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Is Workers' Compensation Disguised Unemployment Insurance?*

Bernard Fortin[†], *Paul Lanoie*[‡], *Christine Laporte*[†]

Résumé / Abstract

Ce texte examine comment l'assurance contre les lésions professionnelles (ALP) et l'assurance-chômage (AC) interagissent pour influencer la durée des réclamations pour lésions professionnelles. Nous utilisons des micro-données administratives longitudinales couvrant plus de 30 000 travailleurs du secteur de la construction pour la période 1976-1986. Nos résultats montrent qu'une réduction du taux de remplacement salarial de l'assurance-chômage est associée à une augmentation de la durée des accidents majeurs difficiles à diagnostiquer (cette catégorie inclut les maux de dos majeurs). De plus, un accroissement du taux de remplacement salarial de l'ALP entraîne un accroissement de la durée des accidents mineurs difficiles à diagnostiquer (cette catégorie inclut les lombalgies). Enfin, il semble y avoir un effet saisonnier lié à la durée des périodes d'indemnisation est plus élevée de 21.2% lorsqu'un accident survient en décembre plutôt qu'en juillet.

This paper examines how yhe Workers' Compensation (WC) and the Unemployment Insurance (UI) programs interact to influence the duration of claims due to workplace accidents. We use longitudinal WC administrative micro-data on more than 30,000 workers in the Quebec construction industry for the period 1976-1986. Our results show that a reduction in the UI wage replacement ratio is associated with an increase in the duration of claims due to serve accidents that are difficult to diagnose (this category includes severe back-related problems). Also, an increase in the WC replacement ratio leads to an increase in the average duration of claims due to minor accidents that are difficult to diagnose (this category includes minor low-back injuries). Moreover, there seems to be an important seasonal effect in the duration of claims; i.e., the average duration of spells on WC is estimated to be 21.2% higher when an accident occurs in December rather than in July.

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1. Introduction

The social cost of workplace accidents is large. In a typical year in the United States, from one-third to one-half as many working days are lost to work injuries as are lost to unemployment (see Krueger 1988). Recently, economists have paid more attention to the economic issues involved in occupational safety and health (see Lanoie, 1994, for a survey). In particular, they have investigated the different effects of government intervention in this area. For example, it has been found that, in general, safety regulations and their enforcement had little impact on the incidence of workplace accidents. There also seems to be a consensus on the fact that an increase in the generosity of workers'compensation (WC) benefits is associated with an increase in both the frequency and the duration of claims due to workplace accidents. In this connection, some researchers have raised a potentially important issue: the possibility of substitution between WC and unemployment insurance (UI). In most countries, WC is more generous than UI in terms of wage replacement[1], and there may therefore be an incentive for workers about to be laid off to try to benefit from WC instead of UI. Workers suffering from a workplace accident may also be tempted to take action in order to obtain a longer period of recovery compensated by WC. These ex ante and ex post moral hazard problems are likely to be particularly important when injuries are hard-to-diagnose (*e.g.*, low-back injuries) and in industries where the level of unemployment is such that many workers may expect to be unemployed and to receive UI benefits after their recovery (*e.g.*, in seasonal industries such as the Canadian construction industry, which is less active during winter because of weather constraints).

Fortin and Lanoie (1992) were the first authors to investigate these issues and to provide empirical evidence of substitution between the two insurance regimes. In particular, from Quebec aggregate data at the industry level, they found that a reduction in the generosity of UI was associated with an increase in the average duration of work absences compensated by WC following a workplace accident. Before policy makers be convinced to take action in order to alleviate this problem, more investigation is required to provide further support for this initial result. In this paper, we estimate the effect of WC and UI benefits on the expected duration of claims due to workplace accidents using a unique panel dataset that allows us to investigate more fully the issues raised above, and to control for some individual characteristics. The database is composed of

[1]For instance, in most American states, accident victims receive on average 67% of their gross wage (non-taxable), while unemployed workers receive on average 50% of their gross wage (taxable). In Canada, the Quebec Workers' Compensation Board (WCB) provides accident victims with 90% of their net average wage (non-taxable), while the Canadian UI regime provides unemployed workers with 57% of their gross wage (taxable). The reader is referred to Fortin and Lanoie (1992) for more institutional details about these programs in Canada.

longitudinal WC administrative micro-data on more than 30,000 workers in the Quebec construction industry for the period 1976-1986.

Our theoretical discussion focuses on the hazard of returning to the labor market after an absence due to a workplace accident[2]. One prediction of the reduced form model is that the UI replacement ratio has a positive impact on the hazard of leaving WC. Another is that this hazard is likely to be lower during months preceding the end of the construction season.

The empirical work tests our theoretical predictions. For this matter, we use a mixed proportional hazard model, devised by Meyer (1990), that does not impose a parametric form on the baseline hazard and that takes unobserved heterogeneity into account using a gamma distribution. Most previous work in this area has relied on parametric methods which assume, despite a lack of theoretical support, a specific form for the baseline hazard (*e.g.*, Weibull). This approach provides inconsistent estimates when the assumed baseline is incorrect. Examples of such work include Johnson and Ondrich (1990) who present three hazard models (a Weibull hazard without random effects, a Weibull hazard with a gamma-distributed random effect, and a Weibull hazard with a non parametric random effect) to analyze the duration of work absences in three American states among WC clients with diverse permanent partial disabilities. Furthermore, Butler and Worrall (1991) used the family of generalized gamma distributions, with parametric or non-parametric controls for heterogeneity, to investigate the duration of claims for temporary and permanent partial disability due to low-back injuries in twelve states (see also Butler and Worrall, 1985). These studies focused on the relation between WC benefits and duration, and in general found the benefit elasticities of duration to be positive but small.

The rest of the paper is organized as follows. Section II presents our theoretical discussion and its implications. Section III discusses the estimation strategy and the data. Section IV analyzes the results. In brief, these results support the hypothesis that a reduction in the UI replacement ratio and an increase in that of the WC program are both associated with an increase in the duration of claims, especially in the case of workers with injuries that are difficult to diagnose. Furthermore, there seems to be an important seasonal effect in the duration of claims; *i.e.*, those occurring in December (at the end of the construction season for a majority of workers) are likely to last longer, *ceteris paribus*. This result is also consistent with the presence of a significant interaction between WC and UI in determining individuals' risk of accident. Section V provides concluding remarks and a discussion of the policy implications of our results.

[2]Fortin and Lanoie (1992) provide evidence that UI affects the duration but not the probability of claims. The present paper therefore focuses only on duration analysis.

2. Theoretical discussion

The reduced form model to be estimated explains the instantaneous rate of exit from WC at time t, given that absence from work has lasted until t, *i.e.* the hazard, λ_t, of the post-injury absence, as a function of a number of covariates[3]:

$$\lambda_t = \lambda(WCR_t, UIR_t, UNEMP_t, WAGE_t, D, MONTH, t, X_t) \tag{2.1}$$

The following theoretical predictions are intuitive. A higher WC wage replacement ratio, WCR_t, increases a worker's utility on WC. This effect reduces the hazard of leaving WC (or increases the expected duration of the recovery period). Note that, in Quebec, in contrast to other jurisdictions, an injured worker is free to choose any physician to establish the validity of the accident. This tends to accentuate this moral hazard problem. On the other hand, when taking into account the employers' behavior, one can also argue that an increase in WC benefits raises the opportunity cost of an accident for employers, who pay WC insurance premia. This may induce them to devote resources to reduce the risk of accidents. This effect may be especially acute when the degree of experience rating is high[4]. Therefore, the net effect of a change in the WC wage replacement ratio is ambiguous and needs to be resolved empirically.

An increase in the UI wage replacement ratio, UIR_t, (which is assumed to be smaller than the corresponding WC ratio) raises a worker's expected utility in the labor market, given that he faces some risk of unemployment upon his return to work, thus increasing his incentive to leave WC[5]. In the same vein, an increase in the level of unemployment ($UNEMP_t$) reduces a worker's probability of finding a job, thus reducing the expected utility in the labor market and the probability of leaving WC. Moreover, the impact of the pre-injury $WAGE_t$ on the probability of leaving WC is ambiguous, since it produces a substitution and an income effect of opposite signs on the demand for non-market time. Moreover, for given levels of the wage replacement ratios, it is associated with an identical percentage increase in WC and UI benefits, which raises both the worker's utility on WC and on the labor market.

Furthermore, Dionne and St-Michel (1991) have convincingly argued that the level of difficulty in making the diagnosis (D) is likely to amplify the effect of variables such

[3]This section summarizes a structural model developed in Fortin et al. (1994).

[4]In Quebec, as in the rest of North America, firms are considered liable for workplace accidents and pay WC insurance premia. Experience rating refers to the adjustment of WC insurance premia to reflect claim experience.

[5]It is assumed that the individual has accumulated right to claim UI benefits. In our empirical work, we also tried variables reflecting the regional minimum work requirement for eligibility to UI and the regional maximum duration of UI benefits, as additional measures of UI generosity. However they were not statistically significant in any specification and have thus been excluded from the model. Note also that there is no experience rating under the Canadian UI scheme.

as the wage rate and the generosity of WC and UI benefits on the rate of exit from WC. In terms of our model, these hypotheses imply the following inequalities:

$$|\lambda_{WAGE_t,D}|,\ |\lambda_{UIR_t,D}|,\ |\lambda_{WCR_t,D}| \geq 0.$$

One other implication of our model is that λ_t is likely to be lower, *ceteris paribus*, in the months preceding winter, the lay-off season in the construction sector. The reason is that a worker facing a workplace accident in December, for example, has a higher probability of being unemployed upon his return to the labor market after his recovery, than if the same accident occurs, say, in July. This is the case since the duration of claims is usually relatively short (31 days on average, in our sample). Therefore, he has a greater incentive to try to prolong his period on WC, given that the WC wage replacement ratio is greater than the UI ratio.

Lastly, the effect of the elapsed spell duration (t) on the rate of exit from WC is expected to be positive for two reasons: First, as time goes by, the worker is likely to have recovered from his injury, which increases his utility at work relative to his utility on WC. Second, the cost of obtaining extra days on WC is likely to be increasing, since a worker is less likely to find an accommodating doctor as this number increases. Note finally that X_t denotes a vector of control variables.

3. Empirical methods and data

To estimate the reduced form model given by equation (2.1), we adopt the following mixed proportional hazard specification with unrestricted baseline hazard:

$$\lambda(z,t) = \lambda_0(t)\exp(z'\beta + \epsilon), \tag{3.1}$$

where $\lambda_0(t)$ is the baseline hazard (*i.e.*, the part common to all individuals), z is a vector of covariates, β is a vector of parameters to be estimated and ϵ is a random variable reflecting unobserved heterogeneity.

As argued by Meyer (1990), this semi-parametric procedure has the advantage of avoiding inconsistent estimation of parameters due to misspecified baseline hazard. Moreover, contrary to Cox's (1972) approach, it simultaneously provides a non-parametric estimate of the baseline hazard and easily accounts for the presence of unobserved heterogeneity. As is well known, ignoring unobserved heterogeneity may lead to a dynamic selection bias in the parameter estimates and in the estimate of the baseline hazard. For example, as time goes by, it is possible that workers who do not return to the labor market are those with an intrinsic bad health condition. If one does not account for this unobserved heterogeneity, one may end up having the false impression that the hazard declines over time (negative bias on the estimate of the baseline hazard).

A convenient and commonly used distribution for $exp(\epsilon)$ is the gamma distribution[6] with the mean normalized to one and variance σ^2. Under this assumption and given observations of failure times over the discrete periods $t = t_0, t_1, t_2, \ldots, t_{T-1}$ for individuals $i = 1, \ldots, N$, the resulting log-likelihood is given by:

$$\begin{aligned} L(\gamma, \beta, \sigma^2) &= \sum_{i=1}^{N} \log \left\{ \left[1 + \sigma^2 \cdot \sum_{j=0}^{k_i - 1} \exp\{\gamma_{t_j} + z_i'\beta\} \right]^{-\sigma^{-2}} \right. \\ & \left. -\delta_i \left[1 + \sigma^2 \cdot \sum_{j=0}^{k_i} \exp\{\gamma_{t_j} + z_i'\beta\} \right]^{-\sigma^{-2}} \right\}. \qquad (3.2) \end{aligned}$$

where $\gamma_{t_j} = ln(\int_{t_j}^{t_{j+1}} \lambda_0(v)dv)$, $\gamma = [\gamma_{t_0}, \gamma_{t_1}, \ldots, \gamma_{t_{T-1}}]$, z_i is a vector of covariates for individual i, $\delta_i = 1$ if the duration for individual i is not (right) censored, $= 0$ otherwise, t_{k_i} is the observed (censored or not) duration of individual i's spell on WC. The expression $exp(\gamma_{t_j})/(t_{j+1} - t_j)$ represents the average baseline hazard over the interval $[t_j, t_{j+1}]$.

3.1. Data and variables

The original data source for this study follows the evolution of 30,341 workers in the construction industry who worked at least one hour at the James Bay hydro-electric project (a major dam construction plan in northern Quebec) during the period 1976-86. We are able to track the work pattern (number of hours worked, occurrence of an accident, etc.) of the workers throughout this period as long as they were working in the construction industry. During the period, we count 8,523 workplace accidents with time lost in the sample involving 6,067 workers[7]. Descriptive statistics for the sample are given in Table 1 at the end of the text. From this table, one can see that the average duration of claims is 31.2 days with a relatively large standard deviation of 66.1.

We now turn to the explanatory variables (the vector z) used in the estimations. Note that since the duration of spells on WC is usually short, we did not allow for time-dependent covariates. The WC replacement ratio (benefits divided by the pre-WC net

[6]Other distributions could be used; in particular, the heterogeneity distribution could be non-parametric, following the approach suggested by Heckman and Singer (1984). However, as conjectured by Meyer (1990, p.771), if the baseline hazard takes a flexible form, the choice of the heterogeneity distribution may be unimportant.

[7]An accident with time lost is an accident involving more than one day off work. The permanent disability cases are excluded from the sample because their duration is calculated in an arbitrary fashion.

marginal wage) has been calculated individually using information on the WC parameters and on the provincial and federal income tax systems in place in each year (see the details in Fortin et al., 1994, Appendix 1). In the literature, only Moore and Viscusi (1990) have a replacement ratio calculated individually. The mean WC replacement ratio in our sample (see Table 1) is 114 percent. One reason why it is greater than 100 percent is that, under the Quebec WC regulations, benefits calculation is based on the earnings of the twelve months preceding the accident. In this calculation, the worker is imputed the same average weekly income for his period off construction as the one he earns during the construction season, which tends to increase the numerator of the ratio. Second, since WC benefits are not taxable and are based either on gross wages (before 1979) or on net *average* wages (after 1979), the (marginal) replacement ratio may be higher than 100 percent for workers with a marginal tax rate higher than a critical level. The expected effect of the WC replacement ratio on the hazard of leaving WC is ambiguous and depends on the level of experience rating that has been introduced in Quebec on a firm-by-firm basis in 1979 (CSST, 1982). For this reason, in most specifications, we introduce two variables to capture the generosity of WC: Log WCRB79 (log WCR times a dummy equal to one for each observation before 1979) and log WCRA79 (log WCR times a dummy equal to one for each observation in the years 1979-1986 inclusively). Given the introduction of experience rating in 1979, the latter is likely to have a less negative effect (or even a positive effect) on the hazard of leaving WC.

The UI replacement ratio is also calculated individually and is 0.57, at the mean of our sample. The variable UNEMP is defined as the regional unemployment rate as determined by Statistics Canada. The WAGE is measured by the worker's hourly wage rate (in 1981 dollars). There is an important literature showing that the wage and the occupational safety level are the result of a simultaneous decision process (e.g., Garen, 1988), making the wage rate (and the replacement ratios) an endogenous variable[8]. Moreover, individual unobservable characteristics that affect the wage rate (*e.g.*, motivation to work) may also influence the rapidity of return to work after an accident. For these reasons, in most specifications, the wage variable and the two replacement ratios (all in log) have been replaced by generated regressors in the likelihood function, following a two-step approach suggested by Durbin (1970) and Pagan (1984)[9].

The initial severity of the injury and the difficulty of the diagnosis are also part of the analysis. Of course, more severe injuries should be followed by longer recovery periods. Moreover, it may be easier for workers to extend their recovery period when they have a hard-to-diagnose injury (*e.g.*, low-back injury), and one should also account

[8] For instance, according to the theory of compensating wage differentials, more dangerous jobs (with longer duration of claims) lead to a higher wage rate, see Thaler and Rosen (1976).

[9] Note however that the standard errors of the corresponding estimated coefficients are generally inconsistent. While bootstrapping methods could theoretically be used to generate consistent standard errors estimates, such an approach would be highly time-consuming. One way to circumvent this problem is to interpret our analysis as conditional upon the value of the generated regressors.

for the fact that hard-to-diagnose injuries may be intrinsically more or less severe than others. For these reasons, and following Dionne and St-Michel (1991), we consider four categories of accident: 1) minor injuries with easy diagnosis, MINEASY; 2) minor injuries that are hard to diagnose, MINHARD; 3) major injuries with easy diagnosis, MAJEASY; and 4) major injuries that are hard to diagnose, MAJHARD. These are entered as dummy variables (MINEASY is default). As discussed in Dionne and St-Michel, this categorization was established in consultation with a physician specialized in work-related health problems[10].

As discussed in the theoretical section, the difficulty of establishing the diagnosis may also interact with some other explanatory variables in influencing the individual's decision over the possible extension of his recovery period. For instance, it may be more tempting for a worker whose injury is hard to diagnose to seek such an extension when UI becomes less generous and when he knows that he will be laid off if he goes back to work. Interaction terms between the dummies MINHARD and MAJHARD (which, in particular, refer to various categories of back disorders) and other covariates such as the wage rate and the replacement ratios have therefore been introduced in some specifications. Initial investigation showed that four interaction terms persistently had a significant coefficient:(log UIR)*MAJHARD, (log WCRB79)*MINHARD, (log WCRA79)*MINHARD and (log WAGE)*MINHARD.

In keeping with our theoretical discussion, we also consider monthly dummy variables to capture the fact that workers may have more incentive to extend their period of recovery if their accident happened just before the usual lay-off season in the construction industry. Therefore, we expect the coefficients of dummy variables for months like NOVEMBER and DECEMBER to be positive and significant (JULY is default).

Furthermore, three personal characteristics (X_t) are taken into account. First, the AGE of the accident victim is introduced, since it is generally accepted (*e.g.*, see Butler and Worrall, 1991) that, *ceteris paribus*, the capacity to recover physically from the effects of injuries declines with age. Second, a variable capturing the level of QUALIFICATION of the worker is included. It is defined as the number of working years required to qualify as a registered member of a given occupation within the construction industry (*e.g.*, carpenter). This variable is intended to capture the worker's skill level. As in Johnson and Ondrich (1990), it is expected that more skilled workers are better able to avoid severe accidents, or that they have more flexibility to re-integrate rapidly into the labor force than less skilled workers.

Third, a proxy for the number of dependent children (DEPENDENTS) is included. Our data set does not provide direct information on the latter variable. However, given

[10] The main types of occupational injuries that belong to each group are the following: (1) MINEASY: contusion, poisoning, amputation without permanent partial disability and friction burn; (2) MINHARD: minor low-back injuries, bursitis; (3) MAJEASY: fracture and (4) MAJHARD: spinal disorders, severe low-back problems.

the age of each worker and general demographic information on the workers in the Quebec construction industry (from the 1981 Canadian Census), we were able to calculate a variable approximating the number of dependent children. The expected sign of the coefficient on this variable is positive. Indeed, if the worker provides an income support to a large family, he may be induced to leave WC rapidly.

In addition, we consider REGIONAL dummies to capture the fact that the nature of the construction projects may vary from one region to another (especially at James Bay), leading to different types of accidents. After some experimentation, dummies for the regions of Quebec, James Bay and Abitibi have been used in estimations (other regions are default). Finally, we introduce year dummies to capture omitted fixed influences that vary across time, but not across individuals. These dummies may be useful to account for institutional changes in unemployment insurance and occupational safety policies (like changes in safety-enhancing measures) during the period that may not be captured with our WC and UI variables.

4. Empirical results

Our estimation results are presented in Table 2. The preceding discussion leads to a general specification using the two-step estimating method involving generated regressors for the wage and the replacement ratios [11]. This specification is presented in column (1). In order to assess the validity of the general specification and to test certain specific restrictions, six other specifications are presented. In particular, we examine how our main results are affected when the two-step estimating method is not used (specification (3)), when we do not account for unobserved heterogeneity (specification (4)), and when a standard parametric form (Weibull) is assumed for the baseline hazard (specifications (5) and (6): with and without gamma-distributed heterogeneity effects). Specification (7) presents a slightly modified version of specification (1) where the coefficients of UI and WC replacement ratios (non significant in specification (1)) are constrained to zero[12].

[11] The variables used in the first-step regression to generate predicted wage rates (and, using a tax-transfer computer program, the replacement ratios) are all the exogenous covariates used in the second step and age^2, age^3, qualification^2, qualification^3, plus interaction terms between age and qualification (up to the second degree). These additional variables (which are all statistically significant) and the non-linearities in the tax-transfer system help identifying the model.

[12] Since there are more than 315 different durations in our sample, we were unable to estimate a γ element for each of these durations. After some investigation, 36 elements of γ, corresponding to different time intervals, have been used in all specifications considered. The 36 elements of γ correspond to the following durations (in days): $\gamma_i, i = 0, \ldots, 17 : i + 2\,\text{days}$; $\gamma_{18} : 20 - 21$; $\gamma_{19} : 22 - 23$; $\gamma_{20} : 24 - 25$; $\gamma_{21} : 26 - 27$; $\gamma_{22} : 28 - 29$; $\gamma_{23} : 30 - 31$; $\gamma_{24} : 32 - 33$; $\gamma_{25} : 34 - 35$; $\gamma_{26} : 36 - 37$; $\gamma_{27} : 38 - 39$; $\gamma_{28} : 40 - 44$; $\gamma_{29} : 45 - 49$; $\gamma_{30} : 50 - 59$; $\gamma_{31} : 60 - 69$; $\gamma_{32} : 70 - 79$; $\gamma_{33} : 80 - 99$; $\gamma_{34} : 100 - 149$; $\gamma_{35} : 150 - 199$, $\gamma_{36} : 200 - 2716$. This last interval is censored. Over 10 intervals, the β coefficients appear to be highly robust to the number of elements of γ considered.

In general, the sign, the magnitude and the precision of the covariates coefficients are robust across specifications using the two-step estimation method, although slight differences appear in the Weibull specifications. In discussing the results, we will first focus on specification (1) and (7) which are representative of the set of results.

The UI replacement ratio (when combined with MAJHARD) has a positive and significant coefficient, indicating that a reduction in this variable induces workers with a major hard-to-diagnose injury to stay longer on WC. Using specification (7), our results imply that a reduction of 1% in the UI replacement ratio is associated with a decrease of 2.7% in the conditional hazard of leaving WC for these workers. Over all workers and at the sample mean of the covariates, the resulting elasticity of the expected duration of claims with respect to UI[13] is estimated at -0.54. The magnitude of this result is close to the elasticity of -0.5 reported in Fortin and Lanoie (1992).

Second, the WC replacement ratios before and after 1979, when combined with MINHARD, have negative and significant coefficients, which means that WC induces workers suffering from a minor hard-to-diagnose injury to go back onto the labor market later. Over all workers and based on results from specification (7), an increase of 1% in the generosity of WC is estimated to induce a 0.71% increase in the expected spell duration before 1979, and 1.09% after 1979, at the sample mean of the covariates. The stronger impact after 1979 than before is puzzling, given that an experience rating mechanism has been introduced in 1979. One possible explanation is that, at the same time, the Quebec WCB opened new regional offices to handle local compensation claims, an operation which, until that date, was done centrally in Montreal and Quebec City. According to the WCB officials (see CSST, 1984), this has made accident reporting easier, thus enhancing any potential moral hazard effect that our WC variables are capturing. Note that in specification (2), where the WC replacement ratio is not divided into two components (before and after 1979), its effect on the hazard rate (when combined with MINHARD) is still negative, as in the rest of the literature, but less significant. Moreover, a likelihood ratio test (with a χ^2 statistic of 9.38) rejects the equality of the WCR coefficients before and after 1979. This suggests that it is warranted to account for the important policy changes occurring in 1979-1980.

The estimated impact of the WC ratio on duration for each sub-period is relatively strong compared to the rest of the literature. Typical WC duration elasticities vary between 0.2 and 0.6 (*e.g.*, Meyer et al. 1995), although one study (Krueger 1990) finds duration elasticities over 1.5. Three reasons could be advanced to explain such a result. First, the construction industry presents certain characteristics (regular lay-offs) that make substitution between the two insurance regimes much more likely than in other sectors of economic activity. Moreover, it is likely that the intertemporal labor supply elasticity is higher in the case of individuals who choose to work in a seasonal industry.

[13]The expected spell duration for a given level of the covariates is obtained by evaluating the integral of the corresponding survival function, over all time intervals (see Katz and Meyer 1990).

Second, as mentioned earlier, in Quebec, in contrast with other jurisdictions, the role of the worker's doctor is crucial in determining the duration on WC. Therefore, moral hazard is likely to be more important than elsewhere. Third, the American studies do not account for the possible interaction between the two insurance regimes, which may bias their results[14].

Furthermore, it is interesting to discuss why the two insurance regimes have a significant effect when combined with different types of hard-to-diagnose injuries (WC is significant when combined with MINHARD, while UI is significant when combined with MAJHARD). At least two reasons come to mind to explain this phenomenon. First, since major injuries that are hard-to-diagnose are intrinsically more severe and associated with longer durations than minor injuries of the same type, recovery is statistically more likely to occur during the industry's dead season in which unemployment is higher, giving a stronger incentive to workers with MAJHARD injuries to stay on WC and avoid UI. Second, victims of major injuries that are hard-to-diagnose are more likely to be less employable and to face a period of unemployment at the end of their recovery period, giving them again a stronger incentive to stay on WC and avoid UI.

In addition, as predicted by the theoretical model, the fact that an accident occurs in DECEMBER (rather than in July) induces a reduction in the conditional hazard of leaving WC of 28%, which corresponds to an increase of 21.2% in the expected duration, at the mean of the other covariates. Such a finding, as well as those described in the preceding paragraphs, indicates that financial incentives play a role in the determination of durations, and is consistent with the existence of substitution between the two insurance regimes[15].

It is also noteworthy that the coefficient of the regional unemployment rate (UNEMP) is always significant, but never has the expected negative sign. One explanation for this phenomenon may be that, when unemployment increases, workers who quit the labor force are those with less seniority and experience, leaving in those who are less likely to have accidents involving a long recovery. This explanation is confirmed to some extent by the fact that, in all specifications, our variable QUALIFICATION has a positive and significant coefficient (except in specification (2)). Conversations with officials of the Association des entrepreneurs en construction du Québec (AECQ) have provided us with further insight on this issue. They added that, during cyclical downswings, employers in the construction industry do not lay off their "best" workers (like the foremen), with whom they have built a stronger relationship, and who are more likely

[14] Fortin and Lanoie (1992) actually show that the magnitude of the WC impact is reduced when one does not account for the UI variables. This result is not surprising since, given that the UI and WC replacement ratios are positively correlated, omitting the UI ratio will produce a downward bias in the estimates of the effect of the WC ratio.

[15] None of the other monthly dummies is significant. NOVEMBER is negative and significant in the specification (4) without unobserved heterogeneity, while SEPTEMBER, OCTOBER and NOVEMBER are negative and significant in the Weibull specification (6) without unobserved heterogeneity.

to go back to work quickly after an accident. In line with these arguments, one can add that, when unemployment is high, workers' status is more likely to be precarious, inducing them to re-integrate the labor force more rapidly after an accident.

Concerning the other variables, as in Meyer et al. (1995), the coefficient of log WAGE is negative and significant (except in specification (3)). While a negative sign may partly be explained by the implied increase in the level of WC benefits (since the WC ratio is assumed constant), this may also suggest the existence of an income effect such that high-income individuals can afford to "buy more leisure" and stay longer on WC. This argument may be particularly appealing given that our analysis focuses on an industry where the average wage is relatively high[16], while other studies with converse results (positive sign) are based on more heterogeneous samples (e.g., Butler and Worrall, 1991). Note however that the coefficient of the interaction term (log WAGE)*MINHARD is always positive (except in specifications (2) and (5)) and is significant in specifications (1), (6) and (7), indicating that the negative wage effect is less important for workers experiencing minor hard-to-diagnose injuries.

Moreover, the coefficient of the variable AGE is negative and significant, confirming that the capacity to recover from the effects of injuries declines with age. Furthermore, the behavior of the variable DEPENDENTS is somewhat erratic, its coefficient being positive in certain specifications and negative in others, but rarely significant. This may suggest that the variable is subject to measurement errors (recall that the number of dependents was constructed using general demographic data).

Finally, as regards REGIONAL dummies, it is noteworthy that workers who experienced an accident at JAMES BAY are predicted to have longer duration of claims. Again, officials of the AECQ provided us with a convincing explanation for this phenomenon. Because the project was located in a remote area in Northern Quebec far away from the main urban centers of the province, the construction firms working in the project had adopted an informal policy of reporting only the more severe accidents to the WC board. In doing so, the employers avoided returning workers with minor injuries to their families in southern Quebec, as such trips involve important expenses.

Let us now turn to the alternative specifications ((2) to (6)) and the different tests that can be made to assess the validity of our general specification. First, when the two-step estimating method is not used (specification (3)), the coefficients of the variables associated with the wage and the replacement ratios are relatively different. In particular, the coefficients of the WC replacement ratios become positive and significant, in contrast with all the other specifications and with the rest of the literature. Also, the coefficient of the interaction term between the WC replacement ratio before 1979 and MINHARD becomes positive and significant, again in contrast with all the other specifications and with the rest of the literature, while the coefficient of the variable (log

[16]For instance, during the period 1976-1986, the average weekly wage in the Canadian construction industry was \$ 632.61 (Cdn \$ 1986) versus \$ 458.43 in the rest of the economy.

UI)*MAJHARD is still positive, but much smaller in magnitude[17]. These results suggest that the potential bias related to the endogeneity of the wage rate can be important.

Second, regarding the heterogeneity issue, in all the specifications estimated with the mixed proportional hazard model (specifications (1), (2), (3), and (7)), the estimate of the heterogeneity variance is strongly significant. Specifications (1) and (4) allow a direct comparison of estimates with and without unobserved heterogeneity. All the coefficients in the heterogeneity specification tend to be larger in absolute values which is consistent with the theoretical prediction of Lancaster (1979, pp. 951-952). Furthermore, from Table 3, it is clear that, in specification (1) which accounts for unobserved heterogeneity, the average conditional hazard per day (at the sample mean of the covariates) is almost everywhere increasing with the duration on WC. However, it is noteworthy that the hazard rate is much lower (except in the first intervals) and tends in fact to decrease with duration, when one does not account for unobserved heterogeneity (see specification (4)). This provides particularly strong evidence of the negative heterogeneity bias in the estimate of the effect of duration on the hazard.

Third, along these lines, it is interesting to examine how the results are affected when one assumes a specific form for the baseline hazard often encountered in the literature (Weibull with and without gamma-distributed random effects)[18]. Specifications (5) and (6) impose a Weibull baseline, while specifications (1) and (4) estimate the baseline non-parametrically, allowing a comparison of the techniques with and without account taken for unobserved heterogeneity. In both cases, the likelihood ratio test strongly rejects the null hypothesis of a Weibull baseline, indicating that the Weibull model is misspecified. The χ^2 statistics with 34 degrees of freedom are 1644.04 and 4860.2 for specifications (5) and (6) respectively. Also, it is noteworthy that α is estimated at 2.746 in the Weibull specification (5) in which we allow for unobserved heterogeneity. This means that the hazard is monotonically increasing since α is greater than 1. However, it is estimated at 0.808 when we do not allow for heterogeneity, which implies that the hazard is monotonically decreasing. This provides further evidence of the presence of a strong downward bias in the effect of duration on the baseline hazard when we do not account for unobserved heterogeneity.

[17] Furthermore, the coefficient of QUALIFICATION is negative only in specification (3).

[18] Note that the conventional estimation of the Weibull was modified to allow for a comparison with the Meyer's model, based on the same information set. More precisely, in contrast with usual calculations, in setting the likelihood function, we only take into account the information on the time intervals within which the actual durations in this sample are located. This allows the Weibull specification to be nested in the semi-parametric model. Thus, in the likelihood function (3.2), our Weibull formulation imposes the γ_{t_j}'s to be equal to $\ln(\int_{t_j}^{t_{j+1}}(\eta\alpha)v^{\alpha-1}\,dv) = \ln\eta + \ln(t_{j+1}^{\alpha} - t_j^{\alpha})$, where η and α are the Weibull parameters to be estimated.

5. Concluding remarks and discussion

In this paper, we have estimated the effect of Workers Compensation (WC) and Unemployment Insurance (UI) benefits on the expected duration of claims due to workplace accidents using longitudinal WC administrative micro-data on more than 30,000 workers in the Quebec construction industry for the period 1976-1986. Our results show that increases in the generosity of WC in Quebec led to an increase in the average duration of claims due to minor accidents that are difficult to diagnose. We also provided evidence that a reduction in the generosity of UI was associated with an increase in the duration of claims due to severe accidents that are hard-to-diagnose. Furthermore, there seemed to be a seasonal effect in the duration of claims: those occurring in December (at the end of the construction season for most workers) are likely to last longer. This is another important piece of evidence that there is substitution between WC and UI in the construction industry.

It is clear that this phenomenon produces both efficiency and equity effects on the labor market. Thus, as long as the WC replacement ratio is higher than that of UI, this may induce some injured workers to invest ressources in order to prolong their period on WC, especially when they expect to be unemployed and receive UI benefits upon their return to the labor market. This also can have the effect of increasing the duration of (disguised) unemployment compensated by WC. Furthermore, this may involve higher payroll tax rates in order to fund the WC program, which may create additional distortions on the labor market. On the equity side, the presence of a substitution between WC and UI implies a redistribution of income towards the unemployed who are compensated by WC instead of UI.

How can policy makers attenuate these effects? At first glance, one would consider equalizing the replacement ratios of the two regimes. However, there are at least two reasons why this should not be done. First, on efficiency grounds, the moral hazard problem is likely to be more severe for unemployment insurance than for workers' compensation; workers are more likely to abuse UI than WC, *ceteris paribus*. Indeed, abusing WC probably imposes a higher cost on the individual (*i.e.*, the cost of finding an accommodating physician, etc.) than abusing UI. Basic insurance theory, which states that there exists a trade-off between the extent of insurance coverage and moral hazard, would therefore recommend lower insurance coverage for UI than for WC. Second, on equity grounds, most of the time unemployed workers have the opportunity to get out of the "bad state of the world" by finding another job, whereas this opportunity is more limited for injured workers.

It is difficult, however, to find the optimal difference between the generosity of the two regimes. There exists a literature on the optimal rate of wage replacement under UI (e.g., Bailey, 1978 or Wright and Loberg, 1987) and, to our knowledge, one paper on the optimal rate under WC (Viscusi and Evans, 1991), but none takes into account the interactions between these two insurance schemes. Presumably, one would have to extend the analytical framework present in these studies to account, not only for the trade-off between insurance coverage and moral hazard within one regime, but also for cross-substitution effects between regimes. The literature on insurance of multiple risks (for instance, Arnott, 1991) generally suggests that, in such a case, there should only be one insurer to take into account, for example, the effect of a risk (*e.g.* unemployment) insured under a particular insurance regime (*e.g.* UI) on the behavior of individuals vis-a-vis another insured risk (*e.g.* workplace accidents). This means that the harmonization of both regimes could be useful in order to internalize the cross effects of each regime on the other. In Canada, such a solution would cause problems since, in contrast with the U.S., WC and UI are under different jurisdictions (provincial in the case of WC and federal in the case of UI).

An international comparison of the generosity of the two regimes may serve as a starting point to determine the appropriate gap between the generosity of the two regimes. Lanoie (1994) provides a comparative analysis of the WC and UI systems in 14 countries of the OECD. It turns out that Canada is one of the countries (with Australia) where the gap between the generosity of the two regimes is the largest. This is especially true when one accounts for the waiting periods in the two regimes, and for the precise calculation of WC benefits which allows, under certain circumstances discussed above, for a net replacement rate over 100% (due to the fact that WC benefits are not taxable). Thus, the incentive for substitution should be larger in Canada.

The preceding discussion suggests a number of ways to alleviate this substitution problem. First, holding total expenditures on both regimes constant, there would probably be less distortions and a potential welfare improvement if the replacement ratio under WC was lowered and that under UI increased. As a variation on this theme, one could lower the generosity of WC (*e.g.*, by making benefits taxable), leaving that of UI unchanged[19]. This reform would also allow to reduce the level of payroll taxes used to fund the system, thus reducing tax distortions on the labor market. Lastly, the role of the worker's own physician, especially in Quebec, could be revised. As in other jurisdictions, an official list of physicians specialized in work-related health problems could be used instead of relying on the worker's own choice of a physician. Along the same line, more attention should maybe be paid by the WCB to certain categories of accidents (those with difficult diagnosis) that are more likely to lead to abuses. Altogether, this analysis raises an important issue for policy-makers: the necessity of considering the possible interaction between social insurance programs when studying the impact of key parameters in a given program. A useful extension to our research would be to investigate whether UI parameters influence the frequency of accidents in the construction industry.

[19]Recently, the WCB in two Canadian provinces, Manitoba and New Brunswick, has reduced benefits from 90% of the net income to 80%.

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TABLE 1

DESCRIPTIVE STATISTICS[a,b]

Variable	Mean	Standard Deviation	Minimum	Maximum
Duration (in days) on WC	31.22	66.12	2.00	1869.00
WC replacement ratio	1.14	0.11	0.47	1.43
UI replacement ratio	0.57	0.07	0.20	0.67
Real marginal weekly wage rate	288.99	114.77	5.33	1158.28
Real weekly WC benefits	329.25	135.29	3.81	910.98
Real weekly UI benefits (before tax)	272.48	122.72	4.08	680.65
Regional unemployment rate	12.55	2.81	6.90	21.60
Mineasy	0.40	0.49	0.00	1.00
Majeasy	0.12	0.33	0.00	1.00
Minhard	0.18	0.39	0.00	1.00
Majhard	0.29	0.45	0.00	1.00
Age	37.17	9.75	18.00	66.00
Dependents	0.52	0.17	0.03	0.65
Years of qualification	1.41	1.50	0.00	4.00
Quebec	0.15	0.35	0.00	1.00
Abitibi	0.06	0.24	0.00	1.00
James Bay	0.22	0.42	0.00	1.00
January	0.05	0.22	0.00	1.00
February	0.05	0.23	0.00	1.00
March	0.06	0.23	0.00	1.00
April	0.06	0.23	0.00	1.00
May	0.08	0.26	0.00	1.00
June	0.09	0.28	0.00	1.00
July	0.08	0.26	0.00	1.00
August	0.12	0.32	0.00	1.00
September	0.12	0.32	0.00	1.00
October	0.13	0.33	0.00	1.00
November	0.11	0.31	0.00	1.00
December	0.07	0.25	0.00	1.00
Censored (=1 if right-censored)	0.01	0.11	0.00	1.00

a N=8523.

b Benefits and wage are in 1981 Canadian dollars.

TABLE 2

HAZARD MODEL ESTIMATES
With generated regressors[a]
(Standard errors in parentheses)

	SPECIFICATION						
Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
Log UI	-0.940	0.134	-0.310	-0.579	-0.484	-0.702	.
	(0.906)	(0.828)	(0.208)	(0.535)	(1.017)	(0.471)	.
(Log UI)*MAJHARD	4.333	4.340	0.523	1.273	7.427	0.937	3.766
	(1.494)	(1.494)	(0.317)	(0.867)	(1.803)	(0.743)	(1.396)
Log WC	.	-0.934	.	.	.	.	.
	.	(1.287)	.	.	.	.	.
Log WCBE79	-5.114	.	1.859	-2.094	-6.724	-2.181	-4.205
	(1.896)	.	(0.342)	(0.940)	(2.321)	(0.834)	(1.679)
Log WCAF79	2.606	.	3.612	2.018	3.762	2.238	2.727
	(1.786)	.	(0.264)	(0.924)	(2.208)	(0.747)	(1.758)
Log WAGE	-0.526	-0.620	-0.041	-0.003	-0.457	0.322	.
	(0.490)	(0.489)	(0.080)	(0.232)	(0.623)	(0.180)	.
(Log WAGE)*MAJHARD	0.583	0.587	0.236	0.170	0.792	0.023	0.551
	(0.153)	(0.153)	(0.101)	(0.078)	(0.189)	(0.056)	(0.152)
Log UNEMP	0.437	0.457	0.432	0.204	0.585	0.200	0.395
	(0.150)	(0.150)	(0.130)	(0.069)	(0.195)	(0.058)	(0.146)
MAJHARD	-2.926	-2.951	-2.346	-1.100	-3.032	-0.299	-2.996
	(1.069)	(1.071)	(0.705)	(0.552)	(1.324)	(0.446)	(1.057)
MAJEASY	-2.191	-2.194	-1.881	-0.853	-3.335	-0.869	-2.194
	(0.101)	(0.101)	(0.089)	(0.045)	(0.111)	(0.030)	(0.101)
MINHARD	-0.582	-0.583	-0.542	-0.320	-0.673	-0.382	-0.582
	(0.064)	(0.064)	(0.057)	(0.029)	(0.081)	(0.024)	(0.064)
DEPENDENTS	0.265	0.393	0.114	0.026	0.306	-0.027	0.216
	(0.159)	(0.153)	(0.125)	(0.079)	(0.197)	(0.067)	(0.145)
QUALIFICATION	0.034	0.041	-0.003	0.008	0.028	-0.010	0.015
	(0.025)	(0.025)	(0.014)	(0.012)	(0.032)	(0.009)	(0.015)
Log AGE	-1.056	-0.944	-1.006	-0.595	-1.319	-0.696	-1.105
	(0.120)	(0.112)	(0.085)	(0.058)	(0.147)	(0.046)	(0.105)
QUEBEC	0.171	0.171	0.172	0.122	0.214	0.169	0.167
	(0.070)	(0.070)	(0.063)	(0.031)	(0.091)	(0.027)	(0.070)
ABITIBI	-0.568	-0.562	-0.479	-0.277	-0.618	-0.252	-0.576
	(0.099)	(0.099)	(0.089)	(0.051)	(0.125)	(0.043)	(0.099)
JAMES BAY	-1.473	-1.483	-1.349	-0.667	-2.071	-0.751	-1.570
	(0.133)	(0.133)	(0.081)	(0.064)	(0.162)	(0.052)	(0.090)

a Fixed effects for years have also been introduced.

TABLE 2 (ctd.)

	SPECIFICATION						
Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
JANUARY	0.068	0.071	-0.033	-0.056	0.127	-0.019	0.069
	(0.126)	(0.127)	(0.114)	(0.058)	(0.158)	(0.050)	(0.126)
FEBRUARY	0.066	0.073	-0.034	-0.039	0.175	0.004	0.069
	(0.124)	(0.124)	(0.112)	(0.062)	(0.156)	(0.054)	(0.124)
MARCH	0.050	0.051	-0.060	-0.035	0.120	-0.036	0.055
	(0.123)	(0.123)	(0.111)	(0.060)	(0.153)	(0.046)	(0.123)
APRIL	0.165	0.165	0.071	0.003	0.260	0.055	0.169
	(0.123)	(0.124)	(0.112)	(0.059)	(0.155)	(0.051)	(0.123)
MAY	0.178	0.172	0.096	0.071	0.211	0.087	0.180
	(0.116)	(0.116)	(0.105)	(0.056)	(0.144)	(0.045)	(0.116)
JUNE	-0.053	-0.055	-0.102	-0.033	-0.046	-0.042	-0.051
	(0.111)	(0.111)	(0.101)	(0.056)	(0.137)	(0.041)	(0.111)
AUGUST	0.105	0.102	0.090	0.055	0.154	0.039	0.104
	(0.105)	(0.105)	(0.095)	(0.052)	(0.130)	(0.039)	(0.105)
SEPTEMBER	0.024	0.023	0.002	-0.074	0.069	-0.142	0.022
	(0.104)	(0.105)	(0.095)	(0.050)	(0.130)	(0.037)	(0.104)
OCTOBER	0.020	0.021	-0.001	-0.039	0.057	-0.097	0.019
	(0.103)	(0.103)	(0.093)	(0.051)	(0.128)	(0.039)	(0.103)
NOVEMBER	-0.100	-0.097	-0.151	-0.161	-0.040	-0.203	-0.103
	(0.105)	(0.105)	(0.095)	(0.051)	(0.132)	(0.040)	(0.105)
DECEMBER	-0.276	-0.275	-0.316	-0.206	-0.228	-0.342	-0.279
	(0.116)	(0.117)	(0.106)	(0.060)	(0.145)	(0.040)	(0.116)
η	.	.	.	.	0.003	0.074	.
	.	.	.	.	(0.001)	(0.002)	.
α	.	.	.	.	2.741	0.807	.
	.	.	.	.	(0.073)	(0.006)	.
Heterogeneity Variance	1.511	1.518	1.137	.	2.814	.	1.514
	(0.101)	(0.101)	(0.080)	.	(0.112)	.	(0.101)
Sample size	8523	8523	8523	8523	8523	8523	8523
Log-likelihood Value	-27945.00	-27949.50	-27865.75	-28212.93	-28765.41	-30653.75	-27946.02

TABLE 3

AVERAGE (PER DAY) BASELINE HAZARD ESTIMATES
(Standard errors in parentheses)

	SPECIFICATION					
Intervals (in days)	**(1)**		**(3)**		**(4)**	
2-3	0.066	(0.003)	0.068	(0.003)	0.083	(0.003)
3-4	0.078	(0.003)	0.075	(0.003)	0.079	(0.003)
4-5	0.106	(0.005)	0.099	(0.004)	0.088	(0.004)
5-6	0.143	(0.008)	0.127	(0.006)	0.097	(0.004)
6-7	0.118	(0.008)	0.102	(0.006)	0.068	(0.004)
7-8	0.123	(0.009)	0.103	(0.007)	0.062	(0.003)
8-9	0.104	(0.009)	0.085	(0.007)	0.047	(0.003)
9-10	0.120	(0.010)	0.097	(0.008)	0.050	(0.003)
10-11	0.142	(0.013)	0.112	(0.010)	0.054	(0.004)
11-12	0.152	(0.015)	0.118	(0.011)	0.052	(0.004)
12-13	0.162	(0.017)	0.123	(0.012)	0.051	(0.004)
13-14	0.171	(0.019)	0.128	(0.013)	0.050	(0.004)
14-15	0.267	(0.029)	0.196	(0.019)	0.070	(0.005)
15-16	0.209	(0.025)	0.150	(0.016)	0.050	(0.004)
16-17	0.182	(0.024)	0.128	(0.015)	0.040	(0.004)
17-18	0.185	(0.026)	0.129	(0.016)	0.039	(0.004)
18-19	0.163	(0.024)	0.113	(0.015)	0.032	(0.003)
19-20	0.195	(0.028)	0.133	(0.018)	0.037	(0.004)
20-22	0.225	(0.029)	0.151	(0.017)	0.039	(0.003)
22-24	0.235	(0.032)	0.155	(0.019)	0.037	(0.003)
24-26	0.240	(0.035)	0.155	(0.020)	0.034	(0.003)
26-28	0.245	(0.038)	0.155	(0.021)	0.032	(0.003)
28-30	0.271	(0.043)	0.169	(0.023)	0.032	(0.003)
30-32	0.283	(0.047)	0.173	(0.025)	0.031	(0.003)
32-34	0.317	(0.054)	0.190	(0.028)	0.032	(0.003)
34-36	0.356	(0.063)	0.210	(0.032)	0.033	(0.003)
36-38	0.295	(0.056)	0.172	(0.028)	0.026	(0.003)
38-40	0.392	(0.074)	0.224	(0.036)	0.032	(0.003)
40-45	0.420	(0.076)	0.233	(0.035)	0.030	(0.002)
45-50	0.444	(0.086)	0.237	(0.038)	0.026	(0.002)
50-60	0.576	(0.118)	0.290	(0.048)	0.027	(0.002)
60-70	0.713	(0.161)	0.334	(0.061)	0.024	(0.002)
70-80	0.894	(0.221)	0.393	(0.078)	0.023	(0.002)
80-100	1.110	(0.301)	0.446	(0.096)	0.020	(0.001)
100-150	2.050	(0.672)	0.683	(0.175)	0.017	(0.001)
150-200	3.508	(1.379)	0.956	(0.296)	0.013	(0.001)

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