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Unions and Wage Inequality: The Roles of Gender, Skill and Public Sector Employment — Source link

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ABSTRACT

Unions and Wage Inequality: The Roles of Gender, Skill and Public Sector Employment*

We examine the changing relationship between unionization and wage inequality in Canada and the United States. Our study is motivated by profound recent changes in the composition of the unionized workforce. Historically, union jobs were concentrated among low-skilled men in private sector industries. With the steady decline in private sector unionization and rising influence in the public sector, unionization is now five times higher in the public than the private sector in both countries. Though the public sector represents only 15-20% of employment, half of unionized workers are in the public sector. Accompanying these changes was a remarkable rise in the share of women among unionized workers. Currently, approximately half of unionized employees in North America are women. While early studies of unions and inequality focused on males, recent studies examine both and reveal striking gender differences. A consistent - and puzzling - finding is that unions reduce wage inequality among men but not among women. In both countries we find striking differences between the private and public sectors in the effects of unionization on wage inequality. These differences have become more pronounced over time. At present, unions reduce economy-wide wage inequality by less than 10% in both countries. However, union impacts on wage inequality are much larger in the public sector. Once we disaggregate by sector the effects of unions on male and female wage inequality no longer differ. The key differences in union impacts are between the public and private sectors - not between males and females.

JEL Classification: J31, J45, J51

Keywords: wage inequality, wage structure, unions, collective bargaining, Canada, United States, public sector

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1. Introduction

The relationship between unions and wage inequality continues to attract research and policy interest as analysts struggle to understand the relative importance of market-based and institutional forces in explaining the rise in income inequality throughout the industrialized world. A central issue is whether trends in inequality can be rationalized by technological change and globalization (Acemoglu and Autor, 2011; Helpman, 2016), or whether labour and product market institutions such as minimum wages, collective bargaining and product market deregulation have played some independent role (Fortin and Lemieux, 1997; Card and DiNardo, 2002; Salverda and Checchi, 2015). Numerous studies have concluded that the decline in unions *within* specific countries has contributed to growing inequality.[1] Other research shows that differences in the extent of unionization contribute to cross-country differences in the *level* of wage inequality (e.g. Lemieux, 1993; DiNardo and Lemieux, 1997), and that differences in the rate of de-unionization are correlated with differences in the *growth* of inequality (Card, Lemieux and Riddell, 2004; Gosling and Lemieux, 2004). This evidence has led some pundits (e.g. Stiglitz, 2012) to argue that labour law reforms should be part of any policy response to rising inequality and secular declines in labour's share of national income.

This paper examines the changing nature of the relationship between unionization and wage inequality in Canada and the United States over the past four to five decades. Our study is motivated by profound changes in the composition of the unionized workforce during this period, and the implications of these changes for the wage structure. Historically, union jobs in both countries were largely held by unskilled and semi-skilled men working in private sector industries such as manufacturing, transportation, construction, forestry and mining. With the steady decline in private sector unionization and rising union influence in the public sector, however, union coverage rates are now ***5 times higher*** in the public sector than the private sector in

[1] We present a brief overview of this work in Section II of this paper.

both the U.S. and Canada.[2] As a consequence, though the public sector contains only 15-20% of all jobs in the two economies, approximately half of the unionized workers in both countries are in the public sector. These sectoral changes have been accompanied by a remarkable rise in the share of women in the unionized workforce. Currently, 47% percent of all unionized employees in the U.S., and 53% of those in Canada, are women. A typical union worker today is more likely to be a female teacher or nurse with a university degree than a male factory worker with only a high school education.

A key lesson of earlier research is the importance of accounting for heterogeneity in the rate of union membership and the size of the union effect on wages. The effect of unions on overall wage inequality depends critically on which skill groups are most likely to be represented by unions, and on the extent to which unions raise wages for more versus less highly paid groups (see Card 1992, 1996; Lemieux, 1993). Estimates of this effect when these differentials are taken into account are typically lower than in the simple two sector homogeneous worker model originated by Freeman (1980) in his seminal study of unions and inequality.

While many early studies of unions and inequality focused on male workers, more recent studies examine men and women and reveal striking gender differences. A consistent finding in Canada, the U.S. and the U.K. is that unions tend to reduce wage inequality among men but not among women (e.g., DiNardo, Fortin and Lemieux, 1996; Card, 2001; Card, Lemieux and Riddell, 2004; Gosling and Lemieux, 2004). In this paper we build on this approach and allow for heterogeneity by skill group and gender, but extend the analysis to account for differences across the public and private sectors. Disaggregating by sector also allows us to explore the extent to which differences in union impacts between men and women found in earlier studies are attributable to

[2] See Appendix Table 1 for summary statistics on unionization and public/private sector employment. We use the definitions of the public sector employed by the U.S. Bureau of Labor Statistics and Statistics Canada. In both countries public sector employment includes those who work for a federal, state/provincial or municipal government, for a government service or agency or a government owned public establishment such as a school, college, university or hospital. In Canada this also includes employees of Crown corporations.

public-private differences in the coverage patterns and wage effects, rather than to a gender-specific effect.

We focus on the U.S. and Canada for several reasons. The two countries share a common legal framework that results in a sharp distinction between union and non-union workplaces. In this setting workers have the right to form and join unions, and if a majority of employees agree their chosen union will be certified as the exclusive representative of all workers in the bargaining unit, whether they join the union or not.[3] This framework creates a highly decentralized form of collective bargaining in which most non-managerial employees in a given enterprise are either covered by a union contract, and pay membership dues to the union, or have their wages set by the employer with no influence of unions. In contrast, in Australia and many European countries collective agreements between unions and employer associations are often extended to all workers in a sector, creating large gaps between the fraction of workers whose wages are set by collective bargaining and the fraction of union members (Visser, 2015) and a fuzzy boundary between the union and non-union sectors.[4] In addition, in both Canada and the U.S. the non-union sector is large, yielding a good approximation to the wage structure that would prevail in the absence of unions.

Canada and the U.S. also provide contrasting experiences that are relevant for our study. Although their economies and labour markets have many common features, the level of inequality is lower in Canada than the U.S. during our sample period, and also rose more slowly. Unionization rates in the two countries followed similar trends in the immediate post-war period, and were approximately equal in the early 1960s. Since then, however, unionization rates have diverged – rising in Canada until the early 1980s, followed by a gradual decline to 30% in 2016, but falling steadily in the U.S. to a rate of

[3] Key features of this legal framework were introduced with the passage of the U.S. Wagner Act in 1935. During World War II the Canadian federal government passed the National War Labour Order, Order-in-Council P.C. 1003, that incorporated Wagner Act-type provisions relating to union recognition and formation together with long-standing Canadian laws designed to prevent and settle disputes such as compulsory conciliation and "cooling off" periods. After the war ended, jurisdiction over labour matters reverted back to the provinces, and most provinces passed similar legislation.

[4] The extreme example is France where 8 percent of workers in 2010 were union members but 98 percent were covered by collective agreements (Visser, 2015).

only 12% in 2016. These differences in the timing of union growth and decline and in levels and movements in wage inequality provide an opportunity to further assess the contribution of institutional change to trends in income inequality.

An additional objective of this paper is to provide estimates of the impact unions exert on the current wage structure. Previous studies conclude that the decline of unionization has made a significant contribution to rising inequality, but what is the impact of unions on wage inequality at the present time? Because labour law reform – if ever enacted -- would likely influence the extent of union coverage at the margin, estimates of the consequences of such reforms should reflect the dramatic changes that have already taken place in the structure of the unionized workforce.

The next section of the paper reviews the economics literature on unions and wage inequality. We then describe our empirical strategy. Section four describes our data sources and provides our results. The final section concludes.

2. Unions and Wage Inequality: An Overview of the Previous Literature

The impacts of unions on the wage structure and the distribution of earnings have long intrigued social scientists. This section briefly reviews this large literature, focusing on Canada and the U.S.[5]

Prior to the 1980s, most economists agreed that by creating an average pay gap between the union and non-union sectors, unions tended to *increase* overall inequality (e.g., Johnson, 1975).[6] This view was challenged by Freeman (1980), who argued that union wage setting tends to lower wage dispersion among more and less skilled workers, and between higher and lower-paying establishments, leading to a "within sector" inequality effect that may offset the "between sector" effect arising from the average union pay gap. Freeman's (1980) analysis showed that the within-sector effect was large and negative, particularly in the manufacturing sector. For other sectors he concluded that the net impact of unions was smaller, reflecting both a smaller within-sector effect and

[5] See Card, Lemieux and Riddell (2004) for a more detailed review up to the early 2000s.
[6] An interesting exception to this general view was Ashenfelter's (1972) study showing that unions tended to lower the black-white pay gap.

larger between-sector effect. Subsequent work (e.g., Freeman, 1982 for the U.S.; and Meng, 1990 for Canada) confirmed that wages tend to be less disperse in the union sector, and that the net effect of unions was to lower wage inequality relative to the level in the non-union sector. Freeman (1993) used this framework to study the impact of falling union coverage rates on the trend in male wage inequality, concluding that de-unionization could explain 20-25% of the rise in wage inequality over the 1980s.

Freeman's (1980) approach was generalized by DiNardo, Fortin and Lemieux (1996) and DiNardo and Lemieux (1997), who introduced a reweighting technique that takes into account potential differences in the union-nonunion pay gap for workers with different observed skill characteristics, and differences in the probability of union coverage for different skill groups. Using this method DiNardo and Lemieux (1997) estimated that in the early 1980s the presence of unions reduced the variance of male wages by 6 percent in the U.S. and 10 percent in Canada. By the late 1980s, their estimates showed that the equalizing effect of unions had fallen to 3 percent in the U.S., but had risen to 13 percent in Canada, contributing to the divergence in pay inequality in the two countries.

DiNardo, Fortin and Lemieux (1996) (hereinafter, DFL) examined wage distributions for both male and female workers in the U.S. in 1979 and 1988. For men, their estimates suggest that shifts in unionization accounted for 10-15 percent of the overall rise in wage dispersion in the 1980s. For women, on the other hand, the estimated contribution of changing unionization is very small.[7]

Card (2001) uses a related skill-grouping approach to compare the effects of unions on trends in male and female wage inequality in the U.S. public and private sectors. His findings suggest that unions had similar equalizing effects on male workers in the two sectors and similarly small effects on wage dispersion for female workers in both sectors. Interestingly, by the end of Card's sample period (1993) there is some evidence that

[7] Gosling and Lemieux (2004) use the DFL method to compare the effects of unions on trends in wage inequality in the U.S. and the U.K. between 1983 and 1998. Their estimates confirmed that in both countries unions have a much smaller equalizing effect on females than males. They estimate that declining unionization can explain up to 33% of the rise in male wage inequality in the U.K. between 1983 and 1998, and up to 40% of the corresponding rise in the U.S., but very little of the rise in female wage inequality.

public sector unions were starting to have a modest dampening effect on wage inequality for female workers in the public sector – a result we re-examine below.

Finally, Card, Lemieux and Riddell (2004) use a DFL style technique to analyse the link between unions and wage inequality in Canada, the U.K. and the U.S. They show that unions have an equalizing effect on male wage dispersion in all three countries, and that declines in male union coverage can explain a modest share of the rises in male pay inequality between the early 1970s and the early 2000s in the U.S., and between the early 1980s and the early 2000's in the U.K. and Canada. As in earlier studies they find that the effects of unions on female wage inequality are substantially smaller than for men, a result that arises from the tendency for unionized jobs held by women to be relatively high-paying, and from the fact that union-nonunion wage gaps do not seem to be any larger for lower-skilled than higher-skilled women.

A longstanding concern with simple comparisons based on the structure of wages for union and nonunion workers is that unionized workers may have different *unobserved* characteristics than nonunion workers with the same observed skills (e.g., Lewis, 1986). Depending on the selection process determining union status, this could lead to an overstatement of the equalizing effect of unions. One way of addressing this concern is to use longitudinal data to follow workers as they move between sectors. Freeman (1984, 1993) showed that wage dispersion falls when workers enter the union sector and rise when they leave, confirming that unions have an equalizing effect even on the same worker. Nevertheless, his estimates suggested that this effect is smaller than would be inferred from a simple cross-sectional approach.

A problem with longitudinal comparisons is that union status can be mis-measured, leading to a high rate of falsely measured transitions between sectors. Using estimates of this mismeasurement rate to correct the observed longitudinal data Card (1992) concluded that unions still have a significant equalizing effect on male wages, and that the fall in unionization from the early 1970s to the late 1980s explained around 20 percent of the increase in U.S. male wage inequality, not far off Freeman's earlier estimate.

Lemieux (1993) used the 1986-87 Labour Market Activity Survey to study the effects of union sector changes in Canada. His results showed that unionized workers from the lowest skill group are positively selected, whereas those in the upper skill groups are negatively selected – a pattern similar to the one found in the U.S. by Card (1992). An implication of this pattern is that the between-group "flattening effect" of unions apparent in the raw data is somewhat exaggerated, although there is still evidence that unions raise wages of low-skilled men more than those of high-skilled men. Overall, Lemieux concluded that the presence of unions lowers the variance of male wages in Canada in the late 1980s by about 15 percent.

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

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

To summarize, we believe the existing evidence points to four main conclusions.

1. Unions tend to reduce wage dispersion for male workers. As a result, declining unionization has contributed to rising wage inequality among men since the 1970s or early 1980s.

2. Unions have little impact -- or even a small disequalizing effect -- on female wage inequality. An unresolved puzzle is whether this is a pure gender effect or due to other factors such as differences between male-dominated and female-dominated types of

employment.

3. There is some evidence of non-random selection into union and nonunion jobs. Existing evidence suggests that this selectivity leads to a relatively small over-estimate of the equalizing effect of unions on male workers.

4. Relatively little is known about the impacts of unions on the wage structure in the public sector, and how this differs from the private sector.

3. Estimating the Effects of Unions on Wage Inequality

In this section we present a simple framework for studying the effect of unions on wage inequality. We begin with a conceptual overview that clarifies the relationship between what we can actually measure with available data and the potential channels through which the presence of unions alters the degree of wage inequality in the overall labour market. We then discuss our measurement approach, which is based on the reweighting methods developed by DFL and DiNardo and Lemieux (1997), and develop a simple within- and between-skill group decomposition that helps illustrate the differences in the effect of unions on males versus females, and between the public and private sectors.

Building on the potential outcomes approach that is now widely used in the program evaluation literature (e.g., Angrist and Krueger, 1999) we assume that a given worker could earn a wage in either the union or nonunion sectors. We depart slightly from the standard approach, however, and allow these two potential wages to depend on the overall level of unionization in the economy.[8] Specifically, let

$W_i^U(U)$ and $W_i^N(U)$ represent the potential wages that worker *i* could earn at a union or non-union job, respectively, assuming that a fraction *U* of all jobs are unionized. In general, these two potential wages could depend not only on the average level of unionization but on which specific jobs were incorporated in the union sector. Since our interest is in comparing the economy with the current observed level of unionism to a counterfactual with *U=0*, however, we abstract from this concern.

[8] The idea that wages of a given worker in the two sectors could potentially depend on the economy-wide level of unionization was first formalized in Lewis (1963).

Let $V^U(U)$ and $V^N(U)$ denote the variances of potential union and nonunion wages, respectively, given a level of unionization, and let $V(U)$ denote the variance of wages observed in the economy given U. With this notation, the "effect" of unionization on the overall dispersion of wages, given the present level of unionization U, is $V(U) - V^N(0)$. Notice that this effect can be written as the sum of two terms:

$$V(U) - V^N(0) = [\, V(U) - V^N(U) \,] + [\, V^N(U) - V^N(0) \,] \quad (1)$$

The first, $V(U) - V^N(U)$, represents the difference in the current variance of wages and the variance that would prevail if we held constant the current level of unionization but all workers were paid their nonunion wage, $W_i^N(U)$. The second term is a "general equilibrium" effect, representing the difference in dispersion of non-union wages that would prevail with the current level of unionization and in the counterfactual world of $U=0$.

As noted by DFL and DiNardo and Lemieux (1997), although we do not directly observe $V^N(U)$, under strong assumptions we can form an estimate by extrapolating from the non-union wages we do observe (at the current level of unionization) being paid to workers in the nonunion sector. Assuming this problem can be solved we can estimate the first term in equation (1).

The second term is more elusive. Researchers since Lewis (1963) have argued that two main factors determine its size: (i) labour supply spillovers – if unions raise wages, doing so may reduce the number of jobs in the union sector, increasing labour supply elsewhere and lowering wages in the nonunion sector; and (ii) the threat effect – some nonunion employers may raise wages to reduce the threat of unionization. Interestingly, these two effects operate in opposite directions, and may offset each other to some extent. We note also that the magnitude of the general equilibrium effect $V^N(U)-V^N(0)$ depends on the size of U. Given currently low levels of unionization in the North American private sector, we suspect that the magnitude of $V^N(U)-V^N(0)$ is small for that sector, though it may have been larger in the past.[9] For the public sector, U is

[9] Fortin, Lemieux, and Lloyd (2018) find that general equilibrium effects of unionization reduce wage dispersion in the non-union sector, indicating that threat effects are more important than spillover effects in the U.S. labour market. This suggests that our estimates of union effects, $V(U)-V^N(U)$, understate the

higher but one might argue that spillover effects are small in the public sector (because of inelastic demand for labour) and that threat effects are also small (because the extent of unionization is determined largely by the political process, rather than by the preferences of workers to support or not support unionization). Thus, in the rest of this analysis we focus on comparisons between $V(U)$, the observed variance of wages, and $V^N(U)$, the variance that would prevail if everyone were paid according to the *current* nonunion wage structure.

Estimating the Variance of Potential Nonunion Wages.

In order to estimate $V^N(U)$, which from now on we denote as V^N, we have to make an assumption about how current *union* workers would be paid if they worked in the nonunion sector. One starting point is the assumption that union status is "as good as randomly assigned," conditional on observed skill characteristics, gender and sector of employment. In this case, the variance V^N can be estimated as the variance of wages for a suitably reweighted sample of nonunion workers. In this section we show how the resulting calculations are related to three key factors: the variation in the union coverage rate by wage level in the absence of unions, the size of the union wage effect for different skill groups, and the union-nonunion difference in the variance of wages within skill categories.[10]

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

$W_i^N(c) = W^N(c) + e_i^N$ and

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

where $W^N(c)$ and $W^U(c)$ are the mean nonunion and union log wages for individuals in

effect of unions on wage dispersion (in absolute terms) since $V^N(U)-V^N(0)<0$. Fortin, Lemieux, and Lloyd (2018) also find that, for U.S. men, general equilibrium effects of unionization $V^N(U)-V^N(0)$ are about three times as small in 2017 as they were in 1979, and that general equilibrium effects are substantially smaller for women than men.

[10] Card, Lemieux and Riddell (2004) discuss how the assumption that union status is independent of unobserved productivity factors can be relaxed.

skill group c, respectively, and the deviations e_i^N and e_i^U are independent of actual union status (conditional on gender, sector of employment and the observed skill level c). Let $V^U(c)$ and $V^N(c)$ denote the variances of potential wage outcomes for individuals in skill group c in the union and nonunion sectors, respectively, where implicitly we are now holding constant the overall level of unionization in the workforce at its current level. The union-nonunion gap in average wages for workers in skill group c is

$$D_W(c) = W^U(c) - W^N(c),$$

while the corresponding variance gap is

$$D_V(c) = V^U(c) - V^N(c)$$

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

Under the conditional independence assumption, the reweighted variance V^N provides an unbiased estimate of the variance of log wages in the absence of unions. The variance of log wages across all groups is the sum of the variance of mean wages across groups and the average variance within groups

$$V = Var(W(c)) + E[V(c)] \tag{2}$$

where $W(c)$ is the mean wage in cell *c* and $V(c)$ is the within cell variance. Similarly, the counterfactual variance V^N is the sum of "between group" and "within group" components

$$V^N = Var(W^N(c)) + E[V^N(c)] \tag{3}$$

where $W^N(c)$ is the mean nonunion wage in cell *c*, $V^N(c)$ is the within cell variance of nonunion wages, and cells are weighted by the fractions of all workers in each cell. Thus the impact of unions on the variance of log wages can be written as the sum of "between group" and "within group" components

$$V - V^N = [\, Var(W(c)) - Var(W^N(c)) \,] + [\, E[V(c)] - E[V^N(c)] \,] \qquad (4)$$

Card, Lemieux and Riddell (2004) further discuss how the between and within group components depend on the wage gap, variance gap, and unionization rate at the cell level. Since $W(c) = W^N(c) + U(c)D_W(c)$, the magnitude of the between group effect depends on how the effect of unions on average wages, or union wage gain $U(c)D_W(c)$, is distributed across the different skill groups. For example, unions will reduce the between group component if the wage gain is larger for lower than higher skill workers.[11] This can either happen because unions "flatten" the wage structure ($D_W(c)$ declines with the skill level) or because unionization is relatively concentrated among lower skill workers.

In the next section we present evidence on the pattern of union coverage and the magnitude and of the flattening effect. We then report estimates of the total effect of unions on the variance of log wages, including the contributions of the between and within group components.

4. Data and Estimation Results

Data Sources. We use two micro data files – the U.S. Current Population Survey (CPS) and the Canadian Labour Force Survey (LFS) together with supplements to these surveys to study the effects of unions on the wage structure since the early 1970s in the U.S. and the early 1980s in Canada. When suitably weighted these very similar household surveys provide representative samples of the adult population in the two countries. The CPS has been

[11] Since $Var[W(c)] = Var[W^N(c) + U(c)D_W(c)] = Var[W^N(c)] + 2Cov[W^N(c),U(c)D_W(c)] + Var[U(c)D_W(c)]$, the union effect on the between group component is $Var(W(c)) - Var(W^N(c)) = 2Cov[W^N(c),U(c)D_W(c)] + Var[U(c)D_W(c)]$. As the variance of the union wage gain is positive, unions can only reduce the between group component if the covariance between the non-union wage and the union wage gain is negative enough to offset the variance term. As Tables 1 and 2 indicate, the covariance term is negative enough in most specifications to yield a negative 'between group' component. While the covariance term can be readily computed (as in Card, Lemieux, and Riddell, 2004), it is not reported in the tables to simplify the exposition.

collecting data on wages and union status annually since 1973. We use the pooled May 1973 and May 1974 CPS samples as our first U.S. observation. For later years, we use the monthly earnings supplements data (the "outgoing rotation group" files) for 1984, 1993 and 2015.

The Canadian LFS added questions on wages and union status in 1997. Prior to that time several surveys carried out as supplements to the regular LFS included questions that provide this information. We use the 1984 Survey of Union Membership as a source of information for the early 1980s and combine two smaller surveys – the 1991 and 1995 Surveys on Work Arrangements (SWA) – as a source of information for the early 1990s. For the most recent period we use the 2015 LFS. In both the CPS and LFS data the earnings and union status information pertain to an individual's main job as of the survey week.

A key variable for our analysis is the measurement of union status. The 1984 and later CPS files as well as LFS data since 1997 include questions on both union membership and union coverage. The 1973 and 1974 May CPS surveys, however, only ask about union membership. For comparability over time, we therefore focus on union membership as our measure of union status in the U.S. In the case of Canada, we use union coverage as our measure of unionization because consistent information on union membership cannot be recovered from the 1991 and 1995 SWA's. We believe that this choice has little effect on the results, since only about two percent of Canadian employees are covered by collective agreements but are not union members.[12]

Another difference between the Canadian and U.S. data is that public sector affiliation is not measured in the same way in the two countries. In the CPS workers are asked directly whether they work for the government. [13] In the LFS workers are simply asked "who they work for", with options including the "name of business, government

[12] For example, in 2015 union membership was 28.6% versus coverage of 30.6%. The two different measures of union density result in nearly identical estimates of the union wage premium in a conventional regression of log wages on union status, education and experience.

[13] The CPS question used to determine the class of worker is: "Were you employed by government, by a private company, a nonprofit organization, or were you self –employed (or working in the family business)?"

department or agency, or person" and "what kind of business, industry or service this is." Statistics Canada then uses this information to code workers as working in the private or public sector.[14] Despite these measurement differences, Appendix Table 1 indicates that, as expected, a slightly higher fraction of Canadian workers (20%) are classified as working in the public sector than their U.S. counterparts (15.4%). This difference likely reflects true differences in the scope of the public sector, as opposed to the way public sector employment is measured in the two countries. [15]

In the data appendix we explain in detail how we process the various data sets to arrive at our final estimation samples. Generally speaking, our samples include only wage and salary workers age 16 to 64 (15 to 64 in Canada) with non-allocated wages and earnings (except in 1984 and 2015 in Canada). We use hourly wages for workers who are paid by the hour and compute average hourly earnings for other workers by dividing weekly earnings by weekly hours (or earnings for a longer time period divided by the corresponding measure of hours). We also exclude workers with very low or very high hourly wage values. Sample weights are used throughout.

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

[14] Individuals reporting working for a local, provincial or federal government, for a government service or agency, a crown corporation, or an organization that receives the majority of its funding from government such as a school, college, university or hospital are coded as working in the public sector.

[15] The most important difference in the scope of the public sector in the two countries is the health care sector which is overwhelmingly public in Canada, but not in the U.S. This is confirmed by Appendix Table 2 that shows that 9.7% of public sector employees are health care professionals in Canada, compared to only 0.8% in the private sector. By contrast, only 4.7% of public sector employees are health care professionals in the U.S., compared to 6.4% in the private sector. Unlike health care, the education sector is mostly public in both countries, and Appendix Table 2 shows that teachers and other education-related occupations represent a much larger share of the workforce in the public than the private sector in both countries.

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

Patterns of Union Coverage. Summary measures of union coverage are shown in the first row of Tables 1 and 2. Tables 1(a) and 2(a) report results for the private and public sectors pooled for U.S. and Canada respectively, while Tables 1(b) and 2(b) contain results separately by sector of employment.

Table 1(a) shows the steady and dramatic decline in union density among men in the U.S. over the past four decades -- from 31% in 1973/74 to 12% in 2015. In contrast, little change in coverage took place among U.S. women – unionization was stable at 13%-14% during this extended period of time. For the U.S. workforce as a whole, male and female union density rates are now approximately equal. Perhaps the most striking feature is the dramatic divergence between the private and public sectors that is evident for the U.S. in Table 1(b). In the private sector the male unionization rate fell from 31% in 1973/74 to 9% in 2015, while it rose in the public sector from 29% to 43%, with most of the increase occurring between 1973/74 and 1984. Similarly, union coverage among U.S. private sector women, already low at 13% in 1973/74, declined to 6% in 2015, while the rate in the public sector rose even more than that for men – from 18% to 41%.

The changes over time in Canada between 1984 and 2015 exhibit similar patterns although coverage levels are very different. For the labour force as a whole, male union coverage declined by 17 percentage points (from 46% to 29%) between 1984 and 2015 – an even larger decrease than in the U.S. over this time period. The decline for women was much smaller although again was greater than in the U.S. – from 37% to 32%. Private sector unionization fell by almost one-half for both groups – from 37% to 19% for males and 25% to 13% for females. In contrast, rates of union coverage in the public sector were stable for both men and women at 75-80%.

The divergence in union coverage rates between the public and private sectors in both Canada and the U.S. raises questions about the comparability of the various control groups of non-union workers used to compute the potential (or counterfactual) variance

V^N. For example, since over 90% of U.S. private sector workers are in the non-union sector in 2015, the wage distribution among these workers is likely very close to the distribution that would prevail if all private sector workers were non-union. By contrast, less than 25% of Canadian public sector workers are in the non-union sector in 2015, which raises potential challenges attempting to extrapolate what the wage distribution would look like if all public sector workers were non-union.

Although we partly address these issues by controlling for observed skills (education and experience) when computing the potential variance, there may still be differences in other observed or unobserved characteristics over different groups of workers. Appendix Table 2 helps alleviate some of these concerns by showing that the pattern of unionization across occupations is surprisingly similar in the public and private sectors in both countries. In particular, the unionization rate in management occupations is always lower than in other occupations in the same sector, while the opposite is true among trade and other blue-collar occupations. Looking across all occupations, the correlation coefficient between the unionization rate in the private and public sector is 0.63 in Canada, and 0.73 in the U.S.[16]

The framework developed in Section 3 suggests that the effect of unions on wage inequality depends in part on how union coverage varies by skill level. Figures 1 and 2 show the unionization rates of men and women in the private and public sectors in the U.S. and Canada, by the level of real hourly wages. These graphs are constructed by calculating union membership/coverage rates for workers in narrow wage bins and smoothing across bins.[17] The wage values reported on the x-axis range from the 1st to the 99th percentile of the distribution of wages for U.S. males in 2015 (expressed in 2001 dollars). Vertical lines identify the 10th, 50th and 90th percentiles of the wage distribution. As the wage distribution is approximately log normal, the mid-point on the x-axis is close to the median, while the $6.69 and $34.81 markers are close to the 10th and 90th wage

[16] Also, in many instances the public and private sectors hire from the same pool of potential workers, so the private sector nonunion wage structure exerts some influence on its public sector counterpart.
[17] The densities are estimated using a bandwidth of 0.05. See DiNardo et. al (1996) for more detail.

percentiles, respectively.[18]

Several noteworthy features of the differences between Canada and the U.S. and among the four groups within each country, as well as trends over time, are evident. One prominent development is that declines in coverage in the private sector have been largest in the middle of the wage distribution for both men and women, resulting in the gradual disappearance over time of the “hump- shaped” distribution of union coverage in both countries. In the early part of our sample period, unionization rates of men employed in the private sector tended to be low at the bottom and top of the wage distribution and to peak for workers near the middle or upper middle of the distribution. However, by 2015 this hump-shaped pattern had largely disappeared in the U.S. and substantially disappeared in Canada, yielding a distribution of union coverage that is much more uniform across the wage distribution. To the extent that a peak still exists it has also shifted to the right, higher up the wage distribution. A similar, but less dramatic, change has also taken place among women in the private sector. Also noteworthy are the decreases in coverage among private sector women at the very top of the distribution in both countries.

In contrast, the shape of the distribution has changed much less over time in the public sector. Unionization rates of public sector women in Canada are low at the bottom of the wage distribution (although there have been increases over time at the very bottom), rise to a peak around the middle, and remain almost as high for workers at the top of the distribution as for those in the middle. This pattern is driven in part by relatively high rates of unionization for teachers, nurses, and other highly skilled public sector workers, who are near the top of the female wage distribution. The U.S. pattern is similar. Among public sector male employees there is a slight hump-shape to the coverage distribution, but the decline in unionization at higher wage levels is modest in both countries and has become less pronounced over time in the U.S. due to further increases in the upper part

[18] The 10th, 50th, and 90th percentiles of the 2015 wage distribution for U.S. males are $6.9, $14.2, and $35.2, respectively (2001 dollars). The wage distribution is shifted to the left for U.S. women, and to the right (in Canadian dollars) for Canadian men and women.

of the earnings distribution. In Canada public sector unionization has been very stable over time in the upper half of the male distribution. The same pattern holds for women except for recent declines in coverage at the very top. In the lower half of the distribution decreases in coverage just below the peak have been offset by increases at the very bottom, a pattern evident for men and women.

Union Effects on Mean Wages. The top section in Tables 1 and 2 shows union wage effects for men and women in the two countries, both for the labour force as a whole (Tables 1a and 2a) and separately for the private and public sectors (Tables 1b and 2b). The unadjusted wage gap refers to the mean difference between union and nonunion log wages in the raw data, while the adjusted gap controls for compositional differences in the distribution of skills between the union and nonunion sectors. The adjusted gap is computed by averaging (over the distribution of union workers) the cell-specific wage gap $D_W(c)$. The composition effect reported in the tables represents the difference between the adjusted and unadjusted gap.[19]

Looking first at the estimates for the pooled samples, the unadjusted wage gaps are larger for women than for men in both countries. The adjusted gaps are generally smaller than the unadjusted gaps, and in the U.S. the difference between the two has increased over time, reflecting the decline in union coverage in the middle and lower end of the wage distribution relative to the top, as was seen earlier in Figure 1. The increase over time in the skill composition adjustment is especially large for U.S. women. In Canada the composition effect adjustment is more stable over time, but larger for women than for men. Nonetheless, the adjusted wage gaps remain larger for Canadian females than for males in all sample years.

Over time, there is a large decline in the union wage impact in the U.S., especially for female employees (Table 1a). Almost all of this change took place between 1993 and

[19] The mean wage in the union sector is $\overline{W}^U = \Sigma\theta^U(c)W^U(c)$, where $\theta^U(c)$ is the fraction of union workers in cell c. Similarly, the mean non-union wage is $\overline{W}^N = \Sigma\theta^N(c)W^N(c)$. The unadjusted union wage gap can be rewritten as $\overline{W}^U - \overline{W}^N = \Sigma\theta^U(c)\big(W^U(c) - W^N(c)\big) + \Sigma\big(\theta^U(c) - \theta^N(c)\big)W^N(c)$. The first term in the equation is the adjusted union wage gap, while the second term is the composition effect.

2015; for example, the adjusted wage gap for men fell from 23% to 16% and that for women from 20% to 9% over this time period. Interestingly, the unadjusted wage gaps actually rose over the 1973/74 to 2015 period for both men and women (although they fell from 1993 to 2015).

One might be tempted to attribute this drop in the union wage impact to the declining importance of unions in the economy. However, Table 1(b) indicates that sectoral changes in the composition of the unionized workforce play an important role. Among men the adjusted union wage impact actually rises over time in the public sector and declines only modestly in the private sector; however, the magnitude of the union-nonunion wage gap is much lower in the public sector (12.5% versus 20.3% in 2015). Thus, a key part of the change over time is due to the growing importance of a sector in which unions have a smaller impact on wages. A similar, although less dramatic, compositional effect is occurring over time for women. These results demonstrate the importance of disaggregating by sector of employment for the interpretation of union impacts.

In Canada there is a steady but moderate decline in the union-nonunion wage gap (both unadjusted and adjusted) over the 1984 to 2015 period (Table 2(a)). Disaggregating by sector of employment again reveals noteworthy differences between the public and private sectors. Adjusted union-nonunion wage gaps are much larger in the private sector than in the public sector, especially in 1984 and 1993, but fall sharply between 1993 and 2015. This steep decline in the private sector wage gap is masked by the growing importance of the public sector, where wage gaps are smaller and more stable over time.

Perhaps surprisingly, the union-nonunion wage gap is negative for male public sector workers in 2015, and very small in 1984 and 1993. One likely explanation for this finding is that, as shown in Appendix Table 2b, public sector managers are much less likely to be unionized than other public sector workers.[20] Adding occupation controls

[20] In unionized organizations in the public sector those who rise up the hierarchy to become senior managers (e.g. Deans or above at universities, principals at schools and senior bureaucrats in the public service) typically must relinquish union representation.

(dummies for the occupation categories shown in Appendix Table 2b) to the wage regression indeed increases the union-nonunion wage gap from -0.023 to 0.036 in 2015, and to 0.054 in 1984 and 1993.

The combination of declining unionization and a fall in the adjusted wage differential implies that the average impact of unions on wages – the union wage gain $E[U(c)D_W(c)]$ – has declined substantially in both countries in recent decades. For example, the adjusted impact on male wages in the U.S. went from 6.3 percentage points in 1973/74 (unionization rate of 0.307 times an adjusted wage gap of 0.204) to 2.0 percentage points in 2015. For U.S. female workers the average impact went from 3.2 percentage points to 1.1 percentage points. In Canada the average impacts of unions on wages are much larger due to greater union coverage and adjusted wage gaps that are generally larger than their U.S. counterparts. For Canadian men the average union impacts fell from 10.4 percentage points in 1984 to 3.6 percentage points in 2015, while that for women went from 11.5 percentage points to 7.1 percentage points.

Wage Flattening Effects of Unions. As noted in section 3, the effect of unions on wage inequality depends in part on how the union wage gap $D_W(c)$ varies across the skill distribution. Figures 3a to 3c illustrate the evolution of this dimension of the union and nonunion wage structures by gender and sector of employment for Canada over the 1984 to 2015 period (similar results for the U.S. are available on request). These figures plot mean union wages in a given age – education cell (i.e. $W^U(c)$) by the mean nonunion wage for the same skill group ($W^N(c)$). If union and nonunion workers in the same skill group have the same average wage the points will lie on the 45-degree line (the solid line in the figures). If the mean union wage exceeds its nonunion counterpart, however, the points will lie above the 45-degree line. If, in addition, unions tend to raise the wages of lower skilled workers more than the wages of higher-skilled workers, then the scatter of points will be flatter than the 45 degree line – the so-called "flattening effect" (Lewis, 1986). This tendency of unions to compress skill differences is evident, for example, in the lower left panel of Figure 3a for male workers in the private sector. A similar tendency is evident for males and females in the public sector. For private sector female workers, however, the scatter is nearly parallel to the 45 degree line, suggesting little or no flattening effect.

Comparisons of the graphs for 1984 (Figure 3a), 1993 (Figure 3b) and 2015 (Figure 3c) reveal several interesting facts about the evolution of union wage differentials in Canada. Among private sector men, compression of the union wage structure is strongest in 1984, lessens over time as union coverage among this group declines, and is limited by 2015. Based on the discussion in Section 3, we would expect these changes to result in a substantially lower effect of unions on the between group component of the variance.

In the case of private sector females, there is very little compression by skill level of the union wage structure over the sample period. Consistent with Table 2, the primary change has been a decline over time in the union – nonunion wage gap, a gap that is relatively constant across the skill distribution. By 2015 the gap between union and nonunion women in the private sector is very small. The story is very different in the public sector where there is substantial wage compression in all periods for both men and women. Note that the union – nonunion wage gap is systematically negative for public sector men at the top of the skill distribution. As discussed above, this likely reflects the fact that higher-paid public sector managers are typically not covered by unions, even in organizations in which most employees are represented by unions.

To quantify the extent to which unions compress the wage structure between skill groups we estimate the following regression model:

$$D_W(c) = a + b[W^N(c) - \bar{W}^N] + e(c)$$

where $\bar{W}^N$ is the mean nonunion wage (over all cells *c*) and *e(c)* is a random error term. The parameter *a* estimates the union – nonunion wage gap for the skill group *c* whose average nonunion wage equals the overall average nonunion wage while the parameter *b* provides an estimate of the extent to which unions flatten the wage distribution across skill groups. If *b* equals zero there is no compression of the union wage distribution relative to the nonunion distribution; with flattening *b* < *0*, and the more negative the value of *b* the greater the amount of compression. The results of estimating this equation for the U.S. (approximately 150 skill groups) and Canada (25 skill groups) are shown in Tables 3a and 3b respectively. We also report estimates of the union- nonunion wage gap for low wage workers (*a – 0.5 b*) -- those earning 50 log points below the mean nonunion wage -- and for high wage workers, those earning 50 log points above the nonunion average wage (*a + 0.5 b*).

For male workers in the U.S. private sector we see a steady decline over the 1984 to 2015 period in the union – nonunion wage gap estimated at the mean nonunion wage. In contrast, the wage gap for men in the public sector is relatively stable over this time period. There is substantial evidence that unions flatten the male wage structure in both the private and public sectors, with the amount of compression consistently being greater in the public sector. The story is different for women in the U.S.; compression of the wage structure by skill is lower in the public sector than for men, and close to zero in the private sector after the 1970s. The female union – nonunion wage gap has also fallen dramatically since the early 1990s in both the public and private sectors.

Union impacts on the Canadian male wage structure are broadly similar to those for men in the U.S. The union – nonunion wage gap at the mean of the nonunion wage fell substantially – from 26.4% in 1984 to 7.7% in 2015. As in the U.S. the male wage gap in the public sector is much lower – and not significantly different from zero in 2015. However, unions substantially compress the male wage structure in both sectors, with the extent of flattening consistently greater in the public sector. One consequence of substantial compression combined with a small wage differential at the nonunion average wage is that high-skilled unionized men in the public sector earn less than their nonunion counterparts, as is illustrated in Figures 3a to 3c (see also the estimated implied gaps for high wage workers).[21] As in the U.S., compression of the female wage structure in Canada's private sector is minimal – and essentially zero in 1993 and 2015.

[21] As noted previously this apparent negative effect appears to be driven to an important extent by the fact that public sector managers in both unionized and nonunion organizations are not represented by unions. The propensity to become a senior manager could in turn reflect differences in unobserved skills between highly educated and more experienced public sector 'managers' and 'non-managerial' employees.

This contrasts dramatically with the public sector in which there is substantial flattening of female wages across the skill distribution throughout the sample period. However, as is the case in the U.S., the degree of compression of the female public sector wage structure is lower than that for men.

Effects of Unions on Wage Inequality. Our estimates of the impacts of unionization on wage inequality are reported in the two lower panels of Tables 1 and 2. We begin in the first panel by showing the variance of log wages for all workers, the difference between the overall variance and the variance of nonunion wages, and the 'composition effect' associated with reweighting – the difference between the variance of log wages for nonunion workers and the reweighted variance. The estimated effect of unions on wage inequality is then the sum of the overall variance minus the nonunion variance and the composition effect.

Pooled sample results are reported in Tables 1a and 2a. In the U.S. the overall variance of log wages has increased steadily and substantially for both men and women since the early 1970s, as has been observed in many studies of trends in wage inequality. For men the nonunion variance has risen less rapidly; as a result, the difference between the overall and nonunion variances has steadily increased (became less negative). As was found in previous studies, the impact of unionization on wage inequality among U.S. male workers is consistently negative. The composition effect has also fallen over time, and is essentially zero in 2015, reflecting the growing similarity between the skill distributions in the union and nonunion sectors, as illustrated in Figure 1, as well as the decline in unionization.[22] As expected, given the substantial decline in unionization among U.S. men, the impact of unions on male wage inequality fell sharply – from over 10 percent of the variance in the 1970s (-0.027 / 0.261) to 4 percent in 2015 (-0.014 / 0.399).

In the bottom panel we report the results of the decomposition into the between and within group components illustrated in equation (3). For U.S. males both effects

[22] As the unionization rate gets very low, the skill composition in the nonunion sector has to get increasingly close to the one for the whole workforce, limiting the scope of potential composition effects.

have the same sign and contribute to lower wage inequality in each year. The impact of unions on the variance of average wages across skill groups is generally the more important of the two -- contributing about 60% of the reduction in inequality. But the effect of unions on the average variance of wages within skill groups is also quantitatively important.

Among U.S. female workers the estimated impact of unions on wage dispersion is close to zero from the early 1970s to early 1990s. However, by 2015 unionization among women has a small equalizing effect, reducing the variance by 3.4 percent (-0.012 / 0.349). Most of this estimated impact is associated with differences in the skill composition of the union and nonunion female workforces in the U.S. Although the magnitude of this effect is small it is not much different from the comparable estimate for U.S. males in 2015. Both the between and within group effects contribute, but most of the reduction in inequality arises from the impact of unions on lowering the average variance within skill groups.

The trends for Canadian men are broadly similar to those in the U.S., but there are some noteworthy differences. The overall variance of male wages has been relatively stable, in contrast to the U.S. experience, while the variance of nonunion wages has declined sharply, resulting in an even larger decline in the actual – nonunion variance gap over the 1984 to 2015 period than for U.S. men.[23] As in the U.S., the composition effect has also fallen as the union and nonunion male skill distributions have become more similar. The result is a steady decline in the equalizing impact of unions on male wage inequality – from 20 percent of the variance in 1984 to 8 percent in 2015. The within group effect is the dominant factor accounting for the reduction in male wage inequality, especially in the 1990s and 2000s. Consistent with the large decline in the flattening effect illustrated in Table 3b, the between group effect becomes much less important over time, and is close to zero by 2015.

Among Canadian women, the total variance of log wages exceeds the nonunion variance -- in contrast to the case for men. This small difference is offset by a

[23] Note that there was a sharp increase in wage inequality in Canada in the early 1980s during the 1981-82 recession. Our sample period begins after this increase.

composition effect of similar magnitude working in the opposite direction. As a result, unions have a very small effect on female wage inequality, similar in size to that found in Card, Lemieux and Riddell (2004) for the 1984 to 2001 period. In 1984 and 2015 the overall effect is slightly inequality-increasing, whereas the opposite holds in the early 1990s, a period of sustained economic weakness in Canada's economy (Fortin, 1996). The decomposition in the bottom panel provides insights into this gender difference: among women, unionization increases the variance in average wages across skill groups and this inequality-increasing effect approximately offsets the reduction in inequality coming from the within-sector effect that contributes to lower wage inequality.

The findings from our analysis of the impacts of unions on wage inequality by gender and sector of employment are reported in Tables 1b and 2b. There are striking differences between the public and private sectors for men and women in both Canada and the U.S. These differences have generally become more pronounced over time.

In the U.S., the variance of log wages has increased over time among all four groups and the extent of the increase has been greater in the private sector where union coverage has fallen sharply. The variance of wages is also higher among males in both sectors. In the private sector, unions reduce male wage inequality in each sample year, but after the 1970s the effect is modest in size –- a reduction of 4.0% of the variance in 1984, declining to 1.7% in 2015. In contrast, in the U.S. public sector unionization reduces male wage inequality substantially, especially since the 1970s, with impacts ranging from 16% to 20% of the overall variance for that group. Both the between cell and within cell effects contribute, with the between sector effect being the more important of the two (about 2/3 of the total over the 1984 to 2015 period).

Among U.S. female employees, differences in the impacts of unions on the wage structure in the public and private sectors are equally striking. In the private sector unions have essentially no impact on female wage inequality, a feature that has not changed over the sample period. However, since 1984 unionization has reduced wage inequality among women in the public sector and the magnitude of this effect has grown over time – to 10.7% of the overall variance in 2015. Clearly, the evidence of a modest reduction in female wage inequality that appeared in the pooled sample in 2015 was driven by the public sector. Both the between and within sector effects contribute to this impact, with the within cell effect being the more important of the two.

In Canada the variance of log wages is more stable over time than in the U.S. for all four groups, with the variance of male wages being larger than that of female employees in the private sector, but about the same in the public sector. Unions reduce male wage inequality in both sectors, but the magnitude of the impact is much lower in the private sector. With the decline in union coverage among men in the private sector, the size of the union impact has also fallen – from 11.3% of the variance in 1984 to 4.8% in 2015. In most sample years the between cell and within cell effects contribute about equally to this reduction. There is also little evidence of union impacts on female wage inequality in the private sector – estimated effects are close to zero in each sample year. In contrast, the effects of unionization in the public sector are very large, and the magnitudes of the estimated impacts are similar for men and women. The size of the effect has fallen over time – from a reduction in the variance of over 70% in 1984 to almost 60% in 1993 to 50% in 2015. A large part of the decline comes from the between cell effect that has fallen by almost half between 1984 and 2015. This is again consistent with the large decline in the flattening effect summarized by the slope parameter *b* that went from -0.569 to -0.343 between 1984 and 2015 for men, and from -0.423 to -0.298 for women.

As we noted earlier, the unionization rates among men and women have converged over time due, in large part, to the large decline in unionization in the private sector that used to be concentrated among men. One interesting pattern that emerges from Tables 1b and 2b is that the effect of unions on inequality is also increasingly similar for men and women, especially in the U.S. For instance, by 2015 the effect of unions on inequality is negative, and generally comparable in magnitude, for men and women in the public and private sectors in both countries. In contrast, the equalizing effect of unions on inequality was generally smaller for women in 1984, and went in the "wrong" direction (inequality enhancing) in the private sector in both countries. Thus, one key finding of the paper is that the well documented difference in the direction of the equalizing effects of unions for men and women no longer prevails in recent data for both the U.S. and Canada, and the magnitudes of the gender differences are now quite small.

5. Conclusions

The relationship between unionization of the workforce and the distribution of labour income has long been of interest to economists and other social scientists. The nature of this relationship has received renewed attention recently as wage inequality has increased in many developed countries. In these countries, a salient development has been the decline of union influence and a natural question to investigate is whether the rise in inequality and the decline in unionization (as well as other changes in labour market institutions) are linked. Numerous studies reviewed previously in this paper conclude that the decline in union coverage in Canada and the U.S. -- the countries we focus on in this study -- did indeed contribute to increasing wage inequality, although this institutional change was not the only (or even the dominant) factor. Initially most research focused on men employed in the private sector. Subsequent studies examined both male and female employees and reached a puzzling conclusion – that unionization decreases male wage inequality but has no or even a positive effect on female wage inequality.

Another salient development has been the divergence of unionization in the private and public sectors. Over the past four or five decades, unionization has grown or remained stable in the public sector but declined substantially in the private sector. The gap in union coverage is now enormous – 39% in the U.S. public sector vs 7% in the private sector and 76% vs 17% in Canada (Appendix Table 1). As a consequence, almost one-half of unionized workers in the U.S. and close to 60% in Canada are employed in the public sector even though that sector accounts for only 15% (U.S.) to 20% (Canada) of total employment.

A central objective of this paper is to investigate the implications of this dramatic divergence for the impacts of unions on the wage structure. With the exception of Card's (2001) study of the U.S., previous studies have either focused on the private sector or pooled together the public and private sectors. A second objective of the paper is to assess whether distinguishing between the public and private sectors might shed light on the puzzling gender difference found in previous studies of the relationship between unionization and wage inequality.

In both countries we find that there are striking differences between the private and public sectors in the effects of unionization on male and female wage inequality. These differences have become more pronounced over time as private and private sector unionization have diverged. In 2015, the overall effects of unions on the economy-wide wage structure are modest in size – reductions in male wage inequality of 3.5% in the U.S. and 7.9% in Canada, and a reduction in female inequality of 3.4% in the U.S. and an increase in inequality of 0.4% in Canada. However, disaggregating by sector of employment yields striking differences: reductions in male wage inequality in the private sector of 1.7% in the U.S. and 4.8% in Canada versus reductions in male wage inequality in the respective public sectors of 16.2% and 48.5%. Similarly, our estimates imply that unionization reduces female wage inequality by 0.6% and 2.5% in the U.S. and Canadian private sectors respectively but 10.7% and 50% in those countries' public sectors. Note also that once we disaggregate by sector the effects of unions on male and female wage inequality no longer differ – union coverage reduces wage inequality among women and men to a similar extent in both sectors and in both countries. The key differences in the impacts of unions are between the public and private sectors – not between male and female employees.

In both Canada and the U.S. the impact of unions on wage inequality has fallen substantially in the private sector. Differences in wage levels between union and nonunion workers are also much smaller now than in the past. With the distributions of union and nonunion wages now very similar in the private sector, it seems unlikely that marginal changes in the extent of private sector unionization will have much impact on inequality.[24] With the decline in employment in traditional areas of union strength such as manufacturing and primary industries, major gains in collective bargaining coverage require making inroads in the service sector. This has proven very difficult in the Wagner Act framework with its emphasis on making collective representation decisions at the enterprise level. Although many service sector firms such as Starbucks, Walmart and retail banks are large employers, individual outlets have few employees and relatively high turnover, both obstacles to union organization. Achieving significant increases in unionization in the service sector probably requires moving outside the Wagner Act framework.

[24] Legree, Skuterud and Schirle (2016) examine this issue for Canada and reach a similar conclusion. They analyse the potential impact of more union-friendly laws on union density and, through that, on the degree of inequality in the wage distribution. Their estimates suggest that if all Canadian jurisdictions adopted the most union-friendly laws among those currently in place, union coverage would increase substantially among male college and university graduates employed in the public sector but would change little among women and less-educated men in either sector. Their simulations suggest that impacts of wage inequality would be small – a 2% reduction in the 90-10 differential for males and no change for females.

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DATA APPENDIX

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

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

Note, however, that weekly earnings are not measured consistently over time. From 1973 to 1993, this variable was collected by asking individuals directly about their earnings on a weekly basis. Since 1994, individuals have the option of reporting their usual earnings on the base period of their choice (weekly, bi-weekly, monthly, or annually). Weekly earnings are then obtained by normalizing the earnings reported by workers to a weekly basis. The available evidence does not suggest, however, that this change in the way earnings are collected had a significant impact on the distribution of wages (see Card and DiNardo (2002) and Gosling and Lemieux (2004) for more detail).

Another potential problem is that weekly earnings are top-coded at different values for different years throughout the sample period. Before 1988, weekly earnings were top-coded at \$999. The top-code was later increased to \$1,923 in 1988 and \$2,884 in 1998 where it has remained since then. For an individual working 40 hours a week, the weekly earnings top-code corresponds to an hourly wage ranging from \$57.0 in 1984 (\$2015) to \$133.3 in 1973 (\$2015). While very few workers are top-coded in 1973/74, in the case of men over 3% of workers are top-coded in 1984 and 2015. Following most of the literature (e.g. Lemieux (2006)), we adjust for this problem by multiplying the wages of top-coded workers by a factor of 1.4. We also trim a few outliers with wages below \$3.35 (\$2015) and above \$172 (\$2015). The wage deflator used is the Consumer Price Index (CPI-U). All the U.S. wage statistics reported herein are also weighted using the CPS earnings weights.

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

Finally, in the ORG supplements of the CPS, wages or earnings of workers who refuse to answer the wage/earnings questions were allocated using a "hot deck" procedure. We exclude observations with allocated wages and earnings for two reasons. First, wages and earnings were not allocated in the May 1973-1978 CPS. We thus need to exclude allocated observations from the 1984, 1993, and 2015 ORG supplement data to maintain a consistent sample over time. Second, union status is not one of the characteristics used to match observations with missing earnings to observations with non-missing earnings in the imputation procedure (hot deck) used by the U.S. Census Bureau. As a result, estimates of union wage effects obtained from a sample with allocation observations included can be severely biased downward (see Hirsch and Schumacher, 2004 for more details).

Canadian Data: As mentioned in the text, for Canada we use the 2015 Labour Force Survey (LFS), the 1991 and 1995 Surveys on Work Arrangements (SWA), and the 1984 Survey of Union Membership (SUM). These data sets are all relatively comparable since both the SUM and the SWA were conducted as supplements to the Labour Force Survey. Relative to the U.S. data however, there are some important limitations in the Canadian data. First, as mentioned in the text, it is not possible to distinguish union membership from union coverage in the SWA. For the sake of consistency over time, we thus use union coverage as our measure of unionization in Canada.

A second limitation is that in the 1984 SUM and the 2015 CLFS missing wages and earnings were allocated but no allocation flags are provided. We thus have to include observations with allocated wages and earnings in the analysis which generates an inconsistency relative to the SWA (where missing wages and earnings are not allocated) and the U.S. data. This likely understates the effect of unions on wages in 1984 and 2015, though it is not possible to quantify the extent of the bias. Another limitation is that age is only provided in broad categories, unlike in the U.S. data where age is reported in years. In particular, it is not possible to separate workers age 15 from those age 16. This explains why we use all wage and salary workers age 15 to 64 in Canada, compared to workers age 16 to 64 in the U.S.

A further limitation is that hourly wages are top-coded at a relatively low level in the 1991 SWA ($50) and the 1995 SWA ($40). More than 1% percent of workers are top-coded in these two years. By contrast, only about 0.1% or workers are top-coded in the 1984 SUM (top-code of $45) and in the 1995 CLFS (top-code of $100). For the sake of consistency, we top-code the 1984 SUM and the 2015 CLFS data at $94 (in $2015) which corresponds to the 1984 top-code of $45 expressed in 2015 dollars. We then use the 1997 CLFS data to compute average wages among workers earning between the lowest SWA top-code ($40 in 1995) and the consistent top-code used for the other years ($94 in $2015).[1] This yields an adjustment factor of 1.143. 1991 and 1995 wages adjusted for top-coding are obtained by multiplying wages at the top-code by the adjustment factor. Wages are deflated using the Canadian CPI for all items. We also trim observations with wages below $3.2 in $2015.

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

[1] 1997 is the closest year available for computing the top-coding adjustment factor for the 1991 and 1995 SWA since the CLFS only started collecting wage information in that year.

Table 1a: Summary of Effect of Unions on U.S. Wage Inequality -- Private and Public Sectors Combined

	1973/74		1984		1993		2015	
	male	female	male	female	male	female	male	female
Union Membership Rate	0.307	0.141	0.236	0.141	0.185	0.132	0.119	0.128
Union Effect on Mean Wages:								
Unadjusted Union Wage Gap	0.195	0.230	0.282	0.328	0.299	0.349	0.240	0.294
Composition Effect[a]	-0.009	0.003	0.064	0.108	0.073	0.150	0.083	0.208
Adjusted Union Wage Gap	0.204	0.227	0.218	0.221	0.226	0.199	0.156	0.086
Union Effect on Variance of Wages:								
Actual Variance	0.261	0.195	0.316	0.227	0.346	0.273	0.399	0.349
Actual - Nonunion Variance	-0.048	0.000	-0.036	0.006	-0.025	0.005	-0.016	-0.002
Composition Effect[b]	0.022	0.000	0.012	-0.005	0.008	-0.007	0.002	-0.010
Total union effect	-0.027	0.000	-0.024	0.000	-0.016	-0.002	-0.014	-0.012
Decomposition of Union Effect on Variance:								
Between cell effect	-0.016	-0.002	-0.014	0.001	-0.007	0.002	-0.011	-0.004
Within cell effect	-0.011	0.001	-0.010	0.000	-0.009	-0.004	-0.003	-0.009

Note: samples include wage and salary workers age 16-64 with non-allocated hourly or weekly pay, and hourly wages between $3.35 and $172 per hour in 2015 dollars. Wages for weekly workers with topcoded earnings are adjusted using 1.4 factor. Wages are deflated by CPI-U. Calculations are weighted by CPS Earnings Supplement sample weights.

[a]Composition effect measures component of the union non-union wage gap attributable to non-random coverage, and is measured by difference between fraction of union workers and fraction of nonunion workers in each age/education cell, multiplied by mean wage of non-union workers in the cell.

[b]Composition effect measures component of the variance of nonunion wages that is due to nonrandom coverage, and represents the difference between the actual variance of wages for non-union workers and a reweighted variance that weights nonunion workers in each age/education cell by the fraction of the overall labor force in the cell.

Table 1b: Summary of Effect of Unions on US Wage Inequality -- by Gender and Private and Public Sectors

	1973-1974				1984				1993				2015			
	Public Sector		Private Sector		Public Sector		Private Sector		Public Sector		Private Sector		Public Sector		Private Sector	
	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female
Union Membership Rate	0.289	0.180	0.311	0.130	0.396	0.334	0.206	0.094	0.392	0.369	0.147	0.071	0.428	0.413	0.087	0.062
Union Effect on Mean Wages:																
Unadjusted Union Gap	0.095	0.230	0.216	0.202	0.156	0.272	0.297	0.275	0.193	0.297	0.283	0.237	0.160	0.237	0.188	0.192
Composition Effect[a]	-0.008	0.018	-0.009	-0.028	0.058	0.134	0.021	-0.002	0.040	0.122	0.015	0.007	0.035	0.125	-0.014	0.074
Adjusted Union Wage Gap	0.103	0.211	0.225	0.230	0.099	0.138	0.276	0.277	0.153	0.175	0.269	0.230	0.125	0.111	0.203	0.118
Union Effect on Variance of Wages:																
Actual Variance	0.233	0.204	0.263	0.173	0.249	0.204	0.325	0.222	0.272	0.237	0.350	0.267	0.296	0.289	0.411	0.352
Actual - Nonunion Variance	-0.041	0.000	-0.048	-0.003	-0.068	-0.003	-0.029	0.002	-0.066	-0.009	-0.020	0.000	-0.063	-0.026	-0.011	-0.001
Composition Effect[b]	0.014	0.002	0.024	0.002	0.019	-0.008	0.016	0.002	0.022	-0.003	0.012	0.001	0.015	-0.005	0.005	-0.001
Total union effect	-0.027	0.002	-0.023	-0.001	-0.049	-0.010	-0.013	0.004	-0.044	-0.012	-0.007	0.001	-0.048	-0.031	-0.007	-0.002
Decomposition of Total Union Effect on Variance:																
Between cell effect	-0.016	-0.001	-0.014	-0.003	-0.032	-0.002	-0.008	-0.001	-0.028	-0.002	-0.004	0.000	-0.033	-0.012	-0.007	-0.001
Within cell effect	-0.011	0.002	-0.009	0.002	-0.017	-0.008	-0.005	0.005	-0.017	-0.010	-0.004	0.001	-0.016	-0.018	0.000	-0.001

Note: samples include wage and salary workers age 16-64 with non-allocated hourly or weekly pay, and hourly wages between $3.35 and $172 per hour in 2015 dollars. Wages for weekly workers with topcoded earnings are adjusted using 1.4 factor. Wages are deflated by CPI-U. Calculations are weighted by CPS Earnings Supplement sample weights.

[a]Composition effect measures component of the union non-union wage gap attributable to non-random coverage, and is measured by difference between fraction of union workers and fraction of nonunion workers in each age/education cell, multiplied by mean wage of non-union workers in the cell

[b]Composition effect measures component of the variance of nonunion wages that is due to nonrandom coverage, and represents the difference between the actual variance of wages for non-union workers and a reweighted variance that weights nonunion workers in each age/education cell by the fraction of the overall labour force in the cell.

Table 2a: Summary of Effect of Unions on Canadian Wage Inequality -- Private and Public Sectors Combined

	1984		1993		2015	
	male	female	male	female	male	female
Union Coverage Rate	0.464	0.369	0.408	0.352	0.288	0.320
Union Effect on Mean Wages:						
Unadjusted Union Wage Gap	0.312	0.426	0.300	0.388	0.200	0.331
Composition Effect[a]	0.088	0.115	0.117	0.127	0.075	0.111
Adjusted Union Wage Gap	0.224	0.311	0.184	0.262	0.125	0.221
Union Effect on Variance of Wages:						
Actual Variance	0.246	0.226	0.249	0.245	0.253	0.230
Actual - Nonunion Variance	-0.060	0.017	-0.038	0.003	-0.022	0.007
Composition Effect[b]	0.012	-0.011	0.007	-0.012	0.003	-0.007
Total union effect	-0.049	0.006	-0.031	-0.009	-0.020	0.001
Decomposition of Union Effect on Variance:						
Between cell effect	-0.020	0.014	-0.006	0.012	-0.002	0.014
Within cell effect	-0.028	-0.008	-0.025	-0.021	-0.018	-0.013

Note: samples include wage and salary workers age 15-64 with hourly wages between $3.20 and $92.00 per hour in 2015 dollars. Topcoded wages in 1993 (1991 and 1995 SWA) are adjusted using a 1.143 factor based on mean upper tail wages in the 1997 LFS. Calculations are weighted by LFS sample weights.

[a]Composition effect measures component of the union non-union wage gap attributable to non-random coverage, and is measured by difference between fraction of union workers and fraction of nonunion workers in each age/education cell, multiplied by mean wage of non-union workers in the cell.

[b]Composition effect measures component of the variance of nonunion wages that is due to nonrandom coverage, and represents the difference between the actual variance of wages for non-union workers and a reweighted variance that weights nonunion workers in each age/education cell by the fraction of the overall labor force in the cell.

Table 2b: Summary of Effect of Unions on Canadian Wage Inequality -- by Gender and Private and Public Sectors

	1984				1993				2015			
	Public Sector		Private Sector		Public Sector		Private Sector		Public Sector		Private Sector	
	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female
Union Coverage Rate	0.799	0.780	0.371	0.251	0.790	0.769	0.319	0.238	0.737	0.767	0.194	0.132
Union Effect on Mean Wages:												
Unadjusted Union Gap	0.049	0.290	0.293	0.336	0.078	0.219	0.260	0.300	0.040	0.119	0.120	0.084
Composition Effect[a]	0.036	0.080	0.044	0.059	0.045	0.074	0.069	0.065	0.062	0.032	0.021	0.021
Adjusted Union Wage Gap	0.013	0.210	0.249	0.277	0.033	0.145	0.191	0.236	-0.023	0.086	0.099	0.063
Union Effect on Variance of Wages:												
Actual Variance	0.164	0.155	0.248	0.206	0.159	0.165	0.249	0.229	0.194	0.164	0.249	0.199
Actual - Nonunion Variance	-0.190	-0.107	-0.042	0.007	-0.143	-0.090	-0.029	-0.003	-0.139	-0.098	-0.017	-0.005
Composition Effect[b]	0.055	-0.005	0.014	-0.002	0.049	-0.005	0.010	-0.001	0.045	0.017	0.005	0.000
Total union effect	-0.135	-0.112	-0.028	0.005	-0.094	-0.096	-0.019	-0.005	-0.094	-0.082	-0.012	-0.005
Decomposition of Total Union Effect on Variance:												
Between cell effect	-0.074	-0.035	-0.015	0.005	-0.048	-0.013	-0.005	0.005	-0.030	-0.019	-0.005	0.000
Within cell effect	-0.061	-0.077	-0.013	0.000	-0.046	-0.083	-0.014	-0.010	-0.064	-0.062	-0.007	-0.005

Note: samples include wage and salary workers age 15-64 with hourly wages between $3.20 and $92.00 per hour in 2015 dollars. Topcoded wages in 1993 (1991 and 1995 SWA) are adjusted using a 1.143 factor based on mean upper tail wages in the 1997 LFS. Calculations are weighted by LFS sample weights.

[a]Composition effect measures component of the union non-union wage gap attributable to non-random coverage, and is measured by difference between fraction of union workers and fraction of nonunion workers in each age/education cell, multiplied by mean wage of non-union workers in the cell.

[b]Composition effect measures component of the variance of nonunion wages that is due to nonrandom coverage, and represents the difference between the actual variance of wages for non-union workers and a reweighted variance that weights nonunion workers in each age/education cell by the fraction of the overall labour force in the cell.

Table 3a: Summary of Flattening Effect of Unions on Between-Group Wages in the U.S., by Gender and Sector

	1973-74		1984		1993		2015	
	Private	Public	Private	Public	Private	Public	Private	Public
Models for Male Workers:								
Intercept[a]	0.184	0.110	0.215	0.117	0.213	0.160	0.169	0.117
(standard error)	(0.007)	(0.009)	(0.009)	(0.009)	(0.010)	(0.008)	(0.013)	(0.012)
Coefficient on Non-Union Wage[b]	-0.392	-0.435	-0.336	-0.480	-0.258	-0.423	-0.337	-0.407
(standard error)	(0.021)	(0.030)	(0.023)	(0.024)	(0.024)	(0.022)	(0.033)	(0.033)
R-squared	0.531	0.449	0.393	0.588	0.323	0.664	0.365	0.565
Implied Gap: 0.5 below mean[c]	0.380	0.328	0.383	0.357	0.342	0.371	0.337	0.320
Implied Gap: 0.5 above mean[d]	-0.012	-0.107	0.047	-0.124	0.084	-0.052	0.000	-0.087
Models for Female Workers:								
Intercept[a]	0.213	0.197	0.251	0.171	0.209	0.203	0.130	0.143
(standard error)	(0.008)	(0.013)	(0.007)	(0.008)	(0.009)	(0.007)	(0.012)	(0.012)
Coefficient on Non-Union Wage[b]	-0.261	-0.154	-0.050	-0.304	-0.027	-0.283	-0.084	-0.275
(standard error)	(0.040)	(0.049)	(0.029)	(0.030)	(0.029)	(0.026)	(0.034)	(0.039)
R-squared	0.137	0.045	0.010	0.290	0.004	0.402	0.035	0.292
Implied Gap: 0.5 below mean[c]	0.343	0.274	0.276	0.323	0.223	0.344	0.172	0.281
Implied Gap: 0.5 above mean[d]	0.082	0.120	0.226	0.019	0.196	0.061	0.088	0.006

Note: See Tables 1a and 1b for description of samples. Entries are estimated intercept and slope coefficient (plus R-squared and standard errors) from regressions of mean union-non-union wage gap for a particular skill group on the deviation of the mean nonunion log wage for the skill group from the mean log wage for all nonunion workers. Each skill group is based on single year of education (up to 14 categories) and single year of age (up to 48 categories). Regressions are estimated by weighted least squares using number of nonunion workers in cell as weight.

[a]Estimated intercept from between-skill-group regression is interpretable as union-nonunion wage gap for workers in skill group earning mean nonunion wage.

[b]Estimated coefficient on deviation of skill-group-specific mean nonunion wage from overall mean wage for nonunion workers of same gender and sector.

[c]Implied union-nonunion wage gap for workers in skill group with mean nonunion wage 50 log points ***below*** overall mean nonunion wage.

[d]Implied union-nonunion wage gap for workers in skill group with mean nonunion wage 50 log points ***above*** overall mean nonunion wage.

Table 3b: Summary of Flattening Effect of Unions on Between-Group Wages in Canada, by Gender and Sector

	1984		1993		2015	
	Private	Public	Private	Public	Private	Public
Models for Male Workers:						
Intercept[a]	0.264	0.031	0.191	0.060	0.077	-0.009
(standard error)	(0.009)	(0.015)	(0.014)	(0.020)	(0.021)	(0.021)
Coefficient on Non-Union Wage[b]	-0.479	-0.569	-0.319	-0.559	-0.324	-0.343
(standard error)	(0.027)	(0.036)	(0.044)	(0.055)	(0.079)	(0.060)
R-squared	0.931	0.916	0.696	0.823	0.422	0.620
Implied Gap: 0.5 below mean[c]	0.504	0.315	0.350	0.340	0.239	0.162
Implied Gap: 0.5 above mean[d]	0.025	-0.253	0.032	-0.220	-0.085	-0.181
Models for Female Workers:						
Intercept[a]	0.282	0.222	0.229	0.153	0.056	0.092
(standard error)	(0.012)	(0.014)	(0.012)	(0.017)	(0.013)	(0.017)
Coefficient on Non-Union Wage[b]	-0.173	-0.423	-0.053	-0.321	-0.052	-0.298
(standard error)	(0.059)	(0.050)	(0.051)	(0.071)	(0.061)	(0.061)
R-squared	0.273	0.765	0.044	0.482	0.031	0.531
Implied Gap: 0.5 below mean[c]	0.368	0.433	0.256	0.314	0.082	0.241
Implied Gap: 0.5 above mean[d]	0.195	0.011	0.203	-0.008	0.031	-0.058

Note: See Tables 2a and 2b for description of samples. Entries are estimated intercept and slope coefficient (plus R-squared and standard errors) from regressions of mean union-non-union wage gap for a particular skill group on the deviation of the mean nonunion log wage for the skill group from the mean log wage for all nonunion workers. Skill groups are based on five education categories and five age groups. Regressions are estimated by weighted least squares using number of nonunion workers in cell as weight.

[a]Estimated intercept from between-skill-group regression is interpretable as union-nonunion wage gap for workers in skill group earning mean nonunion wage.

[b]Estimated coefficient on deviation of skill-group-specific mean nonunion wage from overall mean wage for nonunion workers of same gender and sector.

[c]Implied union-nonunion wage gap for workers in skill group with mean nonunion wage 50 log points ***below*** overall mean nonunion wage.

[d]Implied union-nonunion wage gap for workers in skill group with mean nonunion wage 50 log points ***above*** overall mean nonunion wage.

Appendix Table 1
Union Coverage, Composition of Unionized Workforce and Employment by Sector, Canada and US, 2015

Employment by Sector (%)	*Canada*	*U.S.*
Public Sector	20.0	15.4
Private Sector	80.0	84.6

Composition of Unionized Workforce (%)	*Canada*	*U.S.*
Female	52.2	46.7
Male	47.8	53.3
Public Sector	58.4	48.8
Private Sector	41.6	51.2

Union Coverage (%)	*Canada*	*U.S.*
Female	32.4	11.9
Male	28.9	12.6
Public Sector	75.5	39.0
Private Sector	16.7	7.4

Sources:
Canada: Statistics Canada, Labour Force Survey, Tables 14-10-0027-01 and 14-10-0132-01, available at www.statcan.ca

U.S.: Current Population Survey, available at www.unionstats.com

Appendix Table 2a: Distribution of Employment and Unionization Rate (in %) by Sector and Occupation in 2015, U.S.

Occupation:	Employment		Unionization rate	
	Private	Public	Private	Public
Management	10.3	8.0	3.0	22.0
Business and financial operations	4.9	4.5	2.4	23.8
Computer and mathematical science	3.4	2.6	2.7	23.0
Architecture and engineering	2.2	1.6	5.7	26.3
Life, physical, and social science	0.8	2.1	5.0	22.7
Community and social service	1.4	4.0	6.0	39.8
Legal occupations	1.0	1.7	3.3	18.5
Education, training, and library	2.7	27.4	16.4	53.4
Arts, entertainment, and media	1.7	0.8	7.3	27.0
Healthcare professionals and technicians	6.4	4.7	11.5	37.8
Healthcare support	2.7	1.3	8.1	25.2
Protective service	0.8	10.6	9.1	52.5
Food preparation and serving	6.4	2.0	3.9	23.1
Cleaning and maintenance	3.7	3.3	7.6	31.9
Personal care and service	3.2	2.5	4.8	28.9
Sales and related	11.7	0.8	3.6	20.7
Office and administrative support	12.8	13.9	4.9	39.0
Farming, fishing, and forestry occupations	0.8	0.1	2.1	16.5
Construction and extraction	5.3	2.2	17.0	36.6
Installation, maintenance, and repair	3.8	2.0	13.6	40.1
Production occupations	7.1	1.2	12.9	30.6
Transportation and material moving	7.0	2.8	15.3	41.4

Note: Computed using the 2015 CPS

Appendix Table 2b: Distribution of Employment and Unionization Rate (in %) by Sector and Occupation in 2015, Canada

Occupation:	Employment		Unionization rate	
	Private	Public	Private	Public
Senior management	0.1	0.3	4.0	10.1
Other management	5.9	4.7	3.6	42.4
Professionals in business and finance	3.4	3.2	7.2	57.8
Financial, secretarial and administration	4.5	6.4	7.1	74.3
Clerical, including supervisors	10.2	11.1	11.0	75.3
Natural and applied sciences	8.4	7.3	10.1	69.3
Health professionals	0.8	9.7	31.8	84.4
Technical and assisting health occupations	2.8	8.0	30.2	87.9
Social science, government and religion	4.0	8.6	15.8	72.0
Teachers and professors	0.5	18.3	20.8	86.8
Art, culture, recreation and sport	2.4	2.8	16.2	53.1
Wholesale, insurance and real estate	3.7	0.1	5.7	38.1
Retail salespersons, sales clerks, cashiers	9.4	0.7	10.7	74.8
Cooks and food and beverage service	5.2	0.6	8.3	61.1
Protective services	0.8	4.0	39.1	79.0
Childcare and home support	0.8	3.2	23.6	75.1
Sales and service occupations	10.4	4.7	19.0	80.6
Contractors & superv. in trades & transp.	1.4	0.4	23.9	75.7
Construction trades	2.7	0.4	37.8	92.7
Other trades occupations	6.8	2.0	30.5	86.2
Transport and equipment operators	4.3	2.0	27.0	85.3
Construction and transportation labourers	2.8	0.8	26.8	81.3
Occupations in primary industry	2.2	0.3	13.0	73.6
Machine Operators and Assemblers	5.2	0.4	32.6	73.3
Other labourers	1.3	0.0	30.2	51.0

Note: Computed using the 2015 LFS

Figure 1: Union Coverage by Wage Level, USA

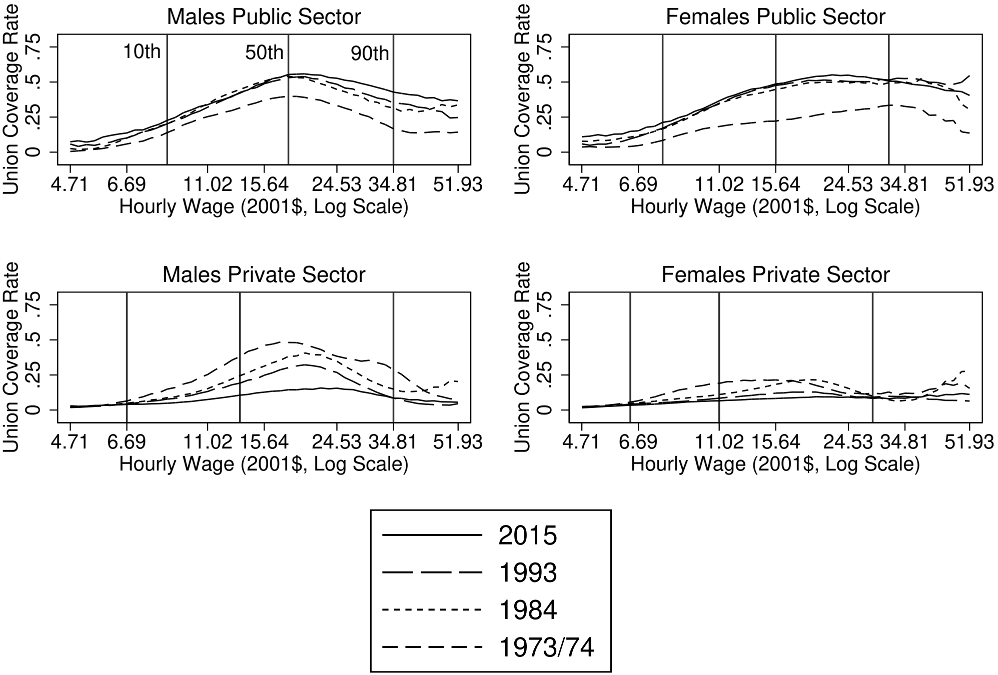


Figure 2: Union Coverage by Wage Level, Canada

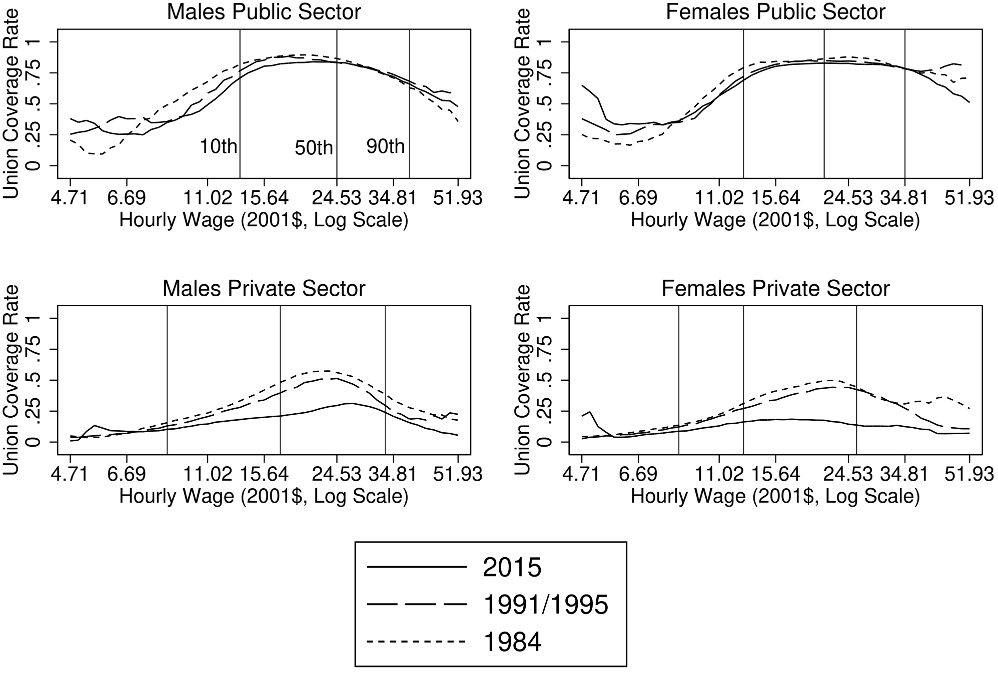
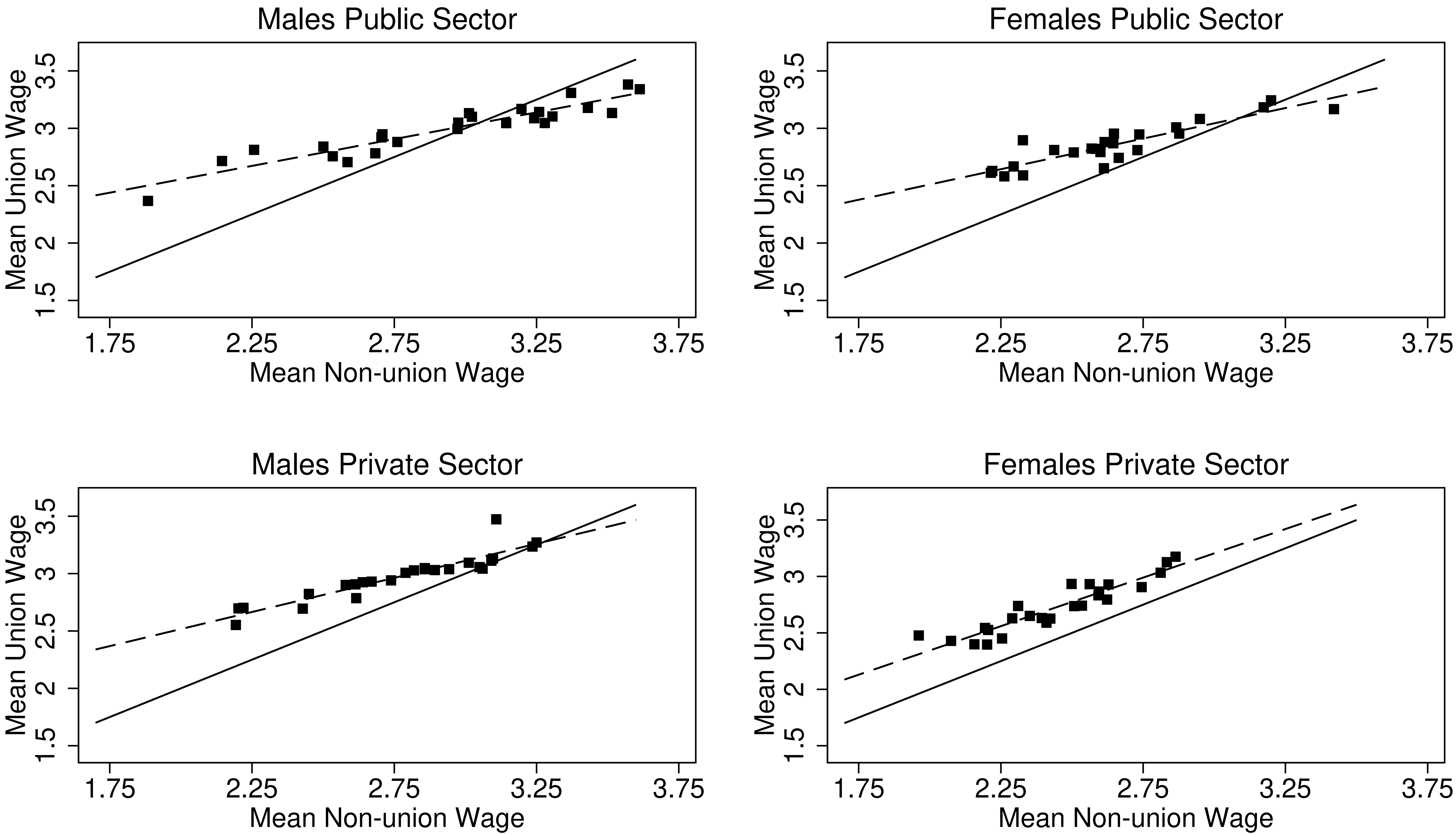
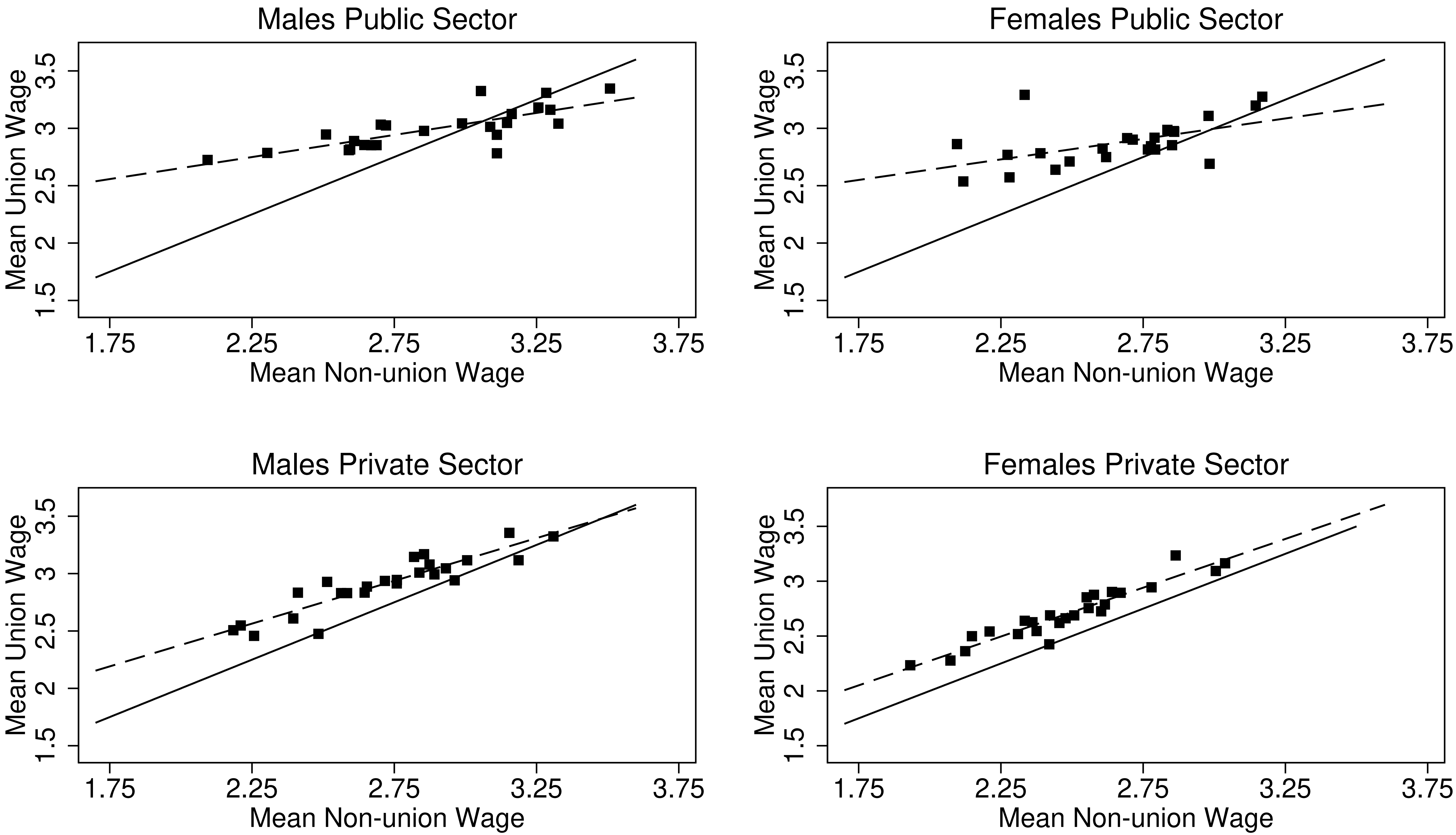


Figure 3a: Union and Nonunion Wage Structures, Canada 1984



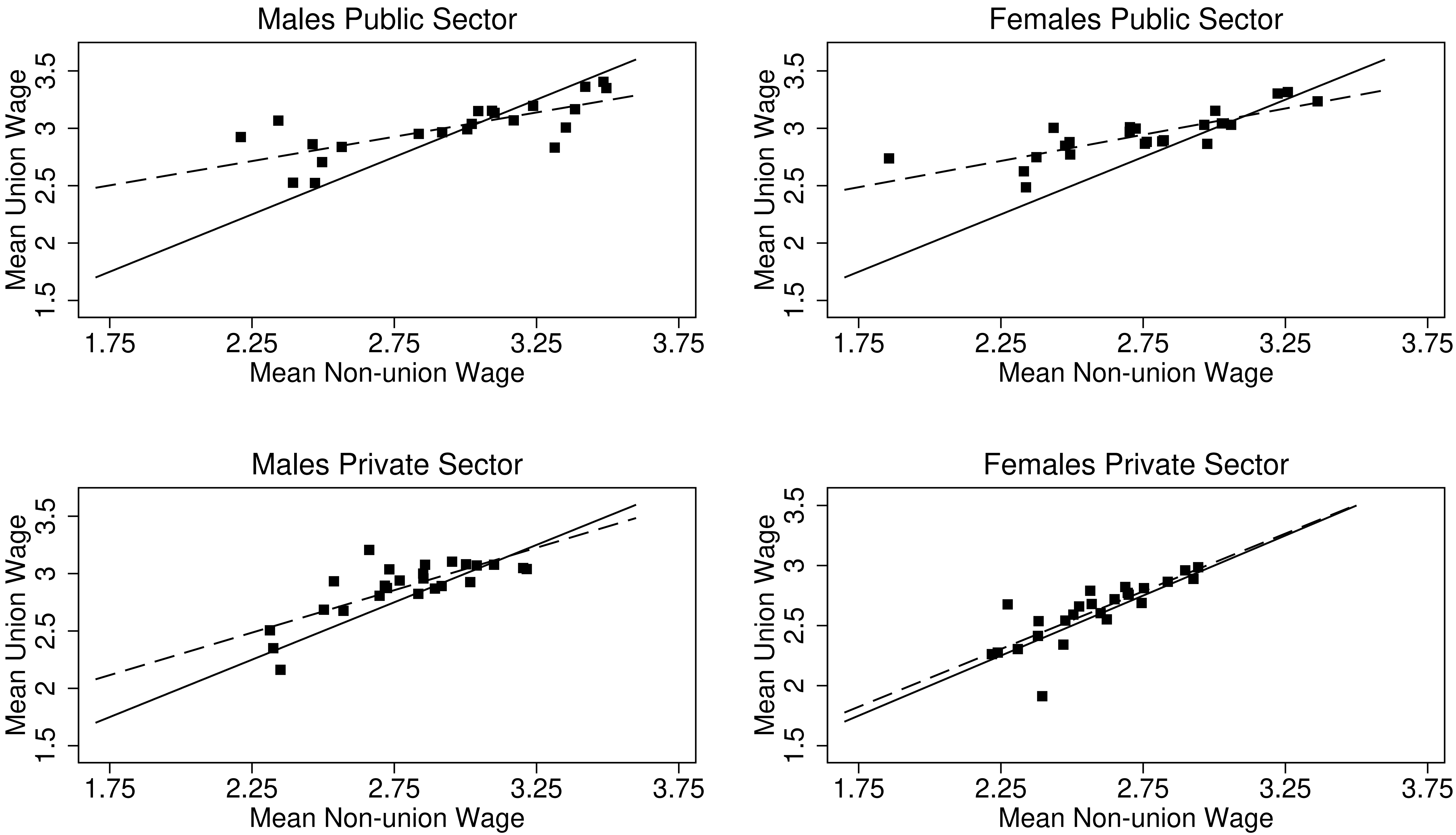
Note: Each point represents an age-education cell. Solid line is the 45 degree line, dashed line is fitted regression line.

Figure 3b: Union and Nonunion Wage Structures, Canada 1993



Note: Each point represents an age-education cell. Solid line is the 45 degree line, dashed line is fitted regression line.

Figure 3c: Union and Nonunion Wage Structures, Canada 2015



Note: Each point represents an age-education cell. Solid line is the 45 degree line, dashed line is fitted regression line.