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Article abstract

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This paper presents estimates of union-nonunion differences in wages, wage dispersion and pension provisions derived from a rich, new Canadian longitudinal micro data base. Canadian empirical research on the economic effects of unions has tended to focus narrowly on union wage effects; research on the nonwage impact of unions (e.g. fringe benefits, wage inequality, job tenure) has been comparatively neglected[1]. However, such a narrow focus may give a misleading impression about the economic effects of unions. For example, given that fringe benefits form a significant portion of total compensation, union effects on total labour compensation derived from estimated union-nonunion wage differentials may be seriously understated if unions have a greater impact on fringe benefits than on wages. Estimates of the union effect on resource allocation and total output would thus be similarly understated.

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** This analysis is based on Statistics Canada microdata tape Labour Market Activity Survey Longitudinal File which contains anonymized data collected in the year 1986 Labour Activity Survey. All computations on these microdata were prepared by the authors and the responsibility for the use and interpretation of these data is entirely that of the authors.

1 There have been several studies of the nonwage effects of unions based on macro data. See, for example, the study by Maki (1983) of the union effect on productivity, and the study by Maki and Meridith (1986) of the union effect on profitability.

Although the exact magnitude is much in dispute, there is considerable empirical evidence to indicate that the union-nonunion wage differential is substantial. Recent Canadian studies by Robinson and Tomes (1984), Simpson (1985), Kumar and Stengos (1985, 1986), and Grant, Swidinsky and Vanderkamp (1987), all of which rely on individual microdata and correct for selectivity bias, yield estimates of the union premium ranging from 12.1 percent (Kumar and Stengos 1986) to 34 percent (Robinson and Tomes 1984). The union effect on wage dispersion and fringe benefits is less well defined both theoretically and empirically. There are few Canadian econometric studies that one can use as a precedent[2], but prominent recent U.S. studies by Freeman (1980, 1981, 1982), Leigh (1981), Hirsch (1982), Freeman and Medoff (1984), and Allen and Clark (1986), suggest collectively that, among other things, unions narrow wage dispersion and increase various fringe benefits, especially pension benefits. The present study provides evidence of the union effect on wages, wage dispersion and pension fringe benefits derived entirely from longitudinal data. To our knowledge this paper is the first attempt to provide such longitudinal estimates for Canada.

UNION EFFECT ON WAGES, WAGE DISPERSION, AND PENSIONS

Unions use their monopoly power which they exercise by withholding, or threatening to withhold, labour to raise the mean wage of union workers. The surplus labour generated in the union sector by this process moves to the nonunion sector where it depresses the wages of nonunion workers. If, as Freeman and Medoff (1984) suggest, workers who became union members were higher paid to begin with, the monopoly effect of unions will widen the gap between union and nonunion wages and increase overall wage inequality. On the other hand, unions can exert an equalizing effect on wages by narrowing wage dispersion among union workers[3]. This narrowing of the relative dispersion of union wage rates is the result of an explicit union wage policy of rate standardization within and among establishments.

The policy of rate standardization across firms is intended to take wages out of competition. Freeman (1980) argues that a common rate

2 The study by Evans and Ondrack (1986) explores the union effect on wage dispersion, but the analysis is limited to a small sample of workers in the petrochemical industry in Sarnia, Ontario.

3 Unions can exert an additional equalizing effect on overall wages by narrowing the diffential between the higher paid white-collar workers and the lower paid blue-collar union workers. See Freeman and Medoff (1984).

across firms operating in the same product market prevents competition based on wages and removes pressure for some unions to undercut wages during economic downturns. Even where firms operate in separate markets unions will attempt to limit wage differentiation in order to minimize divisions within the organization and maintain common policies towards major employers.

According to Freeman (1982), several factors explain the union policy of wage standardization within establishments. First, gross inequities may arise when the wages individual workers are paid depend not on the characteristics of the jobs held but on the worker's wage-determining characteristics as perceived by supervisors, and on employer merit reviews. Second, given the dependence of the union as a political organization on median worker support, when the mean wage within the unionized firm exceeds the median the union is likely to accede to majority preference and favour redistribution towards the lower paid. Accordingly, unions will seek to reduce intra-establishment wage differentials among workers with nominally similar skills and job characteristics by the setting of a single rate of pay for each occupational group and a seniority-based progression of rates up to a maximum. To the extent that union inter and intra-establishment policy on wage structure is enforced, the dispersion of wages among union workers should be narrower than among nonunion workers.

Unions not only raise wages of their members relative to the wages of nonmembers, they may also raise that part of total compensation that is paid in fringe benefits. This, Freeman (1981) argues, is accomplished in two ways; raising the total level of compensation and thus the level of fringe benefits if the division between wages and fringes is maintained, and raising the share of fringes in a given compensation package. On the latter point, Freeman maintains that within the context of the median voter model, unions, as political institutions, must be more responsive to the preferences of the median member than the preferences of the marginal member. Conversely, given that the employer has to attract the marginal worker, it is the preferences of the marginal worker that will determine the composition of total compensation in a competitive market.

Median union members presumably have a greater preference for fringe benefits, especially pension benefits, because they are older and less mobile than marginal nonunion members. These preferences are accentuated if employment tenure in a union firm is more permanent so that the likelihood that the pension benefit will be collected is increased. Moreover, given the complexity of pension plan provisions, workers will have more confidence in such fringe benefits if the pension fund is monitored and explained by union experts. Finally, if there is a misperception by management of the level of pension benefits desired by workers, such misperception is more readily corrected by the union through its use of "voice" than would be the case in a nonunion environment.

SOURCE OF DATA

The source of data is the Labour Market Activity Survey (LMAS) which is designed to complement the stock estimates obtained from the monthly Labour Force Survey (LFS) conducted by Statistics Canada. The LMAS person file provides information on the annual work patterns of Canadians in 1986. For each individual in the sample the LMAS contains information on demographic characteristics, income profiles, work patterns, and job characteristics, including union status, for up to five jobs.

The total sample includes 66,934 individuals, of whom 9,345 held at least two jobs in 1986. Our analysis is based on the latter group of job changers but it is limited to the first two consecutive jobs in the year. Several restrictions were imposed on the sample to make the individuals in the working sample as comparable as possible. The sample was restricted to paid workers who were not students, had a non-zero wage rate, provided complete information on industry, occupation, union status and firm size in both jobs, and had different employers in the two jobs. By restricting the sample to job changers with different employers we avoid the problem common in several earlier studies, but especially Mincer (1983), where job changers report union status change even though they did not change employers.

Two additional restrictions are imposed to contend with measurement error. The first limits the sample to those individuals with an hourly wage rate ranging from a minimum of 4 dollars (slightly below the minimum wage) to a maximum of 50 dollars in both jobs 1 and 2. The second restriction imposes the added constraint of full-time employment (defined as in the monthly LFS to be 30 or more hours per week). These restrictions were necessary because a significant number of individuals reported unreasonable usual hours worked which resulted in unrealistic hourly wage rates[4]. With only the wage restriction the resulting working sample includes 2,438 individuals, but with both wage and hours restrictions the working sample is reduced to 1,560 individuals.

4 Not all of these cases involve measurement error. For example, some individuals reported jobs that required them to be on duty 24 hours a day (e.g. priests, camp counsellors). Also, payment in kind was not included in the reported income and this type of payment may be significant in some cases. See Statistics Canada, *LMAS Mocrodata User's Guide,* Section 9.

Table 1 provides sample means for key variables in both job 1 and job 2 for the larger working sample. Several observations are in order. Unionized workers in our sample include union members as well as nonmembers whose wages are established by collective agreement. But even with this definition of union membership, only 20.5 percent of the job-leavers in our working sample were unionized in job 1 and 22.9 percent in job 2[5]. Furthermore, only 12.3 percent of the individuals in our working sample were unionized in both jobs. This is a considerably lower proportion than in the Grant, Swidinsky, and Vanderkamp (1987) study, but the sample underlying their study does not distinguish between job movers and stayers. However, the proportion of permanent union members in our sample is similar to the proportion of permanent members in Mincer's (1983) NLS (National Longitudinal Surveys) and MID (Michigan Panel Survey of Income Dynamics) samples of job movers. This similarly is surprising given the difference in overall unionization rates in Canada and the United States. An estimated 18.8 percent of the job-changers also changed their union status. This is a considerably higher proportion than in the sample underlying the Grant, Swidinsky, and Vanderkamp (1987) union wage effects study, but, again, it is similar to the sample composition of movers in the Mincer (1983) analysis of union wage effects.

UNION WAGE EFFECTS

The empirical analysis in this section relies on observing wages received by the same worker in a unionized and nonunionized job. This fixed-effect-function approach to estimate union wage gains has been used in Mellow (1981), Mincer (1983), and, most recently, Grant, Swidinsky and Vanderkamp (1987). The advantage of the fixed-effect-function approach is that it estimates whether the same worker, and not just his statistical surrogate, receives higher wages in union than nonunion employment. In this procedure a wage level equation relating the logarithm of hourly earnings to a set of explanatory variables, including union status, and an error term is specified separately for each of the two sequential jobs held by an individual. The error term can be divided into two components: A fixed-effect error associated with an individual and assumed to be time invariant but which may be correlated with union status, and an uncorrelated random error.

5 In job 1, 36.8 percent of the paid workers in the entire LMAS sample (subject to the same restrictions as the working sample), had their wages determined by collective bargaining.

Table 1

Variable Descriptions and Sample Means of Key Variables by Union Status (2438 Observations)

Variable Description	*Job 1*		*Job 2*		*Union*
	Union	*Nonunion*	*Union*	*Nonunion*	*Status Change*
Union Status (Percent)	20.5	79.5	22.9	77.1	–
Ln Hourly Earnings	2.389	1.978	2.366	2.043	–
Standard Deviation of Ln Hourly Earnings	.423	.432	.399	.435	
Percentage With Pension Coverage	52.5	10.7	50.3	11.4	
Union Status Change (Percent)					
UU	–	–	–	–	12.3
UN	–	–	–	–	8.2
NU	–	–	–	–	10.6
NN	–	–	–	–	68.9

Taking the first differences of the variables in the wage level equations yields a wage-change equation in which the fixed-effect error associated with selectivity as well as certain independent variables (e.g. age, sex) have been eliminated. What remains on the right hand side of the wage-change equation are a select number of variables expressed as first differences and the union status change variables: UU (union to union), UN (union to nonunion), NU (nonunion to union), and NN (nonunion to nonunion). Estimates of the average union-nonunion wage differential can be derived from such a wage-change equation by averaging the absolute values of the coefficients of union leavers (UN) and union joiners (NU).

In principle, according to Mincer (1983), the wage-change regression estimate of the union wage premium (the NU coefficient) is equivalent to the increment in the NU coefficients estimated from a prospective wage level regression (job 1) and a retrospective wage level regression (job 2). The coefficient of NU in the prospective regression estimates the wage of prospective movers from nonunion to union jobs in their initial nonunion job. Similarly, the coefficient of NU in the retrospective equation estimates the wage of new union members in their new union job (i.e. job 2). The difference between the estimated coefficients in the retrospective and prospective equations is the union premium net of selectivity. It should be equivalent to the NU coefficient estimated from a single wage-change equation, although in practice the two values will differ marginally because of a differing structure of errors in level and change equations. Similarly, the difference in the coefficients of UN estimated from the prospective and retrospective equations can be interpreted as the loss of the selectivity corrected union premium resulting from a move to a nonunion job. In both cases the estimated wage gains and losses are relative to the wages of the base group of permanent nonunion workers (NN).

While union wage effects net of selectivity can be estimated using either method, the prospective and retrospective wage level regression method has several advantages. First, wage level equations allow for a more accurate specification: Wage-change equations often include level variables that are not consistent with the derivation of the wage change equation. Second, wage level regressions provide information on more than just union wage effects. The prospective equation would reveal whether unionized employers are selective in hiring more productive labour, given that they are forced to pay the union premium. It thus seems reasonable to rely on prospective and retrospective wage level regressions for our analysis but to verify the derived union wage effects using wage-change regression estimates.

The dependent variable in the wage level equations for the first and second jobs is the log of hourly earnings. Both equations contain the same independent variables: Age, sex, marital status, region, education, industry, occupation, job tenure, full-time/part-time status, firm size, and the three union status change variables UU, UN, NU (NN is the omitted category). The wage-change equation is derived by taking first differences of all the variables in the level equations, except the union status change variables. Thus, in addition to union status change designation the right hand side of the wage-change equation contains industry change, occupation change, job tenure change, full-time/part-time status change, and firm size change variables[6].

6 Normally, industry change, occupation change, and firm size change are indicated by dummy variables. In our study we specify these changes by continuous variables constructed by taking the difference in the average wages of the two industries, occupations, or firm sizes in which the two jobs were located. The results were somewhat better when the latter specification was used.

The complete regression results are presented in Table 1A in the Appendix. Table 2 presents only the coefficients on union status change dummies in the two cross-section wage regressions (job 1 and job 2), the wage gains and losses implied by these coefficients, and, for comparison, the net union wage premiums for joiners and leavers, respectively, estimated from a wage-change regression. Estimates are presented for both total and full-time samples. In the total sample, permanent union members enjoyed a 27.7 percent wage premium over permanent nonunion members and a 9.5 percent premium over prospective union leavers. The wage rate of prospective union joiners was 4.2 percent lower (significant at the 10 percent level) in the nonunion job 1 than the wage of permanent nonunion members. Not only is there no evidence that unionized employers select more productive nonunion workers to compensate for the higher union wage, there is weak evidence (significant at the 10 percent level) that prospective union members had lower wages than permanent nonmembers. The estimated union status coefficients for the full-time sample are in general only marginally different.

Table 2

Estimated Union Status Coefficients and t-values from Wage Level and Wage-Change Equations

Union Status Change	*Level Coefficients* Job 1 (1)	Job 2 (2)	*Implicit Wage Change* (2) - (1) (3)	*Wage-Change Equation Coefficients* (4)
		Total Sample (2438 Observations)		
UU	.277	.238	-.039	-.040
	(12.5)	(10.7)		(2.0)
UN	.182	.048	-.134	-.135
	(7.2)	(1.96)		(5.6)
NU	-.042	.134	.176	.179
	(1.93)	(5.8)		(8.2)
		Full-Time Sample (1560 Observations)		
UU	.269	.243	-.026	-.035
	(10.5)	(9.2)		(1.6)
UN	.166	.065	-.101	-.105
	(5.2)	(2.0)		(3.7)
NU	-.047	.115	.162	.171
	(1.8)	(4.0)		(7.0)

Column (3) of Table 2 shows that the wages of union joiners in the total sample increased by 17.6 percent whereas the wages of union leavers declined by 13.4 percent. Gains and losses for full-time union status changers were 16.2 and 10.1 percent, respectively. These estimates are almost identical to the union status change coefficients estimated from a wage-change regression and reported in column (4). Averaging the absolute gains and losses resulting from a change in union status yields an average union-nonunion wage differential of 15.5 percent for all workers and 13.1 percent for full-time workers. These estimates are well within the overall range of 13-16 percent for 1970 reported in the Grant, Swidinsky and Vanderkamp (1987) study based on a variety of estimating techniques.

UNION EFFECT ON WAGE DISPERSION

Trade unionism can be expected to reduce inequality in wages in the union sector by equalizing rates among establishments and replacing personal rates with formal job rates within establishments. Freeman (1980, 1982) has found that the dispersion (measured by the standard deviation) of the log of union wages is substantially lower than that of nonunion wages, even after controlling for differences between union and nonunion workers' personal characteristics. A significant portion of that difference is due to a lower dispersion within unionized establishments. However, these findings, based on cross-section comparisons, do not test for the possibility of simultaneity between unionism and wage dispersion: The inverse relation between unionism and dispersion may also reflect the greater likelihood that unions organize low-dispersion firms. But even after accounting for simultaneous determination, Hirsch (1982) finds that unions significantly decrease earnings dispersion.

According to Freeman (1982), the problem of inferring causality can be overcome by the use of longitudinal data that would indicate whether the reduction in dispersion for a common group of workers occurred before or after the advent of unionism. If union wage policies are responsible for the observed difference in dispersion, the dispersion of wages would be expected to fall for a group of workers moving from nonunion status to union status, and to rise for workers moving in the other direction. The longitudinal analysis of dispersion in ln wages performed by Freeman (1982) on a small sample of union status changers supports the argument that unions choose wage policies that reduce dispersion.

Table 3 presents the results of our longitudinal analysis of wage dispersion for workers who changed union status. The standard deviation of ln wages falls slightly in both samples for workers moving from nonunion to union status (-.005 and -.008 for all workers and full-time workers, respectively) but increases sharply for workers moving from union to nonunion status (.034 and .037 for all workers and full-time workers, respectively). If we average the NU and UN effects, unions appear to reduce the standard deviation of ln wages by .02 (about 5 percent) for all workers and by a slightly larger amount for full-time workers. Although these results are derived from relatively few observations, they nonetheless support the hypothesis that wage dispersion is lower among union than nonunion workers and that the difference is the result of union wage policies with respect to wage structure.

Table 3

Longitudinal Analysis of Dispersion in ln Wages of Workers who Change Union Status

Union Status Change	*Sample Size*	*Standard Deviation of ln Wages*		*Longitudinal Effect of Change on Dispersion*
		Before Change	*After Change*	
		Total Sample		
NU	259	.399	.394	-.005
UN	199	.420	.454	+.034
		Full-Time Sample		
NU	168	.391	.383	-.008
UN	116	.402	.439	+.037

UNION EFFECT ON PENSION PROVISIONS

Using cross-section regressions and establishment data, Freeman (1981) found that a unionized establishment is 29 percent more likely to provide a pension plan than a nonunionized establishment. Moreover, unions were estimated to increase employer expenditures on such pension benefits by about 3.9 cents per hour, a sizeable increase given the average employer pension expenditure of 9.4 cents per hour. Leigh's (1981) analysis, which relies on individual microdata, also showed that unions have an important impact in extending the coverage of voluntary pension programs. In addition unions have been found to have an effect on workers' valuation of future pension benefits (Leigh, 1981) and on the actual pension wealth of beneficiaries (Allen and Clark 1986).

As in the case of union wage premiums, OLS cross-section analysis of the union effect on pension coverage may contain selectivity bias. Workers who have a high preference for fringe benefits, especially pension benefits, may also have a strong preference for unionism. If these workers are more productive, employers may not only pay higher wages but also increase pension benefit coverage to attract and retain such workers. Longitudinal data which traces changes in pension plan coverage as workers change their union status would eliminate the bias such selectivity might induce.

Applying the procedure used to obtain union wage premiums, we estimated prospective and retrospective equations with pension coverage (a dummy variable equal to 1 if the individual is covered by an employer pension plan and 0 otherwise) as the dependent variable. The independent variables are those used in the wage level equation and the ln hourly wages. This latter variable is necessary to distinguish between the union impact on pension coverage via the level of compensation and via the share of compensation going to pension coverage. Given the nature of the dependent variable, there is no pension-change equation similar to the wage-change equation in the analysis of the union impact on wages.

The estimated prospective and retrospective pension coverage regressions for the total sample are reported in Table 2A of the Appendix. Table 4 presents only the union status coefficients and the implicit impact of unions on pension provisions, holding wage compensation constant. Permanent union members were roughly 35 percent more likely to have pension coverage than permanent nonmembers in both jobs. On the other hand, union workers who became nonunion in their subsequent job were only 21.2 percent more likely to have pension coverage in the unionized job 1. However, at least in the total sample, this group of workers still had a 5.5 percent better chance of a pension in their second, nonunion job than did permanent nonunion workers. Nonunion workers who move to unionized jobs increase the probability of pension coverage by 22.4 percent above that of permanent nonunion workers. The results are very similar for full-time workers.

Averaging the (absolute) resulting changes in the likelihood of pension coverage for UN and NU workers shown in column 3 yields selectivity-corrected union effects of 19 and 22 percent for all workers and full-time workers, respectively. However, since unions raise wages and since wages influence pension fringe benefits, the total impact of unions will exceed the above estimates, which assume wage compensation fixed. Freeman (1981) estimates the total effect of unionism on fringes by adding together the effect of unionism on fringes when compensation is fixed and the effect of unionism on compensation multiplied by the effect of compensation on fringes. Following this procedure, the estimated total union effect on pension coverage for all workers and full-time workers is 22 and 25 percent, respectively, which is consistent with the estimates obtained by Freeman and Medoff (1984). Unions appear to affect pension coverage mostly by changing the share of pension fringe benefits in total compensation rather than by changing the level of compensation.

Table 4

Estimated Union Status Coefficients and t-Values from Pension Coverage Equations

	Coefficient		*Change in Pension*
	Job 1	*Job 2*	*Coverage*
Union Status	*(1)*	*(2)*	*(3) = (2) - (1)*
	Total Sample (2436 Observations)		
UU	.346	.355	.009
	(14.9)	(14.5)	
UN	.212	.055	-.157
	(8.1)	(2.1)	
NU	-.015	.224	.224
	(0.7)	(9.0)	
	Full-Time Sample (1560 Observations)		
UU	.363	.374	.011
	(12.9)	(12.2)	
UN	.223	.042	-.223
	(6.4)	(1.1)	
NU	-.020	.220	.220
	(0.2)	(6.8)	

COMPARISON OF LONGITUDINAL AND CROSS-SECTION ESTIMATES

Our longitudinal estimates show that unions have a substantial and predictable effect on wages, wage dispersion, and pension coverage. However, if longitudinal data contains even modest errors in measurement because of the misclassification of union status these estimates may be seriously downward biased. As Freeman (1984) illustrates, in cross-section studies of union effects random misclassification in union status will be proportionately small so that the bias in the estimated union coefficient should be modest. In the relatively smaller sample of union status changers such misclassification over two periods would be proportionately more significant and the resulting bias in the estimated union status coefficient derived from longitudinal models considerably larger.

There is no direct evidence of the extent of measurement error in our data, but if such error does exist our longitudinal estimates of union effects should be lower than cross-section estimates from the comparable data. To test for such bias, cross-section estimates are derived for all workers 29,702 observations) and full-time workers (23,988 observations) in job 1 (both job changers and stayers) subject to the same restrictions imposed on our working samples of job movers. The independent variables in the wage and pension equations are those used in the panel data regressions (see Tables 1A, 2A) except that the union status variables are reduced to a single union dummy variable.

Table 5 compares the union effects on wages, wage dispersion, and pension fringe benefits derived from longitudinal and cross-section experiments. The union wage effects derived from the two experiments are almost identical, whereas the union pension effects are only marginally higher in the cross-section, especially for full-time workers. However, the union effect on wage dispersion is considerably stronger in the cross-section than longitudinal experiment. In the cross-section estimate unions reduce the standard deviation of ln wages for all workers by .117, but in the panel estimates the reduction in the standard deviation of ln wages is only .02. The result for full-time workers is very similar.

Table 5

Longitudinal and Cross-Section Estimates of Union Effects on Wages, Wage Dispersion and Pension Fringe Benefits

	Longitudinal		*Cross-Section*	
	All Workers	*Full-time Workers*	*All Workers*	*Full-time Workers*
Wage Effects (Percent)	15.5	13.8	15.7	14.2
Wage Dispersion (Standard Deviations)	-.020	-.022	-.117	-.124
Pension Fringe Benefits (Percent)	19.0	22.1	26.2	25.9

Measurement error is not the only factor that might explain the discrepancy in wage dispersion. When comparing standard deviations derived from cross-section data for union and nonunion workers there is no attempt to control for factors that may affect wage dispersion other than union status. The longitudinal estimates, on the other hand, net out all extraneous influences so that they capture only the pure effect of unionism. Our results, therefore, seem to suggest that measurement error is not a serious problem in our longitudinal data, and that our longitudinal estimates of union effects are a fairly accurate representation of the true effects of unions.

CONCLUSIONS

In this paper we employed panel data obtained from the Labour Market Activity Survey to estimate union effects on wages, wage dispersion, and pension coverage. Our selectivity corrected estimate of the union-nonunion wage differential is in the 13.1-15.5 percent range, which is consistent with other selectivity corrected estimates for Canada, but most notably those by Grant, Swidinsky and Vanderkamp (1987). The estimated union effect on wage dispersion is less pronounced: The standard deviation of ln wages falls by .02 (or by roughly 5 percent) when workers become unionized. However, the union effect on the probability of pension coverage is considerably greater. Selectivity corrected estimates show that a worker employed in a union job is 22-25 percent more likely to have pension coverage than if he is employed in a nonunion job.

These estimates may be downward biased since longitudinal data is believed to contain serious measurement error. However, comparisons with cross-section estimates from the same data source show that measurement error is not a problem. This leads us to believe that our longitudinal estimates are likely to be reasonably accurate estimates of the true union effects on wages, wage dispersion and pension coverage.

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Appendix

Table 1A

Regression Results for Wage Level and Wage Change Equations

Independent Variable		*Coefficients and t-Values* *Wage Level Equations* *Job 1*	*Job 2*	*Wage Change Equation*
Constant		2.002 (44.8)	2.006 (41.3)	.067 (7.1)
Age[a]	- 25-34	.144 (8.5)	.156 (9.3)	
	- 35-44	.211 (10.1)	.221 (10.7)	
	- 45.54	.276 (9.7)	.271 (9.6)	
	- 55-64	.204 (4.9)	.317 (7.6)	
Female		-.203 (12.2)	-.220 (13.0)	
Married		.035 (2.4)	.054 (3.7)	
Education[b]	- High School	.043 (1.6)	.030 (1.1)	
	- Some Post Sec.	.115 (3.4)	.100 (2.9)	
	- Diploma	.190 (6.1)	.176 (5.6)	
	- Degree	.269 (7.3)	.305 (8.4)	
Region[c]	- Nfld	-.159 (4.4)	-.149 (4.1)	
	- PEI	-.109 (2.1)	-.034 (0.6)	
	- NS	-.095 (2.9)	-.114 (3.5)	
	- NB	-.070 (2.4)	-.091 (3.1)	
	- Que	-.015 (0.7)	.008 (1.3)	
	- Man	-.064 (2.3)	-.024 (0.8)	
	- Sask	.002 (0.1)	-.022 (0.8)	
	- Alta	.063 (3.0)	.054 (2.5)	
	- BC	.072 (2.9)	.067 (2.7)	
Job Tenure		.024 (6.3)	.297 (3.3)	
Job Tenure Squared		-.0006 (3.1)	-.150 (1.8)	
Part-Time		-.049 (2.9)	-.062 (3.5)	
Firm Size[d]	- 100-499 employees	.028 (1.3)	.051 (2.5)	
	- 500+ employees	.064 (3.6)	.107 (5.9)	
Industry[e]	- Forestry	.123 (2.2)	.122 (2.1)	
	- Mining	.165 (3.5)	.167 (3.2)	
	- Manufacturing			
	- Non Durable	-.064 (1.9)	-.083 (2.4)	
	- Durable	-.021 (0.6)	-.040 (1.1)	
	- Construction	.105 (3.3)	.065 (2.0)	
	- Trade	.113 (4.2)	-.095 (3.3)	
	- Finance	.056 (1.3)	.021 (0.5)	
	- Services	.049 (2.0)	-.059 (2.2)	
Occupation[f]	- Clerical	.183 (7.5)	-.177 (7.1)	
	- Sales	-.220 (7.2)	-.244 (7.9)	
	- Service	-.338 (13.0)	-.349 (13.9)	
	- Primary	-.161 (3.6)	-.210 (4.3)	
	- Processing	-.244 (5.6)	-.221 (5.2)	
	- Machining	-.011 (0.2)	-.054 (1.1)	
	- Fabricating	-.132 (4.2)	-.123 (3.9)	

- Construction	-.089 (2.6)	-.136 (3.9)	
- Transportation	-.160 (4.2)	-.220 (6.0)	
- Material Handling	-.160 (2.8)	-.258 (5.1)	
- Other	.219 (4.5)	-.166 (2.3)	
Union Status[g]			
- UU	.277 (12.5)	.238 (10.7)	-.040 (2.0)
- UN	.182 (7.2)	.048 (1.96)	-.135 (5.6)
- NU	-.042 (1.9)	.134 (5.8)	.179 (8.2)
Change in Tenure			.008 (2.3)
Change in Tenure Squared			-.0002 (1.1)
Change in Part-time to Full-time			-.028 (1.4)
Change in Full-time to Part-time			-.063 (3.2)
Change in Industry			.015 (5.9)
Change in Occupation			.013 (5.0)
Change in Firm Size			.018 (5.9)
R^2	.508	.491	.124
N	2438	2438	2438

[a] Omitted category under 25 years.
[b] Omitted category elementary or no education.
[c] Omitted category Ontario.
[d] Omitted category under 100 employees.
[e] Omitted category government services.
[f] Omitted category managerial and professional.
[g] Omitted category NN.

Appendix

Table 2A

Regression Results for Pension Equations

Independent Variable		*Coefficients and t-Values* *Job 1*		*Job 2*	
Constant		-.224	(3.6)	-.673	(4.0)
Age[a]	- 25-34	.014	(0.8)	-.025	(1.3)
	- 35-44	.007	(0.3)	.009	(5.4)
	- 45.54	-.038	(1.3)	-.029	(0.9)
	- 55-64	-.009	(0.2)	-.030	(0.6)
Female		.030	(1.7)	-.019	(1.0)
Married		-.021	(1.4)	.015	(1.6)
Education[b]	- High School	.021	(0.7)	.054	(1.8)
	- Some Post Sec.	.028	(0.8)	.087	(2.4)
	- Diploma	.030	(0.9)	.053	(1.5)
	- Degree	.050	(1.3)	.122	(3.1)
Region[c]	- Nfld	-.027	(0.7)	-.026	(0.6)
	- PEI	-.042	(0.8)	.001	(0.01)
	- NS	.046	(1.4)	-.019	(1.6)
	- NB	.023	(0.8)	.052	(1.7)
	- Que	-.042	(1.8)	-.022	(0.9)
	- Man	.060	(2.1)	.086	(2.8)
	- Sask	.006	(0.2)	.063	(2.3)
	- Alta	-.028	(1.3)	-.017	(0.7)
	- BC	-.039	(1.5)	-.025	(0.9)
Job Tenure		.022	(5.5)	.193	(2.0)
Job Tenure Squared		-.0003	(1.4)	-.078	(0.8)
Part-Time		-.091	(5.2)	-.042	(2.2)
Firm Size[d]	- 100-499 employees	.074	(3.3)	.072	(3.2)
	- 500+ employees	.114	(6.2)	.122	(6.3)
Industry[e]	- Forestry	-.093	(1.6)	-.044	(0.7)
	- Mining	.001	(0.02)	.132	(2.3)
	- Manufacturing				
	- Non Durable	-.051	(1.4)	.042	(1.1)
	- Durable	-.030	(0.8)	-.026	(0.6)
	- Construction	-.019	(0.6)	-.025	(0.7)
	- Trade	-.010	(0.3)	-.011	(0.3)
	- Finance	.031	(0.7)	.025	(0.6)
	- Services	-.056	(2.2)	-.028	(1.0)
Occupation[f]	- Clerical	-.028	(1.1)	-.017	(0.6)
	- Sales	-.088	(2.8)	-.054	(1.6)
	- Service	-.009	(1.3)	-.007	(0.2)
	- Primary	-.070	(1.5)	-.061	(1.1)
	- Processing	-.040	(0.9)	-.054	(1.2)
	- Machining	-.087	(1.8)	-.095	(1.8)
	- Fabricating	-.071	(2.1)	-.053	(1.6)

- Construction	-.021 (0.6)	-.016 (0.4)
- Transportation	-.108 (2.7)	-.011 (2.9)
- Material Handling	-.051 (0.8)	-.085 (1.6)
- Other	-.028 (0.5)	-.075 (0.9)
Ln Hourly Wage	.168 (8.1)	.138 (6.3)
Union Status[g]		
- UU	.346 (14.9)	.355 (14.5)
- UN	.212 (8.1)	.054 (2.1)
- NU	-.015 (0.7)	.224 (9.0)
R^2	.326	.278
N	2438	2438

[a] Omitted category under 25 years.
[b] Omitted category elementary or no education.
[c] Omitted category Ontario.
[d] Omitted category under 100 employees.
[e] Omitted category government services.
[f] Omitted category managerial and professional.
[g] Omitted category NN.

La mesure longitudinale des effets du syndicalisme sur le niveau et la dispersion des salaires ainsi que sur les régimes de pension

La recherche empirique canadienne sur les effets économiques des syndicats a beaucoup plus porté sur les taux de salaires que sur les avantages sociaux. Cependant, des études américaines récentes ont montré que les effets du syndicalisme sur les avantages sociaux, la dispersion des taux de salaires, la permanence, la productivité et les profits sont importants. L'objet du présent article vise à donner une image plus complète des effets des syndicats canadiens en évaluant les différences qui existent entre syndiqués et non-syndiqués dans les traitements, la dispersion de salaires et les régimes de pension. En théorie, on devrait s'attendre à ce que les syndicats exercent un effet à la hausse sur les salaires relatifs des travailleurs syndiqués et un effet à la baisse sur les écarts de salaires entre ces derniers. D'autres études soutiennent également que les syndicats augmentent la part de la rémunération globale qui est versée sous forme d'avantages sociaux, principalement les prestations de retraite. Alors que l'influence des syndicats sur le niveau des salaires résulte de leur pouvoir monopolistique, la dispersion des salaires et le niveau des prestations de retraite découlent davantage des préférences du syndiqué médian dont l'influence s'exerce selon le modèle de l'électeur médian.

D'une façon générale, l'appréciation des effets d'ordre économique des syndicats, lorsqu'on traite des données individuelles à l'aide de techniques statistiques du

type des moindres carrés ordinaires, a tendance à se compliquer à cause d'erreurs de sélection. En effet, les différences entre syndiqués et non-syndiqués sont estimées à partir de sous-populations différentes malgré les contrôles qu'on peut insérer dans les équations de régressions. Pour contourner ce problème, les présentes estimations portent sur des données longitudinales et font appel à une méthode dite «fonction à effet constant» qui consiste à examiner les salaires touchés par le même travailleur dans un emploi syndiqué et dans un autre qui ne l'est pas pour savoir s'il reçoit un salaire plus élevé dans le premier cas. Les données sont tirées d'une enquête sur l'activité du marché du travail effectuée par Statistique Canada en 1986 afin d'établir un portrait du comportement des Canadiens sur le marché du travail. Les principales estimations proviennent d'un échantillon de 2 438 individus qui ont occupé au moins deux emplois différents en 1986. Pour la majorité des salariés compris dans l'échantillon, il n'y a eu aucune modification à leur statut syndical durant cette période lorsqu'ils sont entrés au service de nouveaux employeurs, mais pour 18,8% d'entre eux, ce changement d'emploi a entraîné un changement de statut syndical.

Dans le cas des niveaux de salaires, les résultats indiquent que les salaires de ceux qui étaient non syndiqués et qui le sont devenus à la suite d'un changement d'emploi, ont augmenté de 17,6% tandis que les salaires de ceux qui ont perdu le statut de syndiqué ont baissé de 13,4%, ce qui laisse un écart net de 15,5% entre les deux catégories. L'influence des syndicats sur la dispersion des salaires, telle que mesurée par l'écart-type de l'équation des salaires sous forme logarithmique, présente moins d'uniformité. L'écart-type baisse légèrement pour les travailleurs dont le statut de non-syndiqué change pour celui de syndiqué mais il augmente subitement pour ceux dont le statut passe de syndiqué à non-syndiqué. En moyenne, il semble que les syndicats réduisent l'écart-type d'environ 5%. En contrôlant l'effet sur le niveau des salaires, l'impact du syndicalisme sur la probabilité de dispositions relatives aux fonds de pension s'établit à 19% une fois corrigé le biais de sélectivité. Toutefois, lorsqu'on y ajoute l'effet indirect des hausses de salaires attribuables au syndicalisme, l'effet global des syndicats sur les dispositions relatives aux fonds de pension s'élève à 22%.

Même si ces mesures longitudinales montrent que les syndicats ont un effet marqué et prévisible sur les niveaux et la dispersion des salaires ainsi que sur les dispositions conventionnelles relatives aux caisses de retraite, celles-ci peuvent être sous-estimées si les données longitudinales contiennent des erreurs de mesure. Pour contrer de tels biais, les effets des syndicats ont été estimés en se basant cette fois sur des données comparables en coupes instantannées. L'influence des syndicats, tirée tant de l'analyse des données longitudinales que de celle des données en coupes instantannées, s'avère presque identique en ce qui se rapporte aux fonds de pension, n'étant que marginalement plus élevée dans le cas des données des coupes instantannées. Par contre, les effets des syndicats sur la dispersion des taux de salaires sont beaucoup plus grands dans l'analyse en coupes que dans l'analyse longitudinale. Toutefois, les erreurs de mesure ne sont pas le seul facteur qui puisse expliquer cet écart dans l'estimation des effets sur la dispersion des salaires. Ces résultats incitent les auteurs à penser que leurs mesures longitudinales sont vraisemblablement correctes en ce qui a trait aux effets réels des syndicats sur les niveaux et la dispersion des salaires ainsi que sur les dispositions conventionnelles relatives aux caisses de retraite.