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Abstract:

We provide evidence on the potential for reforms in labour law to reverse deunionization trends by relating an index of the favorability to unions of Canadian provincial labour relations statutes to changes in provincial union density rates between 1981 and 2012. The results suggest that shifting every province's 2012 legal regime to the most union-friendly possible could raise the national union density by up to 7 percentage points in the long run. This effect appears driven by regulations related to the certification of new bargaining units, the negotiation of first contracts and the recruitment of replacement workers. The effects of reform are largest for women, particularly university-educated women employed as professionals in public services. Overall, the results suggest a limited potential for labour relations reforms to address growing concerns about labour market inequality.

1. Introduction

According to data from the OECD, union membership as a proportion of the workforce declined in all but five OECD countries between 1980 and 2010.[1] In Australia, New Zealand, the U.K. and the U.S. the declines were particularly dramatic; in each case the national rate more than halved over the three decades. While there are sharply diverging views on whether a smaller role for unions in labour markets is desirable, there is little disagreement that it matters. On the one hand, evidence suggests that unions reduce corporate profits, investments and dampen employment growth. On the other hand, unions have clear beneficial impacts on the wages, fringe benefits and working conditions of unionized workers.[2] Moreover, in a recent analysis of U.S. data, Western and Rosenfeld (2011) consider the possibility of spillover effects of union wage outcomes on nonunion workers and find that once these spillovers are accounted for, the long-term decline of unions in the U.S. can account for up to one-third of U.S. growth in wage inequality.

For the union movement itself, there is tremendous interest in understanding to what extent deunionization has been a consequence of government policies influencing the balance of power in union-management relations, as opposed to an inevitable development driven by broad globalization and deindustrialization trends. The relative stability of unionization rates in Canada, despite its legal, political and cultural similarities and close economic ties to the U.S., suggests that the phenomenon was not inevitable. Indeed, comparing survey and opinion poll data, Riddell (1993) finds that the vast majority of the Canada-U.S. gap in unionization rates cannot be accounted for by structural economic differences or social attitudes and infers that the gap is most consistent with differences in legal regimes. Following up on this idea, Johnson (2002) exploits variation in labour relations laws across Canadian jurisdictions and finds that requiring secret-ballot votes in representation elections, as is normally required in the U.S., as opposed to automatic certification through card checks, reduces the likelihood that an application is successful by 9 percentage points. Focusing on the British Columbia case, Riddell (2004) finds that this effect of secret-ballot elections is largely attributable to the relative effectiveness of management opposition tactics when elections are required. Most recently, Bartkiw (2008) examines the effect of two substantial legal regime changes in Ontario in the 1990s and consistent with the earlier evidence concludes that labour laws regulating how unions are formed and operate are critically important in influencing the ability of unions to organize new bargaining units.

[1] Exceptions are Belgium, Chile, Iceland, Norway and Spain. The data are from: http://stats.oecd.org/ and measure the proportion of the workforce that are union members.

[2] Despite widespread perceptions that unions reduce workplace productivity and overall earnings inequality in society, the evidence of both effects is, in fact, quite mixed. For a review of the evidence of the economic impacts of unions, see Kuhn (1998).

The obvious question that follows from this evidence is: what is the potential for union – friendly reforms in labour relations laws to reverse deunionization trends? Unfortunately, the current evidence falls short in informing this potential in three key respects. First, the empirical analysis is almost exclusively concerned with the probability that a union certification application is successful, conditional on an application being submitted in the first place. But if legal reforms affect the likelihood of success, they are likely to also affect the flow of applications. Moreover, changes in labour relations laws may affect employers' perceived threat of unionization or their relative bargaining power, thereby influencing employment levels or creating incentives for employers to relocate where laws are seen to be less union friendly. The general equilibrium effects of legal reforms could, therefore, be very different from those implied by certification success rates conditional on application submission. Second, the existing research tends to be focused on particular features of labour relations legislation, for example the distinction between automatic certification of new bargaining units through card checks versus secret-ballot elections. But in informing policymakers, it is critical to know what laws matter most and by how much reforming a series of laws can be expected to raise union density rates. Finally, the current evidence tells us nothing about what types of workers are most likely to be affected by labour relations reforms. The social welfare implications of legal reforms, however, depend critically on who is affected. For example, whether legal reforms lead to an increase or decrease in wage inequality depends critically on whether gains in unionization rates are concentrated in the lower or upper end of the wage distribution.

The only study we are aware of to directly examine the link between labour relations laws and union density rates is Freeman and Pelletier (1990), who relate an index of the favorableness of Britain's labour relations regime between 1945 and 1986 to the British national union density rate. The advantage of the Canadian labour relations system in doing this analysis is that the legislative jurisdiction in Canada primarily lies at the provincial level, rather than the national level as it does in the U.K. and U.S., thereby allowing one to separately identify policy effects from the effects of unobserved events or circumstances correlated with the timing of legal changes. Moreover, given the contentiousness of these laws, changes in the political stripes of governing provincial parties has historically resulted in significant swings across Canadian provinces and over time in the favorableness of provincial laws to unions. In this paper we estimate the effect of a set of 12 labour relations laws on provincial union density rates by patching together a series of nationally-representative household surveys spanning the 1981-2012 period. These data allow us to not only estimate overall union density rates at the provincial level, but also to distinguish these rates between industries, occupations, education groups and gender thereby enabling us to inform what types of workers and workplaces are most likely to be affected by changes in labour relations laws.

Using a dynamic feasible generalized least squares (FGLS) estimator that conditions on a full set of province and year fixed effects, as well as provincial-level measures of unemployment, inflation, the manufacturing share of employment, and popular opinions of unions, the estimates suggest that shifting every Canadian province's current legal regime to the most union-friendly possible (within the set of 12 laws considered) could, in the long run, raise the national union density rate by up to 7 percentage points, from its current value of 30%. Distinguishing between laws, we find that rules regarding the certification of new bargaining units, the negotiation of first contracts and the recruitment of replacement workers following legal strikes are particularly influential. Distinguishing the effects across industries, occupations, and education and gender groups, we find that legal reforms are likely to have the greatest impact among relatively educated women employed as professionals in public services, broadly defined. However, examining our data at a finer level of detail, we also find evidence of significant gains among relatively uneducated women in private service-producing industries. Overall, the results suggest a limited potential for labour relations laws to counter rising wage inequality.

The remainder of the paper is organized as follows. In the following section we describe in detail our empirical methodology for estimating the effects of legal reforms on provincial-level union density rates. In the third section, we describe the data we use to estimate the model and in the fourth section we discuss our findings. The paper concludes with a discussion about the practical policy relevance of our findings.

2. Methodology

Modelling the decision of a union to invest the resources necessary to organize a new bargaining unit involves an optimization problem in which unions compare the relative marginal costs and benefits of additional membership (Pencavel 1971). By influencing these costs and benefits, small changes in the legal environment can potentially alter optimal behaviour, thereby initiating organizing activities in a particular workplace and, in turn, the per-period flow of workers transitioning from the nonunion to union sector.[3] Ideally, we could estimate the effect of legal changes directly on these worker-level flows. However, this requires large samples of longitudinal microdata with information on workers' union status going back to at least the early 1990s, when the key historical variation in laws began. To our knowledge,

[3] Similarly, legal changes could influence the marginal cost of decertifying an existing bargaining unit, which would instead increase union-to-nonunion transitions. However, since decertifications are relatively rare, we focus our discussion on certifications.

suitable data do not exist for Canada.[4] We can, however, estimate provincial union density rates for particular types of workers using repeated cross-sections of nationally-representative household survey data. But this requires that we think carefully about how changes in the per-period flows of workers in and out of the union sector resulting from changes in labour relations laws affect unionization rates in the long-run.

Assuming for simplicity a two-state national labour market in which all workers are employed in either the union or nonunion sector, the union density rate in any year t can be expressed as:

$$U_t = (1 - p_{un})U_{t-1} + p_{nu}(1 - U_{t-1}) \tag{1.1}$$

where p_{un} and p_{nu} are the union-to-nonunion and nonunion-to-union transition probabilities, respectively. That is, in a world with no possibility of non-employment, the union density rate is equal to proportion of the previous year's union members that maintain their union status to the next year plus the proportion of the previous year's nonunion members that switch to the union sector. Rearranging terms, equation (1.1) can be rewritten as the first-order Markov process:

$$U_t = (1 - p_{un} - p_{nu})U_{t-1} + p_{nu}\,. \tag{1.2}$$

Assuming the per-period flows p_{un} and p_{nu} are constant over time and sufficiently small so that $1- p_{un} - p_{nu} > 0$, this process implies a steady-state union density rate given by:

$$U^* = \frac{p_{nu}}{1 - p_{un} - p_{nu}}\,, \tag{1.3}$$

which is strictly increasing in the nonunion-to-union transition rate p_{nu} and strictly decreasing in the union-to-nonunion transition rate p_{un}.[5]

Equation (1.2) implies that one can recover the underlying transition probabilities by regressing aggregate union density rates on their own lagged values. The intercept in the model identifies the numerator in equation (1.3); the coefficient on the lagged dependent variable identifies the denominator; and together this provides two equations to solve for p_{un} and p_{nu}. Moreover, assuming that labour law reforms favorable to unions raise union density rates by

[4] An exception may be the Longitudinal Administrative Databank (LAD), which links T1 tax returns of individuals going back to the early 1980s. Since the Rand Formula insures all individuals covered by the terms of a collective agreement must pay union dues and these dues are tax deductible, one could use these deductions to identify union status and individual-level changes in union status across years. Unfortunately, these data are not, however, readily accessible to researchers outside of Statistics Canada.

[5] This can be derived by either solving the infinite geometric series obtained by substituting in for U_{t-1} or from simply equating U_t=U_{t-1}.

permanently increasing the nonunion-to-union transition rate p_{nu}, one could identify this effect on the long-run union density rate by allowing the legal reform variable to interact with both the overall intercept and the lagged dependent variable (since p_{nu} appears in both the intercept and the lagged dependent variable terms in equation (1.2)).

Of course, changes in union density rates over time are driven by numerous factors, some of which may be correlated with the timing of provincial changes to labour relations laws. The key empirical challenge is therefore to separately identify the effects of the laws from other factors. To do so, we extend the model implied by equation (1.2) by controlling for province and year fixed effects, as well as a set of province-level covariates intended to capture province-specific trends in union density rates that may be correlated with legislative changes. Specifically, we estimate the linear model:

$$U_{pt} = \alpha U_{p,t-1} + \delta R_{pt} + \theta(U_{p,t-1} \cdot R_{pt}) + x'_{pt}\beta + c_p + y_t + \varepsilon_{pt} \qquad (1.4)$$

where R_{pt} is an index of the favorableness to unions of the provincial labour relations regime that exists in province p in year t; x_{pt} is a vector of control variables intended to capture underlying province-specific trends in unionization, which includes the inflation rate (based on the all-items CPI), the unemployment rate (age 25 and over), the manufacturing share of employment, and a measure of popular tastes for trade unions; c_p and y_t are province and year fixed effects, respectively; and ε_{pt} is an error term with an expected value of 0, but potentially non-spherical variance-covariance matrix.[6] Given variation over time in R_{pt} within at least one province, all the parameters of equation (1.4) are identified. Equating U_{pt} and $U_{p,t-1}$, the estimates of equation (1.4) imply an expected steady-state union density rate U_p^*, which depends on all the parameters of the model.[7] Moreover, using union density rates estimated for different subgroups of the labour force, such as more or less educated workers, we obtain evidence of the social welfare implications of legal reform.

Freeman and Pelletier (1990) estimate a specification similar to equation (1.4) using a single time-series of British union density rates and index of labour laws. A difference, however, is that their model does not include the term containing the interaction of the lagged dependent variable and legal index. As it turns out, the coefficient θ on this term is poorly

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[6] See Section 3.3 for detailed descriptions of each of the control variables.

[7] Equating U_{pt} and $U_{p,t-1}$ in equation (1.4), we obtain the expected steady-state union density rate:

$$U_p^* = \frac{\delta R + W}{(1-\alpha-\theta R)}$$

where $W = x'_{pt}\beta + c_p + y_t$. Taking the derivative of this term with respect to the legal index R implies an effect on the steady-state union density rate given by:

$$\frac{\partial U^*}{\partial R} = \frac{\delta(1-\alpha)+\theta W}{(1-\alpha-\theta R)^2}.$$

identified in our data. To address this problem, we compare our estimates of the long-run policy effect at the provincial level to those obtained when we impose the restriction $\theta = 0$, so that legislation only affects the intercept through δ.[8] Having shown that the implied steady-state effects are quite similar whether the interaction term effect θ is estimated or not, we estimate the effects for particular subgroups of the population using the restricted model.

It is well known that a consequence of including the lagged union density rate in equation (1.4) is that the ordinary least squares estimates are biased. They are, however, consistent if the error term ε_{pt} contains no serial correlation. Using a Breusch-Godfrey test of autocorrelation based on the OLS fitted errors from estimating equation (1.4) we are unable to reject the null hypothesis of no serial correlation.[9] However, efficiency gains can be made using a feasible generalized least squares (FGLS) estimator that estimates the structure of the variance-covariance matrix of the error term. We therefore begin by comparing the estimates across four estimators: OLS; FGLS with province-specific heteroskedastiscty; FGLS with province-specific heteroskedasticity and spatial correlation; and FGLS with province-specific heteroskedasticity, spatial correlation, and province-specific autocorrelation.[10] Reporting separate results for the models with and without the θ interaction term discussed above, we obtain eight sets of estimates. As it turns out the estimated steady-state effects of policy reform are remarkably robust across specifications. Given the statistical challenge of identifying these effects for particular subgroups of the population, we take as our preferred specification the estimator with a smallest variance and then examine the robustness of the estimates to: (i) including province-specific linear time trends to capture any possible remaining latent provincial trends correlated with legal reforms; (ii) sample weights based on the underlying number of observations used to estimate the provincial union density rates; and (iii) an alternative source of data on union density rates based on administrative data on union membership. We conclude our analysis by estimating long-run policy reform effects across

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[8] In this case, the effect of a marginal change in the legal index on the steady-state union density rate is simply $\partial U^* / \partial R = \delta / (1-\alpha)$.

[9] We also performed tests of: (i) the poolability of the parameters across provinces; (ii) heteroskedasticity; and (iii) stationarity. The results are discussed in the notes of the Tables 4 and 5.

[10] If the variance-covariance matrix of the error term ε_{pt} is given by Ω, then in the most flexible case we estimate:

$$\Omega = \begin{bmatrix} \sigma_1^2\Omega_1 & \sigma_{1,2}I & \cdots & \sigma_{1,10}I \\ \sigma_{2,1}I & \sigma_2^2\Omega_2 & \cdots & \sigma_{2,10}I \\ \vdots & \vdots & \ddots & \vdots \\ \sigma_{10,1}I & \sigma_{10,2}I & \ldots & \sigma_{10}^2\Omega_{10} \end{bmatrix}.$$

Not allowing province-specific serial correlation imposes that the diagonal matrices Ω_j are all equal to a $T \times T$ identity matrix; not allowing spatial correlation imposes that all the off-diagonal elements $\sigma_{i,j}$ are zero; and not allowing for heteroskedasticity imposes that σ_j^2 is a constant equal to σ^2.

industry, occupation, education and gender, as well as for the individual laws comprising the legal reform index.

3. Data

3.1. Union density rates

Obtaining estimates of the legal regime effects in equation (1.4) with enough statistical precision to be informative requires time-series data on provincial union density rates going back to at least the early 1990s. Moreover, to identify variation in the magnitude of these effects across different types of workers requires union density rates for subsamples of workers within each province. To obtain these rates, we patch together a series of nationally-representative household surveys spanning the 1981-2012 period. Specifically, we use the Survey of Work History (SWH) for 1981; the Survey of Union Membership (SUM) for 1984; the Labour Market Activities Survey (LMAS) from 1986 through 1990; the Survey of Work Arrangements (SWA) for 1991 and 1995; the Survey of Labour and Income Dynamics (SLID) for 1993, 1994 and 1996; and the monthly Labour Force Survey (LFS) from 1997 through to 2012. With the exception of the 1981 SWH, all these files identify not only union membership, but also nonunionized workers whose terms of employment are covered by union contracts.[11] In addition, these data allow us to consistently distinguish four industries of employment: primary, manufacturing, private services, public services (where "public" services are defined broadly including not only provincial and municipal government employees, but also all education and health sector workers, as well as electric power, gas and water utilities workers); three occupations: blue-collar, administrative and professionals; three education levels: high school or less, post-secondary diploma/certificate and university degree; and gender.[12] Employees of the federal government are dropped for all years except 1991 and 1996, where it is not possible to separately identify them, as labour relations for these workers is governed by a separate federal statute. This complication is discussed in more detail in the following subsection.

In Table 1 we consider long-term declines in union density rates across provinces and worker types by comparing rates in 1981 and 2012. The estimates point to relatively large declines in New Brunswick, British Columbia and Alberta; in manufacturing and private services; and among men. In most cases the three-decade decline in male union density rates is more

[11] See Table A.1 in the appendix for detailed descriptions of the household survey data sources. For the 1981 SWH rates we use the 1984 SUM to estimate the difference between union density rates using the coverage and membership definitions and add this difference to the 1981 rates.
[12] See Table B for detailed definitions.

than twice as large for men than women, whether measured in terms of the change in the level of the rate or the proportionate change. There, however, appears relatively little difference in deunionization trends across broad occupation groups, although in the two western-most provinces – Alberta and British Columbia – the overall declines have clearly been much larger among blue-collar workers. The iniquitousness of these trends across provinces, as well as the large gender difference, emphasizes that an important part of the deunionization trends are driven by factors beyond labour relations laws. The empirical challenge is to determine to what extent the declines in Table 1 reflect changes in provincial labour relations laws.

The are two significant limitations of the household survey data we employ: (i) there are missing years (specifically 1982, 1983, 1985, and 1992); and (ii) there are substantial sampling biases in the estimation of union density rates arising from the limited sample sizes, particularly prior to 1997 when the Canada's monthly Labour Force Survey (LFS) first introduced a question identifying union status. To provide ourselves with some confidence in the accuracy of our estimated provincial time-series prior to 1997, we compare our estimates to those obtained using comparable provincial time-series data based on mandatory union filings under the *Corporations and Labour Unions Returns Act* (CALURA). Specifically, prior to 1996 all unions with members in Canada were required to file an annual return in December of each year reporting the total number of union members within each union local. These counts were then aggregated at the provincial level and published annually by Statistics Canada. To obtain provincial union density rates we divide these membership levels by estimates of provincial employment from the LFS. This provides us with union density rates from 1976 to 1995, which can be combined with the 1997 to 2012 LFS data to provide a complete series. However, to make the LFS series consistent with the CALURA, for this comparison series we exclude nonunionized workers covered by union contracts in the LFS union density rates.[13]

The resulting provincial time-series of union density rates using both the household survey data (labeled HS-LFS) and CALURA (labeled CALURA-LFS) are plotted in Figure 1.[14] Consistent with Table 1, both data sources point to larger declines in New Brunswick, Alberta

[13] There are two significant complications in comparing the LFS and CALURA rates. First, unions with less than 100 members did not have to provide information in the CALURA. This will tend to underestimate union density rates in the CALURA relative to the LFS. On the other hand, CALURA membership counts include union members who are not currently employed, such as workers on temporary layoff, and are recorded as of December 31 of each year, when seasonal layoffs are typically highest. Consequently, dividing by December employment levels tends to overestimate union density rates, particularly for the Atlantic Provinces where seasonal layoffs are most prevalent. To limit this measurement error, we instead use estimate employment levels using the July Labour Force Surveys. For detailed information on the comparability of the CALURA and LFS data, see Table A.3.

[14] Note that we are missing some years in both time series. The CALURA are missing 1996 and with the series based on survey data are missing 1982, 1983, 1985, and 1992. To fill in these gaps we use a simple linear interpolation of the neighbouring years. For 1985, 1992, and 1996, this is simply an average of the values for the years on either side of the missing year. For 1982 and 1983 we use a weighted average (e.g. 1982 is two-thirds of the 1981 value and one-third of the 1984 value).

and British Columbia. However, in all provinces the long-term declines are smaller in the CALURA-LFS series. In fact, in Prince Edward Island, Nova Scotia, Quebec, Manitoba and Saskatchewan there is little or no evidence of a long-term secular decline in union density in the administrative data. One possible explanation is that deunionization has occurred primarily through a decline in workers covered by union contracts, as opposed to union membership. Indeed, to some extent, this has been the experience in Australia, the United Kingdom and New Zealand, where declines in union coverage rates since the early 1980s have exceeded declines in union membership rates (Schmitt and Mitukiewicz 2011).[15]

The key advantage of the survey data is that it allows us to estimate union density rates for particular subgroups of the population. Before considering the role of labour relations laws, we examine to what extent Canadian deunionization trends can be accounted for by compositional shifts in employment across provinces, industries, occupations, education groups and gender. For example, unionization rates have always been higher in the manufacturing sector than in private services. Consequently, employment shifts away from manufacturing towards services, will push aggregate union density rates downwards for reasons unrelated to labour relations laws. To quantify the role of these compositional shifts more generally, we compare the estimates from two different regressions, the results of which are reported in Table 2. In the first, we pool the aggregate provincial-level HS-LFS union density rates plotted in Figure 1 and regress them on linear (specification 1) or quadratic (specification 2) time trends. In the second, we do the same thing using union density rates estimated at the province-industry-occupation-education-gender level. With 32 years of data this gives us 320 observations in the first case (32 x 10 provinces) and 23,040 in the second (32 x 10 provinces x 4 industries x 3 occupations x 3 education groups x 2 genders). Estimating the union density rates at this level of detail compromises the precision of the estimates significantly. However, since there is no reason to believe that the expected value of this measurement error is correlated with a trend (although its variance is decreasing due to larger sample sizes beginning with the LFS in 1997), it should not bias our estimates.

The first two columns of Table 2 point to a downward trend in union density rates when the rates from all provinces are pooled. The linear specification points to an annual decrease of 0.37 percentage points, while the quadratic specification suggests this rate of decrease is in decline such that by the end of our sample period, rates have stabilized (the slope of the time trend is -0.0065 x 0.0002**time*, where *time* is equal to 32 in 2012). To the extent that this trend reflects employment shifts across groups, it should not be evident within groups. However, the third and fourth columns of Table 2 suggest only slightly smaller rates of decline when we use

[15] Another difference with the CALURA data series is that professional organizations certified as unions, such as teachers federations and nurses associations, were not included prior to 1983 (Mainville and Olineck 1999). This will tend to understate union density rates in the early 1980s, resulting in flatter profiles over time.

the group-specific union density rates. The linear specification now suggests an annual decline of 0.31 percentage points, while the quadratic suggests rates stabilized by 2009. These results imply that something more than structural economic shifts are responsible for declining Canadian union density over the past 30 years.

3.2. Labour relations index

The rules governing the formation, operation and destruction of union bargaining units in Canada is specified by the labour relations code of the province in which an employee works. However, not all workplaces within a province are governed by these provincial statutes. For example, labour relations for employees of the federal government are governed by the *Federal Public Service Staff Relations Act (PSSRA)*, while employees in federally-regulated industries, such as air transportation and uranium mining, are regulated by the *Canada Labour Code*. In addition, provincial civil servants, police, firefighters, teachers, and hospital workers are, in some cases, but not all, governed by separate statutes.

Ideally, one could separately identify each of these exceptional cases in the data in order to relate the relevant legislation to unionization rates of each employee group. However, with the exception of the federal government employees, the level of industry and occupation detail provided in the data is inadequate to do this in any reasonably consistent way. Fortunately, there is reason to believe that changes in the provincial statutes that we measure may have effects that spill over to these excluded groups. First, the special statutes typically exist primarily to regulate the right to strike where employees are providing services deemed essential. Consequently, key regulations affecting unionization rates, such as rules for certifying new bargaining units, are taken from the overriding provincial statutes. For example, Ontario civil servants are governed by the *Ontario Crown Employees Collective Bargaining Act*, which states that the *Ontario Labour Relations Act* forms part of the Act except where the Act stipulates otherwise. Second, often amendments to provincial statues coincide with comparable changes in the special statutes. And third, it may be that political swings that result in legislative changes lead to broad changes in the labour relations environment within a province. To take a particular example, a change in government to a relatively labour-friendly administration, may lead to both a more union-friendly legal regime and an increase reluctance of the government to force, through legislation, public sector workers in a legal strike back to work, which could influence subsequent employment growth and thereby membership. The key point is that in not excluding employees (with the exception of federal civil servants) from our analysis, we are primarily interested in determining the potential for changes in the overriding provincial labour relations statutes, which typically coincide with broader changes in the labour relations climate within a province, to affect provincial union density rates. Given that this is how changes in labour relations actually happen, as opposed to independent changes in statutes particular to an exceptional group of workers, we think this is the correct question to ask. However, we also report estimates separately for private-sector workers who unambiguously fall under the provincial statutes, providing us with some indication of the importance of legal jurisdiction.

To capture the extent to which the labour relations climate within a province is favourable to unions, we construct a labour relations index based on twelve specific features of the overriding provincial statutes. This set of laws closely follows those examined by Johnson (2010). Listed in no particular order, they include:

- *Secret ballot certification vote*: certification of new bargaining units requires majority support in a mandatory secret-ballot vote;
- *First contract arbitration:* the union or employer can request that a third-party arbitrator be assigned to impose the terms and conditions of the collective agreement;
- *Anti-temporary replacement laws*: prohibits employers from hiring temporary replacement workers during a work stoppage and limits use of existing employees;
- *Ban on permanent replacements*: prohibits employers from hiring permanent replacement workers during a work stoppage;
- *Ban on strike-breakers*: prohibits employers from hiring professional strikebreakers (individuals not involved in a dispute who are employed primarily to "interfere with, obstruct, prevent, restrain or disrupt" a legal strike;
- *Re-instatement rights*: grants striking workers the right to reinstatement at the conclusion of a strike with priority over temporary replacement workers;
- *Compulsory dues check-off*: permits, at the union's request, that a clause be included in the collective agreement that requires employers to automatically deduct union dues from employees' pay and remit them to the union;
- *Mandatory strike vote*: union must demonstrate, through a secret-ballot vote, that it has the majority support of the bargaining unit before it can legally strike;
- *Employer-initiated strike vote*: employer may request that a secret-ballot vote be held to determine if bargaining unit is willing to accept the employer's last offer;
- *Compulsory conciliation*: requires some form of third-party intervention to encourage a contract settlement before a legal work stoppage can occur;
- *Cool-off period*: mandates a number of days, after other legal requirements have been fulfilled, before a legal work stoppage can begin;
- *Technology "re-opener"*: permits, at the union's request, that a clause be included in the collective agreement that allows the contract to be re-opened before its expiry in the event that the union is concerned about the consequences of technological change.

The data on each of these laws across Canada's 10 provinces is presented in Table 3. Since our earliest data begins in 1976, we indicate with "76" instances where the law was already effective in 1976. Otherwise, the dates in the table indicate years in which the law was effective. If no second date is given, the law continues to be effective. We assign to each of the 12 laws a value of zero (one) if the law is viewed as unfavorable (favorable) to unions. These values are shown in the final column of Table 3. The final index we employ is simply the unweighted average of the [0,1] values in each province in each year.[16] As is evident in Table 1, these laws are more often than not passed together with other laws. In fact, 5 of the 12 laws never change independently, while another 3 only change in isolation once. However, the data matrix based on these 12 laws does have full rank, such that we are also able to estimate a version of equation (1.4) in which R_{pt} is replaced with a 12-element vector containing separate dummy variables for each of the laws.

In Figure 1, we plot the labour relations index for each province alongside the union density rate time series based on household and administrative data. Three features of the index stand out. First, and most important, there is considerable variation both across provinces and within provinces over time, allowing us to identify the legal regime on provincial union density rates, while conditioning on both year and province fixed effects, as well as a set of province-level controls intended to capture province-specific trends. Second, provinces that have historically had legislation more favorable to unions – Quebec, Manitoba and British Columbia – have tended to have higher rates of unionization, whereas provinces with historically unfavorable laws -- Nova Scotia and Alberta – have had lower rates. On the other hand, the relationship is complicated, as Manitoba and Saskatchewan have historically had very similar rates of unionization, despite Manitoba always having a more union-friendly labour relations environment. Third, and perhaps most striking, Canadian labour relations laws have clearly not tended to become less favorable to unions over time, despite secular downward trends in union density rates in nearly all provinces (at least based on the survey data that counts workers covered by union contracts). This, of course, does not mean that laws do not matter; but it does suggest that something beyond laws is primarily responsible for the Canadian deunionization experience of the past three decades.

3.3 Control variables:

To control for the broader trends that are common across provinces we include a full set of year fixed effects. However, as is evident in Figure 1 and Table 1, deunionization has clearly been stronger in some provinces – New Brunswick, Alberta and British Columbia – than in others – Newfoundland, Manitoba and Saskatchewan. We, therefore, also include a set of

[16] In years in which a law is introduced, the law instead contributes a fraction (as opposed to 0 or 1), where the fraction represents the proportion of the year that the law was effective.

control variables that employ province-specific data, as well as examine the robustness of the estimates to including province-specific linear trends. Below we justify our choice of controls and briefly describe the data we employ.

Inflation rate:

In periods of high inflation workers' real wages are often eroded. An important benefit of unionization is that unions typically negotiate clauses in collective agreements providing members with automatic cost of living wage adjustments. Since the demand for these COLA clauses, and therefore unionization, is expected to be higher in situations where inflation is high and the legal regime itself may be influenced by levels of inflation, we control for provincial-level inflation throughout our analysis. To do this, we use the all-items Consumer Price Index (Basket 2009, Year=2002). Note, that we use the inflation rate (year-over-year change in CPI), and not the *level* of the CPI.[17] Data is only available at the provincial level starting in 1979, so for the years before then we use the Canadian national CPI to calculate the inflation rate.

Unemployment rate:

Another key benefit of unionization is that it provides its members with increased job security, through seniority rules and restrictions on employers' use of technology to replace workers. Therefore, we would expect the demand for unionization to be increasing in provincial unemployment rates. In addition, job destruction during a recession may occur differentially in unionized workplaces, due primarily to higher fixed labour costs and therefore greater incentives for labour hoarding. Since provincial government initiatives to augment the labour relations environment may itself be influenced by business cycle fluctuations, it is important to condition on the unemployment rate. To do this we include the provincial unemployment rate among individuals aged 25 and over in all the estimated regressions.

Manufacturing share of employment:

There is considerable evidence that an important component of the long-term secular decline of unions in Canada and other OECD countries has been driven by structural economic shifts, in particular the shift from manufacturing to service-producing employment beginning in the 1980s. Since these trends are likely to have occurred differentially across provinces, and may be themselves correlated with changes in labour laws, we follow Bartkiw (2008) and Freeman and Pelletier (1990) and control for the manufacturing share of paid employment in our estimated

[17] Provincial CPI series begin in 1979, so for the regressions using the CALURA-LFS data series, which begins in 1976, we use the national CPI for 1976-1978.

regressions. These annual shares are estimated using the industry codes in the 1976 through 2012 Labour Force Survey (LFS) microdata files that we use to estimate union density rates.

Popular preferences for unions:

Changes in union density rates are driven by individual preferences for unionization in the population, but these preferences are, in turn, likely to be correlated with political preferences and the decisions of politicians to augment labour relations laws. To capture changes in preferences that may be correlated with both union density rates and our legal index, we exploit two sources of public opinion poll data – the Canadian Gallup Poll and the Canadian Election Study. The Canadian Gallup Poll surveyed individuals about their perceptions of unions between 1976 and 1989, and again between 1991 and 2000, while the Canadian Election Study contained questions about perceptions of unions between 1993 and 2008. Given the changes in the exact wording of poll questions over time and missing years, a separate model is estimated to obtain consistent provincial time-series measuring popular tastes for unions.[18]

4. Results

In this section we examine the results from estimating the lagged dependent variable (LDV) model of union density rates presented in Section 2. In Table 4, we begin by comparing the results with and without the interaction of the LDV and legal index ($U_{p,t-1} \cdot R_{pt}$ term in equation (1.4)) and across 4 alternative specifications of the error variance-covariance matrix. We then choose our preferred estimator and in Table 5 examine the sensitivity of the estimates to: (i) using the administrative CALURA-LFS data based on union membership counts; (ii) including province-specific quadratic trends[19]; and (iii) weighting observations by the underlying sample sizes used to estimate the union density rates. In Table 6 and 7, we then compare the estimates across industry, occupation, education and gender groups. Finally in Table 8 we present the results from replacing the labour relations index with separate dummy variables for each of the 12 laws comprising the index.

In the absence of the LDV-labour relations index interaction (columns "a"), the coefficients on the LDV vary between 0.64 and 0.71. In terms of the underlying dynamics

[18] Specifically, we map the categorical responses in each poll regarding support for unions into a binary variable: 1 for a favourable perception of unions and zero for a neutral or negative opinion. We then regress, using a probit model, this variable on a quadratic time trend; a set of province dummies; a set of province dummies interacted with both time and time-squared; and survey indicators to control for survey effects (in particular, changes in exact wording of questions). We then use the parameters from the probit to fit the model across 1976-2012 by province, generating the "tastes" variable used in equation (1.4).
[19] We restrict the quadratic term across provinces, but allow the linear term in the polynomial to vary across provinces.

defined by equation (1.2), this implies considerable annual job flows in and out of the union sector and a gradual adjustment of union density rates following legal reforms. The interaction terms (columns "b") are generally not well identified, although the point estimates are negative in all cases. This is consistent with our expectation that a shift towards a legal environment more favourable to unions will serve to increase the nonunion-to-union transition rate p_{nu}. Similarly, the positive and significant coefficients on the legal index itself across all specifications are, in terms of the structure given by equation (1.2), consistent with more favourable laws increasing nonunion-to-union transitions. To obtain an estimate of the long-run effect of legal reform, we predict the effect of increasing the labour relations index from the average provincial value observed in 2012 (weighted by the population of each province) to one. Given the dynamic structure implied by equation (1.3), the estimates in Table 4 imply a long-run increase in the national union density rate ranging from 5.5 to 7.6 percentage points. Given an actual national rate of 30.6% in 2012, this represents roughly a 20 percent increase.

With regard to the control variables, the unemployment rate effect estimates imply a countercyclical relationship with union density rates, which is consistent with evidence elsewhere (Freeman and Pelletier 1990) and the idea that the demand for unionization and the job protection unions provide increase in recessions. All the point estimates also suggest that union density rates are increasing in inflation, consistent with the demand for unionization and COLA clauses rising with inflation, although this effect is estimated much less precisely. As for the manufacturing share of employment, all the estimates are positive and in 6 of the 8 cases not statistically different from zero at the 5% level. However, to some extent deindustrialization trends have been common across provinces, in which case their influence on unionization will be captured by the year fixed effects. Finally, and most surprisingly, we find no evidence that population perceptions of unions captured in opinion poll data have any influence on union density rates; all the estimates are insignificant at the 5% level. One explanation may be that the tastes variable is itself partially determined by union density rates, in the sense that more union-friendly laws that lead to a greater union presence and power result in a more negative view of unions by the general public.

Given the similarity of the estimated long-run effects in Table 4, we subsequently restrict our attention to the estimator with the lowest variance – the FGLS estimator allowing for province-specific heteroskedasticity and autocorrelation, as well as contemporaneous spatial correlation. In addition, we restrict the interaction effect θ to be zero. The results from this case are reported in column (4a) of Table 4. The first column of Table 5 reports these results again to enable comparison with the results using the same estimator and specification, but with the CALURA-LFS union density rates (see fifth column of Table 5). The additional specifications in Table 5 add province-specific trends (2); or sample weights (3); or both (4).

The estimated long-run effects of legal reform are remarkably similar using the CALURA-LFS data based on union membership. In three of the four cases the CALURA-LFS point estimates are slightly larger, but the differences are never statistically distinguishable. What is more different is the adjustment process. The coefficient on the LDV in the CALURA-LFS is substantially larger in all cases. The structural interpretation of this result, based on equation (1.2), is that transition rates in and out of union coverage exceed the transitions in and out of union membership; as one would expect. However, it is likely also the case that the difference reflects greater measurement (sampling) error in the HS-LFS data. The greater noise in the union density rates estimated using survey data is evident in Figure 1. Given that this measurement error is random, we know it will serve to attenuate the estimated LDV effect, which in turn will bias (or "smear") all the estimates in the model. Fortunately, the similarity of the long-run effects provides us with some assurance that the bias using the HS-LFS is modest, and if anything tends to underestimate the true effects.

Including province-specific trends and sample weights produces larger differences, particularly using the HS-LFS data. In both cases, the estimates of the long-run legal reform effect are diminished, although including province-specific trends seems to matter more than sampling weights; the long-run estimate declines from 7.6 percentage points to 4.5 in the former case, but to 6.6 percentage points in the latter case. The difference appears to primarily reflect a decrease in the coefficient on the LDV, which is now less than 0.49 suggesting that the sum of the union-to-nonunion and nonunion-to-union annual transition rates is about one-half, which is clearly implausibly large. A possible explanation is that including province trends means that more of the remaining variation in the data to be explained is noise, which once again attenuates the estimated coefficient on the LDV. When we include the province trends and the sampling weights in specification (4), the long-run estimate is 3.1 percentage points; less than half the magnitude of the original estimate, but still statistically different from zero.

In Table 6 we report the results using the industry-, occupation-, education-, and gender-specific union density rates. Once again, we present the results from the preferred specification in Table 4 (column 4(a)) and do not include provincial trends or sampling weights. These results should, therefore, be thought of as upper bound estimates; although of primary interest are the relative magnitudes of the estimates across groups in the population.

The industry estimates suggest a relatively large effect of legal reform in primary industries (8 percentage points), followed by public services (5.6) and manufacturing (4.6), and little or no effect in private services. Given that an important part of public services are governed by separate statutes not included in our labour relations index, the relatively large impact for this group is the perhaps the most unexpected result. To obtain some perspective on this finding, we examined data on the number of employees in bargaining units newly certified under the *Ontario Labour Relations Act* between 1985 and 2011.[20] The most striking feature of the data is a clear shift away from manufacturing towards private services. However, in every year up to 2005, more than one-quarter of employees in newly certified bargaining units were employed in health services, education services, local government or utilities (with the vast majority in health and education). Moreover, through the latter half of the 1990s and early 2000s, this proportion always exceeded 40%. This emphasizes that the provincial labour relations statutes on which we base our labour relations index apply to large proportions of public (or parapublic) sector workers within provinces.

Given that public services tend to employ workers with higher average skill levels than private services, an implication of the industry results is that increases in unionization resulting from legal reforms will be concentrated among relatively skilled workers. The occupation and education results in Table 6 appear consistent with this. Although the differences are small, the point estimates suggest larger effects among professionals and university-educated workers. In addition, given that female employment is relatively concentrated in public services, it is not surprising that the estimate for women exceeds that for men; in this case the difference is more than two-fold (10.3 percentage points for women compared to 4.4 for men). The obvious question is why would legal reforms more favourable to unions have its largest impact among relatively highly-educated women employed in health and education services where wages, benefits and working conditions are advantaged even in the absence of unions? One possible explanation is that this finding reflects the optimizing decisions of unions and their organizers seeking secure union dues, as opposed to workers. An alternative explanation is that legal changes will primarily affect workplaces where the net marginal benefit of unionization is close to zero. The reason is that where net benefits are large, workers will already be unionized; and where they are small, small changes in the marginal cost of unionization resulting from legal reforms will be insufficient to alter optimal behaviour. Rather, it is where the net benefit becomes positive as the result of legal reforms that changes in union status will occur. From this perspective, what the results seem to suggest is that legal reforms, at least within the scope of laws in the Canadian experience, are insufficient to affect unionization for many unskilled workers employed in the private services, where the risks inherent in organizing unions are too great. In contrast, in public services where profit incentives tend to be weaker, it is more likely that small changes in the costs of union organizing brought about by legal reforms are sufficient alter organizing decisions.

Richer insight into the types of workplaces where legal reforms are likely to be most influential could be obtained by estimating the effects within the 72 industry-occupation-education-gender cells. For example, for men employed in blue-collar manufacturing jobs.

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[20] These data can be found in the Annual Reports of the Ontario Labour Relations Board between 1985 and 2011.

Unfortunately, in the vast majority of cases the survey data sample sizes are too small to estimate provincial union density rates at this level of detail with sufficient precision. Alternatively, in Table 7 we report the results from the largest 10 of these 72 cells, in terms of the total provincial sample sizes provided in the HS-LFS data. Once again, the results point to relatively large effects among highly-educated public service professionals (although the estimate for university-educated men is not statistically different from zero at the 5% level). However, they also suggest even larger effects for relatively uneducated women, but not men, employed in private services. Specifically, for both blue-collar women with high school diploma or less and college-educated women in administrative jobs, the estimates suggest an increase in the steady-state union density rate that exceeds 10 percentage points. The results, therefore, seem to suggest that, at least for women, the effects of reforms in provincial labour relations laws may, in fact, be quite widespread.

We complete our analysis by examining the independent effects of the individual laws comprising our labour relations index. The results, presented in Table 8, suggest that our previous results primarily reflect three laws: (i) bans on permanent replacement workers; (ii) first contract arbitration; and (iii) mandatory secret-ballot certification votes. We expect that secret-ballot certification votes and first-contract arbitration primarily influence the formation of unions – votes increase the ability of employers to influence employee preferences and thereby their voting behaviour, while ensuring access to interest arbitration reduces the incentives for employers to avoid negotiating an initial collective agreement. In contrast, bans on replacement workers should primarily affect the destruction of unions as employers undermine union-management relations that may be well established, through recruitment of permanent nonunionized replacement workers. Since in the latter case, bargaining units are never officially decertified, these outcomes would be missed in the administrative data that are examined elsewhere in the literature. Interestingly, the estimates in Table 8 suggest that, of the three laws that matter, secret ballot votes, which have been the emphasis of the existing literature, are the least influential, while bans on permanent replacement workers are most. This relative emphasis on union recognition procedures in the literature likely reflects the fact that changes in these laws have historically been the most common, making them particularly visible and contentious. In contrast, bans on permanent replacement workers have only been introduced on two occasions, and never independently of changes in other laws (see third and fourth column of Table 8).

6. Conclusions

Overall, our analysis suggests that changes in labour relations legislation can have substantial impacts on union density rates. Specifically, our preferred estimates suggest that making legislation fully supportive of unions could raise Canada's current national union density rate of 30% by up to 7 percentage points in the long run. With regard to particular laws, our estimates suggest that banning permanent replacement workers, providing certification based union card checks, and guarantees of interest arbitration to ensure first contracts are reached are most effective. While shifting labour relations statutes towards these types of laws is clearly insufficient to reverse long-term deunionization trends, which have been relatively modest in Canada, it is worth emphasizing that the range of Canadian laws that we examine are also limited when compared to laws seen elsewhere. In particular, unlike the U.S., all Canadian statutes permit unions to negotiate union security clauses requiring employers to deduct union dues from employees' pay. Consequently, we are unable to examine the effect of introducing right-to-work laws. Nonetheless, the sizable effects we identify within the relatively narrow range of Canadian laws, contribute to the existing evidence emphasizing the importance of legal structures in determining union density rates.

A key advantage of the survey data we employ is that it allows us to obtain evidence on what types of workplaces and workers are most likely to be affected by legal reforms. Our results indicate that the benefits of shifting to a more union-friendly legal environment are likely to benefit highly-educated professional women employed in public services (broadly defined) relatively more. However, when estimated at finer levels of detail, the estimates also point to significant unionization gains among relatively uneducated women in private service-producing industries. In contrast, the gains among men appear consistently modest in magnitude. The social welfare implications of these findings are mixed. The large gains among women who, in the absence of unions, enjoy relatively high wages and benefits and better job security, does not suggest that labour relations laws are an effective policy instrument addressing labour market inequality concerns. However, it may also be the case that the spillover effects of union outcomes on nonunion workers are more important in labour markets where union density rates are lower, such as the unskilled service sector. A better understanding of these externalities of unions would be a fruitful area of future research.

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Table 1: Provincial union density rates, 1981 and 2012

		NL	PE	NS	NB	QC	ON	MB	SK	AB	BC
All Workers	1981	0.45	0.40	0.36	0.41	0.49	0.35	0.40	0.40	0.32	0.44
	2012	0.38	0.30	0.29	0.28	0.39	0.27	0.35	0.35	0.23	0.30
Industry											
primary	1981	0.51	0.06	0.35	0.37	0.48	0.31	0.34	0.31	0.16	0.60
	2012	0.38	0.06	0.19	0.21	0.23	0.17	0.20	0.27	0.11	0.29
manufacturing	1981	0.69	0.39	0.46	0.43	0.57	0.47	0.45	0.42	0.40	0.63
	2012	0.43	0.26	0.17	0.24	0.36	0.21	0.31	0.25	0.17	0.25
private services	1981	0.25	0.25	0.22	0.28	0.38	0.22	0.27	0.27	0.23	0.30
	2012	0.19	0.10	0.12	0.10	0.26	0.14	0.18	0.18	0.12	0.18
public services[a]	1981	0.73	0.82	0.72	0.78	0.89	0.67	0.77	0.79	0.73	0.78
	2012	0.67	0.69	0.64	0.62	0.70	0.59	0.68	0.68	0.56	0.63
Occupation											
blue collar	1981	0.50	0.35	0.41	0.44	0.60	0.46	0.45	0.42	0.38	0.58
	2012	0.37	0.23	0.26	0.25	0.44	0.30	0.33	0.31	0.20	0.31
administrative	1981	0.26	0.28	0.25	0.35	0.40	0.26	0.33	0.32	0.26	0.29
	2012	0.25	0.20	0.17	0.17	0.26	0.15	0.23	0.24	0.16	0.20
professionals	1981	0.62	0.73	0.58	0.57	0.64	0.41	0.53	0.63	0.44	0.51
	2012	0.47	0.46	0.41	0.41	0.44	0.31	0.46	0.48	0.31	0.38
Education											
high school or less	1981	0.46	0.35	0.36	0.4	0.53	0.38	0.4	0.4	0.32	0.46
	2012	0.25	0.17	0.18	0.18	0.33	0.22	0.27	0.26	0.17	0.23
post-secondary degree	1981	0.46	0.6	0.5	0.56	0.59	0.44	0.52	0.59	0.46	0.55
	2012	0.43	0.36	0.34	0.31	0.43	0.3	0.39	0.4	0.25	0.36
university degree	1981	0.63	0.79	0.58	0.61	0.68	0.41	0.61	0.58	0.42	0.52
	2012	0.48	0.46	0.37	0.43	0.41	0.28	0.45	0.45	0.31	0.34
Gender											
male	1981	0.51	0.40	0.43	0.46	0.59	0.45	0.47	0.46	0.38	0.55
	2012	0.37	0.24	0.25	0.26	0.40	0.26	0.32	0.29	0.20	0.28
female	1981	0.43	0.46	0.37	0.43	0.50	0.32	0.39	0.42	0.34	0.38
	2012	0.38	0.36	0.32	0.30	0.38	0.27	0.38	0.40	0.26	0.32

Notes: Union density rates are from the HS-LFS series and therefore exclude federal government employees. All other relevant sample restrictions are described in Table A1. The definition of union density includes those who are covered by a collective agreement, but who are not a member of the union. Sources: SWH (1981), LFS(2012).

[a] Public services is broadly defined including provincial and municipal government employees, education and related services, health and welfare services and utilities.

Table 2: Union density rates regressed on linear and quadratic time trends

	Union density rates:			
	Provincial-level		Province-industry-occupation-education-gender-level	
Independent variables	(1)	(2)	(1)	(2)
time	-0.0037***	-0.0065***	-0.0031***	-0.0056***
	(0.0003)	(0.0006)	(0.0003)	(0.0005)
time squared		0.0001***		0.0001***
		(0.0000)		(0.0000)
constant	0.4011***	0.4150***	0.3924***	0.4052***
	(0.0220)	(0.0236)	(0.0188)	(0.0186)
observations	320	320	23040	23040
R^2	0.284	0.296	0.014	0.014

Standard errors in parentheses

* p < 0.10, ** p < 0.05, *** p < 0.01

Note: All linear regressions are weighted by sample sizes of underlying survey data. Standard errors are clustered; (1) and (2) at province level, (3) and (4) at unit level.

Table 3: Changes in Canadian labour relations legislation, 1976-2012

Law	NL	PE	NS	NB	QC	ON	MB	SK	AB	BC	Index
First contract arbitration[i]	85:06		11:12[g]		77:12	86:05	82:02	94:10		76	=1
Anti-temporary replacement workers					78:02	93:01-95:11				93:01	=1
Ban on permanent replacement		87:05					85:01				=1
Re-instatement rights		87:05			78:02	76-92:12	85:01	94:10	88:11		=1
Ban on strike-breakers						83:06	85:01			76	=1
Mandatory dues check-off	85:07				78:04	80:07	76	76		77:09	=1
Mandatory strike vote		76	76	76	78:04	95:11	85:01	76	76	76	=0
Employer-initiated strike vote			94:05		02:11	80:07	97:02-00:10	83:07	88:12	87:08	=0
Compulsory conciliation	76	76	76	76	76-78:01	76-86:12			76-81:02, 88:12		=0
Cool-off period[h]	76	76	76	76	77:12	76		83:07	76-88:11	76	=0
Technology re-opener				89:04			72:11			74:03	=1
Secret ballot certification election[a]	94:02-12:06[e]		77:05			95:11[f]	97:02-00:09[c]	08:05[d]	88:11	84:06-93:01, 01:08[b]	=0

Notes: All dates are from Johnson (2010) unless otherwise noted by a reference. Date specifies when law comes into effect (may be different from royal assent date). 76 indicates law was in effect in January 1976.

a: Dates are from Johnson (2002) unless otherwise noted by a reference in this row.

b: Highlights of Major Developments in Labour Legislation (2000-2001)

c: Highlights of Major Developments in Labour Legislation (1999-2000)

d: Bill 6: An Act to amend The Trade Union Act, Chapter 26; Royal Assent: May 14, 2008.

e: Bill 37: An Act to amend The Labour Relations Act, Chapter 30; Royal Assent: June 27, 2012.

f: Bill 144: An Act to amend certain statutes relating to Labour Relations; Royal Assent June 13, 2005. Remove mandatory vote below 55% support for construction workers only. Note: we do not exclude construction workers in HS-LFS series.

g: Bill 102: An Act to Prevent Unnecessary Labour Disruptions and Protect the Economy by Amending Chapter 475 of the Revised Statutes, 1989, the Trade Union Act, Chapter 71; Royal Assent: December 15, 2011.

h: We do not specify the number of days of cool-off period in this table – see Johnson (2010) for more detail.

i: Update since Johnson (2002). PEI did not implement first contract arbitration in 95:05; never received Royal Assent.

Table 4: Estimates of the effect of provincial labour relations index on union density rates

	Dependent variable: HS-LFS union density rates							
Independent var.	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(4a)	(4b)
lagged density rate	0.6422***	0.6593***	0.6873***	0.7101***	0.7057***	0.7297***	0.6735***	0.7055***
	(0.0450)	(0.0514)	(0.0407)	(0.0469)	(0.0408)	(0.0436)	(0.0383)	(0.0395)
legal index	0.0427***	0.0636*	0.0301***	0.0568**	0.0308***	0.0565***	0.0422***	0.0815***
	(0.0124)	(0.0326)	(0.0101)	(0.0287)	(0.0085)	(0.0215)	(0.0060)	(0.0198)
interaction term		-0.0610		-0.0764		-0.0743		-0.1164**
		(0.0883)		(0.0769)		(0.0569)		(0.0559)
unemployment rate	0.1709**	0.1752**	0.1563**	0.1632**	0.1036*	0.1102*	0.0499	0.0443
	(0.0742)	(0.0745)	(0.0629)	(0.0634)	(0.0574)	(0.0573)	(0.0526)	(0.0525)
inflation rate	0.1355	0.1527	0.0472	0.0628	0.0260	0.0347	0.0382	0.0425
	(0.1281)	(0.1306)	(0.1078)	(0.1100)	(0.0373)	(0.0388)	(0.0792)	(0.0801)
manufacturing share	0.0975	0.1032*	0.0934*	0.1035**	0.0753	0.0781	0.0752*	0.0797**
	(0.0615)	(0.0621)	(0.0501)	(0.0508)	(0.0491)	(0.0487)	(0.0390)	(0.0385)
tastes	-0.0368	-0.0356	-0.0312*	-0.0276	-0.0166	-0.0120	-0.0218	-0.0192
	(0.0242)	(0.0243)	(0.0188)	(0.0191)	(0.0172)	(0.0178)	(0.0226)	(0.0227)
constant	0.1307***	0.1232***	0.1193***	0.1072***	0.1096***	0.0982***	0.1271***	0.1171***
	(0.0274)	(0.0294)	(0.0253)	(0.0284)	(0.0266)	(0.0279)	(0.0269)	(0.0271)
<u>Error Terms:</u>								
Var[$\epsilon_{p,t}$]=	σ^2	σ^2	σ_p^2	σ_p^2	σ_p^2	σ_p^2	σ_p^2	σ_p^2
Cov[$\epsilon_{p,t}$, $\epsilon_{q,s}$]=	0	0	0	0	$\sigma_{p,q}$	$\sigma_{p,q}$	$\sigma_{p,q}$	$\sigma_{p,q}$
Cov[$\epsilon_{p,t}$, $\epsilon_{p,t-1}$]=	0	0	0	0	0	0	ρ_p	ρ_p
observations	310	310	310	310	310	310	310	310
R^2	0.969	0.969	-	-	-	-	-	-
long run effect	0.0707	0.0671	0.0571	0.0545	0.0619	0.0591	0.0764	0.0689
	(0.0212)	(0.0193)	(0.0197)	(0.0171)	(0.0176)	(0.0151)	(0.0109)	(0.0103)

Notes: Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Year dummies and province dummies are included in all regressions. The variable *tastes* is between (0,1) with 1 being most supportive of unions. The following tests are performed on specification (1): (a) Poolability: Using the Baltagi (2008, p.57) for full poolability (we need to exclude year dummies to do the test), we reject the null of poolability of all parameters. Using the Beck (2001) test for poolability of a single parameter of interest, we fail to reject the null of poolability of the legal index parameter. (b) Heteroskedasticity: Using the Wald Test proposed in Greene (2003, p.323) we reject the null of no groupwise (panel) heteroskedasticity. (c) Serial Correlation: Using the Lagrange multiplier test for serial correlation in time-series-cross-section data as described in Beck and Katz (1996), we do not reject the null of no serial correlation. (d) Stationarity: Using the Levin, Lin, Chu (2002) test for stationarity of time-series-cross-section data, we reject the null that the panels contain unit roots (cross-sectionally-demeaned stationary). The "long run effect" is the difference between the long run value of $U_{p,t}$ evaluated at R_t=1 and evaluated at R_t=R_{2012} where R_{2012} is the average of all provincial values of R in 2012, weighted by population of the province.

Table 5: Robustness analysis of effect of legislative index on union density rates

	Dependent Variable: union density rates:							
	HS-LFS				CALURA-LFS			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
lagged density rate	0.6735***	0.6963***	0.4917***	0.4552***	0.8459***	0.7900***	0.6210***	0.5719***
	(0.0383)	(0.0350)	(0.0484)	(0.0461)	(0.0233)	(0.0279)	(0.0388)	(0.0412)
legal index	0.0422***	0.0339***	0.0389***	0.0288***	0.0220***	0.0198***	0.0366***	0.0342***
	(0.0060)	(0.0066)	(0.0076)	(0.0079)	(0.0046)	(0.0060)	(0.0053)	(0.0071)
unemployment rate	0.0499	0.0510	-0.0348	-0.0470	0.0231	-0.0154	0.0217	0.0578
	(0.0526)	(0.0486)	(0.0601)	(0.0610)	(0.0345)	(0.0376)	(0.0412)	(0.0456)
inflation rate	0.0382	-0.0161	0.0076	-0.0797	0.0116	-0.0018	-0.0497	-0.0189
	(0.0792)	(0.0753)	(0.0825)	(0.0805)	(0.0618)	(0.0472)	(0.0603)	(0.0498)
manufacturing share	0.0752*	0.0892**	-0.1117	-0.0832	0.0907***	0.0569**	-0.0819	0.0453
	(0.0390)	(0.0375)	(0.0780)	(0.0642)	(0.0284)	(0.0264)	(0.0519)	(0.0459)
tastes	-0.0218	-0.0464***	0.0447	0.0154	0.0050	0.0211*	-0.0036	0.0611**
	(0.0226)	(0.0165)	(0.0522)	(0.0457)	(0.0108)	(0.0127)	(0.0190)	(0.0256)
constant	0.1271***	0.1375***	0.2235***	0.2680***	0.0182**	0.0439***	0.1374***	0.0800***
	(0.0269)	(0.0218)	(0.0499)	(0.0445)	(0.0075)	(0.0104)	(0.0234)	(0.0252)
province trends	No	No	Yes	Yes	No	No	Yes	Yes
sample size weights	No	Yes	No	Yes	No	Yes	No	Yes
observations	310	310	310	310	360	360	360	360
long run effect	0.0764	0.0660	0.0453	0.0313	0.0869	0.0572	0.0588	0.0486
	(0.0109)	(0.0128)	(0.0091)	(0.0088)	(0.0185)	(0.0168)	(0.0088)	(0.0102)

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: Year dummies and province dummies are included in all regressions. The variable *tastes* is between [0,1] with 1 being most supportive of unions. All specifications use the same form of GLS as columns 7 and 8 in Table 4, Var[$\epsilon_{p,t}$]=σ_p^2, Cov[$\epsilon_{p,t}$, $\epsilon_{q,s}$]=$\sigma_{p,q}$, Cov[$\epsilon_{p,t}$, $\epsilon_{p,t-1}$]=ρ_p. Sample size weights refer to total cell counts of micro data underlying the data.

Table 6: Estimates of legislative effect by industry, occupation, education and gender

	Dependent Variable: union density rates:											
	Industry				Occupation			Education			Gender	
	primary	priv good	priv serv	public	blue	admin	profes	HS	PS	university	male	female
lag den rate	0.6063***	0.6976***	0.5105***	0.5811***	0.7477***	0.4499***	0.5691***	0.6217***	0.4821***	0.6608***	0.6583***	0.5984***
	(0.0448)	(0.0428)	(0.0492)	(0.0467)	(0.0382)	(0.0466)	(0.0394)	-0.043	-0.0482	-0.0374	(0.0425)	(0.0377)
legal index	0.0413*	0.0202	0.0073	0.0514***	0.0245***	0.0715***	0.0639***	0.0299***	0.0553***	0.0436***	0.0231***	0.0663***
	(0.0237)	(0.0196)	(0.0074)	(0.0126)	(0.0095)	(0.0158)	(0.0120)	-0.0083	-0.0116	-0.0114	(0.0071)	(0.0097)
unem rate	0.0861	-0.0947	0.0177	0.2349**	0.0775	0.2439***	0.1137	0.1353**	0.1397	0.0364	0.0975	0.1375**
	(0.1491)	(0.1224)	(0.0594)	(0.0944)	(0.0798)	(0.0851)	(0.0692)	-0.0657	-0.0869	-0.0865	(0.0690)	(0.0574)
inflation rate	0.1828	-0.5618***	-0.0115	0.3910***	0.2152*	-0.0337	0.1344	-0.0053	0.0422	0.1992	0.0064	0.1147
	(0.2505)	(0.2162)	(0.0882)	(0.1395)	(0.1194)	(0.1311)	(0.1206)	-0.1025	-0.1265	-0.1454	(0.0954)	(0.0977)
manuf share	0.2895**	0.2492**	-0.1297***	0.0790	0.1141**	-0.0612	0.1130**	0.1431***	-0.0395	-0.1240**	0.0905*	0.0848**
	(0.1314)	(0.1142)	(0.0402)	(0.0694)	(0.0546)	(0.0657)	(0.0498)	-0.0516	-0.0784	-0.0631	(0.0493)	(0.0388)
tastes	0.0429	-0.0329	-0.0119	-0.0754**	-0.0084	-0.0062	-0.0839**	-0.0203	-0.1652***	-0.035	0.0036	-0.0579***
	(0.0640)	(0.0387)	(0.0156)	(0.0320)	(0.0239)	(0.0291)	(0.0374)	-0.0206	-0.0335	-0.0318	(0.0217)	(0.0211)
constant	0.0126	0.1589***	0.1855***	0.2946***	0.0845***	0.1448***	0.2151***	0.1329***	0.3581***	0.1966***	0.1412***	0.1356***
	(0.0427)	(0.0450)	(0.0252)	(0.0459)	(0.0303)	(0.0247)	(0.0340)	-0.0277	-0.0425	-0.0359	(0.0288)	(0.0217)
observations	310	310	310	310	310	310	310	310	310	310	310	310
long run effect	0.0803	0.0460	0.0114	0.0563	0.0666	0.0617	0.0840	0.0513	0.0529	0.0781	0.0436	0.1025
	(0.0354)	(0.0380)	(0.0091)	(0.0164)	(0.0211)	(0.0173)	(0.0157)	-0.0125	-0.0125	-0.0189	(0.0120)	(0.0126)

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: Year dummies and province dummies are included in all regressions. The variable *tastes* is between (0,1) with 1 being most supportive of unions. The estimator used for all 12 regressions above is the same is in Column (4a) of Table 4.

Table 7: Estimates of legislative effect for 10 largest industry-education-occupation-gender cells

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
lag den rate	0.5026***	0.4292***	0.4436***	0.5636***	0.4269***	0.3268***	0.4567***	0.4088***	0.3906***	0.4790***
	(0.0485)	(0.0493)	(0.0522)	(0.0453)	(0.0532)	(0.0523)	(0.0535)	(0.0470)	(0.0505)	(0.0454)
legal index	-0.0164	0.0387***	0.0365	0.0609***	0.0067	0.0595***	-0.0079	0.0613***	-0.0049	0.0419
	(0.0143)	(0.0149)	(0.0399)	(0.0161)	(0.0324)	(0.0197)	(0.0254)	(0.0169)	(0.0253)	(0.0270)
unem rate	0.0253	-0.0233	0.1179	0.1951**	0.4092**	0.2562*	-0.0491	-0.1943	0.0711	0.4680**
	(0.1254)	(0.0978)	(0.2375)	(0.0823)	(0.1821)	(0.1528)	(0.1844)	(0.1325)	(0.1642)	(0.1926)
inflation rate	0.2895	-0.2881*	0.3688	0.2556	-0.0487	-0.0689	-0.0941	0.2227	0.4538**	0.1353
	(0.1977)	(0.1506)	(0.3666)	(0.1592)	(0.2770)	(0.2494)	(0.3080)	(0.2167)	(0.2192)	(0.3245)
manuf share	-0.1523*	-0.1046*	0.4019*	-0.0069	0.3450**	-0.1373	-0.8835***	-0.0874	-0.0462	-0.0052
	(0.0781)	(0.0591)	(0.2281)	(0.0616)	(0.1535)	(0.0932)	(0.1592)	(0.0810)	(0.1469)	(0.1320)
tastes	0.0389	0.0211	-0.0223	-0.1049***	-0.1909**	-0.0389	-0.1120	-0.0575*	0.0026	-0.1282***
	(0.0378)	(0.0226)	(0.0698)	(0.0246)	(0.0783)	(0.0449)	(0.0806)	(0.0344)	(0.0428)	(0.0481)
constant	0.2517***	0.1218***	0.2774***	0.0821***	0.5085***	0.5389***	0.5921***	0.1576***	0.1887***	0.4188***
	(0.0388)	(0.0272)	(0.0820)	(0.0228)	(0.0737)	(0.0625)	(0.0821)	(0.0349)	(0.0510)	(0.0648)
sector	services	services	manuf	services	public	public	services	services	services	public
education	high school	high school	high school	high school	college	university	college	college	high school	university
occupation	blue	admin	blue	blue	profes	profes	blue	admin	admin	profes
gender	male	female	male	female	female	female	male	female	male	male
observations	310	310	310	310	310	310	310	310	310	310
long run effect	-0.0330	0.0679	0.0656	0.1396	0.0117	0.0883	-0.0146	0.1037	-0.0081	0.0804
	(0.0292)	(0.0263)	(0.0715)	(0.0369)	(0.0565)	(0.0290)	(0.0466)	(0.0276)	(0.0414)	(0.0512)

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: Year dummies and province dummies are included in all regressions. The variable *tastes* is between (0,1) with 1 being most supportive of unions. The specification used for all 12 regressions above is the same is in Column (4a) of Table 4.

Table 8: Estimate of the independent effect of components of labour relations index

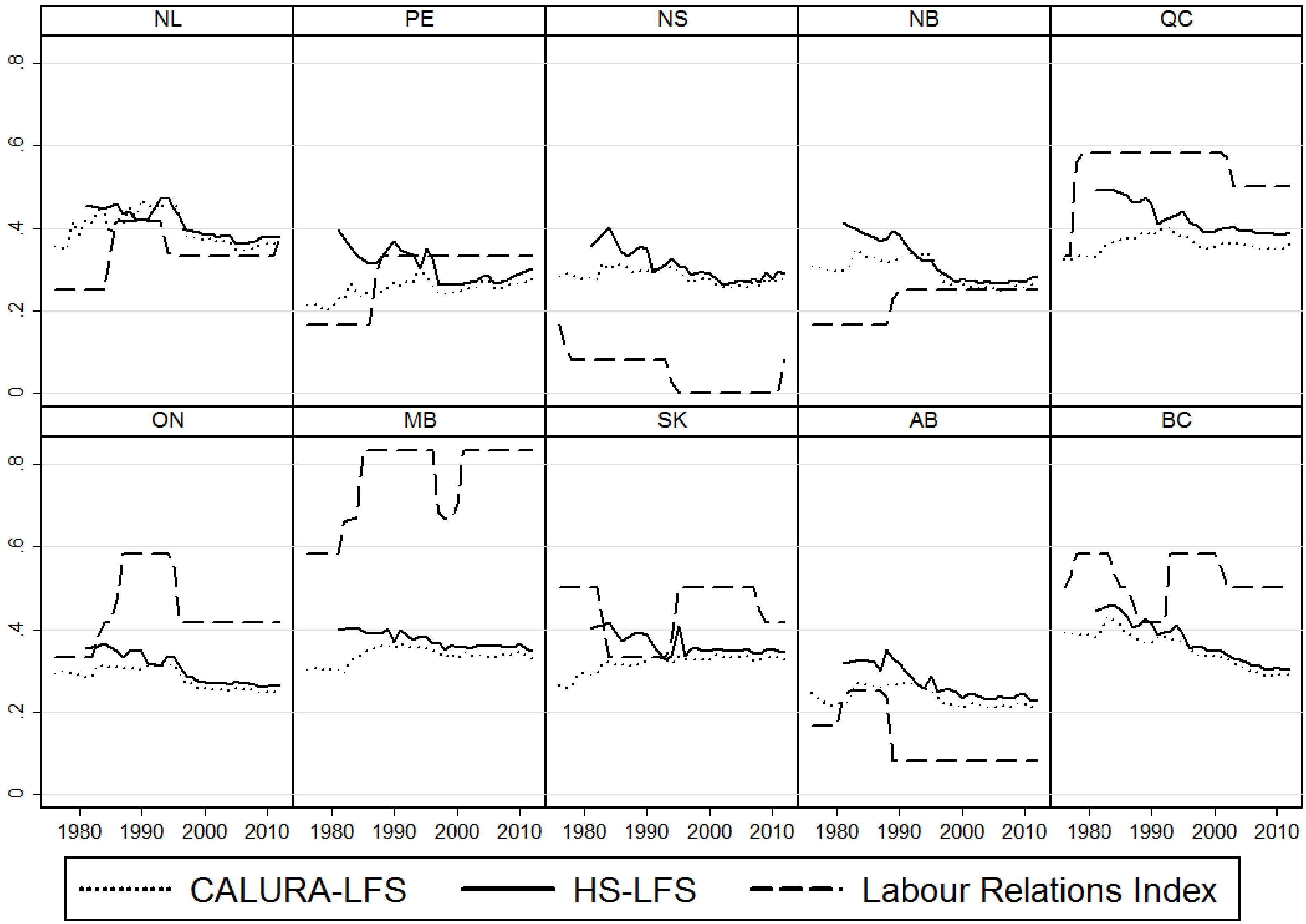
	Coefficient	Standard error	# legal changes	# independent*
lagged density rate	0.6224^{***}	(0.0408)	-	-
secret ballot certification vote	0.0050^{***}	(0.0017)	11	6
first contract arbitration	0.0201^{***}	(0.0047)	6	3
anti-temporary replacement	-0.0105^{***}	(0.0029)	4	0
ban on permanent replacements	0.0307^{***}	(0.0069)	2	0
reinstatement rights	-0.0088^{**}	(0.0039)	7	0
ban on strikebreakers	-0.0032	(0.0056)	2	1
mandatory dues checkoff	0.0013	(0.0089)	4	1
mandatory strike vote	0.0018	(0.0031)	3	0
employer-initiated strike vote	0.0003	(0.0023)	8	3
compulsory conciliation	-0.0065	(0.0047)	4	2
cool off period	0.0048	(0.0077)	2	0
technology re-opener	-0.0056	(0.0048)	1	1
unemployment rate	-0.0233	(0.0642)	-	-
inflation rate	-0.0261	(0.0814)	-	-
manufacturing share	0.0394	(0.0557)	-	-
tastes	-0.0248	(0.0237)	-	-
constant	0.1961^{***}	(0.0340)	-	-

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Notes: Year dummies and province dummies are included in all regressions. The variable *tastes* is between [0,1] with 1 being most supportive of unions. All specifications use the same form of GLS as columns 7 and 8 in Table 2, Var[$\epsilon_{p,t}$]=σ_p^2, Cov[$\epsilon_{p,t}$, $\epsilon_{q,s}$]=$\sigma_{p,q}$, Cov[$\epsilon_{p,t}$, $\epsilon_{p,t-1}$]=ρ_p. * Refers to the number of times this law changed in a given year and province and none of the other 11 laws changed in the same calendar year.

Figure 1: Union density rate and labour relations index by province, 1976-2012



Appendix: Data source descriptions

Table A.1: Household survey descriptions

Survey	**1981 SWH**	**1984 SUM**	**1986-1990 LMAS**	**1991 SWA**	**1993, 1994, 1996 SLID**	**1995 SWA**	**1997-2012 LFS**
Format	Person file	Person File	Person file	Person file	Person (1993,1996), Job (1994)	Person file	Person file
Frequency	One Time (annual)	One Time (annual)	Annual	Two years	Annually	Two years	Monthly
Union status	Monthly	Annually	Weekly	Annually	Monthly	Annually	Monthly
Reference period	Week of 15th of each month	December 1984	Each week	November	Monthly	November	Week of 15th of each month
Variable definitions:							
Class of worker	claswkr: paid worker	clwsker: paid worker	q15cow: paid worker; no distinction of private/public	f05q76: paid worker	clwkr9 (1993,1994), clwkr1 (1996)	cowmain: paid worker	cowmain: public or private
Labour force status	q13: employed.	lfstatus: employed. q11: 'paid worker last week' in reference to reference week	clfs_: employed in week 2 of month	lfstatus: employed q10: 'paid worker last week'	mtwrk1 (1993); mtwr1c (1994); ml*v28 (1996)	lfsstat: employed	lfsstat: employed (at work or absent from work)
Union membership	q26: member only	q13_20; q14_21: member or covered	q112; q113: member or covered	q29: member and covered are combined in one variable	uncoll1 (1993, 1996); uncol1c (1994)	swaq29; swaq30: member or covered	union: member or covered
Industry	siccode: exclude fed gov't employees	sic1_: exclude fed gov't employees	sic`i': exclude fed gov't employees	f05q7374: no way to distinguish federal government	sigc3g10 (1993, 1994); nai3g10, no way to distinguish federal	ind30: exclude fed gov't employees	naics_43: exclude fed gov't employees

				employees	government employees (1996)		
Age	age: < 70 years old	age: < 70 years old	agegrp: < 70 years old	f03q33: < 70 years old	yobg21 (1993); eage26c (1994, 1996)	ageg: < 70 years old	age_12: < 70 years old
Main job	q21 & q22: calculated from data on hours worked per week	Identified by Statistics Canada based on most weekly hours worked	hrs,day: calculated from data on hours worked per week	Job information applies to 'main job'; survey was supplement to LFS. See SWA 1995 codebook	awh (1993, 1994); refers to job #1, no concept of main job in public-use data file (1996)	Job information applies to 'main job'; survey was supplement to LFS	Identified by Statistics Canada based on most weekly hours worked

Table A.2: Comparability of CALURA and LFS union density rates

Issue	CALURA	LFS	COMMENT	SOURCE
100+ members	Only unions (national or international) with 100+ members in Canada reported their union members.	Conditional on being employed, the respondent can answer whether she is in a union or not.	CALURA understates relative to LFS; numerator is smaller.	Mainville/Olinek (1999 p. 11 Table 2). Akyeampong (1998 p. 30.)
Retired / Unemployed	Seasonally unemployed workers with recall rights may be included. Retired very unlikely to be included.	Union question asked conditional on employment. Must be paid worker.	CALURA overstates relative to LFS.	Galarneau (1996 p. 44,46). Table 1 (1970 CALURA report). Mainville/Olinek (1999 p.14). Bill Murnighan (CAW) email July 25, 2013.
Age	All union members. No age limit.	Age ranges from 15 to 70+, each of which has union members in LFS.	CALURA overstates relative to LFS.	Galarneau (1996 p. 44).
\`Employees' denominator	From Dec LFS for each year; conditional on employee.	Data are available for all months of year.	CALURA overstates relative to LFS due to seasonal unemployment in Atlantic Canada. We use July LFS to correct.	Galarneau (1996 p. 44)
Multiple jobholders	Would be counted twice in CALURA.	LFS only asks about main job.	CALURA overstates relative to LFS. LFS only allows main job per respondent so will not double-count.	Akyeampong (1997 p. 45). Historical CALURA data on CANSIM: a note to users.
Union members numerator – report date	Date unions report is as of Dec 31st.	Date report is as of Dec 15th.	No issue.	Galarneau (1996 p. 44). Mainville/Olinek (1999 p. 17 table footnotes). "Historical CALURA data on CANSIM: a note to users".
Union members numerator – new profession	In 1983, teachers, nurses, doctors added based on 1981 legislation.	N/A – these professions included.	CALURA understates relative to LFS (and itself) for pre-1983 SWH.	Mainville/Olinek (1999 p. 3-4, 9). Akyeampong (1998 p.31)
Self-employed	CALURA may include self-employed in (mostly) construction industry	LFS identifies self-employed and we exclude.	CALURA overstates relative to LFS.	"Historical CALURA data on CANSIM: a note to users".