

# Unions and the Wage Structure

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## I. Introduction

This chapter discusses the impact of unions on the wage structure -- the way in which wages vary systematically with characteristics such as education, age, gender, or occupation. Do unions widen or narrow pay differentials between the skilled and unskilled, between men and women, or between blue-collar and white-collar workers? Is the net effect of unions to increase or decrease overall wage inequality? These questions have long intrigued social scientists. Recently, they have attracted renewed interest as analysts have struggled to explain the rise in earnings inequality in several industrialized countries. The fact that two of the countries with the largest declines in unionization -- the US and the UK -- 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?

The impact of unions on the wage structure depends on the industrial relations system -- the social, political, legal, institutional and economic environment in which unions operate. Countries vary widely in their industrial relations systems, and these differences potentially affect both the goals of unions, and their ability to achieve these goals. In some countries unions exert considerable influence on the political process. By supporting minimum wage or pay equity legislation, for example, unions may be able to alter the wage structure in the economy. Unions also affect the wage structure directly through collective bargaining. This influence in turn depends on the extent of union organization in the labour force, and the extent to which collectively bargained wage structures are legally imposed or voluntarily adopted by employers outside of the "covered" sector. For these reasons, the mechanisms through which unions alter the wage structure and the magnitude of these impacts are likely to vary across countries.

Assessing the impact of unions on the wage structure raises the familiar, but nonetheless difficult, challenge of determining an appropriate counterfactual. We can observe the wage structure at a particular point in time, but we cannot observe what the wage structure would look like without unions, or with a different level of union organization and influence. Some form of modeling is required to estimate the counterfactual. There are a number of possible approaches to this problem. Most progress has been made in cases where the non-union wage structure provides an arguably

appropriate benchmark for the wage structure in the absence of unions. Comparisons over time and across countries have also been useful. The former are most compelling when there have been substantial changes in union strength, while the latter are most informative when otherwise similar countries differ substantially in the extent of union organization.

In this chapter we focus on assessing the influence of unions on the wage structures of three countries -- Canada, the UK and the US. These are the countries about which the most is currently known. In part this is because of data availability. More importantly, however, in these countries the nonunion wage structure provides a plausible counterfactual. In all three countries there is a relatively clear distinction between the union and non-union sectors, and there is no legal mechanism to extend collective bargaining provisions to the non-union sector. If the unorganized sector is to serve as a benchmark for wage-setting in the absence of unions, it is crucial to be able to precisely identify those workers whose wages are unaffected by direct union influence. Second, in these countries the non-union sector is relatively large. The relative size of the non-union sector reduces, but does not eliminate, concern that union wage patterns may indirectly influence wages in the non-union sector through market or non-market spillover mechanisms. Third, in these three countries the main way that unions influence wages is through collective bargaining. In other countries where unions have a substantial effect on wage policies through lobbying and direct political involvement, the non-union wage structure is unlikely to provide a good estimate of the wage structure in the absence of unions.

After briefly reviewing patterns of unionism and collective bargaining coverage in industrialized countries, the next section lays out a framework for measuring the effect of unions on wage inequality, under the assumption that the non-union wage structure is an appropriate counterfactual. Then, we present a review of the literature on unionism and wage inequality, focusing on contributions written since 1975. The last section of the paper presents a comprehensive re-analysis of the link between unions and wage inequality in the US, Canada, and the UK, using micro datasets from the three countries.

## II. Collective Bargaining and Union Membership Coverage

Table 1 shows union membership and collective agreement coverage as a percentage of paid (wage and salary) workers in various OECD countries in 1980 and 1994. In Canada and the US, union representation and collective bargaining are regulated by an elaborate legal framework -- the “Wagner Act” model. In this system, workers who meet the statutory definition of an employee have the right to union representation and collective bargaining. The procedures for defining appropriate bargaining units and for certifying bargaining representatives are administered by a quasi-judicial body, often referred to as a Labour Relations Board. Once a group of workers chooses to be represented by a union (usually by majority vote or a card-signing system), the union becomes the exclusive bargaining representative of all employees in the bargaining unit, including those who choose not to formally join the union. Nearly all bargaining units cover a subset of the workers at a single firm – for example, the United Automobile Workers represent the production non-supervisory workers at Ford plants in the US. The UK has also recently adopted the Wagner Act model, although traditionally its system of union recognition and collective bargaining was more informal.<sup>1</sup>

In contrast to the US, Canada, and the UK, where highly decentralized firm-by-firm bargaining is the norm, centralized bargaining between unions and groups of employers in an industry or region is the usual case in Australia and many European countries. In some countries these agreements set legally binding minimum pay levels for all employers. As a consequence, there is no logical connection between union membership and collective agreement coverage, and the gap between the two can be substantial. The extreme case is France, where 95 percent of workers are estimated to be covered by collective agreements, yet union membership is under 10 percent. In other countries, such as Germany, industry wide contracts are not necessarily binding on all employers, but a majority of employers traditionally adhere to the contracts. In either environment it is difficult or impossible to make a meaningful distinction between the union and nonunion sectors.

Furthermore, as shown in Table 1, only a small minority of workers in Continental Europe and Australia has their wages set outside the umbrella of collective bargaining. The case of the Scandinavian countries is particularly instructive. In

Finland, Sweden, and Norway powerful confederations of trade unions have organized a large percentage of the labour force, leaving only a small minority of paid workers outside the covered sector. Wage patterns for these workers are unlikely to be representative of the labour force as a whole, and also are likely to be strongly influenced by patterns in the organized sectors. Thus the nonunion wage structure will not provide a suitable benchmark for comparison with the union wage structure. The Scandinavian trade unions have pursued strongly egalitarian ('solidaristic') wage policies, and these countries are characterized by wage differentials across skill groups and occupations that are small by international standards. One might be tempted to attribute the compressed wage structure to union wage policies, but these countries are also characterized by close ties between union confederations and social democratic political parties with a highly egalitarian bent. Separating the direct and indirect effects of unions on the wage structure would be a difficult task in these circumstances.

Two additional reasons for focusing on Canada, the UK and the US are also evident in Table 1. The UK experienced a substantial decline in union coverage in the 1980s and 1990s. In contrast, the US experienced a moderate decline, while union coverage was nearly constant in Canada over the 1980-94 period. Thus we may be able to use the different experiences of these three countries to learn about the effects of changes in union coverage on the wage structure. Similarly, there is a substantial gap between Canada and the US in union density -- coverage in Canada is approximately double that of the US. Since labor markets are otherwise fairly similar in the two countries, the gap provides another opportunity to examine the impact of unions on the wage structure.

### **III. Unions and the Structure of Relative Wages**

The impact of unions on wages structures in countries such as Canada, the UK and US is determined by two factors: which workers are covered by unions; and how unions alter the pay of those who are covered. To illustrate these forces, consider the effect of unions on the average wage gap between men and women. The economy-wide average wage for either gender is a weighted average of the gender-specific means in the union and nonunion sectors:

$$W_m = \alpha_m W_m^U + (1-\alpha_m) W_m^N \quad (1)$$

$$W_f = \alpha_f W_f^U + (1-\alpha_f) W_f^N \quad (2)$$

where  $W_m$  denotes the average wage of men,  $W_m^U$  is the average wage of men in the union sector,  $W_m^N$  is the average wage of men in the non-union sector,  $\alpha_m$  is the fraction of male workers covered by union agreements, and we employ similar notation for women using the subscript  $f$ . Combining these equations, the gender wage gap is:

$$W_m - W_f = W_m^N - W_f^N + \alpha_m (W_m^U - W_m^N) - \alpha_f (W_f^U - W_f^N) \quad (3)$$

The final two terms in this expression are the average wage gains for men and women associated with the presence of unionism (Lewis, 1963). These are the products of the extent of union representation ( $\alpha_m$  and  $\alpha_f$ ) and the respective union wage gaps ( $W_m^U - W_m^N$ ) and ( $W_f^U - W_f^N$ ).

The influence of these two factors can be seen most clearly by considering the special cases in which union coverage of men equals that of women ( $\alpha_m = \alpha_f = \alpha$ ) and in which the union relative wage effect is the same for both genders. In the case of equal coverage,

$$W_m - W_f = W_m^N - W_f^N + \alpha [(W_m^U - W_m^N) - (W_f^U - W_f^N)] \quad (4)$$

so that unions narrow the gender wage gap if the union wage impact for women exceeds that for men, and vice versa. In the case where the union wage impact is the same for both men and women, we obtain

$$W_m - W_f = W_m^N - W_f^N + (\alpha_m - \alpha_f)[W_m^U - W_m^N] \quad (5)$$

so that unions narrow the gender wage gap if female coverage exceeds that of males, and vice versa. The two effects may operate in the same direction, as would be the case if the union wage gap is greater for men than women and union coverage is also greater for

men, or they may act in opposite directions. Indeed, using Canadian data from the 1980s, Doiron and Riddell (1994) concluded that unions raise the wages of women more than of men but collective agreement coverage is greater for males. In this case the net effect -- which they estimate to be approximately zero for Canada in the 1980s -- depends on the magnitudes of these offsetting factors.

The importance of the two factors may also change over time. For example, the rise of unionism in the public sector in many countries has resulted in more rapid growth (or less decline) of union coverage among women than among men. Other things being equal, this development contributes to narrowing the gender wage differential. Even and Macpherson (1993) estimate that the narrowing of the gender unionization differential between 1973 and 1988 accounts for approximately one-seventh of the narrowing of the male-female earnings gap that took place in the US during that period. Doiron and Riddell (1994) obtain similar results for Canada during the decade of the 1980s.

As this discussion makes clear, the impact of unions on the structure of relative wages depends on both which types of workers tend to be unionized and on how union relative wage impacts vary across different groups of workers. There are large research literatures on both questions, and these are reviewed elsewhere in this volume. Our focus in this chapter is on how these two phenomena combine to influence the structure of relative wages among workers who differ by gender, skill and other characteristics.

### *Unions and the Dispersion of Wages*

In addition to assessing the impact of unions on the relative wages of different groups, we are also interested in the effects of unions on economy-wide wage inequality. Indeed, there has been substantial recent debate about the sources of rising earnings inequality in several industrialized countries, and the extent to which changes in labour market institutions may have contributed to these developments.<sup>2</sup>

A useful starting point for discussing the impacts of unions on earnings inequality is a simple two-sector model. Let  $W_i^N$  be the log wage of individual  $i$  if employed in the non-union sector and let  $W_i^U$  be the log wage of the same individual when employed in the union sector. Assume that

$$W_i^N = W^N + e_i^N \quad (6)$$

$$W_i^U = W^U + e_i^U \quad (7)$$

where  $W^N$  and  $W^U$  denote the mean log wages in the non-union and union sectors, and  $e_i^N$  and  $e_i^U$  are random error terms with conditional means of zero, i.e.  $E(e_i^N | \text{non-union}) = 0$  and  $E(e_i^U | \text{union}) = 0$ . Finally, assume that in the absence of unionization, current union members would receive the same average wage as non-union workers: in other words, that in the absence of unions, the mean wages in both sectors would be the same. The observed union-nonunion differential in mean wages is

$$\Delta_w = W^U - W^N \quad (8)$$

Under the assumptions we have made, this is also the expected wage gain that a nonunion worker would receive if she could obtain a union job, and the expected wage loss that a union worker would suffer if he moved to the nonunion sector. The mean log wage of all workers is

$$W = (1-\alpha) W^N + \alpha W^U = W^N + \alpha \Delta_w \quad (9)$$

As before, the second term in equation (9), the product of the union coverage rate and the union relative wage effect, is the average wage gain associated with unionism.

In addition to affecting the mean level of wages, unions can potentially influence the intra-sectoral distribution of wages. Let  $\text{Var}(e_i^N) = V^N$  and  $\text{Var}(e_i^U) = V^U$  denote the variances of log wage outcomes for individuals in the nonunion and union sectors, respectively. The union-nonunion variance gap is denoted by

$$\Delta_v = V^U - V^N \quad (10)$$

The overall variance of log wages is given by<sup>3</sup>

$$V = V^N + \alpha \Delta_v + \alpha (1-\alpha) \Delta_w^2 \quad (11)$$

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

$$V - V^N = \alpha \Delta_v + \alpha (1-\alpha) \Delta_w^2 \quad (12)$$

The first term on the right hand side of this equation is a “within sector” effect associated the fact that wage dispersion is different in the union and non-union sectors. The sign of this effect depends on the sign of  $\Delta_v$ . The second term is a “between sector” effect, arising because unions insert a wedge between the average pay of union and nonunion workers that is always disequalizing.

Two features of union wage policy contribute to the within sector effect: standardization of pay within establishments and firms and standardization of wages across firms in a common product market. Standardization of pay within firms arises because unions replace wage setting based on managerial discretion with wage rates attached to a job or job classification rather than an individual. This characteristic of union wage policy was noted in many studies carried out by institutional labour economists in the 1940s and 1950s. For example, Slichter, Healy and Livernash (1960, p.602) state that “the influence of unions has clearly been one of minimizing and eliminating judgement-based differentials in pay for individuals employed on the same job” and of 'removing ability and performance judgements as a factor in individual pay for job performance'. Unions have also attempted to achieve standard wage rates across firms or establishments in a common product market -- to 'take wages out of competition'. As noted by Slichter, Healy and Livernash (1960, p. 606) 'wage standardization within an industry or local product market is the most widely heralded union wage policy'. These two features of union wage policy can be expected to reduce wage dispersion in the union sector relative to the nonunion sector.

While this simple model provides a useful starting point for discussing the impact of unions on the distribution of earnings, it does not incorporate differences in the extent of union coverage or the size of the union wage effect across different workers. To incorporate these factors it is useful to assume that workers can be classified into

homogeneous skill groups -- for example, categories based on detailed levels of education and labour market experience.<sup>4</sup> 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. As before, assume that

$$W_i^N(c) = W^N(c) + e_i^N \quad (13)$$

$$W_i^U(c) = W^U(c) + e_i^U \quad (14)$$

where  $W^N(c)$  and  $W^U(c)$  are the mean nonunion and union log wages for individuals in skill group  $c$ , respectively, and that the random terms satisfy the conditions

$$E(e_i^N) = E(e_i^U) = E(e_i^N | \text{nonunion}) = E(e_i^U | \text{union}) = 0 \quad (15)$$

The union-nonunion gap in average wages for workers in skill group  $c$  is

$$\Delta_w(c) = W^U(c) - W^N(c) \quad (16)$$

and the union-nonunion variance gap for skill category  $c$  is

$$\Delta_v(c) = V^U(c) - V^N(c) \quad (17)$$

where  $V^U(c) = \text{Var}(e_i^U | c)$  and  $V^N(c) = \text{Var}(e_i^N | c)$  denote the variances of log wage outcomes for individuals in skill group  $c$  in the union and nonunion sectors, respectively. Denoting the fraction of skill group  $c$  that is covered by union agreements as  $\alpha(c)$ , the mean log wage of all workers in skill category  $c$  is

$$W(c) = W^N(c) + \alpha(c)\Delta_w(c) \quad (18)$$

where the second term is the average wage gain associated with unionism for skill group  $c$ . The variance of log wage outcomes for workers in skill group  $c$  is

$$V(c) = V^N(c) + \alpha(c)\Delta_v(c) + \alpha(c)(1-\alpha(c))\Delta_w(c)^2 \quad (19)$$

As in a model with homogeneous workers, unions exert both a 'within-sector' effect and a 'between-sector' effect on the variance of wages among the subset of workers in skill group  $c$ .

Using equation (19), the variance of log wages across all skill groups can be written as

$$\begin{aligned} V &= \text{Var}[W(c)] + E[V(c)] \\ &= \text{Var}[W^N(c) + \alpha(c)\Delta_w(c)] + E[V^N(c) + \alpha(c)\Delta_v(c) + \alpha(c)(1-\alpha(c))\Delta_w(c)^2] \\ &= \text{Var}[W^N(c)] + \text{Var}[\alpha(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), \alpha(c)\Delta_w(c)] \\ &\quad + E[V^N(c)] + E[\alpha(c)\Delta_v(c)] + E[\alpha(c)(1-\alpha(c))\Delta_w(c)^2] \end{aligned} \quad (20)$$

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

$$V^N = \text{Var}[W^N(c)] + E[V^N(c)] \quad (21)$$

Thus the effect of unions on the variance of wage outcomes, relative to what would be observed if all workers were paid according to the wage structure in the nonunion sector, is<sup>5</sup>

$$\begin{aligned} V - V^N &= \text{Var}[\alpha(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), \alpha(c)\Delta_w(c)] \\ &\quad + E[\alpha(c)\Delta_v(c)] + E[\alpha(c)(1-\alpha(c))\Delta_w(c)^2] \end{aligned} \quad (22)$$

This expression can be compared to equation (12), the equivalent expression when union coverage rates, relative wage differentials, and variance gaps are the same for all skill groups. The final two terms in equation (22) are analogues of the 'within-sector' and 'between-sector' effects discussed previously; when there are many skill groups these are simply averaged across groups. The two additional terms in equation (22) reflect

variation in the union coverage rate  $\alpha(c)$  and/or the union wage effect  $\Delta_w(c)$  across skill groups. The first is a positive component that arises whenever the union wage gain  $\alpha(c)\Delta_w(c)$  varies across groups. The second is a covariance term that may be positive or negative, depending on how the union wage gain varies across the wage distribution. 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.

The magnitude -- indeed, even the direction -- of the effect of unions on wage dispersion may change over time as changes occur in union coverage rates or the relative wage differentials of particular groups. Later in this chapter we examine the changes that have taken place in the components of equation (22) in Canada, the UK and the US during recent decades.

### ***The Effect of Unobserved Skills***

Equation (22) has to be modified if the union and non-union workers in a given skill group have different productivity levels and would earn different wages even in the absence of unionization. Such a phenomenon will arise when workers have productivity-related characteristics that are known to employers but not observed by the researcher, and when the unobserved characteristics are correlated with union status. As before, assume that workers are classified into skill categories on the basis of observed characteristics, and suppose that

$$W_i^N(c) = W^N(c) + a_i + e_i^N \quad (23)$$

$$W_i^U(c) = W^U(c) + a_i + e_i^U \quad (24)$$

where  $a_i$  represents an unobserved skill component, and  $E(e_i^N | \text{nonunion}) = E(e_i^U | \text{union}) = 0$ . Note that  $a_i$  is assumed to shift wages by the same amount in the union and nonunion sectors. Let

$$\theta(c) = E[a_i | \text{union}, c] - E[a_i | \text{nonunion}, c] \quad (25)$$

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

$$E[W_i^U(c) | \text{union}] - E[W_i^N(c) | \text{nonunion}] = \Delta_w(c) + \theta(c) \quad (26)$$

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

$$\begin{aligned} V - V^N = & \text{Var}[\alpha(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), \alpha(c)\Delta_w(c)] \\ & + E[\alpha(c)\Delta_v(c)] + E[\alpha(c)(1-\alpha(c))\{(\theta(c)+\Delta_w(c))^2 - \theta(c)^2\}] \end{aligned} \quad (27)$$

Only the last term of this equation, which reflects the gap in mean wages between union and nonunion workers with the same observed skills in the presence and absence of unions, differs from equation (22), the expression that applies when  $\theta(c)=0$  for all groups.

The relatively simple form of equation (27) depends crucially on the assumption that unobserved skills are rewarded equally in the union and nonunion sectors. The formula needs to be extended if unobserved skills are rewarded differently in the union and non-union sectors.<sup>6</sup>

#### IV. A Review of the Literature on Unions and Inequality

Until recently, most economists believed that unions tended to raise inequality. For example, Friedman (1956) argued that -- principally on the basis of Marshall's laws of derived demand -- 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 identified in the two-sector model -- unions create a 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 union -- i.e., an assumption that the covariance term in equation (22) is positive.

Even economists more sympathetic to unions than Friedman echoed 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 (as of 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 questions was scanty and inconclusive until the widespread availability of microdata in the 1970s. Steiber (1959) examined the effect of unions in the steel industry and concluded that during the 1947-60 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 century, 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 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.'

Since that time a series of studies has substantially altered the prevailing view. Table 2 summarizes the first generation of post-1975 studies. The top panel presents studies based on aggregate data. Hyclak (1979) analysed the determinants of inequality in wage and salary income in urban labour markets and found that higher union coverage is associated with lower earnings inequality, at least for males as a group and also for black males. Hyclak (1980) found a negative relationship between the state mean of union density and the percentage of families with low earnings. These studies suggested that, controlling for other influences, earnings tend to be more equally distributed in more heavily unionized urban areas and states, at least for men. However, they provide no insight into the mechanisms that produce this negative relationship.

Hirsch (1982) carried out a cross-sectional study at the industry level using a model that allows for the joint determination of earnings, earnings dispersion and union coverage. He concluded that the equalizing effects of unions on earning inequality are larger in both manufacturing and non-manufacturing industries 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 UK (without controlling for the joint determination of earnings and union coverage) but concluded that union coverage widened the pay structure across industries. Metcalf also shows, however, that the coefficient of variation of weekly earnings is lower in the union than in the nonunion sector, and that unions narrow pay structure by occupation and race.

The studies in the lower panel of Table 2 are all based on individual micro data. These studies follow the important contribution by Freeman (1980), which first laid out the two-sector framework. Freeman also used establishment-level data to study the impact of unionism on the wage gap between blue-collar and white-collar workers in the organized sector. Since few white collar workers are unionized, this exercise extends 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. He also found that unions substantially narrow the wage differential between blue-collar and more highly paid white-collar employees within the organized sector. These two equalizing effects more than offset the 'between-sector' effect that runs in the other direction. In non-manufacturing industries, Freeman concluded that the net impact of unions was smaller, reflecting both a smaller 'within sector' effect and larger 'between sector' effect.

Meng (1990), using Canadian data, confirmed that wage dispersion is lower in the union sector than the non-union sector under “North American” collective bargaining institutions. Indeed, numerous studies have found that wage differences between different demographic and skill groups are lower, and often much lower, in the union sector than in the non-union sector.<sup>7</sup> 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 union for nonunion jobs and to rise when they move in the opposite direction. The impact 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) reaches the same conclusion that union reduces wage dispersion using more recent longitudinal data from the 1987-88 CPS. On the basis of his longitudinal estimates, he concludes that de-unionization accounts for 21 percent of the increase in the standard deviation of males wages in the US between 1978 and 1988. Using a more sophisticated econometric approach (see the discussion of Card (1996) below), Card (1992) also concludes that de-unionization explains around 20 percent of the increase in wage inequality during the 1980s.

Gosling and Machin (1995) reach a similar conclusion that de-unionization accounts for around 15 percent of the increase in male wage inequality among semi-skilled workers in Britain between 1980 and 1990. They use wage data at the establishment level from the 1980, 1984, and 1990 Workplace Industrial Relations Survey (WIRS). One drawback of the WIRS is that it only provides limited information on workers skills and on within-establishment wage dispersion.

### ***Second Generation Studies***

The studies summarized in Table 2 significantly altered views regarding the relationship between unionization and wage inequality, but they tell an incomplete story. On one hand, the first wave of post-1975 studies focused on male private sector workers.<sup>8</sup> On the other hand, these studies essentially ignored variation in the union coverage rate and the union wage effect across different types of workers.

The studies reported in Table 3 use variants of the framework underlying equation (22) to develop a more complete picture of the effect of unions. To set the stage for these studies, it is helpful to look directly at how the union wage gap and the extent of union coverage vary across the wage distribution. Figures 1-3 provide some simple evidence on the variation in the union wage gap for men and women in the US, Canada, and the UK in the early 1990s. These graphs plot mean wages for unionized

workers in a given skill group (defined by narrow age and education categories) against the corresponding means for non-union workers with the same skill level. Using our earlier notation, the figures plot  $W^U(c)$  against  $W^N(c)$  for skill groups based on age and education.<sup>9</sup>

Observe first that if union workers with a given level of age and education have the same average wages as non-union workers, then all the points in these graphs will lie on the 45-degree line. On the other hand, if the union wage gap  $\Delta_w(c)$  is positive, then the points will lie above the 45-degree line. Moreover, if  $\Delta_w(c)$  is larger for lower wage workers, then 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 US 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 UK, as shown in Figures 2a and 3a. For age-education 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 non-union 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 non-union wage. Thus, in all three countries  $\Delta_w(c)$  is larger for low-wage men than high-wage men, suggesting a potential role for unions to significantly reduce wage inequality. As we discuss in the next section, one caveat to this conclusion is that there may be important unobserved skill differences between union and non-union workers in different age-education groups that tend to exaggerate the apparent negative correlation between wages in the non-union sector and the union wage gap.

For women, the patterns of union wages relative to non-union wages are also remarkably similar in the three countries. Unlike the patterns for men, however, the union wage gaps for women are roughly constant. Thus, unions do not seem to “flatten” the wage differences between older and younger women, or between more and less educated women, relative to the non-union sector.

Although the data in Figures 1-3 pertain to the early 1990s, similar plots from other years suggest that the basic patterns have been very stable in all three countries over the past 20-30 years. Contrary to the predictions of Friedman (1956) and others, the

union-non-union 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.

Another key feature that determines the effect of unions on wage inequality is the variation in union coverage. Figures 4-6 show the fractions of union members among male and female workers in the US, Canada and the UK, by hourly wage. (We plot union membership rates in the US and UK although collective bargaining coverage patterns are generally similar). For the US, we show union densities in 1973-74, 1984, 1993, and 2001. For the UK, the earliest available data are for 1983: thus we show union densities in 1983, 1993, and 2001. Similarly, individual micro data with wages and union status are only available for Canada starting in the 1980s, so we have plotted union densities by wage level for 1984, 1993 (actually, an average of 1991 and 1995), and 2001.

Several important conclusions emerge from these figures. First, in all three countries union membership rates of men tend to be highest for workers near the middle or upper middle of the wage distribution, and lower at the bottom and top of the wage distribution. Despite the higher union wage premiums for men at lower skill levels, low rates of union membership among the least skilled men substantially moderate any potential redistribution effects of unions. Second, unionization rates of women in the US and Canada are not much lower for the highest-wage groups than for those in the middle of the wage distribution. Coupled with the fact that the union wage gaps are roughly constant across different wage groups, these patterns suggest that unions may actually widen wage inequality across skill groups for women. In the UK, there is more of a fall-off in union membership among the highest-paid women, suggesting that unionization may have a greater potential equalization effect for women there.

A third important feature of Figures 4-6 is the obvious decline in unionization rates over time. The declines are most evident for men in the US and UK, but there are also small declines among Canadian men, and among women in the three countries. We discuss the impacts of these changes in more detail later in this chapter.

With this background, we turn to a brief discussion of the studies in Table 3. The first, by DiNardo and Lemieux (1997), uses a reweighting technique to construct estimates of the sum of the terms in equation (22) for men in the US and Canada in 1981

and 1988. DiNardo and Lemieux also present a slightly different decomposition of the net contribution of unionization to the overall variance of wages in each country and year. They estimate that in 1981, the presence of unions reduced the variance of male wages by 6 percent in the US and 10 percent in Canada. The corresponding estimates in 1988 are 3 percent in the US and 13 percent in Canada. Thus, they estimate that changing unionization patterns contributed to the rise in US wage inequality in the 1980s, but worked in the opposite direction in Canada. Their decompositions also show 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, Fortin, and Lemieux (1996), examined both men and women in the US in 1979 and 1988. DiNardo, Fortin, and Lemieux (henceforth, DFL) use the reweighting technique applied by DiNardo and Lemieux. DFL do not report the effects of unions on the levels of wage inequality in either year, but instead focuses on explaining the rise in wage inequality over the 1979-88 period. For men, their methods suggest that shifts in unionization account 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 is very small. DFL also estimate 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 third study in Table 3, by Bell and Pitt (1998), uses DFL's method to analyse the impact of de-unionization on the growth in wage inequality in Britain. Their main analysis is based on Family Expenditure Survey Data (FES) that only contains a proxy for union status (whether there are deductions from pay for union dues). They also analyze data from the National Child Development Study (NCDS), the British Household Panel Survey (BHPS) and GHS that contain direct measures of union membership. Depending on the data source used, they find that between 10 and 25 percent of the increase in male wage inequality (measured by the standard deviation or the 90-10 gap in log wages) can be explained by de-unionization. Machin (1997) reaches similar conclusions using the 1983 GHS and 1991 BHPS data.

The next study in Table 3, by Card (2001), examined the contribution of unions to wage inequality among US men and women in 1973-74, and in 1993. Card reports estimates based on the simple two-sector formula (equation (12)), and on a variant of equation (22) obtained by dividing workers into 10 equally-sized skill groups, based on predicted wages in the non-union sector. Two key findings emerge from this analysis. First, the presence of unions is estimated to have reduced the variance of men's wages by about 12 percent in 1973-74 and 5 percent in 1993. Overall, shifts in unionization can explain about 15-20 percent of the rise in male wage inequality in the 1973-93 period. Second, although the within group variance of wages is lower for women in the union sector than the non-union 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 US 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 1973-74 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 imply 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 1973-74 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. .

The final study in Table 3, by Gosling and Lemieux (2001), examined the effects of unions (and other factors) on the rise in wage inequality in the US and the UK between 1983 and 1998, using the DFL reweighting method. Gosling and Lemieux do not report estimates of the cross-sectional effects of unionization. However, their estimates suggest

that in both the US and the UK, unions have a much smaller equalizing effect on female wage inequality than male inequality. They estimate that shifts in union coverage among men in the UK can explain up to one-third of the rise in wage inequality there between 1983 and 1998, while in the US the decline in unions can explain up to 40 percent of the rise in inequality. Consistent with findings in DFL and Card (2001) they conclude 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 (12) or (22) is that union workers may be more or less productive than otherwise similar non-union workers. In this case, comparisons of the mean and variance of wages for union and non-union 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 non-union 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 is 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 view 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 non-union 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).

The three studies in Table 4 all 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 non-union workers. The studies by Lemieux (1993) and Card (1996) measure the wage outcomes of job changers who move between the union and non-union 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 non-union sectors. Lemieux (1998) adopts a more general approach which allows the union sector to flatten the returns to unobserved ability relative to the nonunion sector.

Lemieux studies men and women in Canada in the late 1980s, and reports separate estimates of the effect of unions for three different observed skill categories (high, medium, and low) in the public and private sectors separately, and in the overall economy. For men, his results show that unionized workers from the lowest skill group are positively selected (i.e., they have higher unobserved skills than non-union 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 US men in the late 1980s.<sup>10</sup> An implication of this pattern is

that the between group “flattening effect” of unions documented in Figures 1a and 2a 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 examines the changes in the variance of wages, and concludes 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 concludes that the presence of unions lower the variance of male wages in Canada in the late 1980s by about 15 percent. A similar calculation for US 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.

As one might suspect given the patterns in Figures 2 and 5, Lemieux’s findings for women in Canada are much different than those for men. In particular, neither the cross-sectional nor longitudinal estimates of the union wage gaps show a systematic flattening effect of unions. Coupled with the fact that union coverage is lower for less-skilled women, these results imply 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 on net Lemieux’s results imply that unions raised the dispersion of Canadian women.

Lemieux (1998) presents 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 concludes 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 imply that unionization reduced the variance of wages among Canadian men by about 17 percent -- not far off the estimate in his 1993 study.

## V. Unions and Wage Inequality in the U.S., U.K. and Canada: An Update

### *Data Sources and Background*

In this section we update and extend the existing literature on the effect of unions on wage inequality for the U.S., U.K. and Canada using data up to 2001. There are several important motives underlying our new analysis. First, as shown in Table 1, unionization rates have declined steeply in the U.S. and U.K. over the past two decades. It is interesting to check whether this decline has resulted in a more modest effect of unions on wage inequality in the early 2000s. On a related point, several studies mentioned in Section IV have shown that de-unionization contributed to the steep increase in wage inequality in the U.S. and U.K. in the 1980s. Wage inequality did not change much, however during the 1990s in the U.S. (Card and DiNardo, 2002). This leads to the natural question of whether the evolution of the impact of unionization on wage inequality can account for some of this slowdown in the growth in wage inequality.

Finally, it is now possible to use large and very comparable micro data sets to look at the impact of unionization on wages in the U.S., U.K., and Canada. All three countries conduct large scale monthly labor force surveys to measure the unemployment rate and other related statistics in a timely fashion. In the U.S., the Current Population Survey (CPS) has been asking question about wages and union status on an annual basis since 1973, and on a monthly basis (for sample members who are rotating out of the sample – the so-called “outgoing rotation group” or ORG) since 1983.

Similar questions were added to the U.K.’s Labour Force Survey (UKLFS) in 1993, and to the Canadian Labour Force Survey (CLFS) in 1997. Since 1997, it is therefore possible to compare much more accurately the extent of wage inequality and the effect of unions on wage inequality in the three countries. Estimates of the role of unionization in cross-country differences in wage inequality are no longer affected by survey differences, or by the limitations of small sample sizes.

In our empirical work we nonetheless want to provide a perspective that is as broad as possible on the role of unions in wage inequality over the last two or three decades for both men and women in the three countries. Our most recent data point is 2001 for which comparable data are available in all three countries.<sup>11</sup> We then go back to 1993, which is the earliest year for which comparable CPS and UKLFS data are

available. For Canada we cover a similar point in time by combining two relatively small Surveys on Work Arrangements (SWA) that were conducted as supplements to the November CLFS in 1991 and 1995. These surveys both ask questions about wages and collective bargaining coverage that are comparable to the questions in the latest CLFS.

In the U.K., the only large survey of individuals that contains information on both wages and unionization prior to 1993 is the 1983 General Household Survey (GHS).<sup>12</sup> A large scale survey on union membership and wages (Survey of Union Membership, SUM) was also conducted for 1984 in Canada as a supplement to the CLFS. We use these two surveys, along with data from the OGR supplements of the 1984 U.S. CPS, as reference points for the early 1980s. Finally, for the U.S. only, we augment the data from the early 1980s, early 1990s, and 2001 with data from the May 1973 and 1974 CPS. From 1973 to 1981, the May CPS supplements asked the same question about wages and union membership that were later used in the ORG supplements. One difference, however, is that unlike the ORG supplement, the May supplement does not ask about union coverage. For the sake of consistency, we look at the effect of union membership on wages in the United States.

Although data on both union membership and “union presence” are available in our U.K. data sources, we focus on the impact of union membership on wages in Britain, because the available measures of “union presence” do not appear to be a satisfactory measure of actual coverage.<sup>13</sup> In the 2001 UKLFS, for example, a high fraction of workers who are not union members and report that a union is “present in their workplace” nevertheless report that their wages and working conditions are not determined by a collective agreement.

Canada is in an opposite situation since the 1991 and 1995 SWA asked only one question about union membership or coverage, instead of asking the questions separately as was done in the 2001 CLFS and the 1984 SUM. For consistency reason we therefore use union coverage as our measure of unionization in Canada. This choice has little effect on the results since only about two percent of wage and salary workers who are not union members are covered by a collective agreement.<sup>14</sup>

To arrive at the final estimation samples, we process the various data sets in the same way as in Card (2001) for the US, Gosling and Lemieux (2001) for the UK, and

DiNardo and Lemieux (1997) for Canada. 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 an 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.<sup>15</sup> Sample weights are used throughout except in the 1983 GHS for which sample weight are not available.

To implement the techniques developed in Section III, we divide workers in each sample into skill groups, based on age and educational attainment. The number of skill groups used varies across data sets, however, depending on the sample size and the ways that age and education are coded in public use files. For example, in Canada age is only reported in 10-year categories in the 1984 SUM and the 1991-95 SWA (a total of 5 categories for workers age 15 to 64) and education can only be consistently coded into 5 categories through time. Thus we only use 25 skill groups for Canada. We use the same number of skill groups for the U.K. (five age and five education groups) to have a reasonable number of observations (in the 100-200 observation range) in each cell. We are able to use a much larger number of cells in the U.S. because of much larger sample sizes and finely coded age and education categories. We have re-analysed the U.S. data using about the same number of skill groups as in Canada and the U.K. and found that this has little impact on our results.

### ***Results***

Some of the main patterns in our data have been noted already in the discussion of Figures 1-6. Tables 5, 6, and 7 summarize a variety of facts about unionization and the structure of wages for the U.S., Canada, and the U.K., respectively. Starting with the U.S., the first row of Table 5 confirms the steep decline in unionization rates documented in Table 1. As illustrated in Figures 4a and 4b, 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 points, from 14 to 12 percent, while it fell much more for men, 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) shows 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 shift of unionization from the private to the public sector was relatively beneficial to women in terms of their unionization rates.

The trends in unionization in Canada between 1984 and 2001 (Table 6) are remarkably 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). These findings are in sharp contrast with the OECD data reported in Table 1, which shows a very modest decline in unionization rates in Canada between 1980 and 1994. They are consistent, however, with a recent paper by Riddell and Riddell (2001) that uses micro data for the 1984-98 period.<sup>16</sup>

Our results for the U.K. are more consistent with the OECD data in Table 1, and reveal a very rapid decline in the unionization rate. Between 1983 and 2001, the unionization rate fell by 27 percentage points (from 57 to 30 percent) for men and by 14 percentage points for women. As in the U.S. and Canada, the differential trends in the male and female unionization rates is closely linked to the relative shift of unionization from the private to the public sector (Gosling and Lemieux, 2001). These changes are compounded by the fact that privatizations moved a significant number of male workers from the unionized public sector (i.e., the former nationalized industries) to the much less organized private sector (Gosling and Lemieux, 2001). By 2001, the male and female unionization rates were more or less equal in all three countries. This unprecedented situation marks a major departure from the historical pattern of greater unionization among men.

The next set of rows in Tables 5 to 7 shows the evolution of both the raw union wage gap and the wage gap adjusted for differences in the relative distribution of characteristics (or skills) in the union and nonunion sectors. In terms of the notation of Section III, the adjusted and unadjusted union wage gaps represent estimates of  $\Delta_w$  and

$E[\Delta_w(c)]$ , respectively. As in the case for the unionization rates, the estimated wage gaps show a remarkably similar pattern across the three countries. In all three countries the unadjusted wage gap is 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.

Like the unadjusted union wage gap, the adjusted wage gap is typically larger for women than for men. The male-female difference in the adjusted gaps is less pronounced than the gender gap in the unadjusted gap. This is consistent with Figures 4 to 6 that 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 workforce reduces the union wage gap more for women than for men. The larger adjusted wage gaps for women and the relative concentration of female union members at the high end of the wage distribution mean that the disequalizing effect of unions on between-group inequality is larger for women than men in all three countries.

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 average impact of unions on wages has declined dramatically over the last two decades. For example, the adjusted union wage gain for UK males went from 9.2 percentage points in 1983 (unionization rate of .57 times the adjusted gap of .162) to 1.7 percentage points in 2001 (.307 times .045). This also means that any disequalizing effect of unions on between-group inequality declined sharply during this period.

The next rows in Tables 5-7 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.<sup>17</sup> For example, Figure 7 displays the wage distribution for US 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 union than nonunion sector. By contrast, the inter-sectoral differences in wage dispersion are much less striking for US 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, it is not clear whether wages are more narrowly distributed in the union or nonunion sector.

An important confounding factor is the minimum wage, which has a strong visual impact in the lower part of the female distribution in the nonunion sector but little impact on (higher wage) union workers.<sup>18</sup> An interesting conjecture is that unions appear to have a more limited effect on the within-sector variance of wages for women than men, in part because minimum wages exert a much stronger equalizing effect on lower-wage non-union women than on higher-wage unionized women.

The wage densities for Canadian men (Figure 9) and women (Figure 10) are qualitatively similar to those in the US. In particular, it is clear that male wages are more narrowly distributed in the union 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 Canada and the US, it is more difficult to see union wage compression effects for UK males (Figure 11) or females (Figure 12). The union-non-union gaps in the standard deviations in Table 7 are nonetheless quite similar to those in Canada or the United States.

The lower parts of Table 5 to 7 shows the various elements of the variance decompositions discussed in Section III. Recall that in the simple two-sector model of equation (12), the effect of unions on the variance of wages is the sum of the within-sector effect  $\alpha \Delta_v$ , and the between sector effect,  $\alpha (1-\alpha) \Delta_w^2$ . Once again, the results are remarkably consistent across countries and time periods. For men, unions always reduce wage dispersion since the within-sector effect always dominates the between-sector effect. Relative to the overall variance, the compression effect ranges from 31 percent in the UK in 1984, when the unionization rate was 57 percent, to 6 percent in the US in

2001 (unionization rate of 15 percent). More generally, the compression effect of unions tends to be positively correlated with the extent of unionization, which is consistent with equation (12).<sup>19</sup>

Relative to men, the within-sector effect for women is smaller for the reasons mentioned earlier. On one hand, since the unionization rate  $\alpha$  is lower for women than for men, unions reduce wage inequality for a smaller fraction of the workforce. Second, the gap between the variances (or standard deviations) in the union and nonunion sectors ( $\Delta_v$ ) is much smaller for women than men. Both elements of the within-sector effect  $\alpha \Delta_v$  are thus lower (in absolute value) for women than men. By contrast, the union wage gap is systematically larger for women than men. This yields a larger between-sector effect  $\alpha (1-\alpha) \Delta_w^2$  that in later years of our analysis dominates the equalizing within-sector effect. Consistent with Card (2001) and Lemieux (1993), unions thus tend to increase the variance of wages among women.

The final set of rows in Tables 5 to 7 show the elements of the variance decomposition when we distinguish between skill groups using the framework of equation (22). Starting with men, controlling for characteristics systematically reduces the magnitude of both the within- and between-sector effect. It is easy to see why this happens in the case of the between-sector effect. As shown previously, adjusting for characteristics reduces the union wage gap and thus 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 the 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 characteristics are controlled for. Recall from Figures 4 to 6 that union workers are more concentrated in the middle and upper part of the wage distribution than nonunion workers. This suggests that union workers have more homogeneously distributed skills than their nonunion counterparts. 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 characteristics also reduces the magnitude of the between-sector effect for women but increases (or leaves unchanged in the US) the

magnitude of the within-group effect. The latter finding means that union women are not more homogenous (in terms of their skills) than their nonunion counterparts, which is consistent with the evidence reported in Figures 4 to 6. Once worker characteristics are taken into account, the within-group effect tends to dominate the between-group effect for both men and women. This suggests that the large male-female differences in the measured effect of unions on wage dispersion from a simple two-sector decomposition are overstated by ignoring differences in the distribution of skill characteristics in the union and nonunion sectors.

Recall from equation (22) that the effect of unions on the variance of wages also depends on the variance and covariance terms  $\text{Var}[\alpha(c)\Delta_w(c)] + 2\text{Cov}[W^N(c), \alpha(c)\Delta_w(c)]$ . Those two terms indicate how unionization changes the distribution of average wages across the different skill groups. As highlighted in our discussion of Figures 1-3, the wage gap  $\Delta_w(c)$  is systematically lower for high-wage men, inducing a negative covariance between  $W^N(c)$  and  $\alpha(c)\Delta_w(c)$ . By contrast, the wage gap for women is not typically lower for high-wage groups, and the higher unionization rate for those groups induces a positive covariance between  $W^N(c)$  and  $\alpha(c)\Delta_w(c)$ .

The results in Tables 5 to 7 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 UK. In the US however, unions have little effect on female wage dispersion across skill groups from 1973 to 1993, and actually reduce wage dispersion in 2001. A natural explanation for the difference between the US on one hand, and Canada and the UK, on the other, is that the union wage gap for US women tends to decline slightly with higher nonunion wages (Figure 1b). This lowers the covariance between  $W^N(c)$  and  $\alpha(c)\Delta_w(c)$  for US 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 effect for women tends to be small and positive (more inequality). This pattern of result is quite similar to what we found with the two-sector model, though the magnitude of the effects tend to be smaller when we control for workers characteristics.

### *Unions and Differences in the Trends in Wage Inequality*

To what extent union wage effects explain the evolution of wage inequality over time and the differences in inequality across countries? In light of the results of Table 5 to 7, we look at this question for men only since unions appears to have little effect on wage inequality for women.<sup>20</sup> Starting with the US, Table 5 shows that the variance of male wages increases from 0.258 to 0.340 (change of 0.082) between 1973/74 and 2001. During the same period, the effect of unions on the variance of wages computed using the simple two sector model declines from  $-0.047$  to  $-0.021$  (change of 0.26). 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 smaller (14 percent) when estimates of wage compression effects that control for characteristics are used instead.

The results for the UK are qualitatively similar. Between 9 percent (model with workers characteristics) and 29 percent (two-sector model) of the 0.087 growth in the variance of wages between 1983 and 2001 can be accounted by the decline in union compression effects. Furthermore, in both the US and UK union wage compression effects remain relatively constant between 1993 and 2001. In particular, the effect computed in the model with workers' characteristics are essentially unchanged during this period. This is consistent with the pattern of change in the overall variance of wages which grew much more rapidly before than after 1993.

As in the US and UK, the union wage compression effect has been steadily declining for Canadian men since 1984. Unlike the US and UK, however, overall inequality remained very stable over time. This suggests that overall inequality would have actually declined if union wage compression effect had remained at their 1984 level. Clearly, union wage effects are not very useful in explaining the evolution of male wage inequality in Canada over the last two decades.

Turning to cross-country differences in wage inequality, first note that in 1983/84 the variance of wages was lowest in the UK (0.216) followed by Canada (0.231) and the US (0.289). By contrast, union wage compression effects (adjusted for differences in characteristics) were highest in the UK ( $-0.050$ ), followed by Canada ( $-0.037$ ) and the US

(-.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 UK-US difference in the variance of wages. By 2001, US-UK difference in the variance of wages is down to 0.037 (0.340-0.303), while the US-UK difference in the union compression effect is 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 that clearly cannot be account for by changes in the union wage compression effect.

## **VI. Conclusion**

What is the effect of unions on pay differentials and wage inequality? Until the late mid-1970s, the consensus among economists was that "...unionism probably has a slight disequalizing effect on the distribution of income." (Johnson, 1975). This prevailing view was substantially altered by the landmark paper by Freeman (1980). Subsequent studies that used different data and more sophisticated econometric methods essentially all confirmed Freeman's finding that, overall, unions tend to reduce wage inequality among men. Our new empirical work indeed indicates that this finding is very robust across countries (U.S., U.K. and Canada) and time periods (from the early 1970s to 2001).

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 union wage structure effects 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. 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 impact of unions on the wage structure in the U.S., Canada, and the U.K. has 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 smaller effects of unions on wage inequality in all three countries that help account for a significant fraction of the growth in wage inequality in the U.S. and U.K.

## Endnotes

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<sup>1</sup> See the chapter by Addison and Siebert in this volume for a review of recent changes in the collective bargaining framework in the UK.

<sup>2</sup> See, for example, the Symposium on Wage Inequality in the Journal of Economic Perspectives, vol. 11, no. 2, Spring 1997.

<sup>3</sup> The equation follows from the standard decomposition of a variance into within-sector and between-sector components.

<sup>4</sup> This presentation follows Card (1992), Lemieux (1993) and Card (2001). Although the exposition is in terms of skill groups, the same principles apply to any situation in which groups of workers can be ordered according to their average wage.

<sup>5</sup> Of course, as noted previously, unions may alter the wage structure in the nonunion sector.

<sup>6</sup> Lemieux (1998) presents a model in which unobserved attributes are rewarded differently in the union and nonunion sectors.

<sup>7</sup> See Lewis (1986) for a review of US studies and Simpson (1985) and Lemieux (1993) for Canadian evidence.

<sup>8</sup> The lone study that included data for women (Hyclak, 1979) found no significant relationship between female earnings inequality and union coverage in urban labour markets.

<sup>9</sup> The US plots are based on skill groups defined by years of education (10 categories) and 2-year intervals of age (25 categories). Due to data limitations, the Canadian and UK plots are based on much broader age and education groups. The data underlying these figures are explained in more detail in the next section.

<sup>10</sup> Hirsch and Schumacher (1998) use data on test scores and find that union members with high measured skills have relatively low test scores.

<sup>11</sup> For the UK, we only use the LFS for the Fall semester since wage and unionization data are not available for other semesters. In Canada, we use the LFS data from November 2001 (all rotation groups have wage and unionization data) since the earlier data sets were collected in November too (December in 1984). Data for all months are used in the 1984, 1993 and 2001 CPS.

<sup>12</sup> Blanchflower (this volume) uses UKLFS data from 1985 to 1991 to estimate the union wage premium during this period. Unfortunately, the available samples (about 1,000 observations a year) are too small to conduct a detailed analysis of the impact of unionization on wage inequality for a large number of skill groups as we do here.

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<sup>13</sup> The UKLFS has asked a question about union coverage similar to the ones in the CLFS and ORG CPS since 1996 only. Both the 1993 UKLFS and the 1983 GHS ask about union membership and union presence in the workplace.

<sup>14</sup> More precisely, 2.4 percent of male workers and 1.9 of female workers are covered but not members of a union in the 2001 CLFS. The two different concepts of unionization also hold very similar union wage gaps and variance gaps. For example, when union membership is used the unadjusted wage gaps are .235 and .361 for men and women, respectively, compared to 0.236 and 0.358 (Table 6) when union coverage is used instead.

<sup>15</sup> The cutoffs points are \$2 and \$90 (1989 dollars) for the US, \$2.5 and \$44 (2001 dollars) for Canada, and £1.50 and £50.00 (2001 pounds) for the UK. Note also that we exclude Northern Ireland from the UK samples because union membership data is not available from Northern Ireland in the 1993 UKLFS.

<sup>16</sup> Note that between 1984 and 1991-95, the unionization rate rates drops by 3-4 percentage points for men and women taken together (6 points drop for men but only 1 for women). By contrast, Table 1 indicates a 1 or 2 percentage point decline, depending on the measure being used. This discrepancy can probably be explained by differences in data sources. Since Table 1 only covers the 1980-94 period, it misses the continuing decline in the unionization rate throughout the 1990s and, thus, understates the extent of the recent decline in the unionization.

<sup>17</sup> The density are estimated using a bandwidth of 0.05. See DiNardo, Fortin and Lemieux (1996) for more detail.

<sup>18</sup> This is similar to DiNardo, Fortin and Lemieux (1996) who show that the minimum wage has a much larger impact on women than on men.

<sup>19</sup> The derivative of the compression effect with respect to the unionization rate is  $\Delta_v + (1 - 2\alpha) \Delta_w^2$ . It is negative (higher negative effect on the variance when the unionization rate increases) as long as the within-group effect ( $\Delta_v$ ) dominates the between-group effect ( $(1 - 2\alpha) \Delta_w^2$ ).

<sup>20</sup> DiNardo, Fortin and Lemieux (1996), Card (2001) and Gosling and Lemieux (2001) all reach the same conclusion that de-unionization explains very little of the increase in wage inequality in the US or UK.

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Table 1  
 Union density and collective agreement coverage in selected OECD countries,  
 1980 and 1994

	<u>Union membership as a percentage of paid workers</u>	1980	1994	1980	1994
				<u>Collective agreement coverage as a percentage of paid workers</u>	
Canada		36	34	37	36
United Kingdom		50	34	70	47
United States		22	16	26	18
<u>Other Countries</u>					
Australia		48	35	88	80
Austria		56	42	98	98
Belgium		56	54	90	90
Denmark		76	76	69	69
Finland		70	81	95	95
France		18	9	85	95
Germany		36	29	91	92
Italy		49	39	85	82
Japan		31	24	28	21
Netherlands		35	26	76	81
New Zealand		56	30	67	31
Norway		57	58	75	74
Portugal		61	32	70	71
Spain		9	19	76	78
Sweden		80	91	86	89
Switzerland		31	27	53	50

Source: Organisation for Economic Co-operation and Development, Employment Outlook, Paris: OECD, July 1997.

Table 2  
Studies of the Impact of Unionization on Wage Inequality

(a) Studies with aggregate data			
Study	Country	Nature of data	Findings
Hyclak (1979)	US	1970 Census data on wage and salary income by SMSA	A 10% increase in union density reduces Gini coefficient by .58% for men and .64% for black men. No significant relationship for women or black women.
Hyclak (1980)	US	1950, 1960, 1970 Census data on family income by state	Each 1% increase in state union density reduces the fraction of families earning below \$3,000 in 1970 by 3%. Similar results for 1950 and 1960.
Hirsch (1982)	US	1970 Census data by 3-digit industries	Each percentage point of union coverage lowered the variance of log earnings by .015 points
Plotnick (1982)	US	Time series of CPS data for men	Each percentage point of unionization lowered variance of log earnings by .065 points.
Metcalfe (1982)	UK	1972, 1973, 1978 New Earnings Survey Data	Union coverage narrows the pay structure by occupation and race but increases it by industry.
(b) Studies using individual data			
Study	Country	Nature of data	Findings

Freeman (1980)	US	CPS data on male wage and salary workers in private sector, May 1973, 1974, 1975	Union workers have 15% lower standard deviation of log earnings than comparable nonunion workers.
		Firm data on expenditures for employee compensation	Unionism reduces blue-collar/ white-collar differential by 10%.
Freeman (1982)	US	BLS Industry Wage Survey data on individuals in 9 industries	The above effects produce a 2-3% reduction in inequality among comparable workers. The net reduction in wage dispersion is greater in manufacturing than non-manufacturing industries
Freeman (1984)	US	Longitudinal data before and after unionization: May 1974-75 CPS NLS of Men 1970-78 PSID 1970-79 Quality of Employment Survey 1973-77	Standard deviation of log wages in union sector is on average 22% lower than in nonunion sector.
Freeman (1993)	US	CPS longitudinal Matched data on Men, 1987-88	Workers who join unions experience declines in standard deviation of earnings of .03 to .09 points relative to workers who leave unions
Meng (1990)	Canada	Labour Market Activity Survey data on adult men in full-time employment in 1986	20 percent of the increase in the standard deviation of male log earnings between 1978 and 1988 is attributable to de-unionization
Gosling-Machin (1995)	UK	1980, 1984, 1990 Workplace Industrial Relations Survey	Standard deviation of hourly wages in union sector is 25% lower than in nonunion sector
			For semi-skilled workers, the decline in unionization explains around 15 percent of the growth in the variance of log wages between 1980 and 1990

Table 3  
 “Second Generation” studies of the impact of unions on the wage structure

Study	Country	Nature of data	Findings
DiNardo-Lemieux (1997)	US and Canada	1981 and 1988 Men only. CPS data (US) and LFS data (Canada)	In 1981 unions reduced the variance of wages by 6 percent in the US and 10 percent in Canada. In 1988 unions reduced the variance of wages by 3 percent in the US and 13 percent in Canada
			Wage dispersion grew faster in the US relative to Canada for age/education groups with bigger relative decline in unionization
DiNardo, Fortin Lemieux (1996)	US	1979-1988 CPS data men, 1987-88	Shifts in unionization explain 15-20 percent of rising wage inequality for men, 3 percent for women. Shifts in unionization explain up to one-half of the rise in the wage gap between high school graduate men and dropouts
Bell and Pitt (1998)	UK	1982-93 FES Supplemented with NCDS, GHS and BHPS	Approximately 20 percent of the increase in the standard deviation of log male wages during the 1980s has been due to declining union density
Card (2001)	US	1973/74 and 1993 CPS data	Unionization rates fell for less educated men and women but were stable (men) or rising (women) for college graduates. Union densities rose in the public sector. Shifts in unionization explain 10-15 percent of the rise in male wage inequality, none of the rise for women

Relative shifts in unionization explain one-half or more of the greater rise in male inequality in the private sector.

Gosling- Lemieux (2001)	US and UK	1983 and 1998 CPS data (US). GHS and LFS data (UK)	<p>Unionization fell faster in the UK than the US.</p> <p>Shifts in unionization can explain up to one-third of the rise in male inequality in the UK and up to 40 percent of the rise in male inequality in the US.</p> <p>Shifts in unionization explain very little of the rise in wage inequality for women in US or UK.</p>
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Table 4  
Longitudinal estimates of the effect of unions on the wage structure

Study	Country	Nature of data	Findings
Lemieux (1993)	Canada	LMAS longitudinal data 1986-87 on men and women in private and public sectors	Unionized men with low observed skills tend to have higher unobserved skills than their nonunion counterparts, whereas unionized men with high observed skills tend to have lower unobserved skills than their nonunion counterparts. This pattern also holds for women in the private sector. Unionized women in the public sector have higher unobserved skills than their nonunion counterparts at all skill levels.
			Unions reduce within sector variance for both men and women.
			Unions reduce the overall variance of male wages by 14.5% but increase the variance of female wages by 4.1%.
Card (1996)	US	CPS longitudinal matched data on men, 1987-88	Unionized men with low observed skills tend to have higher unobserved skills than their nonunion counterparts, whereas unionized men with high observed skills tend to have lower unobserved skills than their nonunion counterparts.
			Unions reduced the overall variance of male earnings in 1987 by 7%.

Lemieux (1998)

Canada

LMAS longitudinal data, 1986-87 on men.

Unions lower the returns to permanent unobserved skill characteristics.

Variance of wages around the expected mean, accounting for observed and permanent unobserved skill characteristics is lower in the union sector

Unions reduced the overall variance of wages by 17 Percent.

Table 5: Effect of Unions on Wage Structure of US Workers, 1973-2001

	1973/74		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

Notes: 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 6: Effects of Unions on Wage Structure of Canadian Workers, 1984-2001

	1983		1993		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 7: 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.