Identifying the Potential of Work-Sharing As a Job-Creation Strategy

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[FINAL DRAFT]

Abstract

Between 1997 and 2000 the Canadian province of Quebec reduced its standard workweek from 44 to 40 hours with the aim of stimulating employment growth. Unlike the European work-sharing policies examined elsewhere, the Quebec policy contained no suggestion or requirement that employers provide wage increases to compensate workers for lost hours. For this reason, among others, the Quebec policy provides a better test of the potential of work-sharing as a job-creation strategy. The evidence suggests that, despite a 20 percent reduction among full-time workers in weekly hours worked beyond 40, the policy failed to raise employment at either the provincial level or within industries where hours of work were affected relatively more.

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1. Introduction

The promise of work-sharing rests on the belief that firms, compelled to reduce their employees' hours of work through either union contract or government regulation, will replace the lost hours of current employees with new employees. On the assumption that marginal increases in hours of work are valued more by the jobless than the employed, initiatives to encourage such a redistribution of work are seen as having the potential to reduce unemployment rates and raise social welfare. Although the theoretical limitations of work-sharing are well established (see in particular Calmfors and Hoel 1988), the empirical evidence has in fact, until very recently, been quite supportive of the potential of this job-creation strategy. Using aggregate time-series data on employment and average weekly hours of work (sometimes instrumented with lags of other variables), econometric research has consistently identified substantial employment-hours substitution elasticities (for reviews see Hart, 1987; Hamermesh, 1993; and Freeman 1998). The problem is, of course, that these historical correlations capture market-driven variation in working hours brought about by the interactions of employers and employees and not the responses of employers and employees to a union or government-imposed reduction in the length of the workweek. Since firms and workers are likely to respond very differently to the latter, particularly since an imposed reduction in working hours will tend to put different pressures on wage rates and productivity, policy inferences drawn from time-series correlations are potentially flawed.

In response, the most recent research on work-sharing has focused on obtaining direct evidence from actual work-sharing initiatives. Hunt (1999) explores job growth in Germany, where between 1985 and 1994 unions negotiated substantial reductions in standard weekly hours of work on an industry-by-industry basis. More recently, Crépon and Kramarz (2002) consider whether François Mitterrand's decision in 1981 to reduce France's nationwide weekly hours standard from 40 to 39 was effective in creating jobs. In sharp contrast to the objective of these policies, the striking finding in both papers is that reduced weekly hours appear to have resulted in job *losses*. In both cases, the primary explanation given for the failure of the policy was its stipulation that employers increase the hourly wage rates of affected workers so as to maintain their weekly earnings.

Even in models where employers are not required to provide employees with wage increases to compensate them for lost hours, economic theory provides reason to doubt the efficacy of work-sharing. For example, workers, particularly those paid on an hourly basis, may be simply unwilling to accept fewer hours. When these complicating factors are combined with compensating wage increases, which must have resulted in substantial negative scale effects (assuming labor is a normal input), it is not surprising that the European initiatives produced job losses. This evidence, however, does not tell us whether work-sharing can work in less regulated labor markets where unions and governments are unable or unwilling to impose full-wage compensation. After all, under the right assumptions, including constant wages, the theory is quite supportive of the potential of reduced hours to raise employment. Dismissal of work-sharing as an effective job-creation strategy should therefore require that it be put to the test under more ideal conditions.

Between October 1997 and October 2000, the Canadian province of Quebec gradually reduced its statutory standard workweek, i.e. the weekly hours beyond which a wage premium of time-and-a-half must be paid, from 44 to 40 hours, with the explicit aim of stimulating employment growth in the province.[1] A significant limitation of the initiative is that the working

[1] One year after the 1995 Quebec referendum on sovereignty, the new premier of Quebec, Lucien Bouchard, brought together government, business and labor leaders in the province in the hope of finding ways to stimulate employment growth. The most significant initiative agreed on was the reduction in the province’s standard workweek from 44 to 40. Besides helping to create 15,000 jobs, the hope was that the policy would provide the impetus for employers and workers in the province to continue efforts to further reduce hours of work, with the goal of creating jobs. For a description of the reasoning behind Quebec’s work-sharing policy see Desmarais (1999).

time regulations in Quebec's *Labour Standards Act* only apply de facto to employees in the province who are paid on an hourly basis and not covered by a union contract.[2] Since these workers account for roughly 30 percent of provincial paid employment and the vast majority already had usual workweeks of 40 hours or less, even in the best-case scenario, the aggregate effect on the provincial unemployment rate would have been negligible and empirically undetectable. However, by comparing changes in hours of work and employment between sectors of the Quebec economy affected relatively more and less by the policy, it is possible to identify the *potential* of work-sharing as a job-creation strategy. For a number of reasons, the Quebec experiment provides a better test of this potential than the European policy evaluations currently found in the literature. First, the legislation contained no requirement or suggestion that employers adjust hourly wage rates to compensate workers for reduced workweeks. Second, nonunionized hourly-paid workers are disproportionately unskilled with relatively high rates of unemployment. Not only do these workers have little bargaining leverage, so that obtaining compensating wage increases without an offsetting improvement in productivity seems unlikely, but fixed costs of employment for these workers, such as training costs and certain fringe benefits, are also likely low, making substitution of jobs for lost hours more cost feasible. Fourth, the argument that the main limitation of work-sharing is the difference in skills between the employed and unemployed, which makes substituting the lost hours of the former with jobs for the latter difficult (Freeman 1998), is arguably less applicable for this relatively unskilled group. Finally, according to survey data, nonunionized hourly-paid workers in Quebec are more likely than workers in any other Canadian province to report a preference for fewer hours of work (as opposed to more or the same) at their current wage rate.[3]

[2] For the reasons for these exclusions, see the discussion in Section 3

[3] Authors' own tabulation of Statistics Canada's 1995 Survey of Work Arrangements.

Using observationally identical hourly-paid workers in Ontario and salaried workers in Quebec as comparison groups, I identify a 20 percent decrease in weekly hours worked beyond 40 among full-time hourly-paid nonunionized workers in Quebec resulting from the reduction in the province's standard workweek. Among men, who work substantially more overtime, these hours account for roughly 2 percent of aggregate work hours in the province. Among women there is little or no evidence of a reduction in aggregate hours resulting from the policy. Overall, the decrease in hours worked beyond 40 accounts for no more than 1 percent of aggregate work hours in the province, implying at the very best an offsetting 1 percent increase in employment. Evidence from the complete 20 percent samples of the 1996 and 2001 population Censuses suggest the policy failed to raise employment at either the provincial level or within industries where there is clear evidence of substantially larger hours effects. In fact, among men the estimates appear more consistent with the European experience of compensating wage increases and job losses. Given the favorable conditions in Quebec, these results suggest there is little, if any, potential for work-sharing as job-creation strategy.

The remainder of the paper is organized as follows. The following section considers some theoretical predictions of work-sharing initiatives under various assumptions about the exogeneity of work hours and wages. The third section discusses the double and triple-difference empirical identification strategies employed throughout the paper. The following three sections of the paper examine the hours, employment and wage effects of Quebec's work-sharing experiment. The final section summarizes the main findings.

2. Theory

The predicted effects of work-sharing from labor demand theory are well documented elsewhere (e.g. Hart 1987, Hamermesh 1993). In its simplest version, optimizing firms choose employment to maximize weekly profits subject to an exogenous hourly wage rate and level of weekly hours. If weekly output is determined by the function $F(N,h)$, where N is the employment level and h are weekly hours, and costs are the product of N, h and the hourly wage rate, w, the first-order condition for profit maximization (normalizing the output price) equates the value of what an additional worker could produce in a week with the cost of hiring an additional worker for a week, i.e. $F_N = wh$. When the length of the workweek, h, is reduced by a government policy, both the value of marginal product of employment (F_N) and the marginal cost of employment (wh) decrease. If the former (the value of an additional worker) falls by less than the latter (the cost of an additional worker), profit-maximizing employers will find it optimal to raise the employment level (thereby maintaining equality of the first-order condition). Since the necessary conditions for this outcome ($F_{Nh} < w$ and $F_{NN} < 0$) are satisfied by all standard production functions, the simplest labor demand theory predicts an unambiguous positive employment response to a reduction in hours of work. Of course, this is a drastic oversimplification of how real labor markets work. Not surprisingly, extending the model to make it more realistic, virtually always makes the desired employment effect less likely.

First, Hart (1987) and Calmfors and Hoel (1988) evaluate the efficacy of work-sharing in a model where firms choose the length of the workweek given some overtime premium which must be paid for hours employed beyond the standard workweek. In special cases, such as where fixed labor costs are particularly high, it may be optimal for firms to employ workers some overtime hours. At this optimum, a decrease in the standard workweek has no effect on the marginal cost of additional hours (since the last hour employed was already paid at the premium rate), but unambiguously raises the marginal cost of employment, due to the new higher price of hours between the old and new standard, all of which continue to be purchased. So, where firms in Quebec were employing workers more than 44 hours per week, we should expect the reduction in the standard from 44 to 40 to *increase* actual weekly hours of work and *lower* employment.[4] In fact, only where workers were employed between 41 and 44 hours per week does the marginal cost of additional hours rise and is there reason to expect any decrease in hours of work.[5]

Secondly, and perhaps most importantly, if the assumption of exogenous and constant wages is relaxed, the predictions of the simplest models of work-sharing become even more dubious. In the most extreme case, the hourly wage is assumed to be perfectly flexible (Trejo 1991). An increase in the price of hours employed over some range due to a reduction of the standard workweek may then have no real effects, because employers simply adjust straight-time wages downward in response, leaving workers' hours of work and weekly earnings unaffected. With no employment-hours substitution, Trejo refers to this model as the "fixed-job" approach, to distinguish it from the more standard "fixed-wage" approach. Of course, it may be that rather than wages adjusting freely, straight-time wages adjust upward, through union contracts or government regulations, so as to maintain workers weekly earnings. Such compensating wage increases, since they affect the marginal cost of all work hours, imply substantial negative scale effects, potentially offsetting any of the intended hours-employment substitution of the policy. Crépon and Kramarz (2002) present such a model, with predictions that are consistent with both the French and German experience.

Thirdly, the partial equilibrium theory considered to this point completely ignores possible labor supply responses to a reduction in the standard workweek. For workers already

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[4] These results assume there is no heterogeneity in hours of work within firms. To the extent that work schedules vary within firms, the marginal cost of employment may be unaffected by a reduction in the standard workweek, since new hires do not necessarily work the inframarginal hours which must now be paid at a premium.

[5] In the year prior to Quebec's reduction in its statutory standard workweek, roughly 7 percent of hourly-paid, nonunionized workers in the province reported actual workweeks between 41 and 44. These hours accounted for 2 percent of the total work hours of hourly-paid nonunionized workers in the province and less than 1 percent of aggregate work hours in the province.

employed overtime, the reduction of the standard workweek produces only an income effect, which will tend to reduce desired hours of work, potentially offsetting the increase in hours coming from the labour-demand side. However, for workers employed between the old and new standard, the policy also produces a substitution effect, which tends to raise desired hours of work and reduces the likelihood that work-sharing leads to employment growth. In addition to Freeman's (1998) concerns regarding the willingness of workers to accept shorter workweeks and the substitutability of employed and unemployed workers, these labor supply adjustments are likely to have significant dampening effects on any potential employment-hours tradeoffs.

3. Identification

The challenge in evaluating the effect of policies such as Quebec's work-sharing initiative is to know the counterfactual: what would have happened, had the policy not been introduced? The usual approach of the popular *difference-in-differences* estimator (see Angrist and Krueger, 1997, for a general discussion) is to compare the change in some outcome of interest following a policy change between individuals affected by the policy (at least potentially) and some similar group of individuals in a neighboring jurisdiction. The complication with this strategy is there may be unobserved differences between jurisdictions that are entirely independent of the policy. Hamermesh and Trejo (2000) address this possibility by adding a comparison group *within* the jurisdiction experiencing the policy change, resulting in their *difference-in-difference-in-differences* estimator.

In this paper I follow the triple-difference approach of Hamermesh and Trejo (2000). Although the raison d'être of Quebec's *Labour Standards Act* is to provide all workers in the province with minimum standards of employment, in practice its working time provisions apply only to employees who are paid on an hourly basis and not covered by a union contract. Salaried workers are excluded because courts in the province have ruled that the requirement to pay an overtime wage premium of time-and-a-half for hours in excess of the standard workweek is inapplicable where an hourly rate is not paid. Unionized workers, on the other hand, virtually always have standard workweeks below the statutory standard, so the reduction in the statutory workweek is unlikely to have been binding in unionized settings.[6] Given these exclusions, the obvious between-jurisdiction comparison group are otherwise similar hourly-paid nonunionized workers in neighboring Ontario since: (i) its industrial structure is similar; (ii) its statutory overtime premium is also 50 percent of the regular wage rate; and (iii) its statutory standard workweek was also 44 hours in October 1997, but was never reduced. As for the within-jurisdiction comparison group, I present results using salaried nonunionized workers in Quebec.[7] A potential complication with this comparison group is that a policy of reducing the working hours of hourly-paid workers could plausibly spill over to salaried workers employed in the same firms. However, managers and supervisors, the group for whom such spillovers seem most likely, are among a number of occupations (and industries) *explicitly* excluded from the legislation's working time provisions.[8] Since these groups of workers are dropped from the analysis from the outset, such spillovers seem less likely.

4. Hours effect

4.1. Data

[6] This information comes from personal communication with the *Commission des Normes du Travail* who are responsible for enforcement of Quebec's labour standards legislation.

[7] All of the analyses have also been done using hourly-paid unionized workers as the within-jurisdiction comparison group. It turns out the hours effect estimates are virtually identical whether salaried or unionized workers are used as the Quebec comparison group. The employment effect estimates, on the other hand, suggest substantially larger job losses when unionized workers are used. This is consistent with the relative employment gains experienced in Ontario's nonunion sector over the period of the policy as a result of a concomitant change in Ontario's rules for certifying new union bargaining units (see Martinello 2000). Due to this complication, the results using salaried workers as a comparison group are preferred. All the results using unionized workers are available on request.

[8] Workers employed in the federal jurisdiction are similarly dropped. See the Appendix for details.

Quebec's statutory standard workweek was first reduced from 44 to 43 in October 1997. It was reduced by an additional hour in each subsequent October, until October 2000 when it reached 40. The only source of microdata on hours of work spanning this period that allows me to identify enough hourly-paid nonunionized workers in Quebec to estimate policy effects with any meaningful precision is the monthly Canadian Labour Force Survey (LFS). I construct a 6-year panel spanning the reduction in Quebec's standard workweek by pooling October-September LFS files over the period from 1996 to 2002.[9] The LFS collects information on both actual hours in the survey reference week and usual weekly hours. Friesen (2002) argues that usual hours are preferred when identifying policy effects on the long-run use of overtime hours. However, it is theoretically possible that employers distribute a fixed number of weekly overtime hours across employees in such a way that in the typical week no employee works any overtime. Since a permanent reduction in these occasional overtime hours could just as plausibly result in employment gains, in this case we would clearly want to examine actual hours worked. Regardless, beginning in January 1997 the LFS usual hours question explicitly asked respondents to exclude any overtime hours worked. For this reason the actual weekly hours data are used throughout. In addition, the decision was made to drop all proxy responses (which account for roughly half of all observations), due to evidence of substantial rounding of the actual weekly hours data. Given the importance of distinguishing hours in the 41-44 range from 40 or 45, this rounding is of greater concern here than in analyses of first or second moments of the hours distribution.[10]

4.2. Nonparametric analysis

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[9] The Canadian LFS did not officially begin identifying whether respondents are paid hourly or their union status until January 1997, but the revised survey was gradually introduced to all new rotations beginning in August 1996.

[10] All of the main findings are robust to the inclusion of the proxy responses, although the magnitudes of the estimated effects tend to be somewhat weaker.

Before estimating the magnitude of the hours effect, it is worth considering whether observed relative changes in the hours of work distribution of hourly-paid nonunionized workers in Quebec over the period of the policy change are consistent with what is arguably the strongest (and certainly the most intuitive) prediction of the static labor demand theory: a decrease in hours of work where workers were employed exactly the standard workweek (i.e. the kink in the firm's isocost curve) prior to its reduction. In this section, I estimate nonparametrically double and triple-difference effects of the policy on the probability of working 40, 41, 42, 43 and 44 hours per week, while controlling for observable differences between workers in the "treatment" and "comparison" groups.

Suppose *f(t)* is the probability density function of some random variable *t* and *F(t)* is its CDF. Then Donald, Green and Paarsch (2000) suggest estimating a hazard model with duration variable *t* and covariates *X*, and using the relationship:

$$f(t) = (1 - F(t \mid X)) \cdot \lambda(t \mid X) = S(t \mid X) \cdot \lambda(t \mid X) \tag{1}$$

to back out the underlying density function, where λ is the hazard function and *S* is the survivor function. By allowing for a perfectly flexible baseline hazard, (1) amounts to a nonparametric estimator of the distribution of *t*.[11] In the case where *t* is actual weekly hours of work, *h*, the conditional hazard function is simply:

$$\lambda(h_j \mid X) = prob(h = j \mid h \geq j; X) \tag{2}$$

i.e. the probability that weekly hours are *j,* given that they are not less than *j*, which since *h* is discrete, can be straightforwardly estimated using logit model (see Jenkins 1995). The likelihood function for the logit hazard model is then:

$$L = \prod_{i \in S} \Omega(h_{ji} \mid X_i \beta) \cdot \prod_{i \notin S} (1 - \Omega(h_{ji} \mid X_i \beta)) \tag{3}$$

[11] To the extent that the conditioning variables, *X*, affect the hazard parametrically, (1), more accurately, amounts to a *semiparametric* estimator.

where Ω is the logistic function, $F(x) = exp(x)/(1+exp(x))$, and i indexes individual observations in an expanded sample which includes j observations for each observation in the original LFS sample and S is the set of observations $h=j$.[12]

To see whether there is any evidence of a policy effect on hours worked between 40 and 44, this logit hazard model is estimated by specifying a 7-element vector of dummy variables V_i (<40, 40, 41, 42, 43, 44, and >44), where all elements, except the one corresponding to observation i's value of h, are equal to 0. If P_i is a vector of dummy variables indicating in which of the 5 annual periods after October 1996-September 1997 individual i is observed, and Q_i and H_i are a Quebec and paid-by-the-hour indicators respectively, the policy effect is given by the parameter vector δ, in the likelihood function:

$$L = \prod_{i \in S} \Omega(g(\cdot)) \cdot \prod_{i \notin S} (1 - \Omega(g(\cdot)) \tag{4}$$

where in the double-difference case

$$g(\cdot) = V_i \alpha_1 + (V_i \otimes P_i)\alpha_2 + (V_i \cdot Q_i)\alpha_3 + (V_i \otimes P_i) Q_i \delta + X_i \beta\,; \tag{5}$$

and in the triple-difference case

$$\begin{aligned} g(\cdot) = V_i \alpha_1 &+ (V_i \otimes P_i)\alpha_2 + (V_i \cdot Q_i)\alpha_3 + (V_i \cdot H_i)\alpha_4 + (V_i \otimes P_i) Q_i \alpha_5 + \\ &(V_i \otimes P_i) H_i \alpha_6 + (V_i \cdot Q_i \cdot H_i)\alpha_7 + (V_i \otimes P_i) Q_i H_i \delta + X_i \beta\,. \end{aligned} \tag{6}$$

Specification (5) is estimated using a sample of hourly-paid nonunionized workers in Quebec and Ontario; and specification (6) adds the sample of nonunionized salaried workers in the two provinces. To sharpen the estimates both samples are restricted to full-time workers (defined as usual weekly hours of 30 or more). Finally, in all specifications the vector X_i includes controls for month, sex, age, education (high school diploma, some post-secondary, post-secondary credential, university degree, and graduate degree), establishment size (<20, 20-99, and ≥ 100), industry (44 categories) and occupation (25 categories).

[12] Friesen (2002) similarly uses a logit hazard model to estimate the distribution of overtime hours.

The estimated policy effects in Table 1 are obvious and entirely consistent with the simple static labor demand theory. In both specifications, there is a significant decrease in the hazard at 44 in the first period of the policy change. The double-difference estimate also shows a significant increase in the hazard at 43 (although this does not appear in the triple-difference specification). In the following period, when the statutory standard is 42, both specifications show significantly lower hazards at 44 and 43 (relative to when the standard is 44) and a significant increase in the hazard at 42. And when the standard falls to 41, the hazards remain lower at 44 and 43, but now increase significantly at 41. Finally when the standard has reached 40 in the final two periods, the hazards at 44, 43 and 42 all appear permanently lower and the hazard at 40 significantly and permanently higher. The timing of this "wave" in the Quebec hours distribution is robust to the double or triple-difference specification.[13] This pattern undoubtedly reflects the province's work-sharing initiative. The interesting question is whether there were enough workers employed over this range of the hours distribution to produce any meaningful reduction in working hours. And if there was, to what extent might it have been offset by increased hours of work where workers were already employed beyond the statutory standard (i.e. more than 44 hours per week prior to October 1997).

4.3. Linear model estimates

The policy effect on hours worked over any range can be estimated by specifying the baseline hazard vector, V_i, appropriately. If h_j^k is defined as the number of hours worked between j and k (so a 44-hour workweek is given by $h_{41}^{44} = 4$), we can, for example, estimate the effect of the policy separately on $E(h_{41}^{44})$ and $E(h_{45}^{60})$, where the distribution must be censored at

[13] It is also robust to whether salaried or unionized workers within Quebec are used as the within-jurisdiction comparison group.

60 for identification.[14] Unfortunately, the nonparametric approach quickly becomes unwieldy when estimating over broader ranges of the hours distribution. For example, in identifying the policy effect on overall hours of work, $E(h_1^{60})$, V_i is a 60-element vector, which in the triple-difference case involves estimating 420 baseline hazards. Instead, in Table 2 I present predictions based on OLS estimates of a marginal standard workweek effect using all 6 years of the data (nonparametric results for the relatively narrow ranges $E(h_{41}^{44})$ and $E(h_{45}^{60})$ give similar results and are available on request). Specifically, in the triple-difference case I estimate:

$$h_{ji}^k = \alpha_1 + h_i^s\alpha_2 + Q_i\alpha_3 + H_i\alpha_4 + h_i^sQ_i\alpha_5 + h_i^sH_i\alpha_6 + Q_iH_i\alpha_7 + h_i^sQ_iH_i\delta + X_i\beta + e_i \tag{7}$$

where h_i^s is the standard workweek in Quebec in the year in which individual i is observed and Q_i, H_i and X_i are, respectively, a Quebec dummy, a paid-by-the-hour dummy, and the vector of observable characteristics defined in (5) and (6) above. Expected weekly hours of nonunionized workers in Quebec before and after the policy change are then predicted at $\overline{X}$ assuming 44 and 40-hour standard workweeks respectively. Double-difference estimates are also reported which are based on the restricted form of (7) in which $\alpha_4 = \alpha_6 = \alpha_7 = \delta = 0$ and a policy effect given by α_5 instead of δ. Finally, unlike the nonparametric model, equation (7) is easily identified separately for men and women.[15] Since the incidence of working more than 40 hours per week varies substantially between men and women, Table 2 also presents separate results for men and women.

With respect to hours worked between 41 and 44, all the double-difference results in Table 2 suggest decreases exceeding 20 percent. However, when salaried workers in Quebec are

[14] The double and triple-difference baseline hazards at t are only identified if there is at least one observation with $h_j^k = t$ from each of the treatment and comparison groups. Weekly hours beyond 60 account for roughly 1 percent of the full sample.

[15] Identification of the hazard model requires at least one observation with $h_j^k = t$ from each of the treatment and comparison samples. This requirement becomes a problem for the female hours distribution beyond 50.

added as an additional comparison group, the female result becomes highly insignificant, whereas the male result is remarkably robust, continuing to suggest a decrease of nearly 30 percent. Clearly an important part of the explanation for this difference is that women worked only 1/3 of the hours men did over this range before the policy change (see before levels). In contrast to the standard labor demand theory, the estimates for men also suggest decreases in hours worked between 45 and 60, although both estimates are statistically insignificant. Part of the explanation for this result may be heterogeneity in work schedules within firms, so that the marginal cost of employment does not necessarily rise in firms where some workers were employed beyond the original standard workweek. For women, on the other hand, the double and triple-difference result are of opposite sign, but again both, particularly the triple-difference estimate, are insignificant. Using the property $E(h_{41}^{60}) = E(h_{41}^{44}) + E(h_{45}^{60})$, the double and triple-difference results in columns (5) and (6) of Table 2 suggest overall decreases in hours worked beyond 40 of 25 (from 1.42 to 1.07) and 15 (from 1.5 to 1.27) percent respectively.[16]

What is more consistent with the theory is that none of the results suggest changes in hours worked between 1 and 40 resulting from the reduction in the standard workweek from 44 to 40. All the estimates are close to zero and measured with considerable precision. When combined with the reductions in hours worked beyond 40, the results for men suggest a decrease in total hours worked of somewhere between 1 and 2 percent. In contrast, for women both estimates are close to 0 and statistically insignificant. Overall, the reduction in Quebec's standard workweek appears to have reduced the working hours of full-time workers in the province by no more than 1 percent.[17] So even with an employment-hours substitution elasticity of -1, the results imply no more than a 1 percent increase in full-time employment of hourly-paid nonunionized

[16] In comparison, when hourly-paid unionized workers are used as the within-Quebec comparison group, the estimates suggest an overall 12 percent decrease in hours worked beyond 40.

[17] This result is robust to whether salaried or unionized workers are used as the within-Quebec comparison group.

workers in the province. However, to the extent that the work-sharing has potential to raise employment, and women and men are not easily substitutable, the sharp difference in hours effects between men and women implies we should see relatively large employment gains among men over the period of the policy change. This comparison therefore provides an additional source of variation to identify the potential of work-sharing.

5. Employment effect

5.1. Data and identification

Given the analysis to this point, it would seem natural to use an industry-level analysis and the double and triple-difference identification strategies above to estimate the policy effect on employment at the provincial level. The obvious difficulty with this approach is the aggregate employment response is likely too small to be identified with any meaningful precision. This is particularly true if estimated with the survey data used to identify the hours effect of the policy.[18] Instead, I focus on identifying the *potential* of Quebec's work-sharing initiative by comparing double-difference employment effect estimates between industries with larger and smaller shares of workers paid on an hourly basis. Since the policy primarily affected workers paid hourly, to the extent that the policy achieved its intended objective, we should observe relatively large employment gains among industries with larger shares of hourly-paid workers. Moreover, since this estimation does not require individual-level information on whether workers are paid hourly or their union status, in this section I instead use the 20 percent samples of the population from the Canadian Census, which provide industry-level employment counts with essentially no

[18] Even ignoring this issue, the triple-difference estimate of the employment effect is problematic since it amounts to comparing relative employment changes of hourly-paid and salaried workers within detailed industries in Quebec. A key assumption in this identification is that salaried employment levels were unaffected by the policy. To the extent that new salaried jobs replaced the lost hours of hourly-paid workers, this assumption is clearly problematic.

sampling error.[19] Fortunately, for its present purpose the timing of the Canadian Censuses could not have been better. Conducted in May of 1996 and 2001, these files capture employment 17 months before the process of reducing Quebec's standard workweek began and 7 months after it was complete.

Specifically, the double-difference estimate of the employment effect of Quebec's work-sharing policy is given by δ in the equation:

$$\log E_{ipt} = P_t I_i \alpha_{1i} + P_t Q_p \delta + u_{ip} + e_{ipt} \tag{8}$$

where E_{ipt} is the employment level of all wage and salary workers employed in industry i, province p, and year t; P_t is a dummy variable indicating a year 2001 observation; I_i is a vector of industry dummies; Q_p is a Quebec dummy; u_{ip} is a vector of province-specific industry fixed effects; and e_{ipt} is an error term.[20] Since measured employment is not restricted to hourly-paid nonunionized workers, in the sample of all industries δ identifies the aggregate provincial-level employment effect of Quebec's work-sharing policy. However, by comparing the magnitude of this effect between subsamples of industries with larger and smaller shares of workers paid on an hourly basis, information on the *potential* of Quebec's policy to raise employment is obtained. I construct a dataset containing employment levels within 4-digit industries and merge industry-level shares of workers paid on an hourly basis in Quebec estimated using common industry codes from the Labor Force Survey. The Census also contains data on actual hours of work in the reference week. To confirm the results in Section 4 and to provide some assurance that industries with larger shares of hourly-paid workers were in fact affected relatively more by the policy, equation (8) is also estimated using mean actual weekly hours beyond 40 and total hours as its

[19] These files include a sampling weight (with a mean of 5), which corrects for non-response bias. The variance of this weight is however very low, so it has essentially effect on the estimates reported.

[20] As usual, all workers employed in the federal jurisdiction, defined in the appendix, are excluded from the counts.

dependent variable.[21] Finally, since we expect larger employment responses among men, results are presented separately for men and women.

The first two columns of Table 3 report the estimated effect on hours worked beyond 40 per week. Since the Census sample includes salaried and unionized workers, we should expect somewhat smaller effects here than those identified using the LFS data. Indeed, using the sample of all industries (middle row of Table 3) the results suggest decreases in weekly hours worked beyond 40 of 14 (men) and 12 (women) percent, compared to 26 (men) and 16 percent (women) in Table 2. More importantly, for both men and women the estimated overtime effect clearly increases, as the sample is restricted to industries with larger shares of workers paid on an hourly basis. For example, in the sample of 156 industries where more than 60 percent of Quebec employees are hourly paid (first row of Table 3), the double-difference estimates suggest 18 (men) and 19 (women) percent decreases in hours worked over 40. In contrast, in the sample of 37 industries where less than 40 percent of employees are paid by the hour (last row of Table 3), the double-difference estimates suggest 6 (men) and 3 (women), and statistically insignificant, percent decreases.

When the dependent variable in (8) is instead total weekly hours, the effects are, as in the LFS data, much smaller than the effects on hours beyond 40. They also continue to suggest relatively large effects for men (although the difference here is substantially smaller than in Table 2). So for men, the full sample estimate (middle row) suggests a 1 percent decrease in total hours of work, compared to a 0.6 percent decrease for women. Most importantly, the estimates once again monotonically increase, as the sample is restricted to industries with larger shares of workers paid by the hour and smaller among industries with smaller proportions of hourly paid workers. Given this clear pattern in hours of work effects across industries, to the extent that

[21] To make the estimates more comparable to those estimated using the LFS data, the samples used to estimate the hours effects are restricted to full-time actual weekly hours (i.e. 30 or more).

Quebec's work-sharing policy had any job-creation potential we should observe relative employment gains in industries with larger shares of employees paid hourly.

The fifth and sixth columns of Table 3 present the employment effect estimates for men and women respectively. In the full sample of all industries, the point estimate for men suggests a 0.5 percent decrease in employment resulting from the 4-hour reduction in Quebec's standard workweek. In contrast, the full sample point estimate for women, whose hours of work were relatively unaffected by the policy, suggests, if anything, employment gains. Although it is theoretically possible that women replaced the reduced hours of men, the amount of substitution needed to generate these results seems highly unlikely given the very different occupations that women work in. More importantly – for identification of the policy's potential – comparison of the double-difference point estimates between industries with larger and smaller shares of hourly-paid employees does not suggest relative employment gains in industries affected more by the policy. So for example, in the sample of 156 industries where the share of employees paid hourly exceeds 60 percent (first row), the point estimate implies a 0.4 percent reduction in male employment. In contrast, in the 37 industries with fewer than 40 percent paid hourly (last row), the estimate suggests virtually no change in male employment. Given the clear differences in hours of work estimates between these two samples, these results imply, if anything, job losses resulting from Quebec's work-sharing policy. The employment estimates for women also exhibit no clear monotonic pattern comparable to the hours of work effect estimates. All of the estimated employment effects are, however, estimated somewhat imprecisely, despite being based on a 20 percent sample of the population. So, in all cases, substantial employment gains fall within 95% confidence intervals of the estimated effects. However, the standard errors are estimated conservatively (robust and clustered within each 6-year industry panel). The fact that none of the male point estimates, using a one-in-five sample of the population, imply employment gains, despite the hours of work effects of the policy being concentrated among men, does suggest that Quebec's work-sharing policy had little, or no, potential to raise employment levels in the province. In the following section, I consider whether the apparent failure of Quebec's policy to create jobs can be explained, at least in part, by the European experience of compensating wage increases.

6. Wage effect

6.1. Data and identification

Beginning with new sample rotations in September 1996, the LFS provides data on individual hourly-wage rates. Table 4 reports the results of estimating the same double and triple-difference linear models used to estimate the hours effects, but with the log hourly wage as the dependent variable. The double-difference estimator now captures how the wages of nonunionized hourly-paid workers in Quebec changed differently over the period of the policy than that of observationally comparable workers in Ontario. The triple-difference estimator, on the other hand, adds nonunionized salaried workers in Quebec to this comparison, to insure that any between-province difference does not reflect province-specific factors unrelated to Quebec's work-sharing initiative. In addition, since there is strong evidence that workers were affected differently across the hours distribution, both specifications are estimated using the full sample and separately for the sample of workers employed less than 40 hours per week, between 40 and 44 hours per week, and more than 44 hours per week. Assuming no compositional changes within these groups, the estimates of δ reflect changes in the wages of individual workers. Given the differences identified for men and women, wage effects are also estimated separately for men and women. Wages are adjusted for inflation using the national monthly consumer price index. Finally, to reduce measurement error, observations are again restricted to non-proxy responses.

6.2. Results

With the exception of the double-difference estimate for men, all the unconditional (on hours) estimates are very close to zero, suggesting a very small or no effect of the policy on wage rates at the provincial level. This is consistent with the absence of any policy requirements or suggestions that employers provide compensating wage increases and, at least theoretically, should have increased the likelihood of the policy achieving its intended job-creation objective. However, when the sample is conditioned on hours of work, there is some evidence of wage effects. In particular, both the double and triple-difference estimates for men suggest falling hourly wage rates for men with workweeks above 44 hours and increased wage rates for men employed between 40 and 44 hours per week. These results are remarkably consistent with the labor demand theory and previous findings. For workers employed more than 44 hour per week, the increase in the price of inframarginal hours can be entirely absorbed by reducing the straight-time wage (Trejo 1991). The estimated reductions in wage rates for men employed more than 44 hours per week before and after the policy change appear to be somewhere between one-third and one-half of what would have been needed to entirely absorb the increased cost of hours worked between 41 and 44.[22] Rising wages for workers employed between 40 and 44 hour per week, on the other hand, is consistent with the European experience of compensating wage

[22]If weekly earnings are:

$$E = wh + pw(h - h_s)$$

where w is the hourly wage rate; p is the overtime premium; h are weekly hours of work; and h_s is the statutory standard workweek; then:

$$dE = 0 \iff \frac{dh_s}{h_s} = \frac{h - p(h - h_s)}{ph_s}\frac{dw}{w}$$

if $dh = 0$. Assuming $h = 50$ (mean weekly hours conditional on $h > 44$ in the Quebec sample), $p = 0.5$, and h_s drops by 9 percent (44 to 40), mean weekly earnings would have been maintained if wages rates had fallen by 4 percent. The double and triple-difference male estimate for $h > 44$ suggest decreases of 3.3 and 2.4 percent respectively.

increases. In this case, the magnitudes of the estimated wage increases are roughly one-quarter to one-half of what would have been needed to maintain workers' weekly earnings.[23]

As would be expected, given the relatively weak hours effects estimated for women, the results in Table 4 do not suggest compensating wage increases for women. However, just as for men, there is some evidence of falling wages for women who worked more than 44 hours per week both before and after the policy change. The latter result is in fact stronger for women, and is in the double-difference case statistically significant, which is perhaps not surprising given the hours results for women are more suggestive of no change, or even an increase, in hours worked beyond 44 per week. Comparing the wage estimates for men and women, the results are remarkably consistent with the finding of relative employment losses for men. However, it remains a possibility that the estimated wage gains for men with workweeks between 40 and 44 reflect productivity improvements. After all, given that they were not required by the policy, obtaining compensating wage gains seems somewhat unlikely for this group of relatively unskilled workers, who presumably have little bargaining power. Yet, even if the apparent wage gains for men experiencing reduced workweeks reflect productivity gains, the evidence suggests that these gains were smaller than the decrease in hours of work. Taken together with the employment effect estimates then, it would appear that any productivity improvements resulting from the policy were not enough to offset the reduction in the scale of production resulting from reduced hours of work.

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[23] Since these workers never received an overtime premium,

$$dE = 0 \iff \frac{dh}{h} = -\frac{dw}{w}.$$

In the period before the policy change, about 75 percent of workers in the 40-44 hours range had exactly 40-hour workweeks. If half of the remaining 25 percent went from 44 to 40-hour workweeks, the decrease in hours would only have been 1.25 percent. The double and triple-difference wage effect estimates for men employed between 40 and 44 hours per week suggest wage increases of 0.3 and 0.7 percent respectively.

7. Summary

Between 1997 and 2000 the Canadian province of Quebec gradually reduced its statutory standard workweek from 44 to 40 hours in the hope of stimulating employment growth in the province. Unlike the European work-sharing policies examined elsewhere, the Quebec policy contained no suggestion or requirement that employers provide wage increases to compensate workers for reduced workweeks. For this reason, among others, the Quebec policy provides a better test of the potential of work-sharing as a job-creation strategy.

Using observationally identical workers in Ontario and nonunionized salaried workers in Quebec as controls, I identify a 20 percent decrease in weekly hours of work beyond 40 among full-time hourly-paid nonunionized workers resulting from this policy. Among men in Quebec, who work substantially more overtime, these hours account for roughly 2 percent of aggregate work hours. Among women there is little or no evidence that Quebec's work-sharing policy reduced total working hours. Overall, the decrease in weekly hours beyond 40 accounts for no more than 1 percent of aggregate work hours, implying at the very best an offsetting 1 percent increase in employment. Evidence from the 20 percent samples of the 1996 and 2001 population Censuses suggest the policy failed to raise employment at either the provincial level or within industries where there is clear evidence that hours of work were affected relatively more. In fact, among men the Census estimates appear more consistent with the European work-sharing experience of compensating wage increases and job losses. Given the relatively unskilled and unregulated setting for Quebec's work-sharing initiative, these findings suggest there is little, if any, potential for work-sharing to raise employment.

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Appendix

Workers in the federal jurisdiction are excluded from all the analyses. Also, workers within Quebec's legal jurisdiction, but not covered by the *Labour Standards Act*'s working-time regulations, are excluded from all the analyses.

Federal jurisdiction workers are identified using the following NAICS codes:

481 – Air transportation
482 – Rail transportation
4831 – Deep sea, coastal and Great Lakes water transportation
484 – Truck transportation
486 – Pipeline transportation
4881 – Support activities for air transportation
4882 – Support activities for rail transportation
4883 – Support activities for water transportation
491 – Postal service
513 – Broadcasting and telecommunications
521 – Monetary authorities, Central Bank
5221 – Depository credit intermediation
911 – Federal government public administration
914 – Aboriginal public administration
919 - International and Other Extra-Territorial Public Administration

Noncovered workers are identified using the following NAICS and 1991 SOC codes:

NAICS
111 – Crop production
112 – Animal production
113 – Forestry and logging
1141 – Fishing
1151 – Support activities for crop production
1152 – Support activities for animal production
1153 – Support activities for forestry
23 - Construction
6216 – Home health care services

1991 SOC
A – Management occupations
B4 – Clerical supervisors
D111 – Head nurses and supervisors
G0 – Sales and service supervisors
G63 – Security guards and related occupations
G814 – Babysitters, nannies and parents' helpers
H01 – Contractors and supervisors, trades and related workers
H02 – Supervisors, railway and motor transportation occupations
J0 – Supervisors in manufacturing
I016 – Supervisors, landscape and horticulture

I111 – Supervisors, logging and forestry
I121 – Supervisors, mining and quarrying
I122 – Supervisors, Oil and Gas Drilling and Service

Table 1: Actual weekly hours hazard estimates (relative to October 1996 – September 1997).

	Double-difference (1)		Triple-difference (2)	
Oct97-Sep98 (standard=43)				
40	0.088	(0.074)	-0.004	(0.111)
41	0.155	(0.227)	0.254	(0.379)
42	-0.031	(0.167)	0.133	(0.277)
43	0.377*	(0.180)	-0.073	(0.282)
44	-0.950*	(0.169)	-0.806*	(0.291)
Oct98 – Sep99 (standard = 42)				
40	0.060	(0.072)	0.109	(0.109)
41	0.207	(0.210)	-0.012	(0.350)
42	0.794*	(0.154)	0.678*	(0.260)
43	-0.332	(0.189)	-0.208	(0.300)
44	-1.206*	(0.172)	-1.059*	(0.289)
Oct99 – Sep00 (standard = 41)				
40	0.019	(0.072)	-0.062	(0.110)
41	1.207*	(0.196)	0.756*	(0.336)
42	0.158	(0.158)	0.331	(0.273)
43	-0.552*	(0.190)	-0.874*	(0.297)
44	-1.161*	(0.173)	-1.096*	(0.289)
Oct00 – Sep01 (standard = 40)				
40	0.507*	(0.073)	0.192	(0.110)
41	0.179	(0.209)	0.175	(0.348)
42	-0.239	(0.167)	-0.157	(0.269)
43	-0.610*	(0.195)	-0.444	(0.301)
44	-1.109*	(0.176)	-0.948*	(0.292)
Oct01 – Sep02 (standard = 40)				
40	0.448*	(0.072)	0.350*	(0.109)
41	0.022	(0.212)	0.295	(0.348)
42	-0.233	(0.163)	0.106	(0.271)
43	-0.237	(0.185)	-0.192	(0.289)
44	-1.273*	(0.174)	-1.064*	(0.282)
No. of observations	384,880		714,329	

Notes: Hazard rates are relative to October 1996 – September 1997 period when Quebec's statutory standard workweek was 44. Hazard model is estimated as a logit function. Each specification includes controls for month, sex, age, education, establishment size, industry and occupation. The double-difference sample is restricted to full-time, hourly-paid, nonunionized workers in Quebec and Ontario and the triple-difference specification adds full-time, nonunionized, salaried workers in Quebec and Ontario. Robust standard errors are in parentheses. * indicates significance at the 5% level.

Table 2: OLS predicted effect of a reduction in statutory standard workweek from 44 to 40 on average weekly hours of full-time workers, Labour Force Survey.

	Men		Women		Both	
	Double-difference (1)	Triple-difference (2)	Double-difference (3)	Triple-difference (4)	Double-difference (5)	Triple-difference (6)
$E(h_1^{60})$						
After	39.63	39.42	36.29	36.49	37.76	37.76
Before	40.17	40.15	36.28	36.43	38.00	38.09
% change	-1.3*	-1.8*	0.03	0.2	-0.6	-0.9
	(0.5)	(0.8)	(0.4)	(0.7)	(0.3)	(0.5)
$E(h_1^{40})$						
After	37.89	37.64	35.74	35.60	36.69	36.49
Before	37.81	37.73	35.61	35.68	36.58	36.58
% change	0.2	-0.2	0.4	-0.2	0.3	-0.2
	(0.3)	(0.5)	(0.4)	(0.5)	(0.3)	(0.4)
$E(h_{41}^{44})$						
After	0.85	0.86	0.30	0.43	0.54	0.62
Before	1.18	1.18	0.38	0.41	0.74	0.75
% change	-28.3*	-26.9*	-21.9	4.6	-27.1*	-18.3*
	(10.0)	(7.8)	(16.2)	(14.3)	(8.8)	(5.9)
$E(h_{45}^{60})$						
After	0.90	0.92	0.26	0.45	0.53	0.65
Before	1.18	1.24	0.29	0.34	0.68	0.75
% change	-23.7	-26.3	-11.2	33.9	-22.0	-13.1
	(18.1)	(17.0)	(24.0)	(75.0)	(15.5)	(11.4)

Notes: Predicted effect based on the marginal standard workweek effect estimated using monthly data from October 1996 to September 2002. $E(h_j^k)$ is expected weekly hours worked between *j* and *k*. Each regression includes controls for month, sex, age, education, firm size, industry and occupation. The double-difference sample is restricted to full-time, hourly-paid, nonunionized workers in Quebec and Ontario and the triple-difference specification adds full-time, nonunionized, salaried workers in Quebec and Ontario. Robust standard errors are in parentheses. * indicates significance at the 5% level.

Table 3: Double-difference overtime, hours and employment effects of reduction in statutory standard workweek from 44 to 40 hours, 1996 and 2001 Census.

Share of industry paid by the hour:	*Δ log* (h^o)		*Δ log (h)*		*Δ log (E)*		No. of industries
	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)	
[0.6 – 1.0]	-0.184* (0.035)	-0.189* (0.039)	-0.014* (0.002)	-0.008* (0.002)	-0.004 (0.023)	0.035 (0.037)	156
[0.4 – 1.0]	-0.157* (0.028)	-0.161* (0.032)	-0.012* (0.002)	-0.009* (0.002)	-0.006 (0.019)	0.019 (0.030)	213
[0.2 – 1.0]	-0.139* (0.024)	-0.124* (0.034)	-0.010* (0.002)	-0.006* (0.003)	-0.007 (0.018)	0.020 (0.024)	248
All industries	-0.138* (0.024)	-0.122* (0.033)	-0.010* (0.002)	-0.006* (0.003)	-0.005 (0.018)	0.020 (0.024)	254
[0.0 – 0.8]	-0.110* (0.025)	-0.103* (0.036)	-0.009* (0.003)	-0.005 (0.003)	-0.012 (0.021)	0.025 (0.028)	196
[0.0 – 0.6]	-0.077* (0.029)	-0.059* (0.040)	-0.006 (0.003)	-0.003 (0.004)	-0.006 (0.027)	0.004 (0.030)	98
[0.0 – 0.4]	-0.055 (0.035)	-0.025 (0.054)	-0.003 (0.005)	0.002 (0.005)	0.00009 (0.044)	0.022 (0.038)	41

Notes: Estimates are obtained by regressing 4-digit industry-level log employment on the interaction of a Quebec and a 2001 year dummy. Each industry is observed at the provincial level in May 1996 and May 2001. Regressions include province-specific industry fixed effects. Results for hours worked beyond 40 in a week (h^0) and total weekly hours (*h*) are based on the sample of workers with actual weekly hours of 30 or more. Observations are weighted by the (unweighted) employment level. Due to mean overtime hours and employment levels of zero within some industries, the female overtime results are based on 242 industries and the hours and employment results on 253 industries. Standard errors reported in parentheses are clustered by industry. * indicates significance at the 5% level.

Table 4: OLS predicted effect of a reduction in the standard workweek from 44 to 40 on log hourly wage rate of full-time workers, Labour Force Survey.

	Men		Women	
	Double-difference (1)	Triple-difference (2)	Double-difference (3)	Triple-difference (4)
$E(w)$	-0.019*	-0.0001	-0.0002	-0.0004
	(0.008)	(0.015)	(0.006)	(0.012)
$E(w \mid 1 \leq h \leq 39)$	-0.048*	0.003	0.007	0.005
	(0.016)	(0.027)	(0.008)	(0.015)
$E(w \mid 40 \leq h \leq 44)$	0.003	0.007	-0.007	-0.018
	(0.011)	(0.023)	(0.010)	(0.022)
$E(w \mid h \geq 45)$	-0.033	-0.024	-0.070*	-0.040
	(0.020)	(0.034)	(0.030)	(0.045)

Notes: Predicted effect based on the marginal standard workweek effect estimated using monthly data from October 1996 to September 2002. $E(w \mid j \leq h \geq k)$ is expected weekly hours worked between j and k. Each regression includes controls for month, sex, age, education, firm size, industry and occupation. The double-difference sample is restricted to full-time, hourly-paid, nonunion workers in Quebec and Ontario and the triple-difference specification adds full-time, nonunion, salaried workers in Quebec and Ontario. Robust standard errors are in parentheses. * indicates significance at the 5% level.