	0.010
Total effect	-0.047	0.000	-0.028	0.006	-0.022	0.005	-0.021	0.001
<i>Model with skill groups</i>								
Within-sector effect	-0.022	-0.006	-0.020	-0.007	-0.018	-0.009	-0.013	-0.009
Between-sector effect	0.007	0.004	0.010	0.008	0.012	0.011	0.004	0.008
Dispersion across groups	-0.011	0.001	-0.007	0.000	-0.008	-0.003	-0.006	-0.005
Total effect	-0.026	0.000	-0.017	0.001	-0.014	-0.001	-0.015	-0.007
Sample Size	43,189	30,500	77,910	69,635	71,719	69,723	55,813	55,167
Number of Skill Groups	180	180	343	343	244	246	245	246

Note: Samples include wage and salary workers age 16-64 with non-allocated hourly or weekly pay, and hourly wages between \$2.00 and \$90.00 per hour in 1989 dollars.

Table 2
Effects of Unions on Wage Structure of Canadian Workers, 1984-2001

	1984		1991-1995		2001	
	male	female	male	female	male	female
Fraction Union Workers	0.467	0.369	0.408	0.353	0.330	0.317
<u>Mean Log Wages (2001\$)</u>						
Nonunion Workers	2.658	2.365	2.661	2.452	2.728	2.495
Union Workers	2.987	2.793	2.972	2.851	2.964	2.853
Union Gap (unadjusted)	0.330	0.428	0.311	0.398	0.236	0.358
Union Gap (adjusted)	0.251	0.321	0.204	0.275	0.153	0.226
<u>Standard Deviation Log Wages:</u>						
Nonunion Workers	0.528	0.446	0.514	0.465	0.501	0.463
Union Workers	0.343	0.368	0.362	0.380	0.386	0.395
Union Gap	-0.185	-0.078	-0.152	-0.084	-0.115	-0.068
<u>Variance Decomposition:</u>						
Overall Variance	0.231	0.218	0.233	0.227	0.229	0.224
<i>Two sector model</i>						
Within-sector effect	-0.075	-0.023	-0.054	-0.025	-0.034	-0.019
Between-sector effect	0.027	0.043	0.023	0.036	0.012	0.028
Total effect	-0.048	0.019	-0.031	0.011	-0.021	0.009
<i>Model with skill groups</i>						
Within-sector effect	-0.041	-0.027	-0.033	-0.028	-0.025	-0.022
Between-sector effect	0.017	0.022	0.010	0.017	0.006	0.012
Dispersion across groups	-0.014	0.014	-0.002	0.014	0.001	0.013
Total effect	-0.037	0.009	-0.025	0.002	-0.017	0.003
Sample size	17,737	15,356	17,981	18,323	24,003	23,703
Number of skill groups	25	25	25	25	25	25

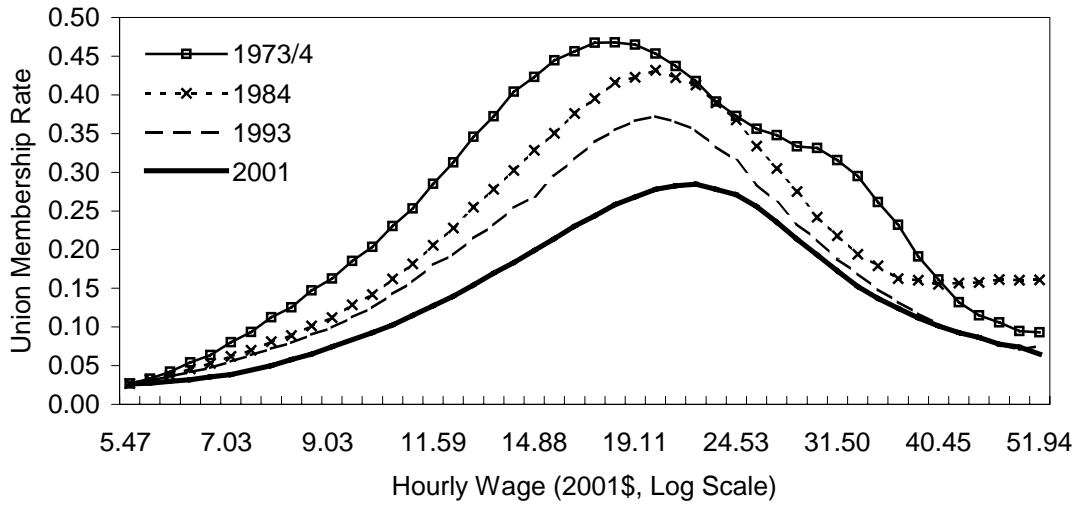
Note: Samples include wage and salary workers age 15-64 with allocated hourly or weekly pay (except in 1991-95), and hourly wages between \$2.50 and \$44.00 per hour in 2001 dollars.

Table 3
Effects of Unions on Wage Structure of U.K. Workers, 1983-2001

	1983		1993		2001	
	male	female	male	female	male	female
Fraction union workers	0.570	0.426	0.392	0.337	0.307	0.285
<u>Mean log wages (2001£):</u>						
Nonunion Workers	1.843	1.416	2.036	1.705	2.170	1.873
Union Workers	2.053	1.685	2.224	2.047	2.306	2.167
Union Gap (unadjusted)	0.210	0.269	0.188	0.342	0.135	0.294
Union Gap (adjusted)	0.162	0.195	0.131	0.184	0.045	0.137
<u>Standard Deviation of Log Wages:</u>						
Nonunion Workers	0.532	0.412	0.586	0.499	0.588	0.510
Union Workers	0.382	0.399	0.438	0.475	0.442	0.468
Union Gap	-0.150	-0.013	-0.148	-0.024	-0.146	-0.043
<u>Variance Decomposition:</u>						
Overall Variance	0.216	0.183	0.293	0.268	0.303	0.266
<i>Two sector model</i>						
Within-sector effect	-0.078	-0.004	-0.059	-0.008	-0.046	-0.012
Between-sector effect	0.011	0.018	0.008	0.026	0.004	0.018
Total effect	-0.067	0.013	-0.051	0.018	-0.042	0.006
<i>Model with skill groups</i>						
Within-sector effect	-0.034	-0.023	-0.031	-0.028	-0.032	-0.025
Between-sector effect	0.009	0.011	0.006	0.008	0.002	0.005
Dispersion across groups	-0.026	0.009	-0.016	0.012	-0.013	0.010
Total effect	-0.050	-0.003	-0.041	-0.008	-0.042	-0.010
Sample size	4,435	3,512	4,009	4,139	7,548	8,113
Number of skill groups	25	25	25	25	25	25

Note: Samples include wage and salary workers age 15-64 with non-missing hourly or weekly pay, and hourly wages between £1.50 and £50.00 per hour in 2001 pounds.

1a. U.S. Men



1b. U.S. Women

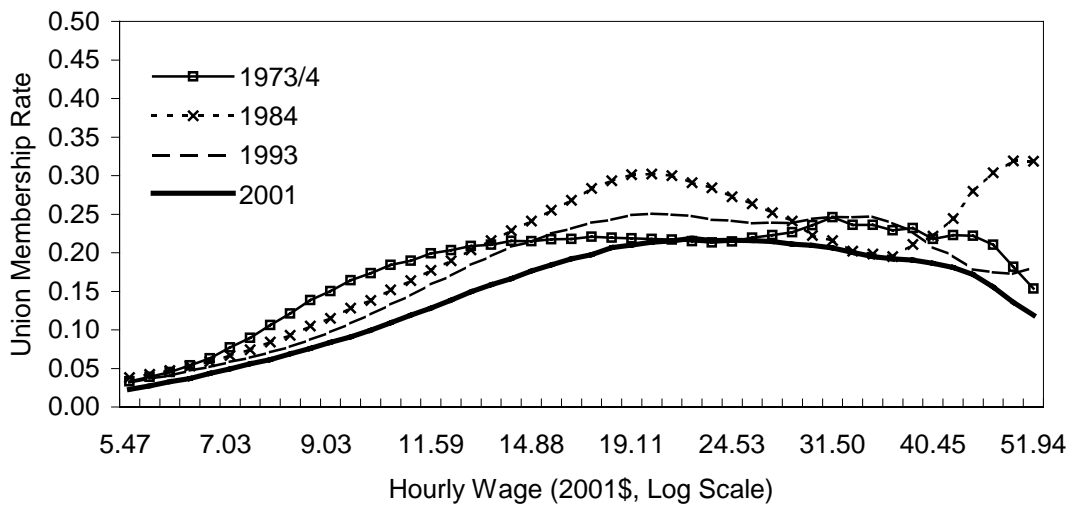
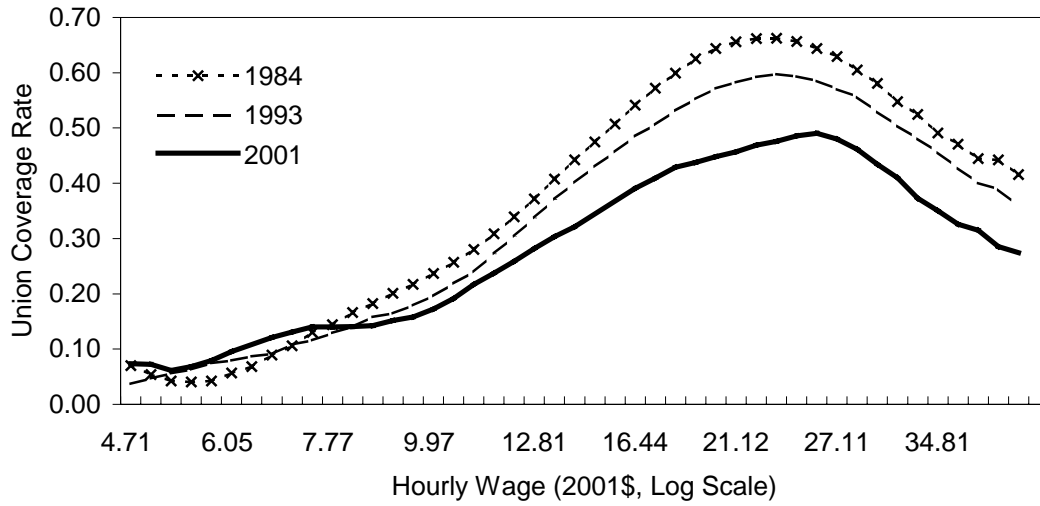


Figure 1. Unionization Rate by Wage Level, United States

2a. Canadian Men



2b. Canadian Women

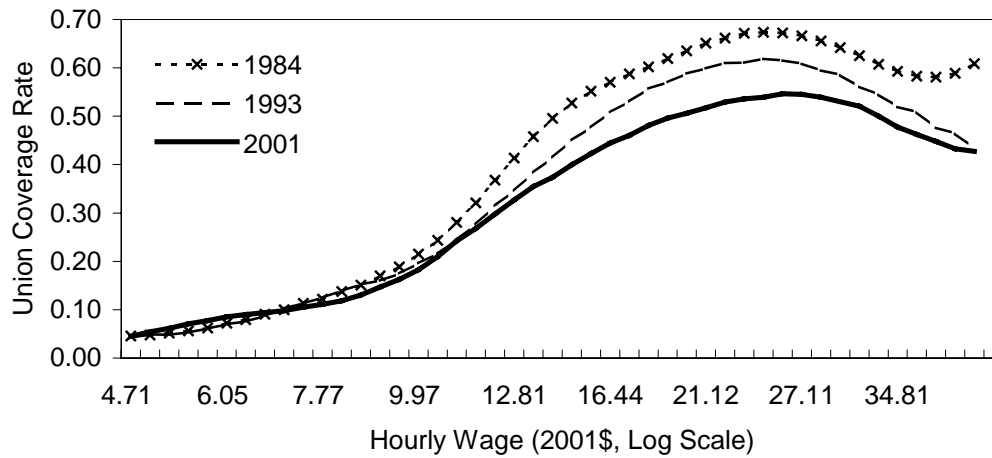
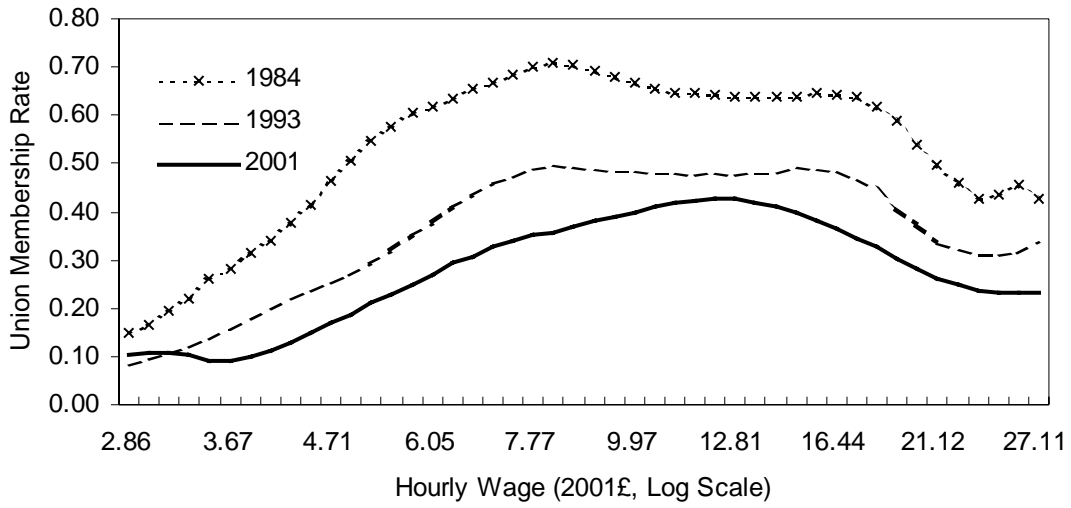


Figure 2. Unionization Rate by Wage Level, Canada

3a. U.K. Men



3b. U.K. Women

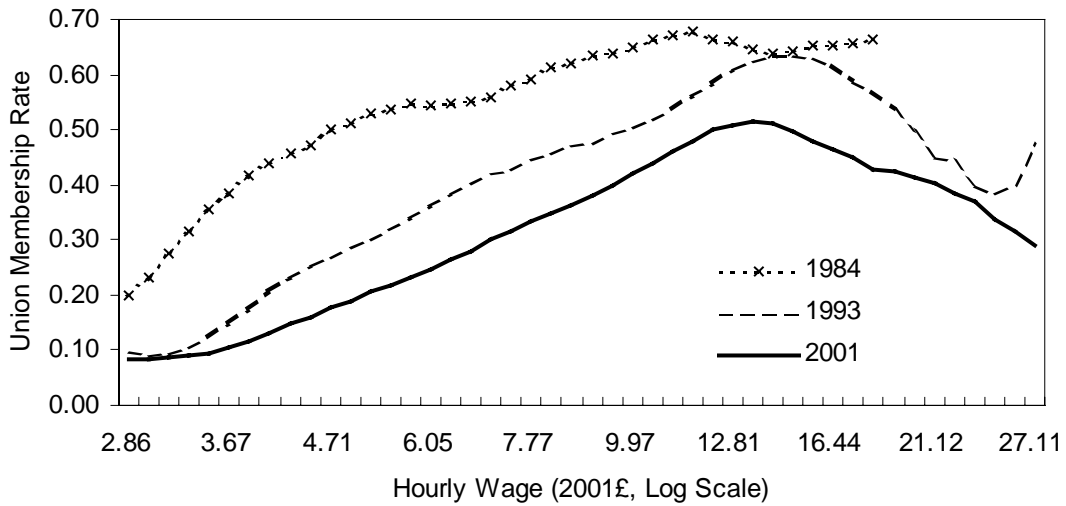
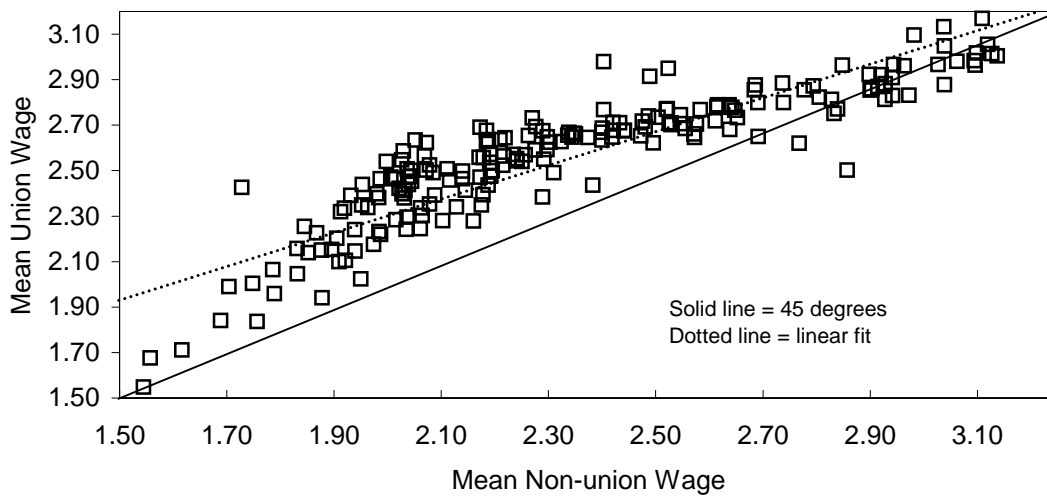


Figure 3. Unionization Rate by Wage Level, United Kingdom

4a. U.S. Men, 1993



4b. U.S. Women, 1993

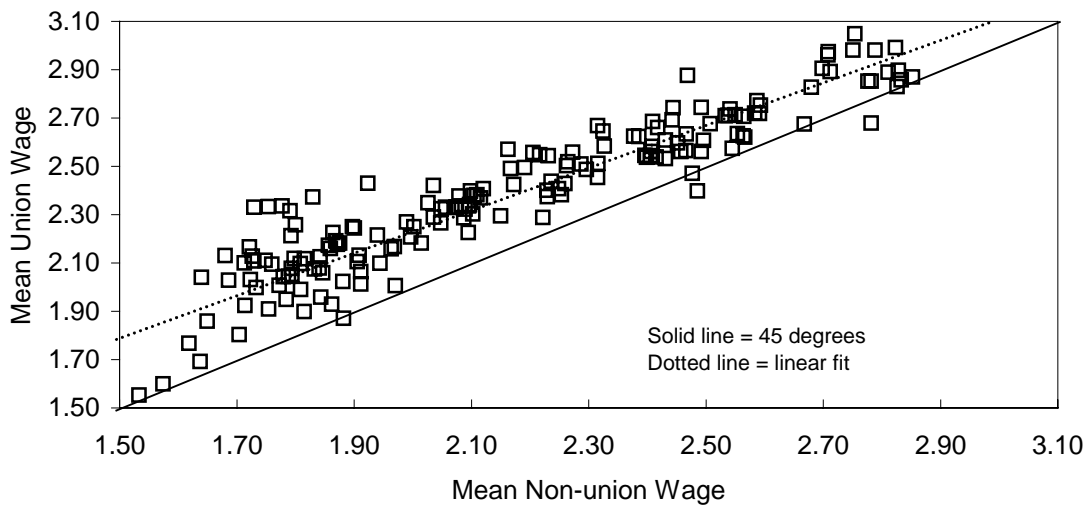
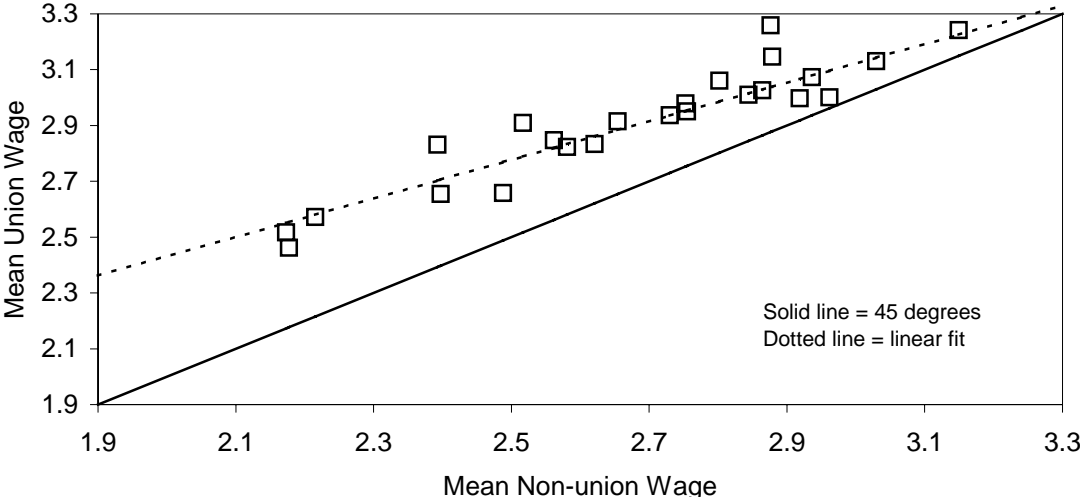


Figure 4. *Union Relative Wage Structure in the United States, 1993*

5a. Canadian Men, 1991-1995



5b. Canadian Women, 1991-1995

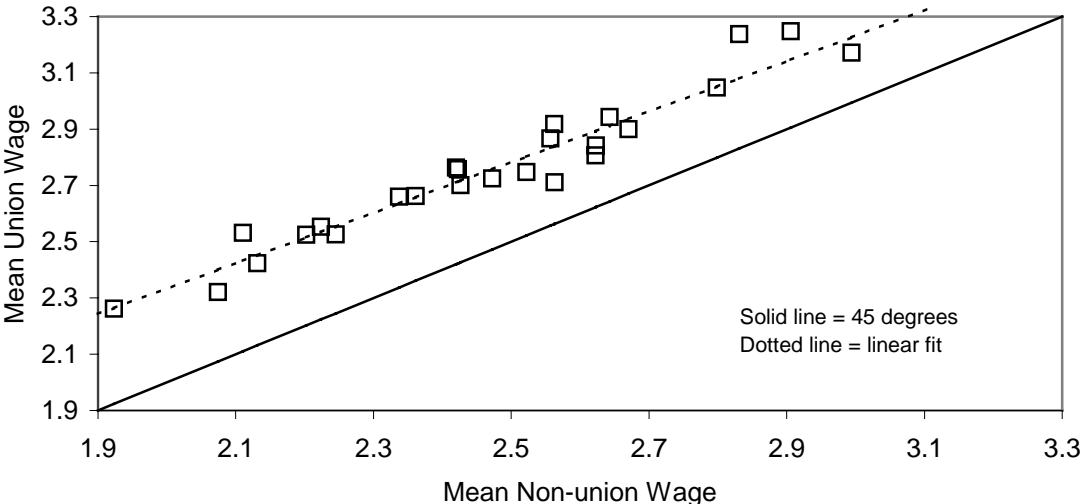
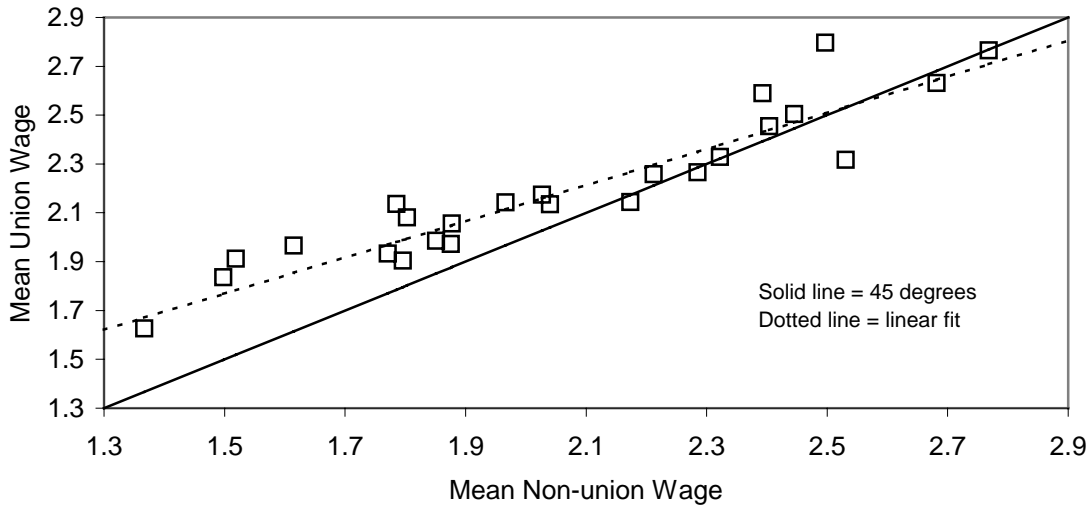


Figure 5. Union Relative Wage Structure in Canada, 1991-1995

6a. U.K. Men, 1993



6b. U.K. Women, 1993

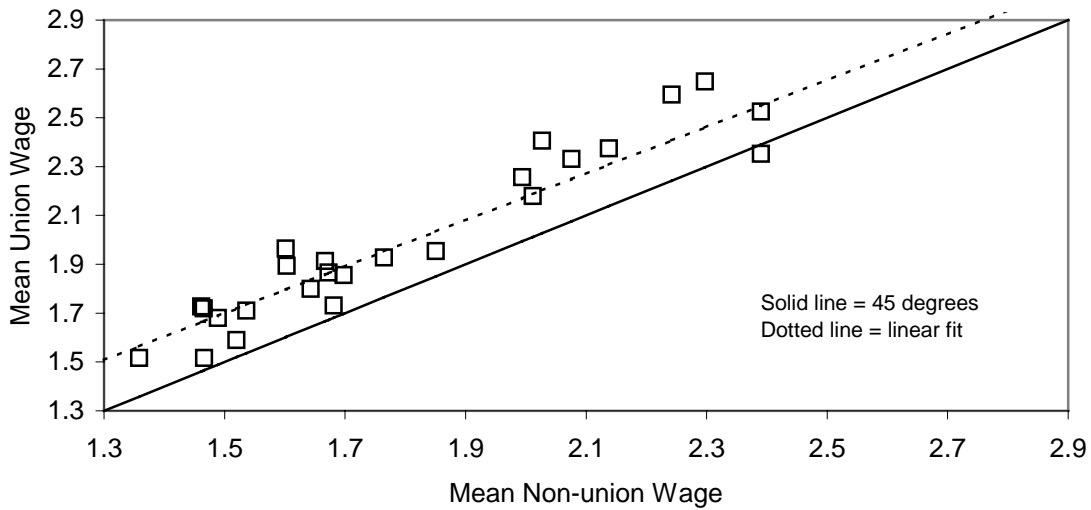
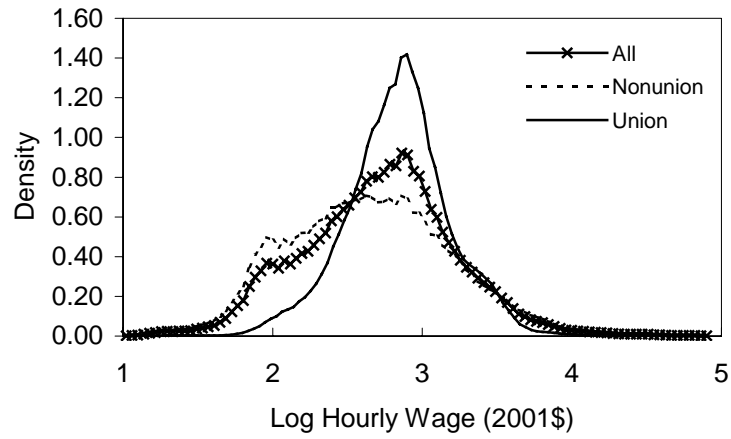
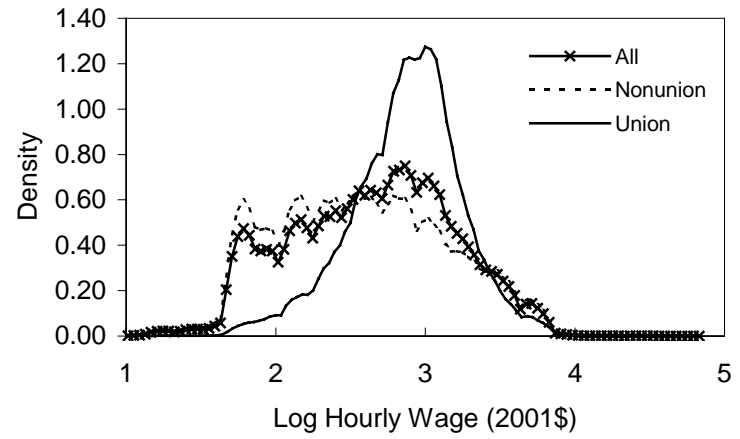


Figure 6. *Union Relative Wage Structure in the United Kingdom, 1993*

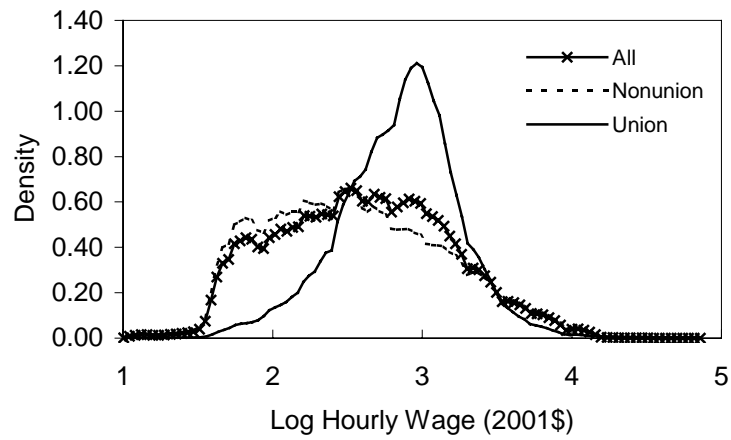
7a. U.S. Men, 1973/1974



7b. U.S. Men, 1984



7c. U.S. Men, 1993



7d. U.S. Men, 2001

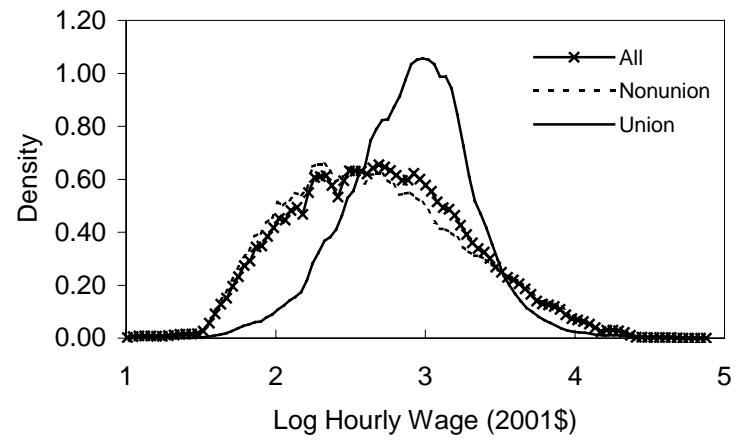
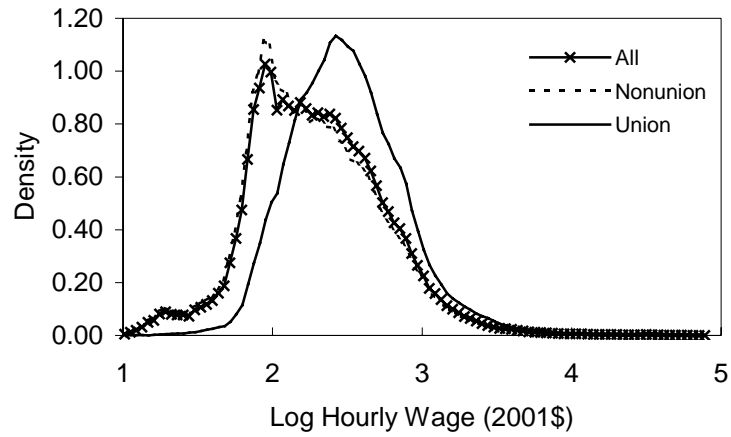
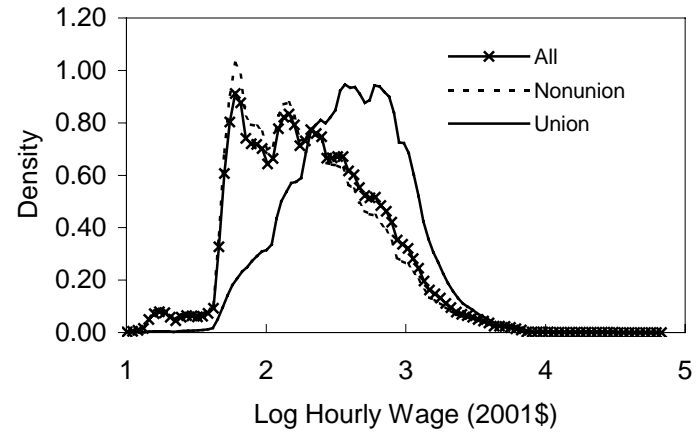


Figure 7. *Density of Wages, U.S. Males*

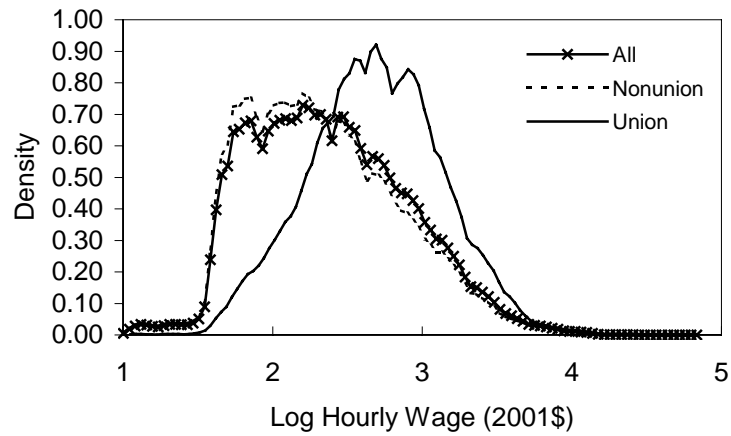
8a. U.S. Women, 1973/1974



8b. U.S. Women, 1984



8c. U.S. Women, 1993



8d. U.S. Women, 2001

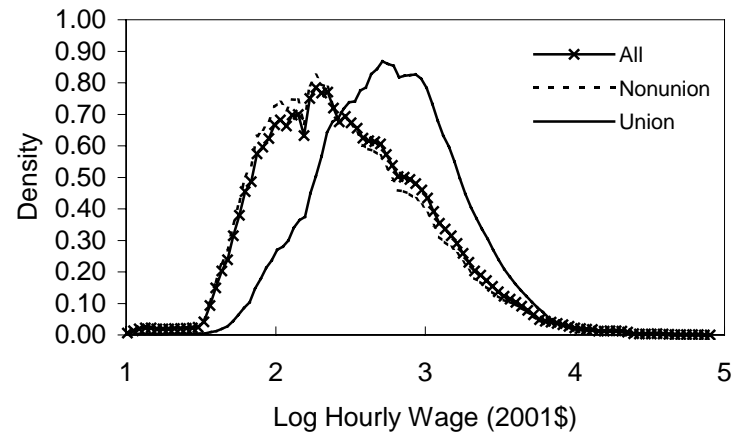
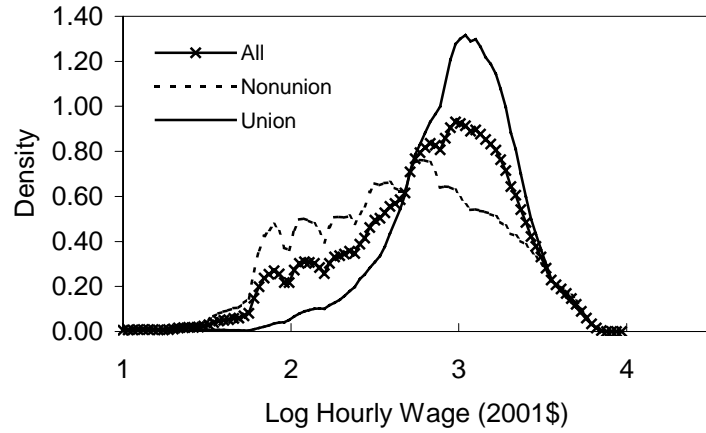
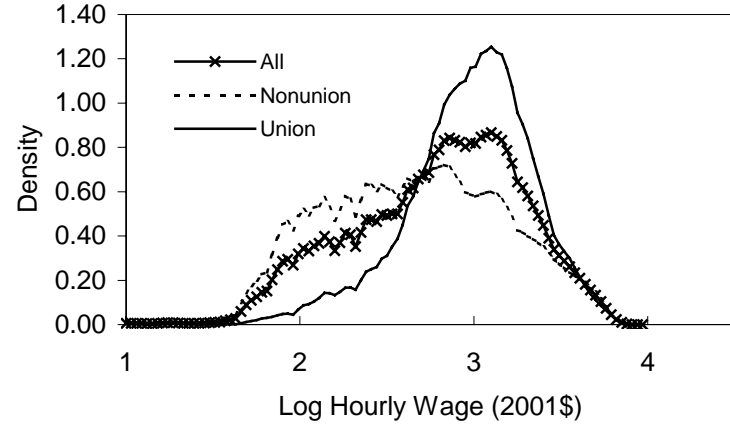


Figure 8. *Density of Wages, U.S. Females*

9a. Canadian Men, 1984



9b. Canadian Men, 1991-1995



9c. Canadian Men, 2001

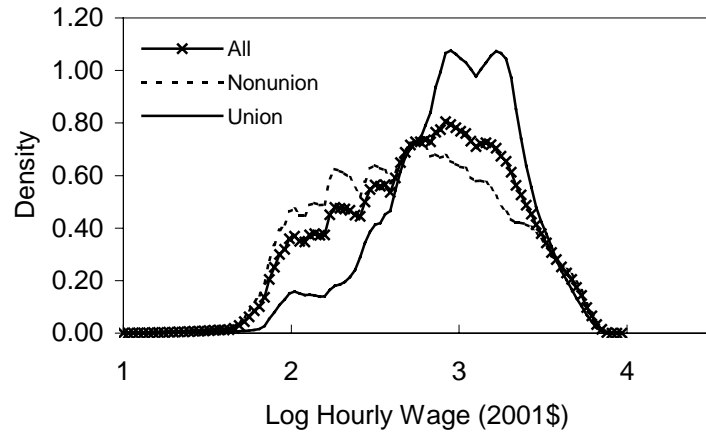
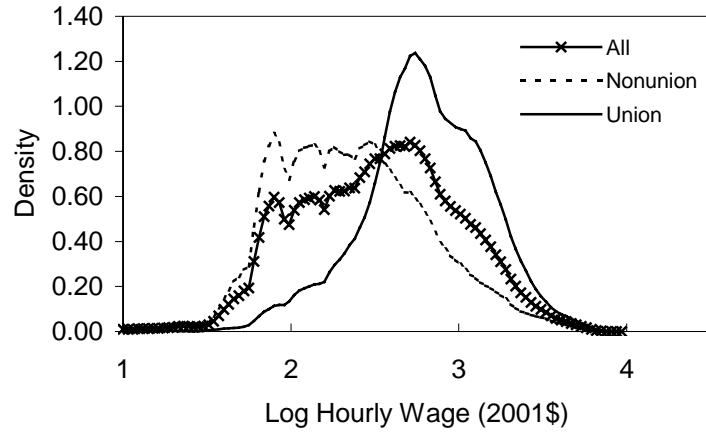
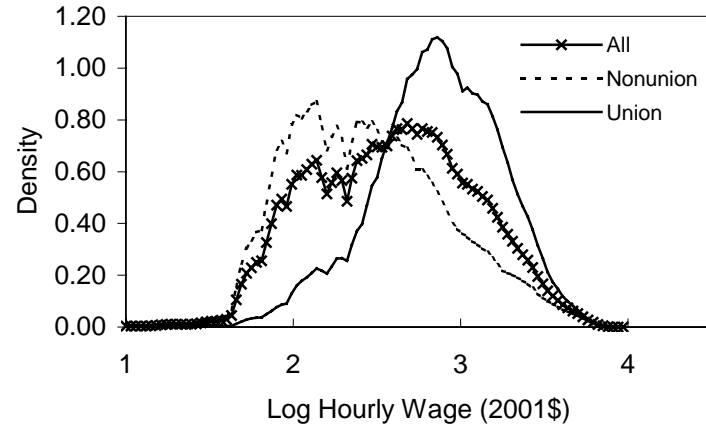


Figure 9. *Density of Wages, Canadian Males*

10a. Canadian Women, 1984



10b. Canadian Women, 1991-1995



10c. Canadian Women, 2001

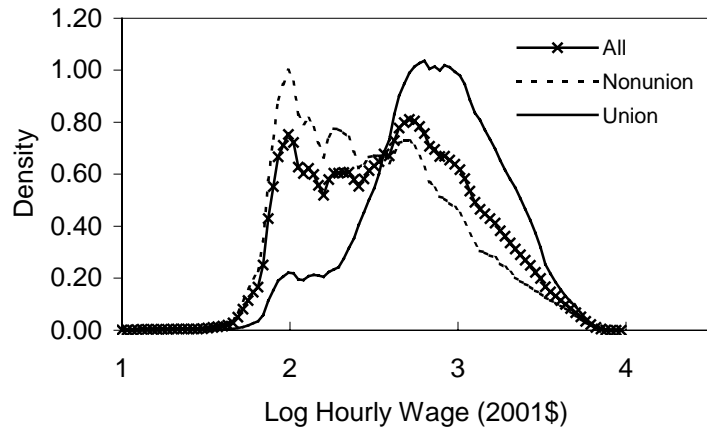
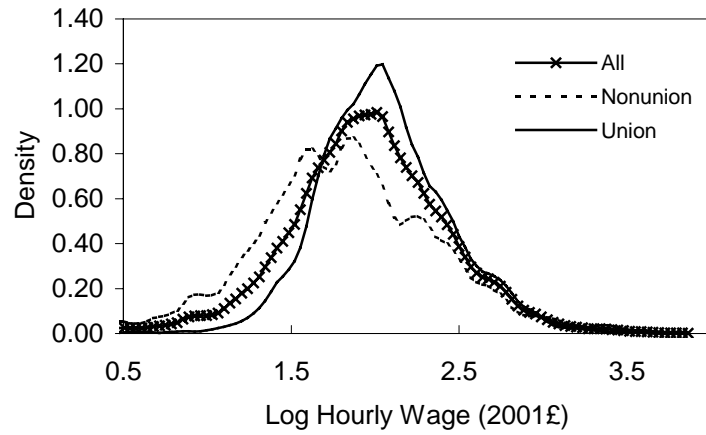
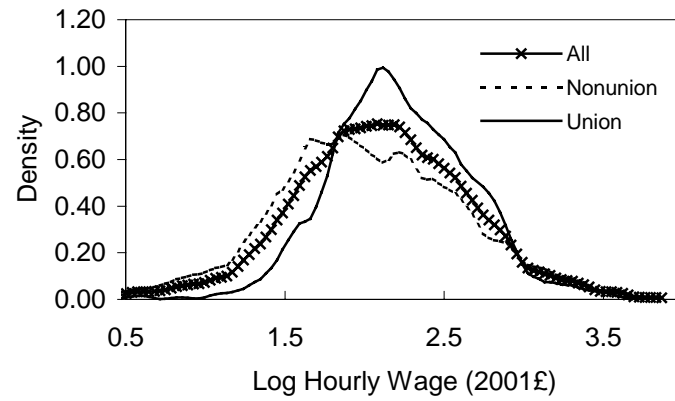


Figure 10. *Density of Wages, Canadian Females*

11a. U.K. Men, 1984



11b. U.K. Men, 1993



11c. U.K. Men, 2001

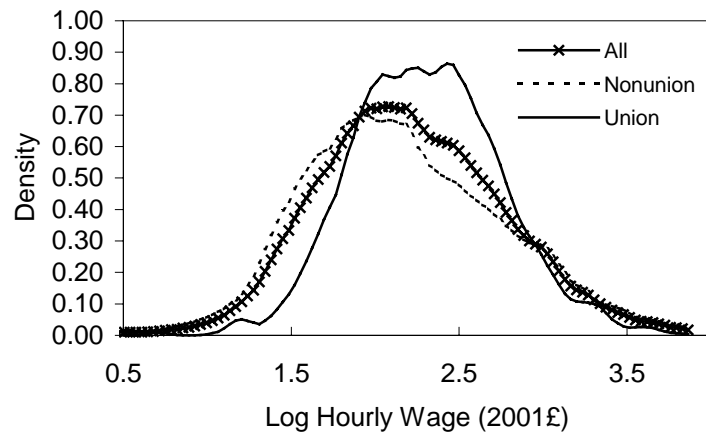
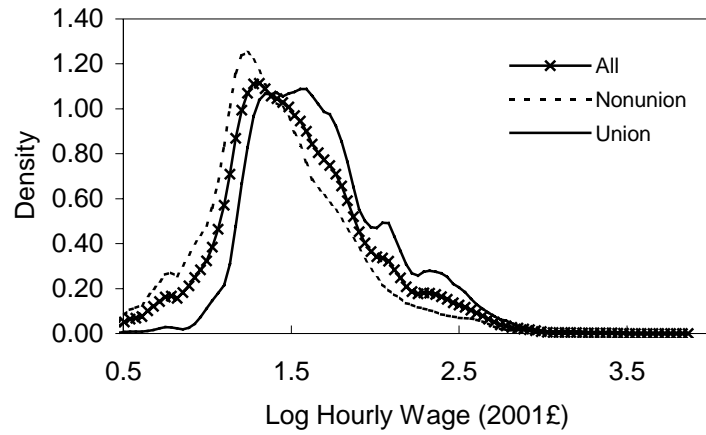
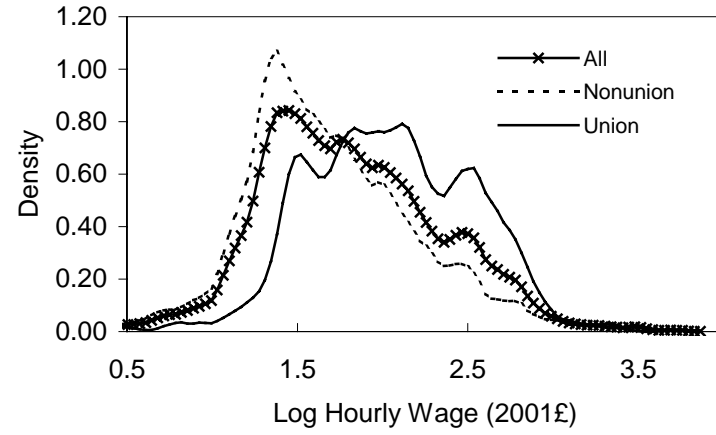


Figure 11. *Density of Wages, U.K. Males*

12a. U.K. Women, 1984



12b. U.K. Women, 1993



12c. U.K. Women, 2001

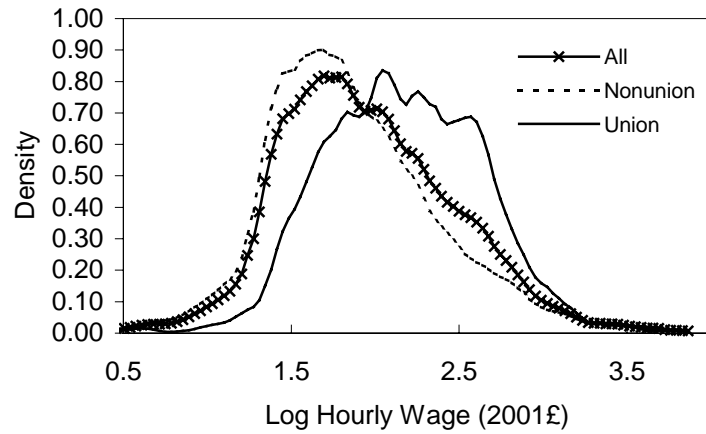


Figure 12. *Density of Wages, U.K. Females*