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Union Wage Differentials in the Public and Private Sectors: A Simultaneous Equations Specification

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The paper attempts to integrate new approaches to estimating union wage effects with the analysis of public-private sector wage differentials. Estimates of the union differential in both public and private sectors, allowing for the endogeneity of union status, are presented. The hypothesis that the recently measured rents to public sector employment primarily reflect the recent increase in unionization in that sector is examined, and receives considerable empirical support. There was evidence of positive selection into the union sector, especially for private sector workers. Union status appears to be strongly influenced by the expected wage gain from joining the union sector.

I. Introduction

The estimation of wage differences associated with unionism has reemerged recently as a controversial topic. A general consensus, based implicitly or explicitly on some form of the monopoly theory of unionism, appeared to have been achieved in the 1960s. Gregg Lewis (1963) summarized and reanalyzed the large body of U.S. literature on the wage

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effects of unions. His finding of a wage difference on the order of 10%–15% represented the consensus view of the effects of unions. Since that time, however, as a result of studies emphasizing the endogenous nature of union status the consensus has disappeared. These studies were recently reviewed by Parsley (1980) and by Freeman and Medoff (1982). The key studies in this area include those by Ashenfelter and Johnson (1972), Duncan and Stafford (1981), Schmidt (1978), and Lee (1978). Both Ashenfelter and Johnson (1972) and Schmidt (1978) demonstrate that the positive effect of unions on wage rates found in single-equation studies can be eliminated in a simultaneous equations framework. Lee (1978) postulated an explicit model of individual choices regarding union status and used this to "correct" for selectivity bias in separate union and nonunion wage equations. The correction reduced the estimated effect of unions, but only by a modest 2 percentage points.

This paper seeks to integrate the new approach to unions with the analysis of public-private sector wage differentials. The literature on public sector wage effects is not as large as that on unions. However, because of the increasing size of the public sector, the importance of this topic is growing. Recent studies have been conducted by Ashenfelter and Ehrenberg (1975), Smith (1977), and Borjas (1980) for the United States and by Cousineau and Lacroix (1977), Gunderson (1979*a*, 1979*b*), and Auld et al. (1980) for Canada. One of the major findings is that the public sector appears to have a less elastic demand curve for labor. Gunderson (1979*a*, 1979*b*) reports that rents are being earned in the Canadian public sector, especially by women and less skilled workers. Similar premia associated with public sector employment are reported by Smith (1977) for the United States.

A natural question that arises from consideration of these recent results is whether the public sector rents merely reflect the usual union wage effects. This is the first study that allows for the determination of union sector status in estimating both potential union-nonunion *and* public-private sector wage differentials. In addition, the hypothesis that the recently measured rents to public sector employment are simply due to greater unionization in that sector is examined. Finally, this is also the first study to provide selectively corrected estimates of union wage differentials for Canada based on individual data.[1]

The plan of the paper is as follows. In Section II the determinants of union status are considered and the union wage differential is estimated

[1] As a by-product of their study of migration, Grant and Vanderkamp (1980) produce an estimate of the payoff to union membership using individual micro data. However, their definition of union status differs from the conventional one by including members of professional organizations. Starr (1973) analyzes wages by establishment in Ontario. The few remaining studies (Kumar 1972; Maki and Christensen 1979; MacDonald and Evans 1981) employ aggregate industry data.

on the assumption that public sector status is exogenous. Estimates of the union wage differential for Canada, which may be compared with recent work on individual data sets for the United States, are presented. These show large differentials, especially for male workers. Union differentials in the public sector are broadly similar to those in the private sector. Thus the hypothesis that union premia are larger in the public sector because of a less elastic demand curve receives little support. Section III examines the public-private sector differential. Controlling for union status there is little evidence of public sector rents. However, because of the high proportion of unionized workers in the public sector, a substantial positive premium is estimated when union status is not controlled for. This provides evidence for the hypothesis that recently estimated rents accruing to public sector workers are in fact union differentials. Finally, Section IV summarizes the conclusions of the analysis.

II. Union-Nonunion Wage Differentials and the Probability of Union Status

If unions are in a monopoly position they must limit entry, thus producing a queue. In order to observe an individual in a union two criteria must be satisfied. First, the individual must have chosen to enter the queue; second, the individual must have been picked out of the job queue by the unionized firm. Because of the lack of information on the relevant populations for the two selection stages (e.g., which workers offered themselves to the union sector but were not chosen, etc.), previous authors have collapsed the process into a single criterion for union status. In this paper we follow the same procedure. Following Lee (1978) we assume that worker *i* has a "reservation wage," ρ_i, for joining the unionized sector. The worker joins the union if

$$\frac{W_{U_i} - W_{N_i}}{W_{N_i}} > \rho_i, \tag{1}$$

where W_{U_i} and W_{N_i} are the wage rates that worker *i* receives in the union and nonunion sectors, respectively. The reservation wage, ρ_i, combines both the monetary costs of unionization implicitly or explicitly borne by the worker and the value an individual attaches to being inside or outside a union because of differences other than the wage levels in the two sectors.

The costs of unionization are assumed to be high for employees with high turnover rates. Thus Lewis (1959) argues that unionism is more likely to be successful for full-time male workers than for females or casual workers. To the extent that unions restrict entry by non-price-rationing other individual characteristics, determining the attractiveness of the individual to the union or unionized firm will affect the implicit cost to the worker. These are assumed to include education, experience, and sex. Since different regulations governing union-management relations differ across provinces, geographical location may also be expected to influence the costs of unionization. If, as has been argued by Ashenfelter and Johnson (1972) and others, union services are normal goods, ρ_i will be smaller the higher the level of the spouses' income. Finally, assuming economies of scale in organizing workers, the costs of unionization will be smaller for workers attached to industries with large plant sizes.

The discussion of the previous paragraph may be summarized by the following criterion for union status:

$$\frac{W_{U_i} - W_{N_i}}{W_{N_i}} \simeq (\ln W_{U_i} - \ln W_{N_i}) > \mathbf{X}_i\beta + \epsilon_i, \tag{2}$$

where $\mathbf{X}_i$ is a vector of characteristics of the individual as follows: region, schooling, experience, language, sex, part-time worker, marital status, income of spouse, and average plant size of the industry in which the individual works.[2] The disturbance ϵ_i represents unobservable characteristics of the individual that affect ρ_i. These are assumed to follow a normal distribution with zero mean and variance σ_ϵ^2. Given this assumption the choice between union and nonunion status may be put in a standard probit form. An individual will be a union member if $I_i > 0$ where

$$I_i = (\ln W_{U_i} - \ln W_{N_i}) - \mathbf{X}_i\beta - \epsilon_i. \tag{3}$$

Wage rates in the union and nonunion sectors, respectively, are assumed to be given by the standard semilogarithmic forms

$$\ln W_{U_i} = \mathbf{X}_{U_i}\gamma_U + \epsilon_{U_i} \tag{4}$$

and

$$\ln W_{N_i} = \mathbf{X}_{N_i}\gamma_N + \epsilon_{N_i}, \tag{5}$$

where $\mathbf{X}_{N_i}$ is a vector of regressors measuring schooling, experience, tenure with the firm, language, sex, and region. Wage coefficients in the union sector are permitted to be different than wage coefficients in the nonunion sector, and the union wage is also permitted to depend on the

[2] The average was used rather than the actual plant size where the individual worked because the latter is endogenous.

percentage of workers that are organized by unions in the industry.[3] Thus $\mathbf{X}_{U_i}$ contains all the regressors in $\mathbf{X}_{N_i}$, plus the percentage of organized workers in the industry as a measure of union power (Rosen 1969). The disturbances in (4) and (5) are assumed to be normal: $\epsilon_U \sim N(0, \sigma_U^2)$; $\epsilon_N \sim N(0, \sigma_N^2)$.

The statistical structure of the model contained in (3)–(5) parallels that of Lee (1978). Estimation procedures for models of this form are by now very familiar and thus will be described only very briefly.[4] First, the reduced-form probability of being in a union is estimated. This follows from substituting for union and nonunion wage rates in (4) and (5) into (3). The results of this stage are used to compute a selectivity correction factor (inverse Mills ratio) for inclusion in the wage regressions (4) and (5), which have to be estimated on censored populations (i.e., either union or nonunion). Finally, the "corrected" wage equation parameter estimates are used to compute a predicted union-nonunion wage difference for each individual, which permits the direct estimation of the structural probit (3).

Union differentials computed in this way are potentially sensitive to the normality assumptions for ϵ_{U_i} and ϵ_{N_i}. If the primary interest is estimating the union differential rather than the process determining union status, the normality assumption may be dispensed with as follows. The separate sector wage equations (4) and (5) may be combined into a wage equation for all workers:

$$\ln W_i = \mathbf{X}_i\gamma_N + \mathbf{X}_i U_i \delta + (\epsilon_{N_i} + \epsilon_i^* U_i), \tag{6}$$

where U_i is a dummy variable equal to one for union members, $\delta = \gamma_U - \gamma_N$, and $\epsilon_i^* = \epsilon_{U_i} - \epsilon_{N_i}$. If normality for ϵ_{U_i} and ϵ_{N_i} fails, the union differential may be estimated by directly estimating (6). This contains the dummy endogenous variable U_i. Following Heckman (1978) this may be consistently estimated by using an instrument, $\hat{U}_i$, for U_i obtained from a simple linear probability model. This approach was used to assess the reliability of the estimates obtained under the normality assumption.[5]

The process determining union status, as noted earlier, is in principle a two-stage process with individuals willing to be chosen and employers

[3] Assuming that the supply of labor to the nonunion sector is infinitely elastic, the level of organization will not affect nonunion wages. In the empirical work we tested for this by including the level of organization in the nonunion wage equation. Its coefficient in that equation was always insignificantly different from zero.

[4] For more details, see Lee (1978) for an application to union choice, Willis and Rosen (1979) for choice of education level, and Robinson and Tomes (1982) for migration decisions.

[5] We are indebted to James Heckman for this suggestion.

(or unions and employers together) choosing from among these individuals. While the former may be the same for both public and private sectors, the latter may well differ if, as is often hypothesized, the objective function of public sector employers differs from that of private sector employers. In addition, there is some evidence (Gunderson 1979*a*) to suggest that the payoffs to individual characteristics such as sex, schooling, and experience are different in the public and private sectors. Hence, the model is estimated separately for public and private sectors.

The variables used in the analysis are defined in table 1. The data source is the Social Change in Canada Survey for 1979. This records union and public sector status of individuals. It also contains a preferred wage measure. This measure consists of a two-part question: first an amount of pay is asked; second the period over which the pay applies is ascertained. Thus for those that give an hour as the pay period a direct observation on the hourly wage rate is available. The sample was restricted to these individuals for two reasons. First, many other studies have confined themselves to "production workers." For comparability with these studies, most production workers are likely to be hourly paid workers. Second,

Table 1
Data Definitions, Means, and Standard Deviations

Name	Definition	Mean	Standard Deviation
Atlantic	Dummy variable = 1 if resident in Atlantic province	.061	.009
Quebec	Dummy variable = 1 if resident in Quebec	.320	.469
Ont	Dummy variable = 1 if resident in Ontario	.410	.492
Prairies	Dummy variable = 1 if resident in Prairies*	.108	.310
Yrssch	Years of schooling	11.375	3.146
Expr	Experience = Age − Yrssch − 6	17.812	14.496
ExprSq	Square of Expr	527.086	670.249
Tenure	Number of years with same firm	5.516	7.389
French	Dummy variable = 1 if French only spoken at home	.264	.441
Male	Dummy variable = 1 if male	.513	.500
POW	Percentage of organized workers in respondent's industry or sector	37.571	21.468
SpInc	Income of spouse in 1978 ($000)	1.545	3.502
Married	Dummy variable = 1 if married	.622	.485
Plantsize	Mean size of plant in individual's industry†	187.815	148.614
Parttime	Dummy variable = 1 if respondent works less than 30 hours/week	.259	.439
Skilled	Dummy variable = 1 if individual has SVP‡ score 7–9	.278	.448
Unskilled	Dummy variable = 1 if individual has SVP‡ score 1–3	.212	.409
Union	Dummy variable = 1 if individual belongs to a union	.454	.498
Public	Dummy variable = 1 if worked in government (federal, provincial, or local)	.234	.423
ln *W*	Natural logarithm of hourly wage rate	1.769	.519

* Reference group is British Columbia.
† Computed from respondents' answers to question on number of employees at respondent's place of work.
‡ Specific Vocational Preparation score. Reference group is semiskilled, SVP 4–6.

Borjas (1980) has recently shown that substantial errors are introduced when wages are obtained indirectly by dividing earnings by hours.[6]

Table 2 presents estimates of union and nonunion wage equations for public and private sector hourly paid workers. Columns 1 and 3 present the selectivity-corrected estimates for the private sector. In the union equation (col. 1) the selectivity correction, $\hat{\lambda}$, is positive and significant at the 5% level.[7] This implies that from the population of private sector

Table 2
Union and Nonunion Wage Equations for Public and Private Sector Hourly Paid Workers

Independent Variables	Dependent Variables							
	Private Sector				Public Sector			
	$\ln W_U$	$\ln W_U$	$\ln W_N$	$\ln W_N$	$\ln W_U$	$\ln W_U$	$\ln W_N$	$\ln W_N$
$\hat{\lambda}$	.2049	...	.3723	...	.1816	...	.1267	...
	(2.04)		(2.78)		(1.01)		(.41)	
Atlantic	−.4544	−.3000	−.2477	−.3652	−.3860	−.3408	−.6362	−.6988
	(2.76)	(2.02)	(1.38)	(2.15)	(3.45)	(3.40)	(1.93)	(2.43)
Quebec	−.5947	−.4661	−.1578	−.2767	−.1729	−.1204	−.3488	−.4117
	(4.49)	(4.06)	(.97)	(1.80)	(1.48)	(1.18)	(1.05)	(1.42)
Ont	−.2999	−.2298	−.1354	−.1911	−.1564	−.1443	−.0347	−.0872
	(4.01)	(3.58)	(1.09)	(1.57)	(1.97)	(1.90)	(.14)	(.43)
Prairies	−.3123	−.2195	−.1790	−.2587	−.2514	−.1908	−.3764	−.4533
	(3.02)	(2.37)	(1.25)	(1.88)	(1.91)	(1.63)	(1.18)	(1.79)
Yrssch	.0194	.0190	.0243	.0194	.0267	.0297	.0249	.0198
	(2.07)	(2.06)	(1.77)	(1.46)	(2.66)	(3.16)	(.93)	(.85)
Expr	.0191	.0118	.0150	.0235	.0092	.0097	.0337	.0321
	(2.66)	(1.91)	(1.76)	(3.06)	(1.57)	(1.68)	(2.65)	(2.69)
ExprSq	−.0004	−.0003	−.0003	−.0005	−.0002	−.0002	−.0007	−.0007
	(2.94)	(2.24)	(1.78)	(3.14)	(1.85)	(1.84)	(2.57)	(2.60)
Tenure	.0134	.0101	−.0032	.0036	.0073	.0049	−.0017	.0025
	(3.69)	(3.21)	(.47)	(.60)	(1.76)	(1.50)	(.09)	(.22)
French	.1781	.0996	−.0291	.0688	.0159	−.0442	.2910	.3248
	(1.58)	(.94)	(.22)	(.54)	(.16)	(.56)	(1.24)	(1.50)
Male	.3474	.2515	.1552	.3119	.2771	.2325	.0355	.0926
	(5.03)	(5.02)	(1.70)	(4.49)	(4.12)	(4.75)	(.18)	(.73)
POW	.0053	.0013	...	...	...	...	...	...
	(2.00)	(.71)						
Skilled	.1533	.1692	.1864	.1936	.1330	.1390	.2934	.2825
	(2.95)	(3.38)	(2.16)	(2.27)	(2.50)	(2.68)	(2.02)	(2.01)
Unskilled	−.0079	−.0240	−.1938	−.1484	−.0677	−.0544	−.2384	−.2592
	(.15)	(.47)	(2.31)	(1.83)	(1.05)	(.88)	(1.28)	(1.47)
Constant	1.2022	1.5923	1.1454	1.2684	1.4636	1.5173	1.1770	1.4188
	(4.63)	(9.27)	(4.80)	(5.51)	(8.18)	(9.02)	(1.57)	(3.10)
N	185	185	312	312	111	111	43	43
R^2	.4395	.4254	.1921	.1711	.4545	.4489	.6055	.6032
F	9.52	9.74	5.45	5.14	6.22	6.65	3.42	3.80

NOTE.—Absolute values of t-statistics, in parentheses, are corrected for the use of $\hat{\lambda}$ in place of the true value, λ, of the inverse Mill's ratio. ln W_U and ln W_N are the natural logarithms of the union and nonunion wage rates, respectively.

[6] These errors are particularly serious when the wage is used as a right-hand-side variable in a labor supply equation. This is the problem treated by Borjas (1980). However, in the present context, problems will also arise, for example, if union status is correlated with being hourly paid, which is true for the sample used in this paper and no doubt for most samples. (In fact the analyses reported here were also applied to a broader sample with similar results except that the estimated differentials were typically smaller.) The other criterion for sample entry was that data were recorded for all the variables used in the analysis.

[7] Hereafter, all references to statistically significant coefficients assume the conventional level of 5%.

workers, those who actually become union members have above-average union-sector-omitted wage-generating characteristics. Schooling has a positive and significant effect, which is comparable to the finding of other studies based on similar data. The experience coefficients are significant and yield the standard concave earnings-experience profiles. For a given experience level, tenure with the same firm has a significant positive effect. No significant language difference was discernible. Skilled workers obtain approximately 16% higher earnings than semiskilled workers. Unskilled workers, however, do not earn significantly less than semiskilled workers. Males earn 42% more than females, holding other characteristics the same. This appears to be a large differential compared with some other Canadian studies[8] but is similar to Lee's U.S. estimate (37%) which is the closest comparable study in methodology. There are significant regional differences in wage rates. The reference province is British Columbia, where wage levels are highest. Next highest are Ontario and the Prairies, followed by Quebec and the Atlantic provinces. This is the standard regional pattern for Canada (Ostry and Zaidi 1979, p. 376). Finally, there is a positive and significant effect for the level of union organization. The magnitude of the effect is comparable with that obtained by Lee (1978). In column 2 the uncorrected union wage equation is presented for comparison.

In the nonunion equation (col. 3), $\hat{\lambda}$ is also positive and significant. Thus those workers who enter the nonunion sector have higher than average nonunion-sector-omitted wage-generating characteristics. Thus workers choosing either union or nonunion status have a wage advantage, other characteristics the same, in the sector of their choice. In the nonunion sector the effect of schooling is higher and the skill differentials are larger than in the union sector. This result—the attenuation of skill differentials in the union sector—is similar to the findings of U.S. studies (Lee 1978). Regional effects have the same pattern but are substantially reduced in magnitude compared with the union case. The male wage differential is smaller in the nonunion sector (17% vs. 42%), closer to the estimates obtained by other authors (see n. 8). Again this result—a smaller sex differential in the nonunion sector—conforms to the pattern found by Lee for the United States. Finally, tenure effects are zero in the nonunion equation, in comparison with a significant positive effect in the union sector. Thus, while unions appear to reduce skill and educational wage differentials, they appear to exacerbate regional, male-female, and job tenure wage differentials. The uncorrected coefficients are presented for comparison in column 4.

The remaining columns in table 2 refer to the public sector.[9] However,

[8] Ostry (1968) and Gunderson (1979*b*) report differentials of approximately 20% for the labor market as a whole.

[9] Since the percentage of organized workers does not vary within the public sector this variable is dropped from the wage equations.

very few workers in the public sector are hourly paid workers, so that the sample sizes in the union and nonunion equations are quite small. Hence the results must be viewed with caution. In the union equation $\hat{\lambda}$ has a coefficient similar to that obtained for the private sector, but it is insignificant. This may be due to the small sample size or to a different mechanism of selection into unions in the public sector. All variables have the same qualitative effects as in the private sector union equation. The nonunion equation in the public sector was estimated on the smallest sample and hence must be viewed with special caution. The coefficient on $\hat{\lambda}$ is positive but insignificant. Again, the reason for this may be the small sample size. The main difference compared to the private sector is that there appears to be no male-female wage differential in the nonunionized public sector.

The relatively small sample sizes for the separate union and nonunion samples, especially for the public sector, make inferences regarding differences in the effects of unions across sectors suspect. In order to lessen this problem, since the basic structures of the wage equations in the public and private sectors appear similar, the two samples were pooled. Estimates from the pooled sample are reported in table 3. All slope coefficients in the pooled sample estimates were constrained to be equal with the exception of the difference between males and females. Since the separate sector regressions suggest a smaller male wage differential in the public sector, especially for the nonunion sector, an interaction between males and public sector status was included.

In both union and nonunion sectors the coefficients on $\hat{\lambda}$ are positive, as in table 2, and the significance level is increased. The additional variation in the percentage of organized workers caused by the addition of 154 public sector observations with high levels of organization increased the level of significance while the point estimate remained the same. The difference in schooling coefficients between the union and nonunion sectors is more marked than in the separate samples. The interaction between male and public sector status is negative in both equations but is only significant at the 20% level. The implied male differential in the public sector for unionized workers is about 30%, as in the estimates in table 2. In the nonunion public sector there is actually a small negative point estimate for the male wage differential compared with a small (but insignificant) positive differential from the table 2 estimates. The regional patterns remain the same across union and nonunion sectors as in table 2. Again the union sector has smaller wage differentials by skill level. Finally, the difference in public and private sector intercepts in both union and nonunion equations is insignificantly different from zero. Thus there is no evidence from the pooled sample of a significant additive public sector wage effect in the logarithmic wage equations.

The coefficients obtained from the wage equations in tables 2 and 3 were used to compute union-nonunion wage differentials for various groups. The results are presented in table 4. The first column presents the percentage union differential for an individual with characteristics corresponding to the averages for the whole sample. This was computed as $\phi_U = \exp\{(\hat{\gamma}_U - \hat{\gamma}_N)\bar{\mathbf{X}}\} - 1$, where $\hat{\gamma}_U$ and $\hat{\gamma}_N$ are the selectivity corrected estimates of γ_U and γ_N from table 2 and $\bar{\mathbf{X}}$ is the vector of average values of the independent variables for the whole sample. The second column (ϕ_U^P) presents analogous estimates using the coefficients from the pooled sample results of table 3. The remaining columns present the results using uncorrected coefficient estimates from tables 2 and 3. A standard decomposition of the overall union differential into the part attributed to differences in coefficients and the part attributed to differences in characteristics was also undertaken, but showed all the difference

Table 3
Union and Nonunion Wage Equations for All Hourly Paid Workers

Independent Variables	Dependent Variables			
	ln W_U	ln W_U	ln W_N	ln W_N
$\hat{\lambda}$	.2386	. . .	.3204	. . .
	(2.47)		(2.44)	
Atlantic	−.5062	−.3885	−.2762	−.3891
	(5.16)	(4.82)	(1.79)	(2.67)
Quebec	−.4449	−.3037	−.1533	−.2700
	(4.50)	(4.02)	(1.04)	(1.98)
Ont	−.2657	−.1987	−.1025	−.1662
	(4.50)	(4.11)	(.92)	(1.54)
Prairies	−.3224	−.2147	−.1749	−.2725
	(3.70)	(3.00)	(1.34)	(2.22)
Yrssch	.0206	.0237	.0294	.0222
	(2.92)	(3.65)	(2.40)	(1.89)
Expr	.0136	.0093	.0205	.0254
	(2.91)	(2.27)	(2.87)	(3.77)
ExprSq	−.0003	−.0002	−.0004	−.0005
	(3.15)	(2.48)	(2.70)	(3.77)
Tenure	.0114	.0077	−.0027	.0044
	(3.94)	(3.43)	(.44)	(.80)
French	.1087	.0130	−.0008	.0894
	(1.37)	(.20)	(.01)	(.80)
Male	.3718	.2561	.1758	.3123
	(5.61)	(5.59)	(2.01)	(4.78)
Public-male	−.0988	−.0500	−.2290	−.2423
	(1.31)	(.73)	(1.29)	(1.33)
Public	.0209	.0223	−.0225	.1979
	(.25)	(.28)	(.16)	(1.79)
POW	.0055	.0012	. . .	. . .
	(2.37)	(.81)		
Skilled	.1359	.1487	.2055	.2059
	(3.52)	(4.18)	(2.76)	(2.78)
Unskilled	−.0296	−.0412	−.1775	−.1511
	(.71)	(1.06)	(2.36)	(2.06)
Constant	1.1668	1.5419	1.0241	1.2032
	(5.83)	(12.57)	(4.68)	(5.91)
N	299	299	360	360
R^2	.4189	.4060	.2208	.2072
F	12.71	12.89	6.50	6.44

NOTE.—Absolute values of t-statistics in parentheses are corrected for the use of $\hat{\lambda}$ in place of the true inverse Mill's ratio.

due to differences in coefficients. Hence the results were similar to those reported in table 4.

The results in table 4 show relatively large union differentials. Using separate samples for the public and private sectors, the average private sector differentials are 40.76% for males and 15.81% for females. Weighted by the proportion of males and females unionized, these result in an overall private sector average of 34.57%. Using the pooled sample estimates the male and female averages are reduced to 33.78% and 9.97%, respectively, and the overall average to 27.83%. Using uncorrected coefficients the average differentials are reduced to some extent—by 2 percentage points and 8 percentage points in the pooled and nonpooled samples, respectively. More important, however, the pattern of differentials by sex is reversed. Using uncorrected estimates the differentials for females become larger than for males. The skill patterns are insensitive to pooling or correction. In all cases the union differentials are inversely related to skill level.

The union-nonunion differentials reported here are large relative to most previous Canadian studies. A recent survey of the Canadian studies in Gunderson (1982) shows most differentials in the range of 10%–23%

Table 4
Union-Nonunion Wage Differentials

	Corrected Coefficients		Uncorrected Coefficients	
	ϕ_u	ϕ_u^p	ϕ_u	ϕ_u^p
Private sector:				
Males:				
Unskilled	64.48	55.16	37.84	37.26
Semiskilled	36.57	33.83	21.71	22.97
Skilled	32.13	24.83	18.78	16.14
Average	40.76	33.78	24.14	23.89
Females:				
Unskilled	35.31	27.55	46.42	45.19
Semiskilled	12.36	10.01	29.29	30.08
Skilled	8.70	2.61	26.17	22.85
Average	15.81	9.97	31.86	31.10
Male and female average	34.52	27.83	26.07	25.69
Public sector:				
Males:				
Unskilled	65.47	84.58	57.81	36.49
Semiskilled	39.50	59.20	28.58	22.29
Skilled	18.82	48.50	11.39	11.74
Average	37.64	59.91	29.05	23.20
Females:				
Unskilled	29.95	33.20	37.20	21.91
Semiskilled	9.56	14.84	11.80	9.22
Skilled	−6.68	7.17	−3.15	3.15
Average	8.67	14.84	12.20	10.03
Male and female average	26.92	43.23	22.82	18.33

NOTE.—$\phi_u \equiv \exp\{(\hat{\gamma}_u - \hat{\gamma}_N)\bar{\mathbf{X}}\} - 1$ is the differential using coefficients from separate private and public sector samples; ϕ_u^p is analogously defined but uses coefficients from the pooled sample. $\bar{\mathbf{X}}$ is the average characteristics of the whole sample. Average skill levels use full sample weights. Male and female averages weight the differentials by the proportions of males and females unionized in the relevant sectors.

(Kumar 1972; Starr 1973; MacDonald and Evans 1981). Some larger differentials around 30% are noted in the study of Christensen and Maki (1981). A possible explanation for the difference is the different samples used in the other studies. In table 4 only hourly paid workers are used. Union differentials may be larger for this group.[10] Using more recent data and a revised methodology, MacDonald (1982) reports substantially larger differentials than those found earlier in MacDonald and Evans (1981). The average differentials for 1979 were 29.2% for unskilled workers, 17.2% for semiskilled workers, and 22.3% for skilled workers. Averaging over skills the differential was 22.8%.

The basic conclusions from table 4 are that the union differential is large and that the selectivity bias correction increases it for males but not for females. As noted earlier, in order to examine the sensitivity of the results to the normality assumptions employed in the probit approach, the union differential was also computed by estimating equation (6) by instrumental variables. The results are similar to the selectivity approach. In particular, estimates of the union differential are higher when union status is treated as endogenous, and are generally high—around 40%.[11] These results may be compared with U.S. studies using a similar methodology. A survey by Freeman and Medoff (1982) lists three studies on union differentials using non-occupation-specific samples with selectivity correction which may be compared with the present study. In Lee (1978), the correction reduces the differential by a modest amount. However, in the other two studies, Heckman and Neumann (1977) and Duncan and Leigh (1980), the differential rises with the correction, and in the former, under one specification, the differential is 40%.

It has been argued that the demand for labor is less elastic in the public sector and hence that union differentials will be larger in the public sector than in the private sector. There is some evidence that the demand for labor is less elastic in the public sector (see Ehrénberg 1973; Ashenfelter and Ehrenberg 1975). However, previous estimates of the union differential for public sector workers in the United States have shown them to be no larger than in the private sector (see, e.g., Ehrenberg and Goldstein 1975). The evidence from the present study on this issue is mixed. The results from the nonpooled sample are consistent with the U.S. results. Union differentials are on average higher in the private sector. The exceptions at the disaggregated level are unskilled and semiskilled males where the differences are very small. Using pooled sample estimates, the public sector union differentials are larger than in the private sector. Since

[10] When the analysis was repeated using a broader sample, the pattern of results was generally the same but the estimated differentials were considerably smaller. (See Robinson and Tomes 1983.)

[11] Full results from the instrumental variables estimation are available on request from the authors.

the public sector status dummy variable was insignificant in the pooled sample, however, the pooled sample results in this respect must be viewed with caution. Moreover, the estimates of equation (6) by instrumental variables reveal a negative (though insignificant) coefficient on the interaction between union status and public sector status.[12]

Most of the estimates of union differentials in recent surveys (Parsley 1980) and almost all of the estimates for Canada (Anderson and Gunderson 1982) are single-equation estimates. Estimates of this type from the present data were computed for purposes of comparison. The inclusion of a dummy variable for union status in a logarithmic wage equation yields a significant coefficient for union status ($t = 6.49$) and a differential of 27.69%. Finally, on the issue of the relative size of union differentials in the public and private sectors, the inclusion of an interaction between union and public sector status, as in the instrumental variables estimates, resulted in an insignificant (negative) point estimate—consistent with the disaggregated samples above.

The final question considered in this section is what determines the probability that an individual will be in a union at the time of the survey. The model structure outlined above included a reduced-form probability equation for union status and a structural equation. The latter is typically of primary interest since it allows investigation of the effect of the union-nonunion wage differential on the probability of union membership. However, care must be taken in interpreting the "structural" equation, since the problem of union membership determination is not well specified and the sample sizes are relatively small. Table 5 below presents reduced-form estimates for both the pooled sample and the separate public sector and private sector samples.

The pooled sample results are very similar to the separate sector results. First, considering the variables excluded from the wage equations, proxying pure "costs" of union membership—part-time worker status and firm or plant size—these are both significantly different from zero in the expected direction. Part-time status reduces the probability of union

[12] More generally, comparison with the U.S. results on differences in union differentials in the public and private sectors must be made with care. In the private sector, comparing union and nonunion workers will generally include comparison of individuals doing a similar job in a similar industry, the only difference being that one is unionized and one is not. In the public sector, a similar experiment may be made if we compare, say, unionized firefighters working for one municipality with nonunionized firefighters working for another. However, in the Canadian sample, the typical experiment carried out in comparing "similar" individuals who differ in their union status will involve a comparison across different kinds of jobs rather than within the same job. This occurs because the sample is dominated by large public sector employers rather than small municipalities. Typically, therefore, "similar" workers who differ only in their union status will not be doing the same "job."

membership. Large plant sizes increase the probability of union membership. The remaining variables excluded from the wage equation were marital status and income of the spouse. The hypothesis that union services are normal goods implies that a higher level of spouses' income will increase the probability of union status. This appears to be the case in the private sector but not in the public sector. This may be due to different types of union services in the two sectors having different income elasticities. Alternatively, it may simply be due to the small sample size for the public sector. The remaining variables are included in the wage equations, hence in general they combine wage effects as well as (in some cases) cost effects.

In table 6 the structural estimates of the process determining union status are presented. The first three columns present estimates using the corrected wage coefficients. The remaining columns present the results using uncorrected wage coefficients for comparison. In both pooled and private sector samples the predicted wage difference has the most significant effect on the probability of union membership. In the public sector the estimated effect is zero. The very strong positive effect in the pooled and private sector samples is consistent with Lee's (1978) results for the United States that worker's union status is responsive to potential wage gains. Because of the very small sample size for the public sector it is not possible to determine whether the structure is similar in public and private sectors. However, pooling the samples actually results in a larger point

Table 5
Probability of Union Status for Hourly Paid Workers: Reduced Forms

Independent Variables	Pooled Sample	Private Sector	Public Sector
Atlantic	−.9654 (3.15)	−1.0823 (2.62)	−.8416 (1.54)
Quebec	−1.0924 (3.71)	−1.1781 (3.17)	−.5915 (1.09)
Ont	−.5805 (2.88)	−.6123 (2.56)	−.3841 (.93)
Prairies	−.8357 (3.23)	−.6950 (2.34)	−1.0339 (1.82)
Yrssch	−.0214 (.84)	.0022 (.07)	−.0318 (.60)
Expr	.0280 (1.76)	.0557 (2.75)	−.0205 (.68)
ExprSq	−.0007 (2.08)	−.0011 (2.86)	.0003 (.43)
Tenure	.0329 (3.30)	.0293 (2.50)	.0398 (1.88)
French	.6770 (2.69)	.7114 (2.18)	.5729 (1.32)
Male	.6018 (3.91)	.5707 (3.49)	.3334 (1.17)
Public-male	−.1648 (.59)	. . .	. . .
Public	.8211 (2.31)	. . .	. . .
POW	.0194 (3.84)	.0222 (4.12)	. . .
Skilled	−.0336 (.23)	−.0127 (.07)	−.1660 (.57)
Unskilled	.0988 (.64)	.0211 (1.21)	−.2841 (.81)
SpInc	.0298 (1.69)	.0565 (2.54)	−.0184 (.56)
Married	.1950 (1.41)	.0459 (.87)	.1947 (.71)
Plantsize	.0017 (2.88)	.0019 (3.19)	. . .
Parttime	−.4432 (2.80)	−.3514 (1.81)	−.6967 (2.29)
Constant	−1.3012 (2.87)	−1.9712 (3.63)	1.4296 (1.54)
$-2 \ln \lambda$	273.26	196.92	30.21
Limits	360	312	43
N	659	497	154

NOTE.—Absolute values of *t*-statistics in parentheses.

Table 6
Probability of Union Status for Hourly Paid Workers: Structure

Independent Variables	Corrected Coefficients			Uncorrected Coefficients		
	Pooled Sample	Private Sector	Public Sector	Pooled Sample	Private Sector	Public Sector
Wage Diff	2.8430 (5.21)	2.6678 (4.91)	−.1232 (.11)	.0136 (.01)	1.7032 (4.28)	−1.4282 (1.26)
Atlantic	−.3035 (.96)	−.4590 (1.14)	−.7794 (1.54)	−.8671 (2.91)	−1.4287 (3.44)	−.3454 (.52)
Quebec	.0207 (.09)	.1235 (.46)	−.2070 (.62)	−.5007 (2.21)	.5767 (1.62)	−.1731 (.39)
Ont	−.1272 (.62)	−.1866 (.79)	−.3700 (1.05)	−.4964 (2.49)	−.3445 (1.44)	−.4370 (1.03)
Prairies	−.4269 (1.61)	−.3423 (1.16)	−1.0937 (2.20)	−.8555 (3.24)	−.9752 (3.22)	−.7122 (1.10)
Yrssch	−.0036 (.14)	.0004 (.01)	−.0387 (.81)	−.0365 (1.48)	−.0179 (.61)	−.0183 (.34)
Expr	.0424 (2.69)	.0367 (1.86)	−.0104 (.29)	.0285 (1.26)	.1261 (4.75)	−.0379 (1.01)
ExprSq	−.0009 (2.87)	−.0008 (2.06)	.0003 (.32)	−.0005 (1.08)	−.0026 (4.63)	.0009 (1.08)
Male	−.0958 (.49)	.0921 (.45)	.4274 (1.06)	.7013 (5.46)	1.0658 (6.20)	.6686 (1.91)
Public	.6157 (2.16)	. . .	. . .	1.7676 (9.05)	. . .	. . .
Skilled	.1944 (1.37)	.0888 (.52)	−.1162 (.38)	.0980 (.63)	.1939 (1.11)	−.2613 (.84)
Unskilled	−.3268 (1.85)	−.2526 (1.21)	−.2864 (.78)	.1924 (.94)	−.6002 (2.15)	.0208 (.04)
SpInc	.0328 (1.87)	.0589 (2.66)	−.0151 (.47)	.0336 (1.95)	.0598 (2.72)	−.0181 (.57)
Married	.2074 (1.51)	.1690 (1.00)	.1991 (.74)	.2218 (1.65)	.1884 (1.13)	.2271 (.84)
Plantsize	.0017 (3.14)	.0021 (3.67)	. . .	.0026 (4.85)	.0022 (3.72)	. . .
Parttime	−.4504 (2.88)	−.3685 (1.91)	−.7138 (2.39)	−.5441 (3.53)	−.4010 (2.09)	−.6673 (2.30)
Constant	−1.4199 (3.23)	−1.7511 (3.46)	1.5770 (2.33)	−.8614 (1.57)	−3.6759 (5.03)	1.4800 (1.63)
$-2 \ln \lambda$	264.80	192.65	23.88	236.49	186.03	25.50
Limits	360	312	43	360	312	44
N	659	497	154	659	497	154

NOTE.—Absolute values of t-statistics are presented in parentheses. These are corrected to account for the computed value of the wage difference via the selectivity correction procedure. Variables that enter only via the wage equations (tenure, French, POW) are excluded here.

estimate for the wage difference and a higher significance level. Therefore there is no strong evidence against similar mechanisms operating in public and private sectors. As expected, the pure "cost" variables, plant size and part-time status, have essentially identical effects in the structure as in the reduced form. In addition, the structural coefficients for marital status and spouse's income are the same as the reduced-form coefficients. Thus the evidence in favor of unions providing normal goods, at least in the private sector, is maintained.

The major differences between the reduced-form and structural coefficients appear for variables that influence wage rates as well as (potentially) costs. Most notably the effect of being male, which is strongly significant in the reduced forms, is eliminated in the structure. This suggests that most of the higher probability of males belonging to unions is due to the larger expected wage gain as compared with females. Similarly, the regional differences apparent in the reduced forms are absent in the structure. Since union legislation is a provincial matter, union costs may potentially vary by province; however, comparison of the results in tables 6 and 7 suggests that any cost differences are dominated by differential

Table 7
Public-Private Sector Wage Differentials for Selected Groups of Hourly Paid Workers

	Corrected Coefficients		Uncorrected Coefficients	
	ϕ_g	ϕ_g^p	ϕ_g	ϕ_g^p
Union sector:				
Males:				
Unskilled	−12.58	−7.49	−2.70	−4.88
Semiskilled	−7.20	−7.49	.30	−4.88
Skilled	−9.06	−7.49	−2.68	−4.88
Average	−9.34	−7.49	−1.18	−4.88
Females:				
Unskilled	−6.22	2.11	−.84	2.34
Semiskilled	−.44	2.11	2.22	2.34
Skilled	2.44	2.11	−.82	2.34
Average	−2.24	2.11	.71	2.34
Male and female average	−6.72	−3.94	−.48	−2.21
Nonunion sector:				
Males:				
Unskilled	2.00	−22.24	−11.61	−4.34
Semiskilled	6.65	−22.24	−1.25	−4.34
Skilled	18.70	−22.24	7.93	−4.34
Average	8.83	−22.24	−1.13	−4.34
Females:				
Unskilled	14.63	−2.22	10.06	21.88
Semiskilled	19.85	−2.22	22.96	21.88
Skilled	33.39	−2.22	34.39	21.88
Average	22.30	−2.22	23.11	21.88
Male and female average	17.59	−9.23	14.63	12.70

NOTE.—Average skill levels use the full sample weights. A level of 67.8% of organized workers is assumed for the private sector in computing ϕ_g. The pooled wage equations automatically hold the level of organization the same across public and private sectors. The differential is computed as $\phi_g \equiv (\ln W_g - \ln W_p) - 1$, where $\ln W_g$ and $\ln W_p$ are the natural logarithms of wage rates in the public and private sectors, respectively. Male and female averages weight the differential by the proportion of public sector males and females in each sector.

wage levels by province. However, public sector status has a positive effect on the probability of union status in both the reduced form and the structure. This suggests that the higher percentage of unionization in that sector is not due solely to larger potential wage gains, but also reflects differences in the costs of organization.

The significant wage difference coefficients in table 6 for private sector workers are quite similar to those obtained by Lee (1978). Lee does not discuss the interpretation of this coefficient in terms of what model parameters it estimates. However, the conformity of the present results with those of Lee suggests that this issue should be pursued further. Substituting (4) and (5) into (3) indicates that the coefficient on the wage equation is the inverse of the standard deviation of a linear combination of the wage equation disturbances, ϵ_U and ϵ_N, and ϵ, the disturbance in the "costs" or reservation wage equation. Thus the coefficient on the wage difference depends on the variances and covariances of ϵ_U, ϵ_N, and ϵ. Since there is a large literature on estimates of variances of disturbances in individual logarithmic wage equations, it would be of interest to know whether the wage difference coefficients in table 6 were consistent with the stylized facts on these variances. Estimating the variances of ϵ_U and ϵ_N is not possible directly because of the problem of censored samples. However, we can obtain predicted values of ϵ_{Ui} and ϵ_{Ni} for *all* members in the sample (see Robinson and Tomes 1983, n. 15). Using these predicted values, the implied variances for ϵ_U and ϵ_N are .02 and .06, respectively. The covariance is $-.03$. The estimated variances are underestimates since they use expected rather than actual values. Inspection of variance estimates in the literature suggests that these results are not inconsistent with that literature. For example, variances of individual components in wage equations are often estimated in the range .10–.15. (See MacDonald and Robinson 1982.)

III. Public-Private Sector Wage Differences

Recent estimates of wage differentials for public versus "comparable" private sector employees suggest positive premiums for public sector workers. Smith (1977) found substantial positive premiums for federal government workers in the United States in 1975. For males the differential was 13%–15%; for females the differential was 18%–20%. For state and local government employees, however, the differentials were only positive for females. Gunderson (1979*a*, 1979*b*) computed wage differences between public and private sector workers in Canada using a methodology similar to that of Smith's U.S. study. Gunderson did not distinguish between different levels of government. He found results similar to those of Smith. The public sector differential was typically positive but larger for females (8.6%) than for males (6.2%). Because of data limitations, Gunderson was unable to take into account the effects of unionization. This is a potentially serious drawback because union coverage is considerably higher in the public than in the private sector. Thus public sector differentials may be confused with union differentials.

Estimates of public sector differentials, controlling for union status, may be obtained from the wage coefficients presented in tables 2 and 3. However, these estimates, like previous studies, assume public sector status is exogenous. If, in fact, the choice of sectors is endogenous these estimates will be subject to bias. In order to avoid this two strategies were pursued. First, a model of public sector choice was specified and estimated to correct the estimates. Second, the instrumental variables approach was pursued by using the linear probability model to obtain an instrument for public sector status in equation (6). There were several problems encountered in undertaking these procedures due to the poor performance of the probit equation for determining public sector choice, or more generally of poor instruments for public sector status. Adopting a compromise solution to these problems yielded wage equations in general very similar to those obtained correcting only for union status (see Robinson and Tomes [1983] for further details). Some confidence may be placed in the equation determining union status, because of its similarity with those obtained by other investigators. Since a satisfactory equation determining public sector status could not be obtained, the analysis of public sector differentials in this section is based on the wage equations of Section II.

The estimated wage equations from the pooled private and public sector sample provide a direct test of significant additive differences in the logarithmic wage equations in the public and private sectors (table 3). The coefficient on public sector status is insignificantly different from zero in both union and nonunion sectors. This holds for both corrected and uncorrected coefficients, though in the latter case a sizable positive point estimate for the differential is significantly different from zero at the 10% level. Thus the direct estimates provide little evidence of significant positive public sector differentials in either union or nonunion sectors. This is reflected in table 7, where the coefficients of tables 2 and 3 are used to compute public sector differentials. Using the coefficients from table 3, all the public sector differentials (ϕ_g^p) are negative except for a small positive differential for unionized females. Without the selectivity correction in the wage equations, the public sector differentials are negative for unionized and nonunionized males and positive for females.

The disaggregated wage equations of table 2 may also be used to compute public-private sector differentials (ϕ_g). However, as noted in the previous section, the disaggregation results in small sample sizes for some of the subsamples, particularly nonunion–public sector workers. Thus, the disaggregated results must be treated with special caution. In computing the public sector differentials implied by the wage equations of table 2, the level of union organization was set at the same level in both public and private sectors. As noted above, the level of organization is 67.8% for all public sector workers and hence is omitted from the public sector wage equation. Its effect is captured in the constant term. In order to hold the level of organization across public and private sectors constant, the private sector wage equation, which explicitly includes the percentage of organized workers, was evaluated at a level of organization of 67.8%. Using the corrected coefficients, the disaggregated wage equations yield public-private sector differentials in the union sector that are similar to those obtained from the pooled sample. In the nonunion sample there is some divergence: the disaggregated wage equations show positive differentials for both males and females, whereas using the pooled sample equations they are both negative. However, the disaggregated sample sizes are small, hence the disaggregated results are subject to potentially large errors. In addition, significance tests on additive public sector effects in the union and nonunion wage equations showed no significant public sector differential. The evidence against a public sector differential was also strengthened by the instrumental variables analysis of equation (6). The coefficient on public sector status, irrespective of whether union status was treated as exogenous or endogenous, was always insignificantly different from zero.

Finally, some estimates of the potential overestimate of public sector differentials from omitting union controls may be made. First, the effect of not setting the level of organization the same in both public and private sectors when the disaggregated equations are used is substantial. For average unionized males a negative public-private sector differential of −9.34% becomes a positive 6.42%. For average unionized females a negative differential of −2.24% becomes a positive 14.75%. Second, if union membership itself is not controlled for, there is a marked increase in the estimated public-private sector differential. For example, using the pooled sample, we estimated that unionized males earn 7.49% less in the public sector than in the private sector and nonunionized males 22.24% less. However, when the estimated wage rates are weighted by the proportions of union and nonunion members in each sector, the differential becomes zero. If females are also included in the calculation, even though three out of four subgroups have negative public sector differentials, the average public-private sector differential without controlling for union status is positive, approximately 5%. This suggests that apparent public sector rents found in the absence of controls for union status (e.g., Gunderson 1979*a*) may in fact be largely union differentials.

IV. Conclusions

Large estimated union differentials for hourly paid workers were obtained in Section II, controlling for union status. There was considerable evidence of positive selection into the union sector, especially for private sector workers. Union status appears to be strongly affected by the expected wage gain from joining the unionized sector. There was some evidence of larger union gains in the public sector than in the private sector from the pooled sample estimates, but this was not replicated in the disaggregated estimates. Estimates of public-private sector wage differentials were presented in Section III. Typically these were negative, though disaggregated results suggested a positive differential for nonunionized workers, particularly females. Controlling for union status was shown to reduce estimates of public-private sector differentials dramatically, suggesting that recently estimated "rents" accruing to public sector employees (Gunderson 1979*a*) may primarily reflect the recent increase in unionization in this sector.

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