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3 Unions and Wage Inequality in Canada and the United States

Thomas Lemieux

3.1 Introduction

Throughout the past decade, Canadian workers were twice as likely to be covered by a collective bargaining agreement as their U.S. counterparts (Riddell, chap. 4 in this volume). Over the same period of time, wages were more equally distributed in Canada than in the United States (Blackburn and Bloom, chap. 7 in this volume). These two observations raise the obvious question of whether the different unionization rates in Canada and the United States can explain the difference in wage inequality between the two countries. Several U.S. studies suggest this might be the case. These studies find that unionization narrows the overall distribution of wages among men.[1] It is thus reasonable to expect that the higher rate of unionization in Canada may narrow even more the distribution of wages.

The purpose of this paper is to compare the effects of unionization on wage inequality in Canada and in the United States. After a short discussion of the role of union wage policies in the distribution of wages in section 3.2, the paper begins by describing the patterns of unionism and wages using compa-

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1. U.S. studies that try to assess the impact of unions on the overall distribution of wages of men include Freeman (1980, 1984) and Card (1992). See also Lewis (1986, chap. 10) for a critical survey of additional studies. Swidinsky and Kupferschmidt (1991) present evidence on the impact of unions on residual wage inequality in Canada.

rable cross-sectional data for Canada (the 1986 Labour Market Activity Survey[LMAS]) and for the United States (the 1986 outgoing rotation group file of the Current Population Survey [CPS]) in section 3.3. Cross-sectional estimates of the effects of unions on wages may be afflicted by selectivity biases since union workers are a nonrandom sample of the population. Panel data methods are thus used to estimate selection-adjusted effects of unions on the level and on the variance of wages in Canada in section 3.4. These estimates are used to measure the overall effects of unions on wage inequality in Canada in section 3.5. The estimated effects of unions on wage inequality in Canada are compared to the effects estimated by Card (1992) for the United States in section 3.6. The main finding of the paper is that, for men, differences in unionization rates account for 40 percent of differences in the variance of wages between Canada and the United States.

3.2 Union Wage Policies and Overall Wage Inequality

3.2.1 Efficiency and Equity Issues

There is a long tradition in public policy analysis of evaluating government interventions in terms of efficiency and equity (see, for example, Okun 1975). Union wage policies can also be studied in such terms. In the traditional analysis of union behavior, it is postulated that unions use their monopoly power to set the wages of their members above their competitive level. By creating a wedge between wages and the opportunity cost of labor, unions thus create a deadweight loss measured by the surface of the usual Harberger triangle (see Harberger 1971).

It has long been recognized, however, that union wage policies also have an important impact on the distribution of wages and thus on welfare inequality. On the one hand, unions tend to reduce wage inequalities by standardizing wages within the workplace. On the other hand, union wage policies may exacerbate existing inequalities, as they benefit union workers at the expense of nonunion workers. Existing studies for the United States suggest that, overall, unions reduce wage inequality among male workers (Freeman 1980, 1984). There may thus be a tradeoff between the efficiency costs of unions and the redistributive aspects of union wage policies.[2]

Most empirical studies for Canada and the United States have implicitly focused on the efficiency aspect of union wage policies by estimating the av-

2. Unions may have other roles aside from their effects on the level and distribution of wages. For instance, unions may also increase productivity by giving a "voice" to workers (Freeman and Medoff 1984). In addition, unions are positively associated with nonpecuniary benefits such as pensions and health insurance. Unfortunately, nonpecuniary workers' benefits and productivity effects are harder to quantify than wage effects in standard household surveys like the CPS. By contrast, several data sets for both Canada and the United States contain good information on wages. For these reasons, the paper focuses on the wage effects of unions.

erage effect of unions on the level of wages. The recent widening of the distribution of wages in the United States, however, has revived interest in the redistributive aspects of union wage policies. For example, Card (1992) concluded that 20 percent of the increase in the variance of wages of men in the United States from 1973 to 1987 was attributable to deunionization. Freeman (1991) reported similar findings for the 1978 to 1987 period.

Measuring the deadweight loss of union wage policies is now a standard textbook case that will not be discussed here.[3] The remainder of the section will explain in some detail how union wage policies affect the overall distribution of wages in the economy.

3.2.2 The Effects of Unions on the Distribution of Wages

The effect of unions on the distribution of wages does not simply depend on the size of the average union wage differential. It depends on the joint distribution of unionization and of wages (union and nonunion) in the work force. To see this, consider the nonunion wage of a worker i with observed characteristics x_i:

$$w_i^N = w^N(x_i) + \varepsilon_i^N,$$

where ε_i^N is an error term with zero conditional mean and a conditional variance $\sigma_N^2(x_i)$. The wage for this same worker in the union sector is given by

$$w_i^U = w^U(x_i) + \varepsilon_i^U,$$

where ε_i^U has zero conditional mean and a conditional variance $\sigma_U^2(x_i)$. Consider $\bar{U}(x_i)$, the probability that worker i is unionized. Consider also the union wage gap $\Delta_w(x_i)$ for that worker,[4]

$$\Delta_w(x_i) = w^U(x_i) - w^N(x_i),$$

and the union variance gap $\Delta_v(x_i)$,

$$\Delta_v(x_i) = \sigma_U^2(x_i) - \sigma_N^2(x_i).$$

3. The standard formula for the deadweight loss is simply $DW \approx (.5\bar{\eta}\bar{\Delta}_w^2)wL$, where $\bar{\eta}$ is the (average) labor demand elasticity, $\bar{\Delta}_w$ is the (average) union wage gap, and wL is the wage bill in the union sector (see Harberger 1971 and Rees 1963 for an application to this particular problem). Note, however, that the simple formula is only valid when Δ_w, η, and the unionization rate are either constant or independently distributed across skill groups. In general, the deadweight loss with heterogeneous skill groups j is $DW = \Sigma_j(.5\eta_j\Delta_{wj}^2)w_jL_j$.

4. Note that it is implicitly assumed in this section that negotiated wages in the union sector have no effect on wages in the nonunion sector. This assumption is unlikely to hold, since general equilibrium considerations (Johnson and Mieszkowski 1970) suggest unions have a negative impact on wages in the nonunion sector, while union threat effects (Rosen 1969) suggest the opposite. The wages that prevail in the nonunion sector might thus be different from the wages that would prevail in the absence of unions. Lewis (1986) has convincingly argued that it was not possible to estimate the wages that would prevail in the absence of unions using standard household-based surveys. The more limited goal of this paper is therefore to compare the actual distribution of wages to the distribution of wages that would prevail if all workers were paid according to the wage schedule observed in the nonunion sector.

For convenience, call the group of workers with a given set of characteristics x a "skill group." The average wage $\bar{w}(x)$ of workers in skill group x is given by

$$(1) \qquad \bar{w}(x) = w^N(x) + \bar{U}(x)\Delta_w(x),$$

while the variance of wages among these workers, $\sigma^2(x)$, is given by[5]

$$(2) \qquad \sigma^2(x) = \left[\sigma_N^2(x) + \bar{U}(x)\Delta_v(x)\right] + \left[\bar{U}(x)(1 - \bar{U}(x))\Delta_w(x)^2\right].$$

The first component in square brackets is the average within-sector (union and nonunion) variance of wages, while the second component in square brackets is the between-sector variance of wages.

Now consider the overall variance of wages, which is a standard measure of wage dispersion. There are two components to the overall variance of wages among workers: the variance of wages among workers with a given set of characteristics x and the variance of wages between workers with different characteristics x. It follows from a standard variance decomposition that the overall variance of wages among nonunion workers is

$$(3) \qquad \mathrm{Var}(w_i^N) = \mathrm{Var}(w^N(x)) + E(\sigma_N^2(x)),$$

while the overall variance of wages among all workers is

$$(4) \qquad \mathrm{Var}(w_i) = \mathrm{Var}(\bar{w}(x)) + E(\sigma^2(x)).$$

Substituting equations (1) and (2) into equation (4) yields

$$(5) \qquad \begin{aligned} \mathrm{Var}(w_i) = {} & \mathrm{Var}[w^N(x)] + \mathrm{Var}[\bar{U}(x)\Delta_w(x)] + 2\mathrm{Cov}[w^N(x), \bar{U}(x)\Delta_w(x)] \\ & + E[\sigma_N^2(x)] + E[\bar{U}(x)\Delta_v(x)] + E[\bar{U}(x)(1 - \bar{U}(x))\Delta_w(x)^2]. \end{aligned}$$

The overall effect of unions on the variance of wages is then obtained by subtracting equation (3) from equation (5):

$$\begin{aligned} \mathrm{Var}(w_i) - \mathrm{Var}(w_i^N) = {} & \mathrm{Var}[\bar{U}(x)\Delta_w(x)] + 2\mathrm{Cov}[w^N(x), \bar{U}(x)\Delta_w(x)] \\ & + E[\bar{U}(x)\Delta_v(x)] + E[\bar{U}(x)(1 - \bar{U}(x))\Delta_w(x)^2]. \end{aligned}$$

The effect of unions on the variance of wages is thus attributable to three separate factors:

1. how unions change the relative position of each skill group in the wage distribution,

$$(6) \qquad \mathrm{Var}[\bar{U}(x)\Delta_w(x)] + 2\mathrm{Cov}[w^N(x), \bar{U}(x)\Delta_w(x)];$$

5. This formula is derived as follows:

$$\begin{aligned} \sigma^2(x) &= \mathrm{Var}\,(w_i|x) = E_U[\mathrm{Var}(w_i|x, U)] + \mathrm{Var}_U[E(w_i|x, U)] \\ &= (1 - \bar{U}(x))\mathrm{Var}(w_i^N|x) + \bar{U}(x)\mathrm{Var}(w_i^U|x) + \mathrm{Var}_U[w^N(x) + \bar{U}\Delta_w(x)] \\ &= (1 - \bar{U}(x))\sigma_N^2(x) + \bar{U}(x)\sigma_U^2(x) + \bar{U}(x)(1 - \bar{U}(x))\Delta_w(x)^2. \end{aligned}$$

2. how unions increase the variance of wages *between* union and nonunion workers in a skill group, averaged over skill groups,

$$E[\bar{U}(x)(1 - \bar{U}(x))\Delta_w(x)^2]; \tag{7}$$

3. how unions affect the residual variance of wages within union workers in a skill group, averaged over skill groups,

$$E[\bar{U}(x)\Delta_v(x)]. \tag{8}$$

A brief examination of effects 1 and 2 indicates that they depend on the *joint* distribution of $w^N(x)$, $\bar{U}(x)$, and $\Delta_w(x)$. It is therefore necessary to estimate this joint distribution to evaluate the overall impact of unions on wage inequality. The sign of effect 1 may be either positive or negative, depending on whether the covariance term is negative enough to offset the variance term. This covariance term is negative whenever the net union wage effect $\bar{U}\Delta_w$ is larger for workers at the low end of the skill distribution (workers with a low w^N) than for workers at the high end of the skill distribution.

Effect 2 is always positive, since union wage policies pull union workers apart from nonunion workers. On the other hand, effect 3 is usually believed to be negative, as unions tend to standardize wages among union workers ($\Delta_v < 0$).

If all the productive characteristics of workers were observed in the data, it would be straightforward to estimate $\Delta_w(x)$, $\Delta_v(x)$, $w^N(x)$, and thus $\text{Var}(w_i) - \text{Var}(w_i^N)$. Most estimation problems arise when some of these characteristics are unobserved in the data. Consistent estimation of $\Delta_w(x)$, $\Delta_v(x)$, and $w^N(x)$ in that context will be addressed in section 3.4. Note also that the effect of unions on the variance of wages between union and nonunion workers (effect 2) is no longer given by formula (7) when x only represents a subset of the relevant productive characteristics. The point is that union and nonunion workers with the same x's could have systematically different wages even in the absence of unions because of differences in unobserved characteristics. That underlying difference in wages would simply be $\bar{\Delta}_w(x) - \Delta_w(x)$, where $\bar{\Delta}_w(x)$ is the observed difference in wages between union and nonunion workers, while $\Delta_w(x)$ is the properly estimated effect of unions on wages. The effect of unions on the variance of wages between union and nonunion workers would thus become

$$\bar{U}(x)(1 - \bar{U}(x))\left[\bar{\Delta}_w(x)^2 - [\bar{\Delta}_w(x) - \Delta_w(x)]^2\right]. \tag{7'}$$

The formulas for the effects 1 and 3 would remain unchanged.

3.3 Data and Basic Empirical Regularities

The effect of unions on the variance of wages can be obtained by estimating the various components of equations (6), (7′), and (8). Before doing so, it is useful to describe the data used in this paper along with the basic empirical regularities in these data. The various components of equations (6), (7′), and (8) will be estimated in section 3.4.

3.3.1 Data

The Canadian data used for this study were obtained by merging the 1986–87 longitudinal file of the Canadian LMAS to the 1986 cross-sectional file of the LMAS. The 1986 LMAS was administered in January, February, and March 1987 to five rotation groups of the Canadian Labour Force Survey (LFS).[6] The public use sample consists of 66,934 people aged 16–69. It contains detailed information on up to five jobs held in 1986, including the usual hourly wage rate on all paid (except self-employed) jobs and the union status on the job.[7] In this paper, workers are classified as union when they are members of the union that collectively bargained with the employer, or are covered by a collective agreement. Otherwise, they are classified as nonunion. The LMAS also contains detailed information on the work history of each individual, including the reason a worker changed jobs. It is thus possible to reconstruct the precise timing of job changes and to know why people did change jobs.

Most of the people who were initially surveyed in 1987 (1986 LMAS) were reinterviewed in 1988 (1987 LMAS).[8] Like the 1986 LMAS, the 1987 LMAS contains information on up to five jobs held during the year. The 1986–87 longitudinal file was created by Statistics Canada by matching the information from the 1986 and 1987 cross-sections. The longitudinal file thus contains information on up to ten jobs held over the 1986–87 period. Since the LMAS is a work history survey, availability of the 1987 LMAS is not crucial for fixed effect estimation. It simply doubles the length of the work history. In addition, the 1987 LMAS contains useful information on ethnic origin, race, immigrant status, and mother tongue. These questions were not asked in the 1986 LMAS.

For the sake of comparability with other studies, this paper uses a sample of men and women aged 20–64 who hold jobs in nonagricultural industries. This subsample is also restricted to people who have worked for at least four weeks on a paid, but not self-employed, job in 1986. Jobs with a usual wage rate of less than $1.00 or more than $75.00 an hour are also excluded from the sample. A total of 34,765 workers satisfied the various sample selection criteria.[9]

6. The LFS sample design is based on six rotation groups including approximately 130,000 people. People remain in the sample for six consecutive months, at which time they are replaced.

7. The definition of a job in the LMAS is "usual duties performed at a usual wage or salary" (Statistics Canada 1988, sec. 4.1).

8. Statistics Canada managed to reinterview more than 90 percent of the people surveyed in the 1986 LMAS, including several thousands who had moved between the two interviews.

9. Of these workers, 32,696 were reinterviewed for the 1987 LMAS. The variables on ethnic origin, immigrant status, race, and mother tongue are only available for these 32,696 workers.

The U.S. data used to describe the basic empirical regularities on unions and wages come from the 1986 merged outgoing rotation group file of the CPS. Both the Canadian and the U.S. data are thus based on earnings supplements to very similar surveys (the LFS and the CPS). The sample selection criteria used to construct the final U.S. data set are also similar to the ones used for the Canadian data.[10] A sample of 161,195 workers satisfied these sample selection criteria. More details on the U.S. data and its comparability with the LMAS data are provided in appendix A.

3.3.2 Empirical Regularities

As mentioned earlier, the effect of unions on the distribution of wages depends on the joint distribution of unionization rates, union wages, and nonunion wages. This joint distribution is analyzed empirically by first tabulating unionization rates and wages over a set of workers and jobs characteristics. Table 3.1 presents the distribution of these characteristics in the sample, along with the unionization rate for workers with these characteristics.

Columns 5 and 6 indicate that the fraction of workers covered by collective bargaining agreements was 45.8 percent among men and 36.4 percent among women in Canada in 1986. In the LMAS sample, women account for 45.5 percent of the work force and for 39.9 percent of all workers covered by collective bargaining agreements. Furthermore, most women holding union jobs work in the public sector, where the union density is 67.3 percent, as opposed to 18.9 percent in the private sector.[11] The union density is also higher in the public sector than in the private sector for men. Overall, 45.1 percent of Canadian workers covered by collective bargaining agreements work in the public sector.

The composition of the Canadian and U.S. samples reported in table 3.1 are similar with a few exceptions. One difference is the race composition of the two samples. There are also some differences in educational achievement in the two countries, in part because of differences in the questions used in the two surveys.[12] Finally, the public sector employs relatively more people (es-

10. The only sample selection criterion that was not used is the condition that the job must last at least four weeks (job duration is not available in the CPS data used). See appendix A for more details.

11. The definition of the public sector used for Canada includes the health and welfare industry in addition to education services and public administration. This definition is used because the LMAS does not contain direct information on whether a job is in the public or in the private sector. Such information is available, however, in the 1984 Survey of Union Membership. Using these data, Riddell (chap. 4 in this volume) estimates the private sector density at 29 percent (men and women together). The definition of the private sector used in this paper would imply a union density of 28 percent (my calculation using table 2 in Kumar 1988), which is a satisfactory approximation to the true union density in the private sector.

12. In the LMAS, educational achievement is classified in five categories: none or elementary, high school (some or completed), some postsecondary, postsecondary certificate or diploma, and university. These five categories were mapped into the five following ranges of years of schooling completed in the CPS: 0–7, 8–12, 13, 14–15, 16 and more.

Table 3.1 Sample Characteristics and Unionization Rates in Canada and the United States

	Sample Composition (%)				Unionization Rate (%)			
	Canada		U.S.		Canada		U.S.	
	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)	Men (7)	Women (8)
Total	54.53	45.47	53.49	46.51	45.78	36.39	25.63	16.60
Age								
20–24	15.12	18.03	14.27	15.63	26.25	20.25	12.90	7.71
25–34	32.19	33.60	33.80	32.30	43.04	37.71	21.74	14.82
35–44	25.50	25.25	24.96	25.60	51.92	43.30	30.60	19.78
45–54	16.44	15.18	16.15	16.19	54.04	40.72	33.18	21.16
55–64	10.75	7.94	10.82	10.28	54.25	37.24	31.88	20.62
Education								
Primary	11.20	7.37	3.30	1.94	57.21	32.17	23.42	15.12
High school	47.38	47.64	50.73	53.54	47.18	30.76	30.94	15.11
More than high school	10.51	11.35	7.06	8.26	38.38	29.31	27.35	12.15
Some post-secondary	14.56	18.80	13.87	14.64	42.27	44.12	22.18	11.91
University degree	16.35	14.84	25.03	21.62	41.79	52.22	16.59	25.31
Marital status								
Single	28.81	33.20	33.18	42.05	35.93	33.30	20.09	16.00
Married	71.19	66.80	66.82	58.95	49.77	37.93	28.39	17.04
Race								
White	93.72	93.69	87.10	85.01	46.78	37.02	24.96	15.37
Nonwhite	6.28	6.31	12.90	14.99	38.74	32.80	30.15	23.57
Part-time status								
Full-time	94.46	75.36	93.48	76.41	46.83	38.57	26.58	18.73
Part-time	5.54	24.64	6.52	23.59	27.95	29.73	12.00	9.70
Private sector	80.93	63.87	83.90	79.97	39.21	18.89	21.55	10.19
Public sector	19.07	36.13	16.10	20.03	73.67	67.33	46.93	42.19
Occupation								
White-collar	53.20	89.89	54.28	87.66	38.15	36.18	18.40	15.23
Blue-collar	46.80	10.11	45.72	12.34	54.45	38.34	34.22	26.35
Mother tongue								
English	59.02	61.50	—	—	42.51	34.15	—	—
French	26.30	24.70	—	—	54.30	45.26	—	—
Others	14.68	13.80	—	—	47.00	33.10	—	—

Sources: Canadian data are from the 1986 cross-sectional file and the 1986–87 longitudinal file of the LMAS. Sample size is 34,765, except for the tabulations for mother tongue and race, which are based on a matched sample of 32,696 observations. U.S. data are from the 1986 merged outgoing rotation group files of the CPS. Sample size is 161,195.

Note: The estimated frequency distributions are all weighted.

pecially women) in Canada, as it is defined to include the health and education sectors.

Table 3.2 presents ordinary least squares (OLS) estimates of standard log hourly wage equations in which the following regressors are included: an indicator variable for union coverage, the set of worker and job characteristics listed in table 3.1, and controls for industry, occupation, and region. The estimated union wage gap is comparable for men in Canada (0.198) and in the United States (0.180). It is much larger, however, for women in Canada (0.287) than for women in the United States (0.156). With the exception of women in Canada, the estimated union wage gaps reported here are consistent with previous findings in the literature.[13] The other estimated wage effects are similar in the two countries except for the effect of part-time employment, which is much larger in absolute value in the United States. Although the estimated returns to education are hard to compare for reasons discussed above, the estimated university–high school wage differentials are similar in Canada (0.257 for men, 0.271 for women) and in the United States (0.247 for men, 0.223 for women).[14]

The average union wage gaps reported in table 3.2 are estimated under the implicit assumption that the wage gap is the same for all workers. As mentioned in section 3.2, differences in the wage gap $\Delta_w(x)$ and in the union density $\bar{U}(x)$ by skill groups may play an important role in the overall impact of unions on the distribution of wages. The relationships among $\Delta_w(x)$, $\bar{U}(x)$, and the nonunion wage $w^N(x)$ are examined graphically by fitting simple index models for these three variables. More specifically, a log wage equation for the sample of nonunion worker is fit to the set of region dummies, indicator variables for marriage and race, and fully interacted age and education dummies. The nonunion wage index for a worker with characteristics x_i is then defined as the predicted wage from that regression (excluding the effect of province, on the assumption that regional wage differences reflect cost of living rather than skill differences). This nonunion wage index can be interpreted as a general skill index. A similar union wage index is constructed by running a wage regression on the sample of union workers. The union wage gap for a

13. Results from American studies are surveyed by Lewis (1986). See also Freeman and Medoff (1984). For Canadian studies that focus on the estimation of the average union wage gap, see Evans and Clark (1986); Grant, Swidinsky, and Vanderkamp (1987); Kumar and Stengos (1985, 1986); Maki and Ng (1990); Robinson (1989); Robinson and Tomes (1984); and Simpson (1985). The finding that the union wage gap is larger for women than for men in Canada is at odds with the results of Maki and Ng (1990) (similar wage gaps for men and women), who used data for 1981. The finding is consistent, however, with the fact that the unadjusted union wage gap was substantially larger for women than for men in Canada in 1984, 1986, and 1987 (Labour Canada 1991).

14. The estimated returns for education are different from those reported by Freeman and Needels (chap. 2 in this volume) because of the inclusion of industry and occupation dummies in the wage regression.

Table 3.2 OLS Estimates of the (Log) Wage Equation

	Canada		U.S.	
	Men (1)	Women (2)	Men (3)	Women (4)
Covered by collective bargaining	0.198 (0.007)	0.287 (0.007)	0.180 (0.004)	0.156 (0.004)
Age 25–34	0.211 (0.010)	0.180 (0.009)	0.209 (0.005)	0.180 (0.005)
Age 35–44	0.340 (0.011)	0.221 (0.010)	0.335 (0.005)	0.231 (0.005)
Age 45–54	0.372 (0.012)	0.234 (0.011)	0.376 (0.006)	0.231 (0.005)
Age 55–64	0.332 (0.013)	0.207 (0.013)	0.355 (0.006)	0.226 (0.006)
High school	0.138 (0.010)	0.090 (0.012)	0.273 (0.009)	0.252 (0.011)
More than high school	0.197 (0.013)	0.148 (0.015)	0.362 (0.010)	0.315 (0.012)
Some postsecondary	0.267 (0.012)	0.228 (0.014)	0.386 (0.009)	0.366 (0.012)
University degree	0.395 (0.014)	0.361 (0.016)	0.520 (0.009)	0.475 (0.012)
Married	0.111 (0.007)	0.012 (0.007)	0.109 (0.003)	0.012 (0.003)
Nonwhite	−0.075 (0.016)	−0.006 (0.017)	−0.093 (0.004)	−0.030 (0.004)
Part-time	−0.124 (0.013)	−0.016 (0.007)	−0.299 (0.006)	−0.169 (0.004)
Mother tongue				
French	−0.014 (0.011)	−0.004 (0.011)	—	—
Not English or French	−0.027 (0.011)	−0.017 (0.011)	—	—
Gender dummy[a] (women = 1)	−0.262 (0.008)		−0.241 (0.002)	
Observations	18,679	16,086	84,275	76,920
R^2	0.368	0.430	0.415	0.402
Root mean squared error	0.388	0.372	0.407	0.395
Mean of dependent variable	2.422	2.088	2.250	1.906

Sources: Canadian data are from the 1986 cross-sectional file and the 1986–87 longitudinal file of the LMAS. U.S. data are from the 1986 outgoing rotation group file of the CPS. The dependent variable is the log of the hourly wage rate. All specifications also include region dummies (ten provinces in Canada, nine regions in the United States), seven industry dummies, and eight occupation dummies. The base group is age 20–24, primary education, single, white, mother tongue English (Canada only).

[a]Estimated from a separate pooled regression for men and women.

given value of the nonunion wage index is simply the difference between the union and the nonunion wage index. Finally, a union density index is constructed by fitting a linear probability model and using the procedure discussed above to predict the probability of union coverage.[15]

The predicted union wage gap and the predicted unionization rate are plotted against the nonunion wage index for the sample of Canadian men in figure 3.1a. The fitted lines in the figure are obtained by regressing the predicted wage gap and the predicted unionization rate on a third-degree polynomial of the nonunion wage index.[16] The graph indicates that the union wage gap declines with skill. It is even negative for workers at the high end of the skill distribution. The graph also indicates that the unionization rate increases with skill among workers at the low end of the skill distribution. The unionization rate then remains more or less constant for workers at the middle and high ends of the skill distribution.

Figure 3.1a thus suggests that, for men in Canada, unions have a mixed impact on the distribution of wages across skill groups. The wage gap for low-skill men is large, but few of these men are unionized, while the opposite is true for high-skill men. This patterns hides important differences, however, between the impact of unions in the private and in the public sector. On the one hand, Figures 3.1b and 3.1c show that the wage gap declines with skill in both sectors. On the other hand, the figures show different patterns of unionization in the two sectors. While the unionization rate rises steadily to reach 80 percent at the high end of the skill distribution in the public sector, it peaks around 40 percent and then declines with skill in the private sector. Union workers are thus concentrated in the middle of the skill distribution in the private sector, but at the high end of the skill distribution in the public sector. The unionization rate is thus high for highly skilled men in Canada because of the pattern of the unionization in the public rather than in the private sector.

Figure 3.1d shows the same plots for men in the United States. Like Canadian union workers in the private sector, union workers in the United States are concentrated in the middle of the skill distribution. The union wage gap also declines with skill, though not monotonically, and is negative for workers at the high end of the skill distribution. Unions thus have similar relative wage effects in Canada and in the United States. Union workers are more skilled in Canada than in the United States, however, because of the high level of unionization among public sector workers.

Figure 3.2 shows analogous plots for women. These figures suggest two

15. Since age and education dummies are fully interacted, the predicted probabilities from a linear probability model are almost identical to the predicted probabilities from a probit or logit model (they would be numerically equivalent if all the regressors were fully interacted).

16. Only the fitted values (from a cubic regression) of the union wage gap and of the unionization rates. as opposed to the predicted values of these variables for each age-education-race-marital status cell, are plotted to simplify the graphs.

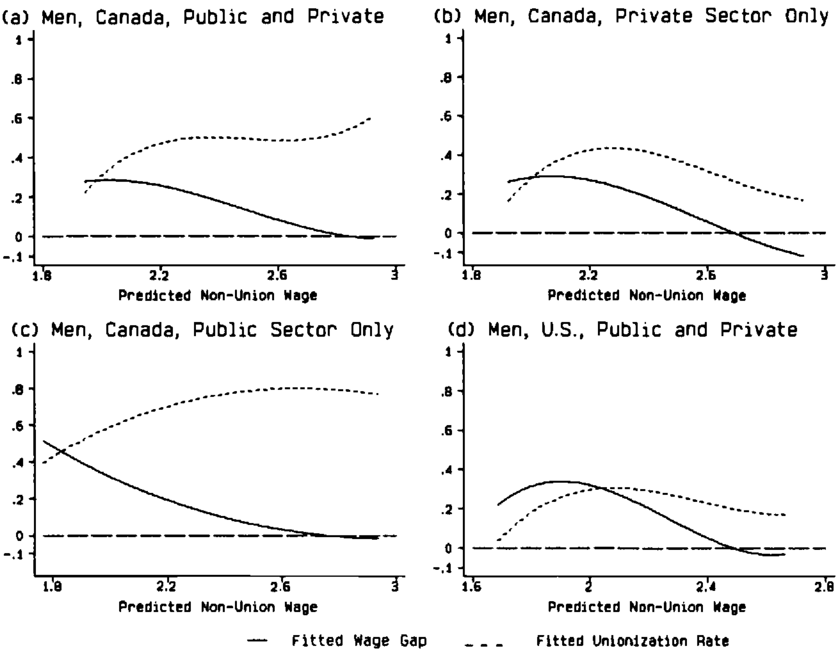

Fig. 3.1 Predicted union wage gap and predicted unionization rate by skill level: men

Note: Fitted values from a cubic regression of the predicted union wage gap (or unionization rate) on the predicted nonunion wage by age-education-race–marital status cell (see text for details).

major differences in the patterns of unionization and wages between men and women. First, the predicted union wage gap declines much less with skill for women than for men. The predicted union wage gap is always at least half as large for women at the high end of the skill distribution as for women at the low end of the skill distribution. It is almost a flat function of skills for women in Canada when private and public sector workers are pooled (figure 3.2a). A second major difference between men and women is that, for women, unionization is concentrated at the high end of the skill distribution in both Canada and the United States. The breakdown between the private and the public sector in Canada (figures 3.2b and 3.2c) suggests this overall pattern is due to concentration of union jobs in the public sector.

The division of union jobs between the public and the private sectors thus goes a long way toward explaining the patterns of unionization along skill lines for men and women in Canada and the United States. The fraction of union workers who hold a public sector job is 29 percent for U.S. men, 31 percent for Canadian men, 51 percent for U.S. women, and 67 percent for Canadian women. Figures 3.1 and 3.2 show that, as this fraction increases,

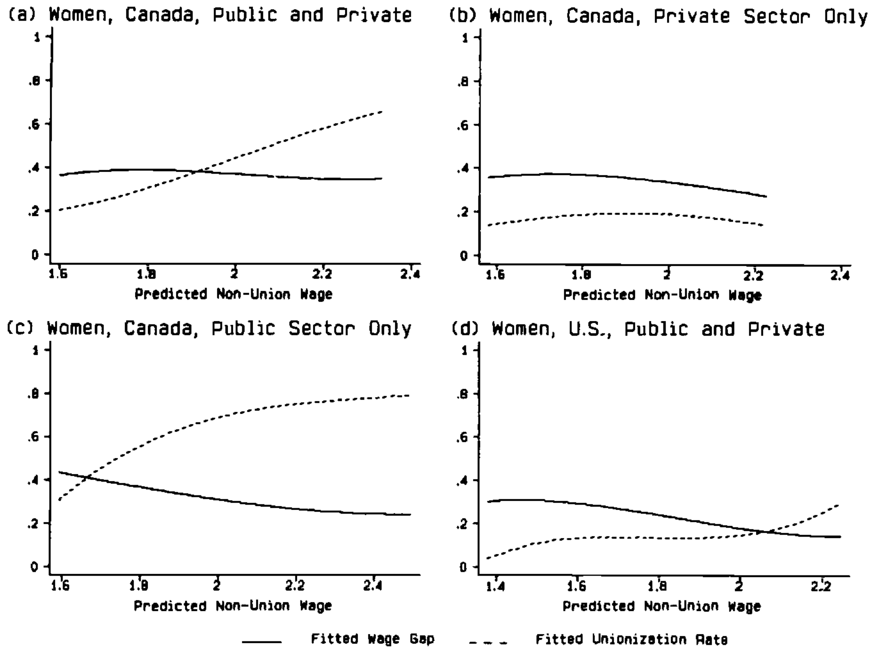

Fig. 3.2 Predicted union wage gap and predicted unionization rate by skill level: women

Note: Fitted values from a cubic regression of the predicted union wage gap (or unionization rate) on the predicted nonunion wage by age-education-race–marital status cell (see text for details).

the distribution of unionization gets more and more skewed to the right of the skill distribution.

The other main conclusion to be drawn from figures 3.1 and 3.2 is that the union wage gap declines in the skill level. This negative relationship is stronger for men than for women. The same basic conclusions are reached using a more standard regression-based approach for Canada (see appendix table 3B.1). Appendix table 3B.2 summarizes the results of that regression-based approach by showing the effects of unions on a selected number of wage differentials such as the university–high school wage differential and the white-collar/blue-collar wage differential.

3.4 Fixed Effect Estimation

The descriptive analysis of section 3.3 indicates substantial diversity in the role of unions for various subgroups of the work force. This section presents detailed estimates of the effects of unions on both the level and the variance of wages for some of these subgroups. These estimates will be used in section

3.5 to calculate the overall impact of unions on wage inequality. This section first discusses the importance of adjusting these estimates for the self-selection of workers into the union sector for each subgroup of the work force. The actual estimates are reported later in the section.

3.4.1 Self-selection and the Fixed Effects Method

The pattern of the union wage gap along skill lines documented in section 3.3 raises the question of whether the estimated gaps represent "true" effects of unionism on wage or merely reflect the selection of workers into the union sector on the basis of their unobserved productive characteristics. For instance, it may be that, among workers with low observed skills, only the most qualified are hired by unionized employers. It may also be that most workers with high observed skills do not get many benefits for joining unions, so only those with low unobserved skills join unions. As a result, union workers might be positively selected at the low end of the observed skill distribution but negatively selected at the high end of the observed skill distribution. This pattern of selection would arise from a model in which (1) unions compress the distribution of wages across skill groups, (2) workers who get the highest wage premium are the most likely to want to join unions, and (3) employers would rather hire workers for which the wage premium is the lowest.[17] In this model, union employers do not want to hire low-skill workers as they command a high wage premium, while high-skill workers do not want union jobs because their wage gain from unionization is low or negative. Unionization is thus concentrated among workers in the middle of the skill distribution. Since total skills are the sum of observed and unobserved skills, union workers with low observed skills tend to have high unobserved skills (positive selection) and vice versa. This model thus generates a pattern of selection that could explain why the estimated wage gap declines with skill.

The fixed effect approach is a standard technique used to consistently estimate the effects of union on wages when workers are selected in the union sector on the basis of unobservable characteristics.[18] The approach has the advantage of being robust to the complicated pattern of selection mentioned above. The goal of this section is thus to consistently estimate the union wage gap $\Delta_w(x)$ and the union variance gap $\Delta_v(x)$ by exploiting the panel data aspect of the LMAS and of the (matched) CPS data. Several recent U.S. studies, including Card (1992), have used fixed effect methods with the CPS data to estimate the effects of unions on wages. Since Card also addresses measure-

17. These are the building blocks of the queuing model of unionization of Abowd and Farber (1982).

18. See, for example, Chamberlain (1982) and Freeman (1984), who use various versions of the fixed effect approach to estimate the average union wage gap in the United States. There is a large debate on whether or not the fixed effect approach appropriately adjusts the wage gap estimates for the selection of union workers. Lemieux (1992) discusses these issues in detail. The results reported there suggest that fixed effect estimates adjust for most of the selection bias.

ment issues specific to the matched CPS data that are beyond the scope of this paper (see appendix A), his results will be directly used here and compared to the results obtained using the LMAS data.

One shortcoming of the fixed effect approach is that, in the LMAS data, it is only applicable to the limited sample of workers who changed jobs at least once during the 1986–87 period. This reduces the precision of the estimates and limits the ability to measure the effect of unions for small subgroups of the work force. It is thus important to use a parsimonious approach to get precise enough estimates while still letting the effect of unions vary along the lines suggested in section 3.3. In light of the results in section 3.3, it was decided to analyze men and women separately and to further break down the data into workers in the public and private sectors. It is also important to allow for some heterogeneity in the effect of unions by skill level. For the sake of comparability with the study of Card (1992) for the United States, the sample will thus be divided into three skill group (tiers) on the basis of workers' predicted wage in the nonunion sector.[19]

3.4.2 Dividing the Sample in Three Tiers

The sample of men and the sample of women are divided into three tiers by first fitting log wages to a regression of province, language, race, and marital status dummies, and age and education dummies fully interacted. A predicted nonunion wage is then constructed from all these variables except the province dummies (see section 3.3.2). This predicted wage is used to separate workers into three skill groups (lower tier, middle tier, and upper tier) of approximately equal sizes.[20]

Table 3.3 reports average nonunion wages, union wages, and unionization rates for the three tiers of the sample of Canadian men (columns 1–3). It also presents OLS estimates of the average union wage gap for each tier. The OLS estimates reported in column 4 are obtained by regressing log hourly wages on a dummy variable for union coverage, province and marital status dummies, age and education dummies fully interacted, and an extensive set of job characteristics available in the LMAS data.[21] Table 3.4 reports analogous estimates for the sample of Canadian women.

19. Card (1992) uses the large samples of the 1987–88 matched files of the CPS and divides his sample of men aged 24–66 in five quintiles. The Canadian sample is divided in three tiers only to improve the precision of the estimates (the Canadian sample is much smaller than the U.S. sample).

20. Dividing the sample in three tiers is simply one method among others to let the effects of unions vary over workers with different skill levels. A different but related method is used by Simpson (1985), who divides workers on the basis of the skill requirements of the occupation they hold. In addition, since workers are divided into tiers on the basis of a *predicted* wage, estimates by tier are not biased, while they would be if workers were divided into tiers on the basis of their *actual* wage (the dependent variable).

21. These job characteristics include a part-time dummy, seven occupation dummies, three tenure dummies, and four firm-size dummies.

Table 3.3 OLS and First-Differenced Wage Gap Estimates by Tier for Men in Canada

	Nonunion Wage (1)	Union Wage (2)	Unionization Rate (3)	Wage Gap Estimates[a] OLS (4)	Wage Gap Estimates[a] First-Differenced (5)	Selection Bias (6)
			Public and private pooled			
Tier 1	2.010	2.407	38.2	0.232 (0.011)	0.207 (0.035)	0.025
Tier 2	2.312	2.534	50.3	0.129 (0.011)	0.220 (0.038)	−0.091
Tier 3	2.597	2.713	50.5	0.044 (0.013)	0.006 (0.059)	0.038
All	2.305	2.550	46.3	0.133 (0.007)	0.163 (0.024)	−0.030
			Private sector only			
Tier 1	2.016	2.417	34.9	0.223 (0.012)	0.190 (0.038)	0.033
Tier 2	2.309	2.547	45.8	0.131 (0.012)	0.229 (0.040)	−0.098
Tier 3	2.577	2.678	35.9	0.049 (0.017)	0.076 (0.088)	−0.027
All	2.300	2.547	39.1	0.139 (0.008)	0.162 (0.026)	−0.023
			Public sector only			
Tier 1	1.950	2.366	59.2	0.308 (0.033)	0.302 (0.084)	0.006
Tier 2	2.346	2.488	76.6	0.106 (0.029)	0.161 (0.101)	−0.055
Tier 3	2.703	2.743	77.6	0.028 (0.022)	0.074 (0.092)	−0.046
All	2.332	2.532	73.6	0.111 (0.015)	0.166 (0.050)	−0.055

Notes: Based on 18,679 observations (14,773 in the private sector, 3,906 in the public sector) divided into tiers on the basis of the predicted nonunion wage (see text). The public sector accounts for 13.5 percent of employment in tier 1, 14.6 percent in tier 2, and 20.9 percent in tier 3.

[a]Both OLS and first-differenced estimates are obtained by fitting log hourly wage regressions that also include controls for age, education, marital status, part-time status, tenure, firm size, industry, and occupation. The first-differenced estimates are based on a sample of 1,559 involuntary job changers (744 in tier 1, 480 in tier 2, and 335 in tier 3).

The results show the same patterns that were observed in figures 3.1 and 3.2. In the case of men, the union wage gap declines in the skill level while the unionization rate first increases and then remains constant in the middle and upper tiers. The results for women indicate that the union wage gap slowly declines in the skill level. For both men and women, public sector unionization increases in the skills of workers, while private sector unioniza-

Table 3.4 OLS and First-Differenced Wage Gap Estimates by Tier for Women in Canada

	Nonunion Wage (1)	Union Wage (2)	Unionization Rate (3)	Wage Gap Estimates[a] OLS (4)	Wage Gap Estimates[a] First-Differenced (5)	Selection Bias (6)
			Public and private pooled			
Tier 1	1.748	2.150	27.0	0.229	0.190	0.039
				(0.012)	(0.037)	
Tier 2	1.904	2.273	33.6	0.230	0.077	0.153
				(0.013)	(0.051)	
Tier 3	2.165	2.538	53.4	0.178	0.191	−0.013
				(0.015)	(0.054)	
All	1.934	2.319	37.9	0.212	0.167	0.045
				(0.008)	(0.026)	
			Private sector only			
Tier 1	1.736	2.078	17.3	0.200	0.146	0.054
				(0.015)	(0.045)	
Tier 2	1.889	2.241	19.8	0.231	0.152	0.079
				(0.017)	(0.066)	
Tier 3	2.103	2.391	17.0	0.170	0.320	−0.150
				(0.027)	(0.107)	
All	1.908	2.236	18.2	0.206	0.186	0.020
				(0.011)	(0.037)	
			Public sector only			
Tier 1	1.814	2.217	56.4	0.299	0.267	0.032
				(0.024)	(0.058)	
Tier 2	1.987	2.296	65.1	0.246	−0.022	0.268
				(0.020)	(0.074)	
Tier 3	2.288	2.558	75.0	0.187	0.153	0.034
				(0.017)	(0.108)	
All	2.028	2.355	68.4	0.231	0.150	0.081
				(0.011)	(0.035)	

Notes: Based on 16,086 observations (9,788 in the private sector, 6,298 in the public sector) divided into tiers on the basis of the predicted nonunion wage (see text). The public sector accounts for 24.8 percent of employment in tier 1, 30.4 percent in tier 2, and 62.8 percent in tier 3.

[a]Both OLS and first-differenced estimates are obtained by fitting log hourly wage regressions that also include controls for age, education, marital status, part-time status, tenure, firm size, industry, and occupation. The first-differenced estimates are based on a sample of 1,268 involuntary job changers (552 in tier 1, 362 in tier 2, and 354 in tier 3).

tion is concentrated in the middle of the skill distribution. The average union wage gaps are similar in the public and in the private sector.

3.4.3 Longitudinal Data for Fixed Effect Estimation

A data set like the LMAS, containing detailed information on work histories of individuals, has several advantages over standard panel data sets, such as matched CPS's, for estimating the union wage gap by fixed effect methods. A first advantage is that it is known from the work history whether a worker changed jobs. This information reduces the odds of misclassification errors in recorded changes in the union status, since union status changes *only* for job changers. Intuitively, observing a change in union status is not surprising when it is known that the worker has changed jobs. By contrast, observing a change in union status is surprising for a worker who has not changed jobs.[22] The probability that a recorded union status change for a non–job changer is due to misclassification errors is thus high. Since true job changes are infrequent events, a large number of the recorded changes in union status are likely to be spurious when job movers and job stayers are pooled. This problem is avoided by limiting the longitudinal analysis to workers who are known to have changed jobs.[23]

A second advantage of the LMAS is that it records the reason a worker changed jobs. It is thus possible to separate workers who quit their jobs voluntarily from workers who did not. In the presence of endogenous job search, fixed effect estimates based on a sample of voluntary quitters are likely to be biased. It is thus useful to estimate the model separately for involuntary job changers to see whether the results are robust to the choice of sample.

To be classified as a job changer, a worker has to hold consecutive jobs for two different employers over the 1986–87 period. The job changer also has to work at least four weeks on each of these jobs. On the one hand, workers holding two jobs simultaneously for more than a week are not classified as job changers. On the other hand, workers who are recorded to hold two jobs simultaneously during the transition week are also classified as job changers, to account for the possibility of job changes during the transition week as opposed to over the weekend. Finally, job changers are divided into a sample of voluntary quitters and involuntary changers on the basis of their response to the question, "What was the main reason . . . left that job or business?"[24] A sample of 5,200 job changers, including 2,826 involuntary changers and 2,374 voluntary quitters, were selected on the basis of their answer to that question. A panel of two jobs is available for both type of job changers.

3.4.4 Fixed Effect Estimates by Tier

The fixed effect estimates of the union wage gap are reported in column 5 of tables 3.3 and 3.4. The estimates are obtained by fitting to the sample of involuntary job changers a first-differenced version of the regressions used to

22. True transitions would only occur when the job became organized or decertified, which is a very unlikely event.

23. See Krueger and Summers (1988) for some evidence on this point in the context of estimating interindustry wage differentials.

24. Voluntary quitters left their job for one of the following reasons: low pay, no opportunity of advancement, no opportunity to use training or skills, working conditions, other reasons for which they were dissatisfied, or a decision to quit for no particular reason.

compute the OLS wage gaps (column 4). The first-differenced regressions also include a dummy variable indicating whether the second job was recorded in the 1987 LMAS, as opposed to the 1986 LMAS, to account for growth in log wages between 1986 and 1987. To improve the precision of the results, the first-differenced wage gap estimates in the private and public sectors are obtained by fitting a regression for the pooled sample in which the union coverage variable is interacted with a public sector dummy (a public sector dummy is also included separately). Note that, since there are only two observations per worker, first-differenced estimates are equivalent to standard within estimates.

The first-differenced wage gap estimates for men reported in table 3.3 are always *larger* than the OLS estimates when the three tiers are pooled. On average, men holding union jobs are thus negatively selected in both the private and the public sectors. The selection bias (the difference between the OLS and the first-differenced wage gap estimates) is reported in column 6. There is also some evidence that men in the lower tier are positively selected, while men in the middle and upper tiers are negatively selected in both private and public sector union jobs. It is nevertheless clear that the selection-adjusted wage gaps decline with skill. The selection mechanism only accentuates this pattern.

The results reported in table 3.4 indicate that, unlike men, women holding union jobs are positively selected in both the public and the private sector. As in the case of men, the selection is negative for lower-tier women and for upper-tier women working in the public sector. Unlike men, however, middle-tier women and upper-tier women in the private sector are positively selected into the union sector. Overall, the selection-adjusted estimates reinforce the conclusion that there is little systematic relationship between the union wage gap and the skill level of women in Canada.

The differences in the pattern of wage differentials for men and women can be restated in terms of selection-adjusted returns to skills in the union and the nonunion sectors. These returns to skills are calculated as the difference between the predicted wage of an average worker in the upper tier and the predicted wage of an average worker in the lower tier. These predicted wages $\hat{w}^N(G)$ and $\hat{w}^U(G)$ for tier G are defined as $\hat{w}^N(G) = \bar{w}(G) - \bar{U}(G)\Delta_w(G)$ and $\hat{w}^U(G) = \hat{w}^N(G) + \Delta_w(G)$, where $\bar{w}(G)$ is the average wage in tier G, $\bar{U}(G)$ is the unionization rate, and $\Delta_w(G)$ is the first-differenced wage gap estimate. Applying these formulas to the estimates reported in tables 3.3 and 3.4 yields an estimated return to skill for men of .37 in the union sector, and of .57 in the nonunion sector. The estimated return to skill is equal to .46 for women in both the union and the nonunion sector.

The union wage gap is thus the same for lower-tier and upper-tier women because the returns to skills for women are lower in the nonunion sector *and* higher in the union sector. Relative to men, the skills of women are thus more rewarded in the union than in the nonunion sector. This explains why high-skill women are relatively more likely than high-skill men to select the union sector. This pattern of self-selection is even stronger in the public sector, suggesting that public sector unions play a very different role for men than they do for women. Controlling for observables, unionized jobs in the public sector seem to attract relatively skilled women and relatively unskilled men. A potentially fruitful area of research would be to explore how differences in both wages and benefits packages, such as maternity leaves, make unionized public sector jobs particularly attractive to high-skill women.

The validity of these findings relies heavily, however, on the assumption that first-differenced wage gap estimates for the sample of involuntary job changers are consistent estimates of the true wage gap. Since first-differenced wage gap estimates are overidentified, it is possible to perform specification tests of these estimates. One straightforward test is to compare the wage gap estimates for union joiners and union leavers. These two wage gaps are estimated by interacting the union coverage variable with a dummy variable indicating whether the worker is a union joiner or a union leaver, and then fitting a first-differenced version of that enlarged wage equation.

The wage gap estimates for union joiners and union leavers are reported in columns 2 and 3 of appendix table 3C.1. The wage gap estimates for joiners and leavers are very similar, especially for men, which suggests the first-differenced model for the sample of involuntary leavers is well specified. The table also shows additional evidence of robustness of the main findings by presenting estimates for the sample of all job changers and for a sample of dual-job holders.

Most of the results presented in table 3C.1 are more directly comparable to the results of Card (1992), as they include demographic and location characteristics but not job characteristics in the wage equations being fitted.[25] Figure 3.3 compares the pattern of the union wage gap and of the unionization rate for men in Canada and the United States. The Canadian wage gap estimates are taken from column 1 of table 3C.1. The U.S. wage gap estimates are Card's (1992) wage gap estimates by quintile averaged in three tiers. The figure indicates similar patterns of selection-adjusted wage differentials in the two countries. The figure also indicates the unionization rates in the two countries diverge at the high end of the skill distribution, as was discussed in section 3.3. Note also that Card finds no evidence of selection bias, on average. He finds some evidence of positive selection at the lower end and negative selection at the upper end of the skill distribution, but the two effects cancel out in the aggregate.

3.4.5 Estimates of the Union Variance Gap in Canada

Of the main components of equations (6), (7′), and (8), only the variance gap $\Delta_v(x)$ remains to be estimated. One estimator of the variance gap is the

25. The covariates used are thus the regressors used in the predicted nonunion wage equation on the basis of which workers are divided in tiers (or quintiles).

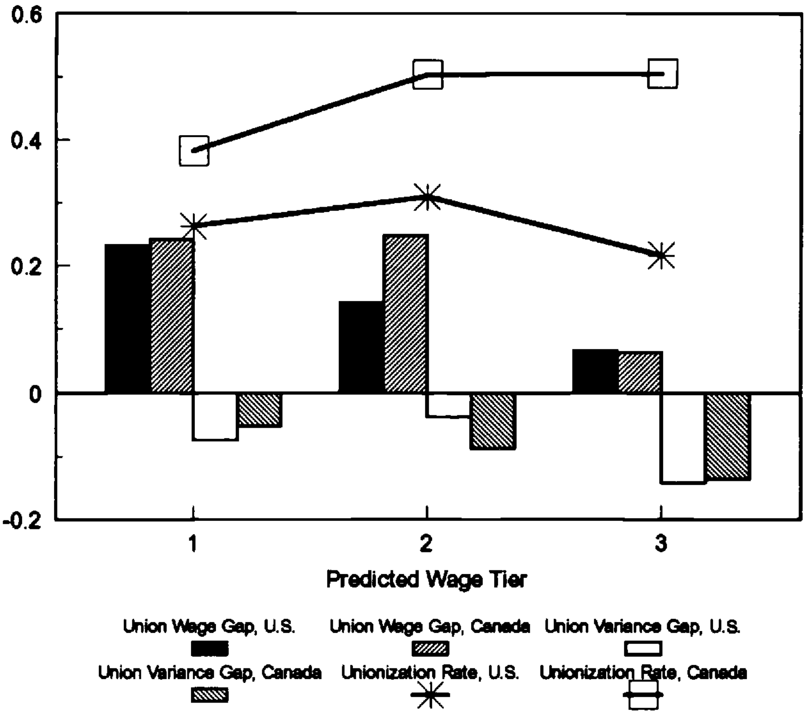

Fig. 3.3 Union wage effects and unionization rates by predicted wage tier for men in Canada and the United States
Sources: Rows 1, 4, and 5 of tables 3.6 and 3.8.

difference between the cross-sectional variance of wages in the union and in the nonunion sector. One problem with that approach is that it fails to distinguish whether unions reduce the variance of wages from whether union workers are more homogeneous than nonunion workers (Freeman 1984). This cross-sectional estimator is thus potentially afflicted by selectivity biases. One alternative panel data estimator of the variance gap that is not afflicted by selectivity biases is obtained by contrasting the change in the variance of wages of union joiners to the change in the variance of wages of nonunion stayers:

$$(V_2^{01} - V_1^{01}) - (V_2^{00} - V_1^{00}),$$

where V_t^{jk} is the variance of wages on job t (t = 1,2) among workers with union history $U_{i1} = j$ and $U_{i2} = k$. Another estimator is obtained by comparing union leavers to union stayers:

$$-[(V_2^{10} - V_1^{10}) - (V_2^{11} - V_1^{11})].$$

Both of these estimators of the variance gap are consistent but inefficient. The efficient longitudinal estimator of the variance gap is obtained by fitting a weighted linear regression of the change in union status (0,1, −1,0) to the change in the variance of wages ($V_2^{00} - V_1^{00}$, $V_2^{01} - V_1^{01}$, $V_2^{10} - V_1^{10}$, $V_2^{11} - V_1^{11}$) for the four union histories. Although this longitudinal estimator has clear advantages over the cross-sectional estimator, it may still be biased if the sample of job changers is small and unrepresentative. Both the cross-sectional and the longitudinal variance gap estimates are thus potentially biased. Both estimates will be presented below.

Changes in the variance of wages, $V_2^{jk} - V_1^{jk}$, for each of the four union histories 00, 01, 10, and 11 are reported in columns 1–4 of table 3.5. The efficient longitudinal estimate of the variance gap is reported in column 5, while the difference in cross-sectional variances is reported in column 6. The results for men indicate that the variance of wages of union joiners (01) decreases when they join the union sector, while the variance of wages of union leavers (10) increases after they leave the union sector. The estimated variance gap is −0.50 for the whole sample of men, which is smaller than the cross-sectional estimate of −0.134. In addition, the estimated union effect is larger (in absolute value) for upper-tier than for lower-tier workers. For women, the longitudinal variance gap estimate (−0.029) is also smaller than the cross-sectional variance gap estimate (−0.064).

On the one hand, these results reject the view that the variance of wages is lower in the union than in the nonunion sector simply because union workers are more homogeneous than nonunion workers. On the other hand, the longitudinal estimates of the variance gap are less than half of the cross-sectional variance gaps.[26] Correcting for selection biases thus has a bigger impact on the variance gap estimates than on the wage gap estimates reported in tables 3.3 and 3.4. There is some evidence, however, that part of the discrepancy between the cross-sectional and the longitudinal estimates of the variance gap is due to the composition of the sample of involuntary job changers. While the cross-sectional estimates of the variance gap are equal to −.134 for men and −.064 for women in the full sample, they are equal to only −.077 and .033 in the sample of involuntary changers. Furthermore, the preferred estimate of the variance gap in Lemieux (1992) is closer to the cross-sectional variance gap for the full sample than to the longitudinal variance gap for the sample of involuntary changers.[27] The cross-sectional estimates of the variance gap (table 3.5, column 6) will thus be used to calculate the overall impact of unions on the variance of wages.

The estimated variance gaps for men in Canada and the United States are

26. This was also noted by Swidinsky and Kupferschmidt (1991).

27. Lemieux (1992) also finds that the composition of the sample of involuntary job changers does not significantly affect the longitudinal estimates of the union wage gap (effect on level of wages). The composition problem occurs because unions flatten the returns to the permanent component of unobservable characteristics and the dispersion of these unobservable characteristics is small among job changers. The flattening effect thus reduces the variance of wages of job changers by less than it reduces the variance of wages of all union workers.

Table 3.5 Change in Variance of Wages for Job Changers in Canada

	Change in Variance by Union History				Estimates of the Union Variance Gap	
	00 (1)	01 (2)	10 (3)	11 (4)	Longitudinal[a] (5)	Cross-sectional (6)
			A. Men			
Public and private sectors						
Tier 1	−0.020	0.009	0.039	0.007	−0.018	−0.052
Tier 2	−0.009	−0.033	0.034	0.029	−0.033	−0.088
Tier 3	0.003	−0.027	0.078	−0.025	−0.057	−0.136
All tiers	−0.011	−0.019	0.074	0.007	−0.050	−0.134
Private sector only						
All tiers	−0.001	−0.004	0.040	−0.004	−0.024	−0.134
Public sector only						
All tiers	−0.107	−0.123	0.210	0.059	−0.181	−0.208
			B. Women			
Public and private sectors						
Tier 1	−0.018	−0.047	0.036	−0.068	−0.042	−0.037
Tier 2	0.031	−0.040	0.049	0.042	−0.044	−0.094
Tier 3	−0.001	−0.122	−0.063	−0.092	−0.024	−0.097
All tiers	−0.002	−0.051	0.008	−0.064	−0.029	−0.064
Private sector only						
All tiers	−0.005	0.038	0.006	−0.063	0.019	−0.054
Public sector only						
All tiers	0.010	−0.180	0.005	−0.075	−0.071	−0.133

[a]Estimated by fitting a weighted linear regression of the change in union status to the change in variance of wages for the four union histories. See text for more details.

compared in figure 3.3 (the estimates for the United States are in row 5 of table 3.8). As in the case of the wage gap, the estimated variance gaps follow similar patterns in Canada and in the United States. In both countries, the estimated variance gaps tend to be larger for high-skill men than for low-skill men.

3.5 The Overall Impact of Unions on Wage Inequality

This section uses the fixed effect, or selection-adjusted, estimates to calculate the impact of unions on the overall variance of wages. Following the discussion in section 3.2, the effect of unions on the overall variance of wages can be divided in three parts: (1) the effect of unions on the relative position of each skill group in the wage distribution; (2) the effect of unions on the between-sector variance of wages in a skill group, averaged over skill groups; and (3) the effect of unions on the within-sector variance of wages in a skill group, averaged over skill groups. The formulas (6), (7′), and (8) can be thus be used directly to compute these effects by replacing the general skill categories x by an index G for the three tiers defined above (G = lower tier, middle tier, and upper tier). These formulas depend on the nonunion wage $w^N(G)$, the unionization rate $\bar{U}(G)$, the union wage gap $\Delta_w(G)$, the union variance gap $\Delta_v(G)$, and the unadjusted wage gap $\bar{\Delta}_w(G)$. For Canada, estimates of $\bar{U}(x)$, $\Delta_w(x)$, $\Delta_v(x)$, and $\bar{\Delta}_w(x)$ are available from tables 3.3–3.5 and 3C.1, while $w^N(G)$ is obtained from the formula $w^N(G) = \bar{w}(G) - \bar{U}(G)\Delta_w(G)$ ($\bar{w}(G)$ is the average wage in the tier).

The calculations of the overall impact of unions on the variance of wages in Canada are reported in table 3.6 for men and in table 3.7 for women. For both men and women, the effect of unions on the within-sector variance of wages (row 6) is smaller than the effect of unions on the between-sector variance (row 7) in the lower tier. The reverse holds in the middle tier and the upper tier. Unions thus reduce the within-tier variance of wages by 0.029 for men and by 0.003 for women (row 8). In the case of men, unions also reduce the between-tier (across skill groups) variance of wages by 0.011 (row 9). The total impact of unions on the variance of wages is thus equal to -0.040, which represents a 14.5 percent reduction in the overall variance of wages.

As mentioned before, the impact of unions on the relative position of the

Table 3.6 Effects of Unions on Wage Inequality: Men in Canada

	Tier 1	Tier 2	Tier 3	All
1. Unionization rate ($\bar{U}_G$)	38.2	40.3	50.5	46.3
2. Mean log wage				
Nonunion	2.010	2.312	2.597	2.305
Union	2.407	2.534	2.713	2.550
Unadjusted wage gap ($\bar{\Delta}_{wG}$)	0.397	0.222	0.116	0.245
3. Standard deviation of log wages				
Nonunion	0.436	0.448	0.532	0.530
Union	0.371	0.335	0.383	0.383
4. Estimated union wage gap (Δ_{wG}, table 3C.1, col. 1)	0.242	0.248	0.064	—
5. Estimated union variance gap (Δ_{vG}, table 3.5, col. 6)	−0.052	−0.088	−0.136	—
Effect of unions on within-tier variance				
6. Effect on within-sector variance (row 1) * (row 5)	−0.020	−0.044	−0.069	−0.044
7. Effect on between-sector variance ($\bar{U}_G(1 - \bar{U}_G)[\bar{\Delta}^2_{wG} - (\bar{\Delta}_{wG} - \Delta_{wG})^2]$)	0.032	0.012	0.003	0.015
8. Total effect (row 6 + row 7)	0.012	−0.032	−0.066	−0.029
Effect of unions on between-tier variance				
9. ($\text{Var}_G(\bar{U}_G\Delta_{wG}) + 2\text{Cov}_G(w^N_G,\bar{U}_G\Delta_{wG})$)	—	—	—	−0.011
Total effect on variance of wages				
10. (row 8 + row 9)	—	—	—	−0.040

Table 3.7 Effects of Unions on Wage Inequality: Women in Canada

	Tier 1	Tier 2	Tier 3	All
1. Unionization rate (U_G)	27.0	33.6	53.4	26.4
2. Mean log wage				
Nonunion	1.748	1.904	2.165	1.934
Union	2.150	2.273	2.548	2.319
Unadjusted wage gap ($\bar{\Delta}_{wG}$)	0.402	0.369	0.373	0.375
3. Standard deviation of log wages				
Nonunion	0.395	0.437	0.496	0.466
Union	0.345	0.312	0.386	0.392
4. Estimated union wage gap (Δ_{wG}, table 3C.1, col. 1)	0.245	0.205	0.264	—
5. Estimated union variance gap (Δ_{wG}, table 3.5, col. 6)	−0.037	−0.094	−0.097	—
Effect of unions on within-tier variance				
6. Effect on within-sector variance (row 1) * (row 5)	−0.010	−0.032	−0.052	−0.031
7. Effect on between-sector variance ($\bar{U}_G(1 - \bar{U}_G)[\bar{\Delta}^2_{wG} - (\bar{\Delta}_{wG} - \Delta_{wG})^2]$)	0.027	0.024	0.031	0.028
8. Total effect (row 6 + row 7)	0.017	−0.008	−0.021	−0.003
Effect of unions on between-tier variance				
9. ($\mathrm{Var}_G(\bar{U}_G\Delta_{wG}) + 2\mathrm{Cov}_G(w^N_G, \bar{U}_G\Delta_{wG})$)	—	—	—	0.013
Total effect on variance of wages				
10. (row 8 + row 9)	—	—	—	0.009

tiers in the overall wage distribution is very different for men and women. For women, the union wage gap is more or less stable across tiers, and upper-tier women are disproportionately represented in the union sector. As a result, unions worsen the relative position of lower-tier women and increase the between-tier variance of wages by 0.013 (table 3.7, row 9). Overall, unions thus increase the variance of wages among women by 0.009, which represent 4.1 percent of the overall variance of wages. The finding that unions reduce the variance of wages of men but increase the variance of wages of women is robust to the choice of estimator of the variance gap. If longitudinal estimates of the variance gap were used instead of cross-sectional estimates, the estimated effect of unions on the variance of wages would become −0.013 (instead of −0.040) for men and 0.022 (instead of 0.009) for women.

The estimates can also be used to compute the overall effect of unions on the variance of wages of Canadian men and women pooled together. This effect depends on (1) the effect of unions on the variance of wages within men and within women, and (2) the effect of unions on the wage differential between men and women. The first component of the overall effect is simply the weighted sum of the effects reported in row 10 of tables 3.6 and 3.7. It is equal to -0.017, which represents a 6.1 percent reduction in the variance of wages. The second, or "between," component is given by

$$[s_m(\bar{w}_m - \bar{w})^2 + s_w(\bar{w}_w - \bar{w})^2] - [s_m(w^N_m - w^N)^2 + s_w(w^N_w - w^N)^2],$$

where s_m and s_w are the proportion of men and women in the work force, and $\bar{w}$ is the average wage for men and women. This component is equal to -0.002, which indicates that unions slightly improve the position of women relative to men in the wage distribution. Overall, unions thus reduce the variance of wages for Canadian men and women by 0.019 (0.017 plus 0.002).

Finally, the results for men in the United States are reported in table 3.8. These results are obtained by transforming the estimates reported in table 8 of Card (1992) by quintiles into estimates by tier. The relationship between the estimates by tier ($T1$ to $T3$) and the estimates by quintile ($Q1$ to $Q5$) is given by the weighted averages $T1 = .6Q1 + .4Q2$, $T2 = .2Q2 + .6Q3 + .2Q4$, and $T3 = .4Q4 + .6Q5$. As in the case of men in Canada, the effect of unions on the within-sector variance (-0.020) is larger than the effect on the between-sector variance (0.009). Unions thus reduce the average within-

Table 3.8 **Effects of Unions on Wage Inequality: Men in the United States**

	Tier 1	Tier 2	Tier 3	All
1. Unionization rate ($\bar{U}_G$)	26.2	30.9	21.7	26.4
2. Mean log wage				
Nonunion	1.973	2.276	2.636	2.274
Union	2.325	2.480	2.602	2.460
Unadjusted wage gap ($\bar{\Delta}_{wG}$)	0.352	0.204	−0.034	0.186
3. Standard deviation of log wages				
Nonunion	0.446	0.483	0.523	0.568
Union	0.362	0.342	0.362	0.380
4. Estimated union wage gap	0.232	0.142	0.067	—
5. Estimated union variance gap	−0.075	−0.038	−0.142	—
Effect of unions on within-tier variance				
6. Effect on within-sector variance (row 1) * (row 5)	−0.019	−0.010	−0.031	−0.020
7. Effect on between-sector variance ($\bar{U}_G(1 - \bar{U}_G)[\bar{\Delta}^2_{wG} - (\bar{\Delta}_{wG} - \Delta_{wG})^2]$)	0.021	0.009	−0.003	0.009
8. Total effect (row 6 + row 7)	0.002	−0.000	−0.034	−0.009
Effect of unions on between-tier variance				
9. ($\mathrm{Var}_G(\bar{U}_G\Delta_{wG}) + 2\mathrm{Cov}_G(w^N_G,\bar{U}_G\Delta_{wG})$)	—	—	—	−0.010
Total effect on variance of wages				
10. (row 8 + row 9)	—	—	—	−0.019

Notes: The estimates were obtained by transforming the estimates reported in Card (1992) by quintiles into three tiers. The relationship between the estimates by tier ($T1$ to $T3$) and the estimates by quintile ($Q1$ to $Q5$) is given by the following weighted averages: $T1 = .6Q1 + .4Q2$, $T2 = .2Q2 + .6Q3 + .2Q4$, and $T3 = .4Q4 + .6Q5$.

tier variance of wages by 0.011. Unions also reduce the between-tier variance of wages by 0.008, for a total effect of −0.019, or 6.3 percent of the overall variance of wages.

3.6 Unions and Relative Wage Inequality in Canada and the United States

The results reported in table 3.8 indicate that unions reduce the variance of men's wages in the United States by 0.019, which is half of the estimated effect for Canada (0.040, table 3.6). The difference is mostly attributable to the larger effect of unions on the within-tier variance in Canada than in the United States. Authors such as Freeman (1991) and Card (1992) have argued that a significant fraction of the increase in wage inequality in the United States over the last two decades is attributable to the decline of unionism in the United States. Does the Canadian evidence support the view that wage inequality among men would be lower in the United States if American unions were "as strong" as Canadian unions? To answer this question, consider what would happen to wage inequality in the United States if the Canadian, as opposed to the U.S., distribution of unionism was to prevail, holding constant the U.S. wage structure. Alternatively, consider what would happen to wage inequality in Canada if the U.S., as opposed to the Canadian, distribution of unionism was to prevail, holding constant the Canadian wage structure. The results of these experiments are reported in table 3.9.

The first row of table 3.9 indicates that there is a gap of 0.050 between the actual variance of wages of men in Canada and in the United States. Row 4 indicates that if the extent of unionization in the United States were the same as in Canada, this gap would be reduced to 0.030. The gap would also be reduced to 0.030 if the extent of unionization in Canada was the same as in the United States (row 3). It would be reduced to 0.029 if there were no unions in either Canada or the United States (row 2). Taken together, these results suggest that differences in the pattern and extent of unionism in Canada and in the United States explain 40 percent of the difference in wage inequality of men between the two countries.

The results reported in column 1 also indicate that the variance of wages of Canadian women, unlike men, would be essentially unchanged if the unionization rate was the same as in the United States. This result is consistent with the overall finding that unions have a small, though positive, effect on the variance of wages of women.

The evidence from the Canada-U.S. comparison for men thus yields similar conclusions to the longitudinal comparison between the United States in the 1970s and in the late 1980s (Card 1992, and Freeman 1991). These studies find that deunionization in the United States between 1973 (or 1978) and 1987 accounts for 20 percent of the increase in wage inequality over that period. The unionization rate was relatively constant in Canada over the same period.

Table 3.9 **Relative Impact of Unions on Wage Inequality in Canada and in the United States**

	Women	Men		
	Canada (1)	Canada (2)	U.S. (3)	Difference between U.S. and Canada (4)
1. Actual variance of wages	0.228	0.234	0.284	0.050
2. Variance of wages that would prevail in the absence of unions[a]	0.219	0.274	0.303	0.029
3. Variance of wages that would prevail with U.S. unionism[b]	0.226	0.254	0.284	0.030
4. Variance of wages that would prevail with Canadian unionism[c]	0.228	0.234	0.264	0.030

[a]Actual variance of wages minus the estimated effect of unions on the variance of wages (tables 3.6–3.8, row 10).

[b]Actual variance of wages minus the effect of unions on the variance of wages calculated by replacing the actual unionization rates by the U.S. unionization rates in row 1 of table 3.5. The U.S. unionization rates for women are calculated from the 1986 CPS data used in tables 3.1 and 3.2 (11.7 percent in the lower tier, 15.4 percent in the middle tier, and 20.1 percent in the upper tier).

[c]Actual variance of wages minus the effect of unions on the variance of wages calculated by replacing the actual unionization rates by the Canadian unionization rates in row 1 of table 3.6.

Can changes in unionization rates between Canada and the United States explain the finding by Blackburn and Bloom (chap. 7 in this volume) that inequality in earnings increased by 0.034 in the United States but only by 0.018 in Canada over the 1979 to 1986 period? Although this paper does not provide direct evidence on that question, some back-of-the-envelope calculations can be made by combining some results from Riddell (chap. 4 in this volume) with the main findings of this paper. Table 4.1 in Riddell shows that the U.S. union density fell by 6 points relative to the Canadian density from 1980 to 1986. These 6 points represent a third of the gap in unionization rates between the two countries in 1986. Since the gap in unionization rates explains 0.020 of the gap in the variance of wages, a third of the unionization rate gap must explain a third of 0.020 (0.006 to 0.007). This represents 40 to 45 percent of the relative increase in earnings inequality of 0.016 (0.034 − 0.018) reported by Blackburn and Bloom. The strength of the union movement in Canada thus seems to be a major factor in explaining why wage inequality did not increase as quickly in Canada as it did in the United States.

From a social welfare perspective, these benefits of unionization do not necessarily come at no cost. As mentioned in section 3.2, unions may also cause efficiency losses by raising wages above their competitive level. Standard calculations indicate these losses are of the order of 0.2 percent of GNP in the United States and 0.5 percent of GNP in Canada.[28] These costs are small and would be even smaller if labor contracts were negotiated efficiently.[29] They nevertheless illustrate the tradeoff Canada would face if it were to move to more "U.S.-like" labor market institutions. GNP per capita would increase by 0.3 percent, but the variance of wages of men would increase by 8.5 percent (40 percent of 0.050/0.234).

3.7 Conclusion

The recent divergence in the extent of unionism in Canada and in the United States yields a unique opportunity to measure the impact of unionism on the distribution of wages using a comparative perspective. The major findings of the paper are the following:

1. Union relative wage effects are similar in Canada and in the United States. In the case of men, the union wage differential is negatively related to skills. This negative relationship is much less accentuated for women.

2. Private sector unionization is concentrated in the middle of the skill distribution, while public sector unionization is concentrated in the upper end of the skill distribution. This explains why unionization in Canada and among women is more skewed toward the upper end of the skill distribution.

3. The selection process into unionized jobs is different for men and women in Canada. For women, the permanent unobservable component of wages is positively correlated with the union status, while it is negatively correlated with the union status for men. This is particularly true in the public sector. There is no evidence of selection bias (on average) for men in the United States.

4. Unions reduce the within-sector variance of wages for both men and women.

5. Unions reduce the overall variance of wages by 14.5 percent for men in Canada and by 6.3 percent for men in the United States, but they increase the variance of wages of Canadian women by 4.1 percent. Differences in the pattern and extent of unionism in Canada and in the United States explain 40 percent of the difference in wage inequality of men between the two countries.

28. The efficiency losses computed over the three tiers are equal to $\Sigma\theta_G(.5\eta_G\Delta^2_{wG})$, where θ_j is a weight that represents the fraction of the total wage bill that goes to union workers in tier G ($\theta_G \approx [\bar{w}(G)/\bar{w}]\bar{U}(G)/3$). The labor demand elasticities η_G chosen for the calculations are .5 in the upper tier, .75 in the middle tier, and 1 in the lower tier.

29. The efficiency loss in the monopoly model of union occurs because the negotiated outcome is not Pareto efficient. This result is very sensitive, however, to the assumption that unions cannot bargain over employment. Labor contracts are said to be efficient when the firm and the union bargain over wage and employment simultaneously. Under the strong version of efficient contracts (Brown and Ashenfelter 1986), the negotiated wage is purely an instrument to redistribute rents between the parties. Unions cause neither efficiency losses nor employment distortions.

These findings shed new light on the role of unions in the relative distribution of wages of men and women in Canada and in the United States. More remains to be learned, however, on why unions have such a different impact on the wage distribution of men and women. A more thorough analysis of the role of unions in the provision of nonwage benefits such as maternity leaves and the role of unions in promoting wage equity in the workplace could shed considerable light on these issues. It would also be interesting to measure more directly the impact of unions on changes in wage inequality in Canada during the eighties.

Appendix A
Data

A Comparison of the LMAS and CPS Samples

The wage data used for Canada and the United States are based on supplements to very similar labor force surveys (the LFS in Canada and the CPS in the United States). The structures of the supplements are quite different, however. The Canadian LMAS is based on a work history that asks workers about all the jobs they held during the previous year. By contrast, the outgoing rotation group supplement of the CPS asks people about the job they held during the week of the survey. In both surveys, the earnings questions refer to usual, as opposed to actual, earnings and hours. The earnings in the LMAS may nevertheless be more noisy than in the CPS because of the recall bias problem.

On the one hand, the sampling frame for *jobs* (but not individuals) is different in the two samples, since the CPS is only a snap shot while the LMAS captures all the jobs held during the year. Short-duration jobs are thus more likely to be captured in the LMAS. On the other hand, the LMAS sample used in the cross-sectional analysis is limited to one job per person, and to jobs lasting at least four weeks. These sample selection criteria reduce the probability of sampling a short-duration job and thus make the LMAS sample more comparable to the CPS sample.

Another difference between the two samples is that earnings are top coded at 999$ a week in the CPS, while there is essentially no top coding in the LMAS.[30] In addition, only unallocated wages are used in the CPS, while all wages are used in the LMAS because there are no allocation flags in the LMAS.

The nature of the longitudinal data used is also quite different in the two

30. The LMAS user's guide indicates that "two records with total earnings from all jobs in 1986 in excess of $150,000 have had their hourly wage rates reduced to values which yield totals close to 150,000$."

surveys. The longitudinal CPS sample is obtained by matching people interviewed twice in one year (rotation groups 4 and 8). The matching is imperfect, and the measurement error in changes in the union status variable is substantial for the reasons mentioned in section 3.4. Card (1992) handles the measurement error problem by using additional information from the CPS validation study. The longitudinal LMAS sample is discussed in the main text.

Description of the Variables Used

The public use sample of the LMAS contains only bracketed information on age and education. This explains why the continuous version of these variables is not used in the analysis. The CPS age and education variables were grouped in these five categories to make them comparable with the LMAS data. The seven industry categories used in table 3.2 are primary industries, manufacturing, construction, transportation and communication, trade, services, and government (including health and education). The eight occupation categories are managers, professionals, nurses, clerical workers, sales workers, service workers, manual workers, and craft workers. The first six categories are considered white-collar workers, while the last two categories are considered blue-collar workers.

Appendix B
A Regression-based Approach to Analyze the Patterns of Unionization and Wages

Figures 3.1 and 3.2 indicate how $\bar{U}(x)$ and $\Delta_w(x)$ depend on a particular function of worker's characteristics, namely the estimated nonunion wage index. Similar findings are obtained using a regression-based approach that describes how $w^N(x)$, $\bar{U}(x)$, and $\Delta_w(x)$ jointly depend on worker, job, and location characteristics. Most of this analysis is limited to the case of Canada.

Table 3B.1 reports the differences in nonunion wages, unionization rates, and the union wage gap associated with changes in various individual characteristics of workers, when all other characteristics are held constant at their sample mean. The first row of the table gives average values of the three outcome variables by gender. Subsequent rows show the deviations from the overall means associated with a particular characteristic (e.g., age 20–24) *holding constant* all other characteristics. For simplicity, these deviations are called excess predicted nonunion wages, union rates, or union wage gaps.

The entries in table 3B.1 are calculated on the basis of separate wage regressions fit to the union and nonunion sectors. These regressions include the explanatory variables listed in table 3B.1 plus a full set of interactions between age and education dummies. The pattern of results for age and educa-

Table 3B.1 Excess Nonunion Wage, Excess Union Density, and Excess Wage Gap by Demographic Characteristics and Job Characteristics in Canada

	Men			Women		
	Excess Nonunion Wage (1)	Excess Density (2)	Excess Wage Gap (3)	Excess Nonunion Wage (4)	Excess Density (5)	Excess Wage Gap (6)
Average	2.361	0.463	0.136	2.006	0.379	0.211
	(0.005)	(0.003)	(0.007)	(0.005)	(0.003)	(0.008)
Age						
20–24	−0.233	−0.063	0.066	−0.127	−0.034	0.014
	(0.012)	(0.010)	(0.019)	(0.009)	(0.008)	(0.016)
25–34	−0.039	−0.002	0.017	0.029	0.008	−0.010
	(0.007)	(0.005)	(0.010)	(0.006)	(0.005)	(0.009)
35–44	0.102	0.011	−0.038	0.051	0.018	−0.012
	(0.008)	(0.006)	(0.010)	(0.007)	(0.005)	(0.010)
45–54	0.101	0.023	−0.039	0.022	0.004	0.006
	(0.011)	(0.008)	(0.015)	(0.010)	(0.008)	(0.016)
55–64	0.014	0.027	0.015	−0.054	−0.028	0.041
	(0.017)	(0.012)	(0.024)	(0.016)	(0.013)	(0.028)
Education						
Primary	−0.159	−0.014	0.041	−0.140	−0.031	0.012
	(0.014)	(0.010)	(0.020)	(0.016)	(0.014)	(0.031)
High school	−0.053	0.007	0.018	−0.057	−0.001	−0.011
	(0.005)	(0.003)	(0.006)	(0.004)	(0.003)	(0.006)
More than high school	0.007	0.007	−0.019	−0.025	−0.006	0.010
	(0.013)	(0.010)	(0.020)	(0.013)	(0.011)	(0.024)
Some postsecondary	0.070	0.002	0.006	0.069	0.006	0.020
	(0.011)	(0.008)	(0.014)	(0.009)	(0.007)	(0.013)
University	0.254	−0.018	−0.094	0.216	0.019	0.001
	(0.013)	(0.010)	(0.018)	(0.013)	(0.010)	(0.019)
Mother tongue						
English	0.007	−0.011	−0.001	0.006	−0.005	−0.005
	(0.005)	(0.004)	(0.007)	(0.005)	(0.004)	(0.007)
French	−0.011	0.011	0.011	−0.003	0.014	0.020
	(0.012)	(0.009)	(0.016)	(0.011)	(0.009)	(0.017)
Others	−0.010	0.023	−0.014	−0.017	−0.005	−0.015
	(0.014)	(0.010)	(0.018)	(0.012)	(0.010)	(0.006)
Race						
White	0.004	0.002	0.000	−0.002	0.002	0.006
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
Nonwhite	−0.060	−0.032	−0.007	0.024	−0.031	−0.093
	(0.014)	(0.016)	(0.030)	(0.020)	(0.016)	(0.031)
Marital status						
Single	−0.080	−0.011	0.043	−0.008	−0.006	0.000
	(0.008)	(0.006)	(0.010)	(0.006)	(0.005)	(0.009)
Married	0.028	0.004	−0.015	0.004	0.003	−0.000
	(0.003)	(0.002)	(0.004)	(0.003)	(0.002)	(0.004)

Table 3B.1 (continued)

	Men			Women		
	Excess Nonunion Wage (1)	Excess Density (2)	Excess Wage Gap (3)	Excess Nonunion Wage (4)	Excess Density (5)	Excess Wage Gap (6)
Province						
Newfoundland	−0.109	0.008	−0.067	−0.112	0.053	0.003
	(0.017)	(0.012)	(0.022)	(0.017)	(0.013)	(0.024)
Prince Edward Island	−0.131	−0.032	−0.019	−0.097	−0.022	0.030
	(0.023)	(0.018)	(0.033)	(0.021)	(0.017)	(0.032)
Nova Scotia	−0.093	−0.046	−0.013	−0.075	−0.043	−0.041
	(0.014)	(0.011)	(0.020)	(0.013)	(0.011)	(0.021)
New Brunswick	−0.090	−0.037	0.026	−0.104	−0.039	0.051
	(0.014)	(0.010)	(0.019)	(0.013)	(0.010)	(0.020)
Quebec	−0.019	0.095	0.020	0.014	0.079	−0.018
	(0.010)	(0.007)	(0.013)	(0.010)	(0.007)	(0.013)
Ontario	0.046	−0.014	−0.008	0.032	−0.060	−0.001
	(0.008)	(0.006)	(0.010)	(0.007)	(0.006)	(0.011)
Manitoba	−0.068	−0.034	0.059	−0.019	−0.001	0.009
	(0.015)	(0.011)	(0.021)	(0.014)	(0.010)	(0.020)
Saskatchewan	0.019	−0.001	−0.014	−0.006	0.043	0.024
	(0.014)	(0.010)	(0.018)	(0.012)	(0.009)	(0.018)
Alberta	0.073	−0.083	−0.020	0.068	−0.028	−0.018
	(0.010)	(0.007)	(0.014)	(0.009)	(0.007)	(0.014)
British Columbia	0.095	0.082	0.022	0.054	0.053	0.008
	(0.014)	(0.009)	(0.017)	(0.012)	(0.009)	(0.017)
Part-time status						
Full-time	0.007	0.004	−0.007	−0.002	0.010	−0.004
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.004)
Part-time	−0.118	−0.072	0.125	0.006	−0.027	0.011
	(0.016)	(0.013)	(0.025)	(0.007)	(0.005)	(0.010)
Firm size						
Less than 20	−0.116	−0.213	0.051	−0.090	−0.172	−0.001
	(0.007)	(0.006)	(0.013)	(0.006)	(0.005)	(0.013)
20–99	−0.054	−0.086	0.021	−0.008	−0.024	0.002
	(0.009)	(0.007)	(0.013)	(0.009)	(0.007)	(0.014)
100–499	0.020	0.047	−0.011	0.044	0.087	−0.018
	(0.012)	(0.008)	(0.015)	(0.012)	(0.008)	(0.015)
500 and more	0.099	0.124	−0.049	0.084	0.121	−0.022
	(0.007)	(0.004)	(0.009)	(0.007)	(0.005)	(0.010)
Don't know	−0.020	0.073	0.026	−0.032	0.027	0.061
	(0.011)	(0.007)	(0.014)	(0.010)	(0.007)	(0.014)
Tenure (years)						
Less than 1	−0.134	−0.084	0.024	−0.131	−0.076	0.039
	(0.009)	(0.007)	(0.014)	(0.008)	(0.006)	(0.013)
1–5	−0.030	−0.045	−0.002	−0.034	−0.032	0.009
	(0.006)	(0.005)	(0.009)	(0.005)	(0.004)	(0.008)
5 and more	0.062	0.054	−0.006	0.098	0.068	−0.028
	(0.005)	(0.003)	(0.006)	(0.005)	(0.004)	(0.008)

(continued)

Table 3B.1 (continued)

	Men			Women		
	Excess Nonunion Wage (1)	Excess Density (2)	Excess Wage Gap (3)	Excess Nonunion Wage (4)	Excess Density (5)	Excess Wage Gap (6)
Occupation						
Managers	0.136	−0.253	−0.048	0.169	−0.168	−0.073
	(0.011)	(0.009)	(0.018)	(0.015)	(0.012)	(0.026)
Professional	0.029	−0.072	−0.006	0.065	0.037	0.050
	(0.015)	(0.011)	(0.021)	(0.019)	(0.012)	(0.026)
Nurses	−0.060	0.024	−0.067	0.165	0.055	−0.088
	(0.059)	(0.029)	(0.066)	(0.020)	(0.012)	(0.026)
Clerical	−0.064	0.010	−0.048	0.027	−0.055	−0.046
	(0.018)	(0.012)	(0.023)	(0.012)	(0.009)	(0.019)
Sales	−0.003	−0.150	−0.033	−0.057	−0.090	0.001
	(0.015)	(0.013)	(0.030)	(0.016)	(0.014)	(0.037)
Service	−0.161	−0.047	0.055	−0.198	−0.026	0.079
	(0.014)	(0.010)	(0.019)	(0.013)	(0.010)	(0.022)
Manual workers	−0.030	0.096	0.029	−0.073	0.353	0.101
	(0.008)	(0.006)	(0.012)	(0.092)	(0.068)	(0.151)
Craft workers	0.061	0.132	−0.014	−0.038	0.387	0.165
	(0.011)	(0.008)	(0.015)	(0.119)	(0.090)	(0.190)
Industry						
Primary	0.165	−0.139	−0.061	0.194	−0.288	0.069
	(0.017)	(0.014)	(0.028)	(0.034)	(0.031)	(0.097)
Manufacturing	0.029	−0.033	−0.036	0.017	−0.251	−0.078
	(0.013)	(0.009)	(0.018)	(0.055)	(0.041)	(0.090)
Construction	0.052	−0.108	0.138	0.126	−0.225	−0.176
	(0.020)	(0.016)	(0.031)	(0.030)	(0.028)	(0.126)
Transportation and communication	−0.008	0.042	0.052	0.078	0.122	0.094
	(0.015)	(0.010)	(0.019)	(0.022)	(0.014)	(0.029)
Trade	−0.096	−0.159	0.015	−0.100	−0.172	0.055
	(0.010)	(0.009)	(0.018)	(0.010)	(0.008)	(0.022)
Services	−0.071	−0.129	−0.061	−0.027	−0.164	−0.031
	(0.016)	(0.013)	(0.029)	(0.011)	(0.008)	(0.020)
Public sector	0.015	0.301	−0.001	0.040	0.256	0.012
	(0.016)	(0.011)	(0.023)	(0.011)	(0.008)	(0.017)

Sources: Data from the 1986 cross-sectional file and the 1986–87 longitudinal file of the LMAS.
Note: The excess nonunion wage for a given value of a characteristic is the difference between the predicted wage of a worker with that value of the characteristic, holding all other characteristics at their observed frequency distributions, and the mean wage in the sample.

tion is similar to the pattern uncovered in figures 3.1 and 3.2. The excess wage gap is inversely proportional to the excess nonunion wage (skills) for men, but remains more or less constant for women. Table 3B.1 also shows the relationship among the excess predicted values of $w^N(x)$, $U(x)$, and $\Delta_w(x)$ for a variety of other worker, job, and location characteristics.

The estimates in table 3B.1 are summarized in table 3B.2 by looking at the impact of unions on a selected number of wage differentials. Similar regressions are estimated using the CPS sample for the United States except that mother tongue is not included in those regressions. These union wage gaps are calculated by adding the average wage gap to excess wage gaps like the ones reported in table 3B.1 (the wage gaps used are similar but not identical to those in table 3B.1, since the underlying regressions are more parsimonious).

The results indicate that unions reduce the university–high school differential for men in Canada and the United States, but increase it for women in Canada. This is simply a restatement in regression terms of the findings illustrated in figures 3.1 and 3.2 (unions help more high-skill women than low-skill women, especially in Canada). Unions also tend to help public sector workers and blue-collar workers relative to private sector and white-collar workers. Unions also revert the small wage disadvantage of French speakers relative to English speakers in Canada, mostly because of the high unionization rates in Quebec. Finally, unions increase the wage differential between

Table 3B.2 Effect of Unions on Selected Wage Differentials

	Canada		U.S.	
	Men (1)	Women (2)	Men (3)	Women (4)
Mature workers (35–44) versus young workers (20–24)				
Wage differential	0.399	0.265	0.378	0.280
Effect of unions	−0.035	−0.005	0.027	0.025
University graduates versus high school graduates				
Wage differential	0.393	0.353	0.399	0.412
Effect of unions	−0.054	0.076	−0.065	0.003
White versus nonwhite				
Wage differential	0.098	−0.023	0.147	0.071
Effect of unions	0.022	0.046	−0.019	−0.021
White-collar versus blue-collar				
Wage differential	−0.042	0.008	0.056	0.096
Effect of unions	−0.040	0.003	−0.060	−0.044
Public sector versus private sector				
Wage differential	0.032	0.117	−0.003	0.073
Effect of unions	0.044	0.132	0.032	0.051
English-speaking versus French-speaking				
Wage differential	0.023	0.027	—	—
Effect of unions	−0.037	−0.057	—	—

Notes: These effects are found by estimating separate wage regressions for union and nonunion workers: the covariates used are the age dummies, the education dummies, the region dummies, and dummy variables for marital status, part-time status, race, blue-collar, public sector, and mother tongue (only for Canada). The regressions are used to calculate a union wage gap Δ_w for each category of worker listed in the table. The effect of unions on the wage differential between two groups A and B is simply $\bar{U}_A\Delta_{wA} - \bar{U}_B\Delta_{wB}$.

whites and nonwhites in Canada. This goes in opposite direction to what is typically found in the United States (here and in Ashenfelter 1972).

Appendix C
Robustness of the First-Differenced Estimates

Detailed first-differenced estimates for union joiners and union leavers are reported in columns 2 and 3 of table 3C.1 for men and women in Canada. The results reported in columns 1–5 are obtained by fitting first-differenced regressions that do not include controls for job characteristics. More precisely, consider the following wage equation in which control variables are omitted for the sake of clarity:

$$w_{it} = \gamma_t + U_{it}\Delta_w + \theta_i + \varepsilon_{it},$$

where γ_t is a time effect, θ_i is a time-invariant person-specific effect (fixed effect), and Δ_w is the union wage gap. The first-differenced version of this equation is

$$\Delta w_{it} = \Delta\gamma_t + \Delta U_{it}\Delta_w + \Delta\varepsilon_{it}.$$

OLS estimates of the first-differenced version of the wage equation yield consistent estimates of Δ_w even when θ_i is correlated with the union coverage variable U_{it}. Separate wage gap estimates Δ_w^{01} for union joiners ($U_{i1} = 0$ and $U_{i2} = 1$) and Δ_w^{10} for union leavers ($U_{i1} = 1$ and $U_{i2} = 0$) are obtained by fitting the following regression:

$$\Delta w_{it} = \Delta\gamma_t + U_i^{01}\Delta_w^{01} - U_i^{10}\Delta_w^{10} + \Delta\varepsilon_{it},$$

where U_i^{01} is an indicator variable equal to one for union joiners, while U_i^{10} is an indicator variable equal to one for union leavers.

Column 6 reproduces the estimates that were reported in tables 3.3 and 3.4, with job characteristics included as regressors. Job characteristics are not used in columns 1–5 for the sake of comparison with the results of Card (1992), and because there are few degrees of freedom available in the small samples of union joiners, union leavers, and dual-job holders. On the one hand, it is preferable to include job characteristics in the regression when we try to measure the union wage gap for the same kind of workers holding the same kind of jobs. This may be particularly important when involuntary job leavers are concentrated in few particular industries, for example, because of industrial restructuring. On the other hand, jobs are choice variables, and it is not clear they should be included in the definition of *skills* used for the analysis by tier. In any case, whether or not job characteristics are included in the regressions does not affect the substance of the results.

Table 3C.1 Robustness of First-Differenced Estimated for Canada

	Demographic Controls Only					Demographic and Job Controls
	Involuntary Job Changers					
	All (1)	Union Joiners[a] (2)	Union Leavers[a] (3)	All Changers (4)	Dual-Job Holders (5)	All Involuntary Job Changers (6)
A. Men						
Public and private pooled						
Tier 1	0.242	0.209	0.270	0.243	0.219	0.207
	(0.034)	(0.054)	(0.048)	(0.025)	(0.079)	(0.035)
Tier 2	0.248	0.288	0.217	0.221	0.580	0.220
	(0.036)	(0.058)	(0.051)	(0.029)	(0.082)	(0.038)
Tier 3	0.064	0.060	0.068	0.064	0.636	0.006
	(0.055)	(0.090)	(0.079)	(0.040)	(0.106)	(0.059)
All tiers	0.204	0.202	0.207	0.196	0.376	0.163
	(0.034)	(0.037)	(0.033)	(0.017)	(0.052)	(0.024)
Private sector only						
All tiers	0.207	0.201	0.206	0.210	0.380	0.162
	(0.026)	(0.042)	(0.036)	(0.019)	(0.067)	(0.026)
Public sector only						
All tiers	0.191	0.195	0.205	0.126	0.347	0.166
	(0.050)	(0.074)	(0.072)	(0.039)	(0.090)	(0.050)
Observations	1,559	1,559		2,789	425	1,559
B. Women						
Public and private pooled						
Tier 1	0.245	0.254	0.236	0.238	0.319	0.190
	(0.035)	(0.053)	(0.052)	(0.026)	(0.057)	(0.037)
Tier 2	0.205	0.198	0.212	0.197	0.371	0.077
	(0.046)	(0.070)	(0.070)	(0.032)	(0.097)	(0.051)
Tier 3	0.264	0.380	0.157	0.254	0.255	0.191
	(0.049)	(0.077)	(0.073)	(0.036)	(0.054)	(0.054)
All tiers	0.240	0.281	0.200	0.233	0.308	0.167
	(0.025)	(0.038)	(0.037)	(0.018)	(0.038)	(0.026)
Private sector only						
All tiers	0.256	0.244	0.223	0.224	0.416	0.186
	(0.035)	(0.052)	(0.055)	(0.024)	(0.063)	(0.037)
Public sector only						
All tiers	0.197	0.289	0.147	0.204	0.182	0.150
	(0.035)	(0.052)	(0.048)	(0.026)	(0.049)	(0.035)
Observations	1,268	1,268		2,789	425	1,268

[a]These estimates are obtained by estimating a first-differenced version of a wage equation in which the union coverage variable is interacted with dummy variables for whether the worker is a union leaver or a union joiner. There are 141 union joiners and 188 union leavers among the sample of men, while there are 137 union joiners and 145 union leavers among the sample of women.

In the case of men, the wage gap estimates for union joiners (0.202) and union leavers (0.207) are nearly identical when all skill groups are pooled. The estimated wage gaps differ more substantially for women (0.281 and 0.200). First-differenced estimates for the sample of all job changers, as opposed to involuntary changers only, are reported in column 4. The wage gap estimates for that sample are very similar to the estimates for the sample of involuntary changers only. Finally, the first-differenced procedure is applied to an alternative sample of dual-job holders. This sample consists of workers who hold two jobs simultaneously over a period of at least four weeks at any time in 1986–87. The job on which the worker usually spends the most hours per week is classified as the main job; the other job is classified as the secondary job. An alternative wage gap estimate is thus obtained by fitting the difference in wages on the two jobs to the differences in the characteristics of the two jobs, including the union coverage status. The results are reported in column 5.

Consider the results for all tiers together. The wage gap estimates based on the sample of dual-job holders are larger (0.376 for men, 0.308 for women) than any of the wage gap estimates based on samples of job changers. They also imply that men are negatively selected into the union sector, while women are positively selected into the union sector, as was found in tables 3.3 and 3.4. More research is nevertheless needed to explain why these wage gap estimates are as large as they are.

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