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Article abstract

The analysis in this paper is based on the 1986-87 Labour Market Activity Survey (LMAS) longitudinal data made available by Statistics Canada. It presents evidence of the effect of unionism on job tenure and job separation rates derived from regressions which control for the effects of wages, pension rights, firm size and other factors. Job separations and job tenure provide different perspectives of job attachment in that the former represents decision-making at a single point in time whereas the latter over extended periods. This paper also addresses the questions of bias inherent in selectivity and the simultaneous determination of wages and tenure.

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This paper presents evidence of the effect of union voice on the exit behaviour of Canadian workers. Theoretically, unions can affect exit behaviour not only through the context of the standard monopoly model of unionism but, more notably, through a distinctive collective voice which provides a mechanism for expressing preferences and resolving grievances. All other things being equal, workers with a voice institution for the resolution of workplace problems should resort to the exit option less frequently and maintain longer attachments with their firms. Thus, a prime prediction of the exit-voice model of unionism is that, even after controlling for the monopoly wage effects of unions, unionized workers should have lower job separation rates and longer job tenure than comparable nonunion workers.

In its most interesting and imposing extension, the exit-voice theory undermines the traditional view that unionism has a negative impact on productivity. The quantitative evidence may not be universally persuasive[1], but the more contemporary assertion is that unionism is associated with productivity gains, and that these gains are largely the outcome of a collective voice operating within structured internal labour markets. As Freeman and Medoff (1984) maintain, union voice enhances

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1 See Hirsch and Addison (1986) for a critical assessment of recent findings.

efficiency within firms not only by inducing more rational management practices but, more critically, by lowering training and recruitment costs associated with turnover. If unionism reduces turnover and if management responds by increasing firm-specific training, the resulting gains in productivity may offset the adverse productivity impact of restrictive work practices negotiated by unions. Thus, the effect of unionism on job separations and tenure is of paramount concern in the economic analysis of unions.

Econometric support for the exit-voice hypothesis has been provided primarily by Freeman (1980a, 1980b), Blau and Kahn (1983), Freeman and Medoff (1984) and Rebitzer (1986). Collectively, these studies, which are based on individual microdata regressions, examine the effects of union status on employee job separation rates or their counterpart, spells of job tenure, holding monopoly wage and nonwage compensation effects constant. A negative union effect on job separations (or a positive effect on job tenure) is interpreted as supportive of the exit-voice hypothesis. For the United States, there is strong evidence that unionism is associated with lower job separations and longer job tenure. Estimates of the impact of union voice on employment stability remain firm even after corrections for simultaneity bias (Rebitzer, 1986) and selectivity bias (Freeman, 1980a, 1980b).

However, it is somewhat difficult to reconcile the exit-voice model of unionism with Allen's (1984) finding that union workers have significantly higher rates of absenteeism than nonunion workers. If absenteeism reflects job dissatisfaction this finding suggests that union workers are more likely to resort to temporary exit rather than collective voice to correct unsatisfactory working conditions. Moreover, Ulman and Sorensen (1984) question both the results of the United States studies and the interpretation assigned to the unionism variable. They see the union variable as capturing not only the voice influence embedded in the grievance process but also the effects of other collective bargaining voices, such as the effects of strikes or strike threats. Indeed, when their aggregated quit equation is augmented by a strike variable, the estimated union coefficient loses its significance. This result, they argue, means that the exit-voice hypothesis is untenable when taken in its strongest form.

Finally, in a study of British labour markets, Shah (1985) shows that when pension rights, housing mobility costs, and the simultaneous determination of tenure status and wages are incorporated into the model there is virtually no support for the exit-voice hypothesis. Furthermore, job tenure in Britain appears unaffected by union monopoly wage gains, although it is strongly influenced by union monopoly effects on pension rights.

The analysis in this paper is based on the 1986-87 Labour Market Activity Survey (LMAS) longitudinal data made available by Statistics Canada. It presents evidence of the effect of unionism on job tenure and job separation rates derived from regressions which control for the effects of wages, pension rights, firm size and other factors. Job separations and job tenure provide different perspectives of job attachment in that the former represents decision-making at a single point in time whereas the latter over extended periods. However, the effect unionism has on job separations will be mirrored in its effect on tenure.[2] Together, the separations and tenure measures should provide a more comprehensive picture detailing the effect union status has on exit behaviour. This paper also addresses the questions of bias inherent in selectivity and the simultaneous determination of wages and tenure. Results will be derived for males and females and disaggregated by age. To our knowledge, this is the first study of the union effect on the exit behaviour of Canadian workers based on individual longitudinal microdata.

UNION VOICE AND EXIT BEHAVIOUR

Unions, according to Freeman (1976), exert their influence on the length of job tenure by reducing worker-initiated job separations. Both the monopoly and voice effects of unionism are expected to induce a reduction in voluntary quits. Workers are less likely to quit jobs that command high wages and related fringe benefits, and union monopoly induced wage and nonwage compensation gains have been found to be substantial. For example, Grant, Swidinsky and Vanderkamp (1987) report selectivity-corrected union wage premiums in the range of 13-16 percent for Canada, and Freeman (1981) reports union pension plan effects of 29 percent for the United States. Thus, even in the absence of a voice mechanism, the union impact on exit behaviour may be quite significant.

When workers do not have a mechanism for resolving industrial relations problems their recourse is to either to accept an unsatisfactory situation or exit the firm. Unionization, however, provides voice mechanisms for the expression and resolution of industrial relations problems. One of the most important voice mechanisms is the grievance/arbitration system, which in the event of a rights dispute offers

2 See Freeman (1980a) for a technical discussion of the relationship between separation probability and tenure.

workers a less costly alternative to quitting.[3] Even if a grievance is not resolved to the worker's advantage, the process may, at the very least, provide some nonpecuniary gratification in terms of equity and participation in an administrative process. That is, by providing a fair hearing the grievance system may diffuse the intensity of the grievance and reduce the incentive to quit.

More generally, to the extent that unions convey worker preferences during negotiations, union voice will be embedded in the collective bargaining process itself. Preferences regarding compensation structures, working conditions and industrial jurisprudence systems are all negotiable in unionized firms. Thus, workers who posses strong preferences for changes in conditions of employment may be compelled to quit if there is no vehicle for affecting the desired changes. However, should a bargaining alternative exist, these workers may remain with the firm and seek to affect these changes through a negotiated solution. Providing they can exert an effective collective voice through grievance procedures and the collective bargaining process, unions will have an influence on exit behaviour that goes beyond the influence they exercise through their monopoly effects on wage and nonwage compensation.

DATA AND MODEL

The Labour Market Activity Survey provides longitudinal information on the work patterns of Canadian workers from January 1, 1986, to December 31, 1987. For each individual in the sample the LMAS data file contains information on personal characteristics, employment and earnings patterns, and job characteristics, including union status, for up to five jobs in each year. The starting date for each job is recorded, and if a job is terminated in either 1986 or 1987, so is the stop date, along with the reason for job termination.

The analysis in this paper, except for the question of selectivity bias, will be confined to tenure in and separation from the first job (job 1). A total of 48,395 individuals were employed in job 1, but several restrictions were imposed on the original sample to make the individuals in the working sample as comparable as possible. The working sample was restricted to paid workers who were not students, had a non-zero wage rate, provided complete information on industry, occupation, union

3 Statutes in every jurisdiction in Canada require that collective agreements contain clauses which specify grievance procedures for the resolution of rights disputes, with arbitration as the final step. See Craig (1990: 131).

status, and firm size, and who held a job as of January 1, 1986. These restrictions reduced the working sample to 25,058 individuals.

Job tenure is defined as the length of the current or completed spell of tenure in job 1. A current, or incomplete, spell of tenure is measured by the number of years a worker has been employed by the same firm as of January 1, 1986, whereas a completed spell of tenure is measured by the number of years a worker whose job had been terminated during the period 1986-1987 remained with the same employer. Although the distributions of current and completed spells of tenure will generally have different means, both will, as Freeman (1980b) notes, embody the same information, and it is therefore immaterial whether the analysis focuses on the distribution of current or completed job tenure. Since there is an advantage in having as large a sample as possible when analysing age-sex subgroups, the emphasis in this study will be current rather than completed spells of job tenure. However, to ensure that the results are independent of the definition of tenure, similar equations will be estimated for both current and completed job tenure.

Job separations are defined as permanent breaks in job 1 attachments between employees and firms at any time during the period 1986-87. These separations are categorized into voluntary quits (or worker-initiated breaks) and *other* separations, a category which includes among other causes firm-initiated breaks in spells of tenure.[4] The distinction between quits and other separations will facilitate the analysis of the precise route by which unions influence job separations and, thus, spells of job tenure.

Table 1 gives the means and standard deviations of job tenure spells and job separation rates for the working sample disaggregated by union status and gender. On average, males have longer spells of tenure and lower separation rates than females within each union status classification, but unionized workers have longer tenure and lower separation rates than nonunionized workers for each gender. The mean spell of current tenure is 9.6 years for unionized workers compared with 5.8 years for nonunion workers. However, the large standard deviations relative to the means suggest that the distribution of current tenure spells is fairly wide. Average lengths of completed spells of tenure are shorter than the average lengths of current spells, indicating that separation rates

4 The distinction between quits and other separations is not very precise. Essentially, a separation is defined as a quit if the individual separated voluntarily from job 1, with the apparent intent of obtaining an alternative job. In practice, a separation was classified as a voluntary quit if the main reason for leaving job 1 was: (i) found a new job; (ii) working conditions; (iii) low pay; (iv) no opportunity for advancement; (v) worried about job security, reduction in hours, or layoff. Individuals who were separated from job 1 for any other reason were placed in the *other* category, even if the separation was voluntary for reasons such as illness or family responsibilities.

decline as tenure increases.[5] Nonunion workers tend to quit about 2.5 times more frequently than union workers, but the rate at which they separate for reasons other than quitting is only 1.3 times that for unionized workers.

TABLE 1

Means and Standard Deviations (in Parentheses) of Tenure and Separation Rates by Union Status

Sample	*Tenure Years*				*Separation Rates*			
	Current		*Completed*		*Quits*		*Other Separations*	
	Union	*Non-union*	*Union*	*Non-union*	*Union*	*Non-union*	*Union*	*Non union*
Total	9.6 (8.3)	5.8 (7.1)	8.6 (9.1)	4.6 (5.8)	.053 (.225)	.137 (.344)	.150 (.357)	.252 (.434)
Males	10.9 (8.8)	6.7 (7.8)	10.1 (10.4)	5.3 (6.8)	.051 (.219)	.136 (.342)	.132 (.338)	.228 (.419)
Females	7.7 (7.0)	4.9 (6.1)	7.0 (7.2)	3.9 (4.6)	.057 (.233)	.139 (.346)	.175 (.380)	.276 (.447)

Since the functional form of the tenure equation can be fairly complex when based on a realistic model of job tenure, this study adopts the common procedure of using the linear functional form to estimate tenure behaviour.[6] The linear job tenure equation is specified as: $T = BX + e$ where X is a vector of explanatory variables, including unionism, B is the corresponding vector of coefficients, and e is a mean zero error term. By contrast, the job separation decision is analysed in the framework of the following probability model: $P(S=1) = F(BX)$ where $S=1$ if a worker separates from job 1 during 1986-87 and zero otherwise, X is a vector of explanatory variables, and B is the corresponding vector of logistic coefficients.

The vector of explanatory variables common to both linear and logistic models includes pecuniary and non-pecuniary compensation, union status, firm size, and personal and industry-level control variables.[7] The distinctive collective voice of unions that makes its mark through the grievance process and collective bargaining is captured by a dichotomous union status variable. The influence of union monopoly compensation on exit behaviour is captured by a natural log of wages variable and a dichotomous pension plan coverage variable. Shah (1985)

5 Only in the case of a constant separation propensity with respect to tenure will the mean of incomplete spells equal the mean of completed spells. See Freeman (1980b), p. 34.

6 The complexities of the functional form are outlined in Freeman (1980a). Both Freeman (1980a, 1980b) and Rebitzer (1986) rely on the linear functional form to estimate job tenure equations.

7 All variables are defined in the appendix . In accordance with LMAS guidelines, the variables were rescaled by the appropriate weights before any analysis was conducted.

argues that since the pension capital of a worker is generally firm specific, workers with pension plan coverage will need greater inducement to leave a firm than workers without such coverage. Regrettably, the LMAS does not contain information on other nonwage compensation factors, and, since union workers have higher fringes than nonunion workers, this omission may generate an upward bias in the estimated union status coefficient. However, Freeman (1980b) shows that under reasonable assumptions the potential upward bias is negligible. Moreover, the influence of some of the nonwage compensation factors on exit behaviour may be captured by other control variables in the equations.

Rebitzer (1986) has argued that large employers tend to have more stable employees because they are more likely to institute internal labour markets that govern employment relations within the firm. Since large firms are more likely to be organized than smaller firms, failure to capture the effects of employer size on exit behaviour may produce a significant upward bias in the estimated union status coefficient. Accordingly, three firm size variables (20-99 employees, 100-499 employees, and 500 or more employees) are included in the list of explanatory variables to control for the impact of employer size on job attachment.

The impact of personal characteristics on exit behaviour is captured by age, sex, marital status, visible minority status, foreign/Canadian born status, and education control variables. Other controls include four regional variables, nine industry variables, nine occupation variables, and a fulltime status variable. Finally, the quit and other separations equations are augmented by current tenure, which Freeman (1980a) argues can be taken to represent unobserved factors that have a bearing on the stability of past and current employment in a given job.

EMPIRICAL ANALYSIS

Space constraints prevent the presentation of the full regression equations used to estimate the impact of unionism on job tenure and job separations. Accordingly, only the coefficients essential to the exit-voice hypothesis will be reported in this section. Full equations for the aggregate sample are reported in Table 1A in the appendix.. The discussion is organized into five subsections: (i) Current tenure results; (ii) Completed tenure results; (iii) Simultaneity corrections; (iv) Quits and other separations results; and v) Selectivity corrections.

Current Tenure Results

Equations 1-9, Table 2, present the union status, Ln wages, and pension plan coverage coefficients estimated from the current tenure equations. Equation 1 shows that, even after accounting for the effects of wages, pension plan coverage and other factors, unionization has a strong, positive effect on the length of current job tenure for all workers. Union voice alone increases tenure by one year, accounting for 26.3 percent of the observed differential in mean years of current tenure between union and nonunion workers.

TABLE 2

Union Status, Ln Wage and Pension Plan Coefficients Estimated from Job Tenure Regressions (t statistics in parentheses)

Equation	*Sample*	*Sample Size*	*Coefficients and t-statistics*					
			Union Status		*Ln Wage*		*Pension Plan*	
1	Total	25 058	1.007	(9.47)**	1.251	(12.48)**	2.152	(20.88)**
2	Males	13 557	1.225	(8.15)**	1.026	(6.98)**	2.004	(13.66)**
3	Females	11 501	.687	(4.73)**	1.263	(9.54)**	2.105	(15.95)**
4	Males 24 and under	1 418	.075	(0.50)	.302	(2.06)*	.311	(2.16)*
5	Males 25-44 years	8 215	.946	(6.25)**	1.706	(11.43)**	1.704	(11.29)**
6	Males 45 and over	3 924	2.216	(5.40)**	.747	(1.96)*	4.138	(9.52)**
7	Females 24 and under	1 625	-.105	(0.77)	.388	(3.21)**	.229	(1.90)
8	Females 25-44 years	6 974	.703	(4.45)**	1.655	(11.10)**	1.911	(13.43)**
9	Females 45 and over	2 902	.443	(1.04)	1.053	(2.91)**	3.645	(9.25)**
10[a]	Total	7 730	.977	(5.41)**	.996	(6.24)**	2.319	(13.84)**
11[a]	Males	3 790	1.198	(4.27)**	.751	(2.98)**	2.599	(9.69)**
12[a]	Females	3 940	.727	(3.28)**	1.098	(5.61)**	1.817	(9.17)**
13[b]	Males	13 557	1.442	(3.99)**	-3.818	(1.05)**	3.207	(4.83)**
14[b]	Females	11 501	1.112	(1.97)*	-.386	(0.11)	2.026	(4.46)**

a Dependent variable is completed tenure. Dependent variable in all other equations is current tenure. b 2SLS estimates

* significant at the 5 percent level **significant at the 1 percent level

However, the positive wage coefficient shows that unions also affect tenure through monopoly wage gains, but the monopoly wage effect is relatively weak when compared to the voice effect. For example, unions would need monopoly wage gains of 80 percent to have the same effect on tenure through the monopoly wage route that they exert through the union voice mechanism. By contrast, the introduction of a retirement pension plan would increase tenure by 2.1 times as much as a move from nonunion to union status. These findings are consistent with those reported by Freeman (1980a, 1980b). Assuming that the union wage and pension plan effects are 15 and 29 percent, respectively,[8] unionization will increase tenure by an additional .78 years through its monopoly wage and nonwage effects. Thus, overall effect of unionization on tenure, including both voice and monopoly effects, is 1.78 years, which accounts for 47 percent of the mean difference in current tenure between union and nonunion workers.

Turning to other variables (see Table 1A), the effect of firm size on job tenure is positive, significant and quite large. For example, workers in the largest sized firms (500 or more employees) remain with their firms 1.57 years longer than identical workers in firms with fewer than 100 employees. This finding is roughly consistent with that reported by Rebitzer (1986) for the United States. Since larger firms are more likely to be unionized than smaller firms, firm size can offer a partial explanation of the remaining observed difference in current tenure between union and nonunion workers.

Schooling is shown to have a significant negative effect on job tenure. This finding appears to support the argument that education increases general skills, enhances alternative job opportunities and thus makes workers more mobile. However, Freeman (1980b) argues that, because the regressions control for age rather than years in the job market, such an inference would be erroneous. At any given age, individuals with more schooling have less tenure than individuals with less schooling simply because they have had fewer years in the labour market since leaving school. If years since leaving school were held constant, the effect of education on tenure would be positive.[9] Not surprising, tenure increases with age. Married workers, females, and, especially, full-time workers all have longer tenure, while workers identified as visible minorities and workers born outside Canada tend to have shorter job tenure.

8 The wage effect estimate is taken from Grant, Swidinsky and Vanderkamp (1987) and the pension plan effect estimate from Freeman (1981).

9 Since the age and schooling variables are collapsed into several groups in the LMAS such an experiment can not be replicated in this study.

There are also regional, industrial and occupational effects on current job tenure. Workers in Western Canada have shorter spells of tenure than those in Eastern Canada, especially in the Atlantic and Québec regions. This finding is counter-intuitive in view of the higher rates of unemployment in the latter two regions. Job mobility is relatively high in finance, education, health, services and public administration, but it is especially high in construction where workers have two years less tenure than workers in the resource and manufacturing industries. Workers in more skilled occupations tend to have longer spells of current tenure, reflecting perhaps a greater investment in specific training.

When the results are disaggregated by gender, the estimated union coefficients in equations 2 and 3 show that union voice has almost twice the effect on male than female tenure (1.22 years versus .68 years). This gender differential is opposite to that reported in Rebitzer (1986), but it is a more plausible result. Since male workers generally have a greater investment in specific human capital and more seniority, quitting may be a more costly option for them, and in this respect union voice embedded in the grievance/arbitration procedure takes on more significance. But more important, in the context of the median voter model of unionism, it is the preferences of male workers that are likely to dominate and form the substance of collective negotiations. It is not surprising then that union voice should be a more important factor in male than female tenure decisions.

Equations 4-9 present estimates disaggregated by gender and age. Job tenure of young males and females (under 25 years) is unaffected by union voice. Any impact unionization has on years of tenure of younger workers is felt exclusively through monopoly wage effects, and even that is fairly modest. Similarly, for older females (45 years and over) tenure is related to unionization only through union monopoly wage and pension plan effects, but these effects are fairly potent, especially those operating through retirement provisions. However, job tenure of older males is strongly related to union voice. Finally, union voice, union monopoly wage effects and pension plan effects all have a positive impact on years of tenure for males and females aged 25-44 years.

Completed Tenure Results

Equations 10-12 show how union voice and union monopoly wage and pension plan effects influence the length of completed tenure for all workers and for males and females, separately. As expected, the results are basically unaffected by the change in the definition of tenure. The union coefficients estimated from current and completed tenure equations are only marginally different. However, the standard errors associated with the union status coefficients are almost twice as large in the completed tenure equations. This observation applies to the wage and pension plan coefficients as well. In the total sample, a move from nonunion to union status would increase the length of completed tenure by .98 years. To have an equivalent effect on completed tenure wages would need to rise by almost 100 percent. However, the introduction of a pension plan would increase completed tenure by 2.3 years, which is more than twice the effect that a change in union status would produce.

Two-Stage Least Squares Results

It has been shown that wages are important determinants of job tenure, but it is also generally recognized that job tenure is an important determinant of wages. Freeman (1980b) shows that, given plausible values for the tenure coefficient in the wage equation, simultaneity should generate an upward bias in the wage coefficient in the tenure equation. Since wages and unionism are correlated, this upward bias in the wage coefficient can be expected to produce a downward bias in the union status coefficient in the tenure estimates. The expected downward bias in the union status coefficient can be eliminated using two-stage least squares (2SLS) estimation procedures.[10]

Tenure equations estimated by 2SLS for males and females separately are reported as equations 13 and 14, respectively. As expected, the coefficients on union status increase in magnitude but become statistically less significant. The effects of pension plan coverage remain strong, but the wage coefficients become negative and insignificant. These results, which generally hold when the sample is disaggregated by age and gender, are consistent with those reported in Rebitzer (1986).

Our findings suggest that it is unlikely that the estimated effects of union status on job tenure reported in equations 1-9 can be attributed to the bias created by the simultaneous determination of wages and tenure. In fact, as Freeman (1980b) notes, failure to adjust for the simultaneity problem imposes an even more exacting test of the effect of union status on job tenure.

10 Essentially, the variables that affect tenure should also affect wages. Additional exogenous variables employed in constructing the instrumental variables include language and training. However, pension plan coverage is excluded.

Quits and Other Job Separations Results

Table 3 presents union status coefficients estimated from logistic quits and other separations equations that include tenure as an independent variable.[11] As Freeman (1980b) notes, the inclusion of tenure in the logistic equations creates a more difficult test of the impact of union voice on job separations. Union and nonunion workers whose job separation behaviour is now being compared not only have the same wages, pensions, firm size, and other industry and personal characteristics, they also have the same job tenure and thus the same stability history. In this respect, the specification of the logistic quits and other separations equations rules out the problem of selectivity bias.

TABLE 3

Union Status Coefficients Estimated from Quit and Other Separations Logistic Equations (t statistics in parentheses)

Population Sample	Quits	Other Separations
Total	-.374 (6.08)**	-.113 (2.49)*
All Males	-.361 (4.39)**	-.140 (2.21)*
All Females	-.377 (3.96)**	-.124 (1.84)
Males - 24 and under	-.486 (2.47)*	.319 (1.73)
Males - 25 to 44 years	-.316 (3.23)**	-.205 (2.36)*
Males - 45 and over	-.591 (2.07)*	-.155 (1.42)
Females - 24 and under	-.433 (2.03)*	-.066 (0.36)
Females - 25 to 44 years	-.311 (2.65)**	-.142 (1.66)
Females - 45 and over	-.682 (2.55)*	-.040 (0.29)

* significant at the 5 percent level
** significant at the 1 percent level

The union status coefficients derived from the quit equations are negative, significant and strongly supportive of the exit-voice hypothesis. Even after union monopoly effects have been taken into consideration, union workers still tend to quit less frequently than nonunion workers. As predicted, wages and pension plan coverage also exert negative effects on the propensity to quit, but these effects are fairly modest when contrasted to the effect of unionism. For example, in the total sample, a move from nonunion to union status would reduce the propensity to quit by 23.7 percent, whereas a 15 percent wage gain would reduce the quit rate by only 3.5 percent and the introduction of a pension plan by 14.0 percent. Union voice significantly reduces the quit probability for each

11 See columns 3 and 4, Table 1A, for total sample equations.

age-sex group in the sample, but as Blau and Kahn (1983) have already noted, the magnitude of this reduction is greatest for older workers. By contrast, the estimated effect of union status on job separations other than voluntary quits is very weak. Although generally negative, the union status coefficients derived from the other separations equations are much smaller in magnitude than those derived from the quit equations and they are typically insignificant. Only the coefficient for prime-aged males is negative and significant. Although a modest negative union effect on other separations cannot be ruled out, the evidence strongly supports Freeman's (1980b) contention that the union effect on job tenure works primarily through a reduction in voluntary quits rather than a reduction in involuntary separations.

SELECTIVITY BIAS

It has been argued that the observed inverse relation between unionism and exit behaviour is the result of the influence unions exert on employee quit behaviour. However, this relationship may simply reflect selectivity in union membership: Unions may prefer to organize workers who are innately more stable. Union and nonunion workers exhibit different exit behaviour not because unions alter quit behaviour but because union workers have unobserved characteristics that lead to lower quit rates.

When longitudinal data is available the fixed-effect-function approach provides a convenient method of eliminating selectivity bias from the estimated union coefficients.[12] In essence, the fixed-effect-function model eliminates selectivity bias by comparing the exit behaviour of the same individual, and not just a statistical surrogate, in union and nonunion employment. Such a comparison removes the impact of unobservable personal characteristics (assumed to be invariant between the two union states) on exit behaviour and isolates the behavioral impact of unionism.

In the context of this study, the fixed-effect-function model relates changes in the length of completed job tenure as individuals move between jobs to changes in union status and a select number of variables expressed as first differences.[13] The simple first difference equation

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12 See the study by Mellow (1981) and the more recent study by Grant, Swidinsky and Vanderkamp (1987) for an application of the fixed-effect-function approach in the estimation of union wage gains.

13 Freeman (1980a, 1980b) tested for selectivity using a fixed effect logit model of quit behaviour. He concluded that the union-quit tradeoff is due largely to the impact of unionism on worker behaviour rather than to the propensity of stable workers to be organized. Unfortunately, the LMAS data does not appear suitable to the fixed effect logit approach. However, a fixed-effect-

which can be estimated by ordinary least squares is specified as: $T_{i2} - T_{i1} = a + b(U_{i2} - U_{i1}) + c(X_{i2} - X_{i1}) + e$ where T_{ij} is the length of completed job tenure for the ith individual in job j (j = 1, 2), U is union status, X is a vector of other variables, a, b, c are coefficients to be estimated, and e is a random error term uncorrelated with union status. Changes in union status as an individual moves from job 1 to job 2 are classified as UN (union to nonunion), NU (nonunion to union), UU (union to union), and the omitted category NN (nonunion to nonunion). Other variables in the first difference equation include changes in ln wages, pension plan provisions, firm size, fulltime status, industry, and occupation. Estimates of the selectivity-adjusted average union effect on completed job tenure can be obtained from this tenure-change equation by averaging the absolute values of the coefficients of union leavers (UN) and union joiners (NU).

TABLE 4

Union Status Change Coefficients Derived from Fixed-Effect Function Estimates of Completed Tenure (t statistics in parentheses)

Sample	*Sample Size*	*Coefficients*				*Average Effect*	*Cross-section Estimate*
		UN		*NU*			
Total	2616	-1.159**	(4.46)	-.024	(0.08)	.579	.977
Male	1 466	-1.347**	(3.45)	-.178	(0.40)	.673	1.198
Female	1 150	-.692*	(2.12)	.166	(0.46)	.346	.727

* significant at the 5 percent level
** significant at the 1 percent level

Estimates of the fixed effect model are based on 2 616 individuals who became separated from job 1, moved to job 2, and also became separated from the second job during the 1986-87 period. The mean length of completed tenure was 2.01 years in job 1 and .66 years in job 2. Most individuals (66.4 percent) were nonunion in both jobs (NN), 12.7 percent were union in both jobs (UU), while 7.6 percent changed from nonunion to union (NU) and 10.7 changed from union to nonunion (UN).

The estimated UN and NU coefficients are presented in Table 4. There is a significant reduction in the length of completed tenure for individuals who move from union to nonunion jobs. However, completed tenure does not change significantly for individuals moving from nonunion to union jobs. Estimates of the selectivity-corrected union effect on completed tenure, obtained by averaging the absolute values of the UN and NU coefficients, are given in column 3. The fixed-effect-function estimates are approximately 50 percent lower than those

function approach applied to completed spells of tenure in two consecutive jobs seems a more simple and equally acceptable alternative for dealing with selectivity.

derived from cross-section regressions, suggesting that selectivity bias may be a significant factor in the estimated effects of unionism on exit behaviour.

However, the fixed-effect-function results should be interpreted with a degree of caution. The number of individuals in the sample who changed union status when moving from job 1 to job 2 is fairly small, and the LMAS may contain considerable *noise* from the various sources of non-sampling error.[14] Moreover, the time frame allowed for separation decisions in job 2 is relatively short, which may mean that the sample of 2 616 individuals who became separated from both jobs is not representative of the population. But even if the fixed-effect-function results are reliable, they do not negate the finding that union voice has a significant effect on workers attachment to their firms.

OBSERVATIONS AND CONCLUSIONS

This study has examined the effect of unionism on the job attachment of Canadian workers. Cross-section regressions derived from information on 25,058 individual workers show that trade unionism is associated with significantly lower probabilities of job separation and significantly longer spells of tenure. Although union monopoly wage and pension plan effects are important contributing factors, the main reason why unionism is associated with increases in worker attachment to firms is the change in worker quit behaviour induced by union voice. This finding is supportive of the conclusions drawn from studies for the United States, especially those by Freeman (1980a, 1980b), Blau and Kahn (1983) and Rebitzer (1986).

In 1986, Canadian workers accumulated an average of 7.4 years of current tenure, with workers in organized firms averaging 3.8 years more than workers in unorganized firms. Our findings show that unionism explains 26.3 percent of the unadjusted differential in average spells of current tenure between union and nonunion workers. Comparatively, union voice is much more effective than monopoly wage gains in prolonging workers attachment to firms. A move from nonunion to union status would increase current tenure by one year. To produce an equivalent effect, wages would have to rise by an estimated 80 percent. However, the introduction of pension plan coverage would increase current tenure by twice as much as a change to union status. The effect of union voice on quits is equally strong, but there is convincing

14 The LMAS Microdata User's Guide urges analysts to approach results involving changes in certain variables with caution because of the possibility of non-sampling error.

evidence that unionism is not a decisive factor in separation decisions which are not voluntary quits.

Correcting for the bias created by the simultaneous determination of wages and tenure does not diminish the estimated impact of union voice on the tenure decision. On the other hand, a fixed-effect-function approach to correct for selectivity bias yields estimates of union voice effects which are significantly lower than estimates derived from cross-section analysis. The selectivity-adjusted estimates, however, should be treated with caution.

Our findings have important implications for the economic assessment of unionism. In addition to their socially undesirable monopoly wage effects and restrictive work practices, unions have a significant socially beneficial voice effect on labour markets. These socially constructive and destructive aspects of unionism must be weighed against each other in any social evaluation of the economic effects of trade unions.

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APPENDIX

Definition of Variables

Current Tenure:	Number of years a worker has been employed by same firm as of January 1, 1986 in job 1.
Completed Tenure:	Number of years a worker separated permanently from job 1 during 1986-87 had been employed in that job.
Quits:	Voluntary separation from job 1 when the main reason for leaving the job was: (i) found a new job; (ii) working conditions; (iii) low pay; (iv) no opportunity for advancement; and (v) worried about job security, reduction in hours, or layoff.
Other Separations:	Separations from job 1 for all reasons other than those listed under "Quits" above.
Union Status:	Coded as one if a union member or if wages covered by collective agreement, otherwise zero.
In Wage:	The natural log of the hourly wage rate.
Pension Plan:	Coded as one if covered by a pension plan connected with the job, otherwise zero.
Firm Size:	A set of four dummy variables representing (i) 19 or less, (2) 20 to 99, (3) 100 to 499, and (4) 500 or over persons employed at all locations in Canada by the firm, with 19 or less serving as the reference firm size.
Female:	Coded as one if female, otherwise zero.
Age:	A set of seven dummy variables representing age categories (1) under 19 years, (2) 20-24, (3) 25-34, (4) 35-44, (5) 45-54, 6) 55-64, 7) 65-69, with 25-34 serving as the reference age category.
Married:	Coded as one if married, otherwise zero.
Schooling:	A set of five dummy variables representing educational categories (1) none or elementary, (2) high school (some or completed), (3) some post-secondary (4) post-secondary certificate or diploma, (5) university, with high school serving as the reference schooling category.

Visible Minority:	Coded as one if identified as visible minority, otherwise zero.
Canadian Born:	Coded as one if born in Canada, otherwise zero.
Full Time:	Coded as one if employed full time, otherwise zero.
Region:	A set of five dummy variables representing regions (1) Atlantic, (2) Québec (3) Ontario, (4) Prairies, (5) B.C., with Ontario serving as the reference region.
Industry:	A set of ten dummy variables representing the following industries (LMAS codes in brackets): (1) Resource [01-08], (2) Manufacturing [09-28], (3) construction [29-30], (4) transportation [31-32], (5) communications and utilities [33-34], (6) trade [35-36], (7) finance [37-39], (8) education and health [40-41], (9) services [42-47, 52], (10) administration [48-51]. The resource sector serves as the reference industry.
Occupation:	A set of ten dummy variables representing the following occupations (LMAS codes in brackets): (1) managerial [01-03], (2) professional [04-08], (3) education [09-12], (4) health [13-15], (5) clerical [16-22], (6) sales [23-24], (7) services [25-28], (8) primary [29-33], (9) processing [34-42], (10) other [43-50], with other serving as the reference occupation.
Language:	Coded as one if English was first language spoken, otherwise zero.

TABLE 1A

Tenure and Job Separation Equations
Estimated for the Total Sample
(t-statistics in parentheses)

Independent Variable	*Equation*			
	Current Tenure (1)	*Completed Tenure (2)*	*Quits (3)*	*Other Separations (4)*
Constant	-6.957** (9.01)	-5.425** (4.45)	1.371** (3.48)	.478 (1.54)
Union Status	1.007** (9.48)	.977** (5.41)	-.374** (6.08)	-.113** (2.49)
Ln Wage	1.251** (12.49)	.996** (6.24)	-.394** (7.81)	-.196** (4.87)
Pension Plan	2.152** (20.88)	2.319** (13.84)	-.218** (3.93)	-.431** (9.94)
Firm Size				
– 20 - 99 employees	-.076 (0.59)	-.127 (0.68)	.134* (2.21)	-.214** (4.24)
– 100 - 499 employees	.560** (3.99)	-.117 (0.54)	-.001 (0.02)	-.181** (3.20)
+ 500 and over employees	1.566** (12.59)	1.443** (7.84)	-.039 (0.62)	-.091 (1.85)
Female	-.769** (7.66)	-.374* (2.34)	-.294** (5.65)	.305* (7.33)
Age				
– under 19	-1.362** (3.47)	-.969* (2.24)	.538** (3.86)	.227 (1.66)
– 20-24	-1.282** (8.87)	-.846** (4.46)	.251** (4.16)	.196** (3.66)
– 35-44	3.021** (28.67)	1.646** (9.10)	-.248** (4.34)	-.378** (8.11)
– 45-54	6.241** (50.58)	3.625** (15.57)	-.763** (9.06)	-.337** (5.94)
– 55-64	10.121** (67.21)	10.990** (44.72)	-.958** (7.89)	.663** (10.62)
– 65-69	11.885** (29.98)	11.907** (23.88)	-1.458** (4.02)	1.401** (10.20)
Married	.297** (3.12)	.060 (0.41)	-.128** (2.68)	0.58 (1.48)

Independent Variable	*Equation* Current Tenure (1)	Completed Tenure (2)	Quits (3)	Other Separations (4)
Schooling				
– elementary or less	.627** (4.14)	.183 (0.73)	-.357** (3.45)	.160** (2.66)
– some post-secondary	-.643** (4.46)	.042 (0.19)	.322** (4.66)	.001 (0.02)
– post-secondary certificate	-1.254** (10.46)	-.771** (4.07)	.294** (4.79)	.084 (1.69)
– university	-1.994** (14.09)	-1.351** (5.89)	.384** (5.20)	.118 (1.94)
Visible Minority	-.464* (2.51)	-.313 (1.03)	-.153 (1.55)	-.076 (0.97)
Canadian Born	.768** (6.72)	.247 (1.32)	.002 (0.03)	.038 (0.79)
Full Time	1.082** (8.03)	.948** (4.97)	-.082 (1.23)	-.145** (2.85)
Region				
– Atlantic	.482** (2.81)	.706* (2.51)	-.701** (7.01)	.160* (2.27)
– Québec	.578** (5.66)	.501** (2.98)	-.607** (10.65)	.102* (2.38)
– Prairie	-.757** (6.44)	-.510** (2.81)	-.414** (6.77)	.256** (5.38)
– B.C.	-.499** (3.61)	.063 (0.29)	-.315** (4.33)	.305** (5.46)
Industry				
– Manufacturing	.198 (0.70)	.542 (1.19)	.260 (1.55)	-.947 (0.82)
– Construction	-1.987** (5.86)	-.760 (1.50)	.194 (1.01)	.594** (4.58)
– Transportation & storage	.813* (2.51)	1.814** (3.43)	.368 (1.90)	.038 (0.29)
– Communication & utilities	.683* (2.07)	1.724** (2.94)	.229 (1.10)	-.248 (1.73)
– Trade	-.457 (1.58)	.172 (0.37)	.386* (2.31)	-.202 (1.75)
– Finance	-1.002** (3.20)	-.026 (0.05)	.161 (0.89)	-.366** (2.84)
– Health & Education	-1.215** (4.07)	-.318 (0.65)	-.100 (0.56)	-.266* (2.19)
– Services	-1.172** (4.02)	-.509 (1.12)	.606** (3.65)	-.005 (0.04)
– Public Administration	-.867** (2.92)	-.001 (0.00)	.169 (0.93)	-.099 (0.81)
Occupation				
– Managerial	.805** (4.41)	1.673** (5.59)	.317** (3.15)	-.183* (2.39)
– Professional	.379 (1.67)	.689 (1.80)	-.076 (0.59)	-.143 (1.47)

Independent Variable	Equation Current Tenure (1)	Completed Tenure (2)	Quits (3)	Other Separations (4)
– Teaching	1.959** (7.39)	1.043* (2.23)	.056 (0.33)	-.036 (0.32)
– Medicine	1.195** (4.70)	.841 (1.89)	.425** (2.86)	-.435** (3.92)
– Clerical	.328 (1.91)	.497 (1.77)	.291** (3.06)	-.185** (2.61)
– Sales	-.641** (2.98)	.173 (0.53)	.387** (3.57)	-.207* (2.35)
– Services	-.100 (0.51)	.569 (1.84)	.393** (3.73)	-.150 (1.88)
– Primary	.862* (2.28)	.566 (1.01)	.323 (1.56)	-.010 (0.06)
– Processing	-.025 (0.15)	.586* (2.07)	.099 (1.02)	-.046 (0.67)
Current Tenure	—	—	-.076** (14.26)	-.024** (8.60)
R^2	.380	.421	—	—
– Ln Likelihood	—	—	8.872	12.475
N	25.058	7.730	25.058	25.058

* significant at the 5 percent level
** significant at the 1 percent level

Le syndicalisme et l'intérêt des travailleurs canadiens envers leur emploi

Cet article présente des résultats confirmant l'effet «voice», ou la voie d'expression offerte par le syndicalisme, sur le comportement de défection des travailleurs canadiens. Théoriquement, les syndicats peuvent influencer le comportement de défection non seulement *via* le modèle de monopole syndical, mais aussi grâce à un mécanisme d'expression des préférences et de résolution des griefs. Toutes autres choses égales par ailleurs, les travailleurs bénéficiant d'un tel mécanisme institutionnel pour la résolution des problèmes vécus sur les lieux de travail auraient moins fréquemment recours à la défection (option «exit») et maintiendraient plus longtemps leur lien d'emploi avec l'entreprise. Ainsi, une première prédiction découlant de ce modèle est à l'effet que les travailleurs syndiqués auraient des taux de défection au travail plus bas et des liens d'emploi plus longs que ceux de travailleurs non syndiqués de même type. et ce, en tenant compte du contrôle des syndicats sur les salaires. Le support économétrique utilisé dans les hypothèses de type «exit-voice» est toujours dominant aux États-Unis, mais il n'y a pas eu, à toutes fins pratiques, de support à la vérification d'hypothèses impliquant l'usage de données britanniques. A ce jour, il n'existe toujours pas d'étude portant sur l'effet des syndicats sur le comportement de défection des travailleurs au Canada.

La présente analyse est basée sur des données longitudinales colligées par Statistiques Canada en 1986-1987. L'échantillon utilisé se limite à 25 058 salariés qui occupaient un emploi en date du 1er janvier 1986. Le comportement de défection vis-à-vis de ce premier emploi est considéré en tenant compte des quatre éléments suivants: (1) la continuité du lien d'emploi; (2) la terminaison du lien d'emploi (1986-1987); (3) la défection volontaire; (4) les autres formes de défection. En moyenne, les travailleurs syndiqués ont des liens d'emploi plus longs et un taux de défection plus bas que les travailleurs non syndiqués.

Les principales variables explicatives renvoient aux statuts des syndicats ainsi qu'aux régimes de retraite et d'assurance salaire. La variable dichotomique du statut syndical englobe le mode particulier d'expression collective des syndicats rendu possible par la procédure de règlement des griefs et la négociation collective, alors que les deux variables suivantes servent à évaluer l'impact du contrôle syndical sur les mécanismes compensateurs en cas de défection. D'autres variables contrôle incluent la taille de l'entreprise, le type d'industrie et d'occupation, la région ainsi qu'une gamme de caractéristiques personnelles. Les équations de régression sont évaluées en fonction de diverses catégories d'âge et de sexe.

En général, les statuts syndicaux, les régimes de retraite et d'assurance salaire ont tous un effet marquant et positif sur la durée du lien d'emploi. L'effet du contrôle des salaires est relativement faible une fois comparé à l'effet «voice» des syndicats. Par exemple, un contrôle de 80% des gains salariaux pour l'ensemble des travailleurs représentés par un syndicat serait nécessaire pour obtenir le même effet sur l'attachement organisationnel que celui obtenu grâce à l'option «voice». En contraste, l'introduction d'un régime de retraite accroîtrait l'attachement organisationnel de 2,1 fois plus qu'à l'occasion du changement du statut non syndiqué pour le statut syndiqué. Un régime de retraite pour employés constitue un moyen plus efficace pour assurer l'attachement organisationnel que tout gain salarial ou que l'obtention du statut syndical.

Bien que demeurent certaines variations, les résultats exposés ci-dessus tendent à conserver leur validité lorsque les données sont ventilées selon l'âge et le sexe. De la même façon, les résultats demeurent fondamentalement les mêmes lorsque la variable dépendante est redéfinie comme la durée du lien d'emploi complété (1986-1987). Comme prévu, le fait d'ajuster les données en fonction de la simultanéité de la détermination des salaires et du lien d'emploi accorde aux coefficients estimés une signification moindre, mais cela ne renverse pas pour autant le résultat premier à l'effet que l'option «voice» a un effet positif marquant sur la durée du lien d'emploi.

Les résultats découlant de l'analyse des défections montrent, même après avoir pris en considération les effets de compensation, que les travailleurs syndiqués ont moins tendance à abandonner leur emploi que leurs vis-à-vis non syndiqués. Les régimes de retraite et d'assurance salaire exercent également des effets négatifs sur la propension à la défection, mais les effets sont passablement modestes lorsque comparés avec ceux résultant de l'option «voice». D'autre part, l'importance du facteur «statut syndical» *dérivé de l'analyse des «autres types de défection» est non seulement moindre* en grandeur, mais s'avère aussi, plus généralement, non significatif. Ces résultats suggèrent que l'effet du syndicalisme sur le lien d'emploi tend principalement à réduire les défections volontaires plutôt que les défections involontaires.

Néanmoins, la relation observée entre le syndicalisme et le comportement de défection peut simplement refléter une forme de sélectivité découlant d'une préférence chez les syndicats vis-à-vis de la syndicalisation d'une main-d'oeuvre foncièrement plus stable. Dans cette étude, le problème du biais induit par la sélectivité est éliminé grâce à un modèle d'«effet fixé» qui relie les changements dans la durée du lien d'emploi (abandonné volontairement pour un autre emploi) avec les changements dans le statut syndical et un nombre restreint de variables exprimant les différences premières. Considérant les données traitées suivant ce modèle, l'effet du statut syndical sur le lien d'emploi complété en 1986-1987 est approximativement à demi moindre que celui évalué par les méthodes de régression. Cela suggère que le biais introduit par la sélectivité peut être un facteur significatif dans l'estimation des effets du syndicalisme sur le *comportement de défection. Toutefois, étant donné la taille relativement* réduite de l'échantillon à partir duquel ils furent déduits, les résultats du modèle d'«effet fixé» devraient être traités avec passablement de réserves. Mais bien que ces derniers résultats soient fiables, ils ne mettent pas en doute le fait que l'option «voice» ait un effet significatif sur l'attachement organisationnel.