UI

Qualifying for
Unemployment Insurance:
An Empirical Analysis of
Canada

by David A. Green
and W. Craig Riddell

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UI Impacts
on Worker
Behaviour

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****This version has been reformated for electronic presentation. For the purpose of reference, the page numbers have been maintained identical to the published report.**

Unemployment Insurance Evaluation Series

Human Resources Development Canada (HRDC), in its policies and programs, is committed to assisting all Canadians in their efforts to live contributing and rewarding lives and to promote a fair and safe workplace, a competitive labour market with equitable access to work, and a strong learning culture.

To ensure that public money is well spent in pursuit of this mission, HRDC rigorously evaluates the extent to which its programs are achieving their objectives. To do this, the Department systematically collects information to evaluate the continuing rationale, net impacts and effects, and alternatives for publicly-funded activities. Such knowledge provides a basis for measuring performance and the retrospective lessons learned for strategic policy and planning purposes.

As part of this program of evaluative research, the Department has developed a major series of studies contributing to an overall evaluation of UI Regular Benefits. These studies involved the best available subject-matter experts from seven Canadian universities, the private sector and Departmental evaluation staff. Although each study represented a stand alone analysis examining specific UI topics, they are all rooted in a common analytical framework. The collective wisdom provides the single most important source of evaluation research on unemployment insurance ever undertaken in Canada and constitutes a major reference.

The Unemployment Insurance Evaluation Series makes the findings of these studies available to inform public discussion on an important part of Canada's social security system.

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Abstract

Little is known about the effect of unemployment insurance (UI) on employment durations. In this paper we take advantage of a unique accidental experiment in the Canadian UI system which created an exogenous increase in the entrance requirement (the number of weeks an individual must work to qualify for benefits) of up to four weeks in some regions. We identify the effects of this increase primarily by comparing the hazard rate out of employment for the experimental year, 1990, with that for the preceding year in regions where other parameters of the UI system do not change. We found that increasing the entrance requirement led to significant increases in employment spell durations near the entrance requirement number of weeks. Low-wage workers in seasonal industries were found to be the most affected by the changes. The fact that employment duration changes were much more pronounced in spells ending in layoffs than in those ending in quits, along with other evidence, suggests that the observed adjustment involves firms as well as workers. Overall, there is strong evidence that the UI entrance requirement does alter employment durations.

Introduction

The behavioural effects of unemployment insurance (UI) have been extensively investigated. However, most of the research has focused on the effects of UI on the duration of unemployment. Much less is known about the impact of UI on the duration of employment, and therefore on the incidence of unemployment.[1] In this paper we use a unique accidental experiment that occurred in Canada's UI program in 1990 to identify the effects of the UI entrance requirement (the requirement that individuals work a certain minimum number of weeks in a specified period to qualify for benefits) on the distribution of employment durations.

We use a unique accidental experiment that occurred in Canada's UI program in 1990 to identify the effects of the UI entrance requirement on the distribution of employment durations.

The entrance requirement may affect behaviour in various ways. A short entrance requirement, especially if combined with long maximum benefit duration, may encourage employment among those who would otherwise remain out of the labour force. In addition, a short entrance requirement may make it less costly for workers and firms to terminate poor job matches—especially if quitters are eligible for UI or layoffs are not fully experience rated. Lengthening the entrance requirement will make it less likely that individuals will enter the work force or lengthen job spells in order to qualify for UI. Setting high entrance requirements may also discourage the use of UI to subsidize seasonal work patterns. More generally, substantial entrance requirements may reduce the insurance value of the UI program by limiting coverage to individuals exhibiting strong labour force attachment. For these reasons, understanding the economic effects of UI requires knowing whether the entrance requirement has a significant impact on employment durations and how individuals adjust to changes in the entrance requirement.

In Canada, beginning in 1978 the UI entrance requirement varied from 10 to 14 weeks of employment in a 52-week period according to a formula based on the local unemployment rate. In the highest unemployment rate regions workers needed the least number of weeks to qualify. The legislation instituting this variable entrance requirement (VER) needed to be periodically renewed. In 1990, because of a dispute between the House of Commons and the Senate over issues other than the entrance requirement, these provisions were not renewed and the entrance requirement reverted to a flat 14 weeks in all regions of the country. By appropriate choice of regions we are able to analyze the effects of the 1990 change in the entrance requirement holding constant all other program features.

The empirical work in the paper uses both the 1989 and 1990 Labour Market Activity Surveys (LMAS), large representative samples of individuals which contains retrospective information on work patterns and job and personal characteristics for the sample year. We use a special version of the LMAS, in which place of residence is coded as the UI region in which the individual lives rather than the province, as is done on the public use data set. Since UI parameters vary

1 See Christofides and McKenna (1993) for an examination of employment durations in Canada.

by UI region and most provinces contain several such regions, accurately characterizing the effects of UI requires information on the UI region.[2]

We analyze the hazard rate out of employment in 1989 and 1990 for a sample of new job starts from UI regions where the suspension of the VER implied a large change in the qualifying period. Comparison of the exit rates for the two years allows us to identify the effects of the entrance requirement. We carry out that comparison using both the raw empirical hazard rates and after controlling for covariates using a duration model discussed in Moffitt (1985) and Meyer (1990). Controlling for covariates is important in order to account for possible differences between the two sample years other than those caused by the change in the entrance requirement. In addition, we make use of the panel data nature of the LMAS to examine how individuals who worked at or near the entrance requirement in 1989 adjusted to the new requirement in 1990.

Our results support several conclusions. First, there is strong evidence of a significant impact of the entrance requirement on employment durations. Secondly, many of the employment spells that would have ended at between 10 and 13 weeks in 1989 were extended to at least the new entrance requirement of 14 weeks in 1990. This suggests that significant moral hazard effects exist with respect to the taking up of UI. Third, the group of individuals who work just enough to qualify for UI in a year are disproportionately low-skilled, resource-sector workers. Fourthly, employers played an important role in adjusting to the change in the eligibility requirement. In comparisons of our empirical results to the predictions of several theoretical models, contracting models that directly incorporate firm and worker interactions appear superior to pure labour supply models in explaining observed patterns. Observed "moral hazard" effects cannot be attributed to worker reactions alone.

The paper is organized as follows. Section 1 describes key features of Canada's UI program, including the system of regional extended benefits and the suspension of the variable entrance requirement in 1990. Section 2 discusses the effect of the entrance requirement on employee and employer behaviour in the context of several theoretical models. The data are described in Section 3 and the empirical results are presented in Sections 4, 5, and 6. Section 7: Conclusion – outlines the implications for future work.

2 In work that was carried out independently of ours, Baker and Rea (1993) have also examined the impact of the 1990 suspension of the *VER*. However, their analysis is handicapped by their not having access to the information on the UI region.

1. Institutional Background

Regional differentiation has been a defining feature of Canada's UI program since 1971, and particularly since 1977-1978, when the number of UI regions was increased from 16 to 48 and a three-phase benefit structure was introduced:

Phase 1: The initial benefit phase provided one week of benefits for each week of insured employment, to a maximum of 25 benefit weeks;
Phase 2: The labour force extended phase provided for an additional week of benefits for each two weeks of insured employment in excess of 26 weeks, to a maximum of 13 weeks of benefits; and
Phase 3: The regional extended benefit phase provided for two weeks of benefits for each 0.5 percentage point increment in the regional unemployment rate in excess of 4.0 percent, to a maximum of 32 weeks of benefits.

The maximum duration of benefits from all three phases was 50 weeks, or 52 weeks from the start of the claim (given the two-week waiting period, when benefits are not payable).

The variable entrance requirement *(VER)* was also introduced at this time. The legislation contained the following provisions:

Regional unemployment rate (per cent)	Minimum number of weeks of work to qualify
6.0 or less	14
6.1 to 7.0	13
7.1 to 8.0	12
8.1 to 9.0	11
more than 9.0	10

The *VER* was to expire (unless renewed by Parliament) at the end of three years, and to revert to a fixed entrance requirement of 14 weeks in all UI regions. This feature continued to be renewed until January 1990. At that time, the House of Commons voted to renew the *VER* provisions as part of a larger package which involved revisions to the UI Act. However, the Senate refused to pass this or any other legislation because of a dispute over the new Goods and Services Tax. The stand-off continued until November 18, 1990, when the UI legislation was finally passed. Thus, for most of 1990, the entrance requirement was a fixed 14 weeks in all parts of the country. As these events were not directly related to the *VER;* the suspension thus represents an exogenous change to the UI system.

The regional extended benefit provisions provide significant incentives for individuals in high unemployment regions to obtain the minimum amount of employment required to qualify for UI. For example, if the regional unemployment rate exceeds 11.5 per cent , then 10 weeks of insured employment qualifies an individual for 42 weeks of benefits, 10 weeks from Phase 1 and 32 weeks from Phase 3.[3] In contrast, if the regional unemployment rate is 4.0 per cent , an individual

The regional extended benefit provisions provide significant incentives for individuals in high unemployment regions to obtain the minimum amount of employment required to qualify for UI.

3 Perhaps reflecting these incentives (which were introduced in the late 1970s), Card and Riddell (1993) found that during the 1980s there was an increased tendency of Canadians to work for from 10 to 12 weeks per year.

with the minimum 14 weeks of employment would be eligible for only 14 weeks of benefits. Relative to other OECD countries, the combination of a relatively brief qualifying period and long maximum duration in high unemployment regions is a distinctive feature of the Canadian system, making it an interesting case study for entrance requirement effects.[4]

Special regulations apply to "repeat users" of UI, individuals who have received UI benefits at some time in the 52 weeks preceding the beginning of a new claim. These individuals face longer entrance requirements with one exception: in regions with unemployment rates above 11.5 per cent, no distinction is made between repeat users and other claimants.

Analysis of the impact of the 1990 change in the entrance requirement is also complicated by the fact that changes in regional unemployment rates caused changes in two key UI parameters—the entrance requirement itself and the maximum benefit duration. Changes in the latter alter the incentives for qualifying for UI and, thus, potentially make the effects of the entrance requirement difficult to separately identify. This is particularly true in 1990, when unemployment in Canada increased significantly. Thus, one might observe differences between 1989 and 1990 in employment durations not because of a behavioural response to the shift in the entrance requirement but because of changes associated with increases in the unemployment rate. To avoid identification problems of this sort, we restrict our sample to UI regions which have unemployment rates of 11.5 per cent or more throughout 1989 and 1990. For these regions, the *VER* suspension implied an increase in the qualifying period from 10 weeks in 1989 to 14 weeks in 1990. These regions, which we call maximum entitlement regions, constitute the top category in terms of regional extended benefits, and within them UI parameters no longer change with the unemployment rate, nor are there different rules for so-called repeat users. Focusing on these regions has the important benefit of allowing us to examine the change in the entrance requirement holding constant all other UI parameters.[5]

The maximum entitlement regions contain the most generous form of the UI system in Canada and thus the strongest incentives to work just enough to qualify for UI benefits. In this sense our estimates of entrance requirement effects will be the upper bounds on the average effects. However, with the national unemployment rate reaching 11.8 per cent in the 1981-1982 recession and 11.3 per cent in the 1990-1992 recession, that upper bound has been relevant for large numbers of Canadians. Most important, any limitations due to the unrepresentativeness of the regions are more than offset by the advantages of a clean "experiment" in which only one UI parameter changes.

4 See OECD Employment Outlook July 1991 for information on qualifying requirements, maximum benefit durations and other UI parameters in member countries. In both the United States and the U.K., a certain level of contributions is needed to qualify. Thus the qualifying period varies across individuals according to their earnings, making analysis of the effects of the entrance requirement more difficult. Among OECD countries, only the U.K. is similar to Canada in that its high unemployment regions have a brief qualifying period and long maximum benefit duration.

5 It is not possible to carry out both a pre (-1989) and post (-1990) experimental analysis for two reasons. First, the LMAS is not available after 1990, and secondly, the UI program, including the number of UI regions and the *VER*, changed at the end of 1990 rather than reverting to its previous form.

2. UI Entrance Requirements in Theory

This section discusses the impact of the UI entrance requirement in the context of three alternative models of employee-employer behaviour: a static supply model, a search model of employment and unemployment, and an implicit contract/lay-offs model. These models provide complementary viewpoints for examining the predicted effects of changes in the minimum amount of employment required to qualify for UI. In order to conserve space in this paper, and because the models are extensions of those in the literature, we simply summarize the implications of this theoretical analysis.[6]

These models provide complementary viewpoints for examining the predicted effects of changes in the minimum amount of employment required to qualify for UI.

In the static labour supply model—as used in the UI context by Moffitt and Nicholson (1982) and Phipps (1990)—an individual chooses the number of weeks of employment and non-employment (leisure) in a given period by selecting the utility maximizing combination of weeks worked and consumption from a feasible budget set.[7] Figure 1 shows the budget sets with and without a UI program for an individual with a one-year planning horizon and a UI system with parameter values associated with a maximum entitlement region. *AB* is the budget set in the absence of UI and *ACDB* with UI.[8] The individual must work at least H_{min} weeks in the year to qualify for benefits; once qualified, he is eligible to receive UI benefits equal to *wb*, (where *w* is weekly employment earnings and *b* is the benefit rate) for the rest of the year.[9]

In this model, a continuous distribution of preferences, as measured by the marginal rate of substitution *(MRS)* between consumption and leisure, will induce a

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6 Details are available from the authors on request.

7 The model is static in the sense that current period decisions are assumed to not affect decisions in other periods. This situation would apply if the individual's lifetime allocation decision could be decomposed into two stages, the number of 'years' to work and the number of weeks to spend working in each 'year'. Note also that the model assumes that the allocation of time to work and non-employment can be made with certainty.

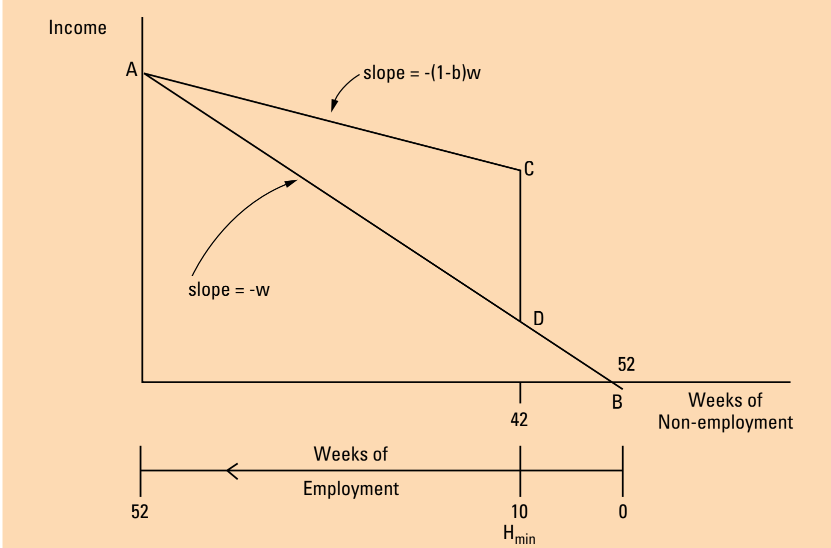
8 The budget constraint does depend on the length of the planning horizon. Previous analyses of UI in the context of the static labor supply model typically assume a one year horizon. For the maximum entitlement regions, with a one-year planning horizon there is a discontinuity in the budget constraint at 10 weeks, corresponding to the point at which the entrance requirement is met. At point *C* the individual works 10 weeks and receives UI benefits for 42 weeks. Along *AC*, the slope of the budget constraint is *(1-b)w*, reflecting the fact that each additional week worked implies one less week of UI benefits.

With a longer planning horizon, there will continue to be a discontinuity in the budget constraint at the entrance requirement H_{min}. In maximum entitlement regions there will be an additional kink in the budget constraint at 18 weeks of work, corresponding to the point at which the maximum 50 weeks of potential UI benefits is attained. In these regions, the slope of the budget constraint between 10 and 18 weeks is *(1+b)w,* reflecting the fact that each additional week of work yields market earnings of *$w* and entitles the individual to an additional week of benefits (from Phase I). For weeks of work exceeding 18 weeks, the budget constraint has slope *(1-b)w,* reflecting the fact that each additional week worked in excess of 18 weeks implies one less week of potential UI benefits, given the upper limit of 50 weeks of benefits.

In regions with unemployment rates below 11.5 percent, the possibilities become more varied, as the end of Phase I benefits and in some cases the end of Phase II benefits produce additional kinks in the budget constraints. Whatever the regional unemployment rate or planning horizon, a discontinuity in the budget constraint will occur at the entrance requirement H_{min}.

9 Apart from the two-week waiting period, this accurately characterizes the UI system in the regions on which we focus. In the maximum entitlement regions the minimum employment of 10 weeks qualifies the worker for up to 42 weeks of benefits; thus the available 52 weeks are fully allocated for any choice of weeks of work.

Figure 1
Budget Constraint: Maximum Entitlement Region, One Year Horizon



continuous distribution of employment durations in the segments *(0,*H_{min}*)* and *(*H_{min}*,52)*. However, a set of *MRS* values will cause an individual to choose $H=H_{min}$, and similarly for $H=0$ and $H=52$. Thus the predicted density function for H will have continuous portions corresponding to $H \varepsilon (0,H_{min})$ and $H \varepsilon (H_{min},52)$, and mass points corresponding to $H=0$, $H=H_{min}$, and $H=52$.

The empirical work in this paper is conducted in a hazard function framework, so it is useful to discuss the impact of the entrance requirement on the hazard rate out of employment. The hazard rate, $h(H)$, is the probability that an employment spell will end at duration H conditional on it being of at least duration H. Thus,

$$h(H) = \frac{\psi(H)}{(1 - \Psi(H0)} \qquad (1)$$

where ψ is the density function for H and Ψ is the cumulative distribution function for H. The hazard rate is a useful tool because it provides an equivalent characterization to the distribution function of employment durations that is both intuitive and empirically tractable. The hazard rate corresponding to $H=H_{min}$ is just the probability that an individual would choose exactly H_{min} weeks divided by the probability of an annual weeks choice of at least H_{min} weeks. Because a range of *MRS* values can generate a choice of H_{min}, a spike will exist in the hazard function at this point.

The shape of the hazard function in the range $H < H_{min}$ depends on the distribution of preferences. Nonetheless, a plausible pattern is that the hazard rate will decline as H approaches H_{min}. Individuals whose preferences would lead them to work less than 10 weeks in the absence of UI are, with UI, more likely to lengthen their employment spell in order to qualify for benefits the closer to H_{min} is their preferred employment duration in the absence of UI.

The search model used here assumes that individuals can occupy one of two states: employment and non-employment.

Note that the predicted spike in the hazard function results from two distinct behavioural responses to the UI program. For those who, in the absence of UI, would work $H>H_{min}$ weeks per year, both income and substitution effects imply a reduction in weeks of employment, causing some individuals to reduce their weeks worked to $H_{min.}$ In addition, some individuals who, in the absence of UI, would perfer to work fewer than H_{min} weeks will, in the presence of the UI program, increase their number of weeks of work to H_{min} in order to qualify for UI benefits when not employed.

What happens to the distribution of weeks worked and the associated hazard function if H_{min} is increased to H_{min}'? It is clear that the mass point formerly located at H_{min} will move to H_{min}', but the height of the new point relative to the old is less clear. Given a fixed decision period of one year, the benefit from working just enough weeks to qualify will be reduced both because the individual can now collect UI benefits for fewer weeks in the year, and because they must work additional weeks in order to qualify. Thus, some of those who formerly worked exactly H_{min} weeks may now reduce their labour supply. At the same time, some of the individuals who formerly worked $H \varepsilon (H_{min},H_{min}')$ may now increase their weeks worked to H_{min}'. The overall result is indeterminate and depends on the distribution of preferences. The same can be said of the hazard function: in response to the increase in the entrance requirement, the spike occurring at the UI qualification point will move to H_{min}' and may become larger or smaller.

Search models provide an alternate theoretical framework for examining these phenomena. These models are set in a dynamic framework and do not assume that individuals choose weeks of work and non-employment with certainty. The search model used here assumes that individuals can occupy one of two states: employment and non-employment.[10] When employed, individuals are paid a wage, w, equal to their marginal product. New, independent draws on productivity occur over the course of an employment spell, with the new draws occurring with probability η in any period. A non-employed individual is assumed to search for a new job with constant search intensity. New job offers arrive with probability $\lambda > \eta$ in any non-employed period; thus unemployed search may be an attractive alternative to employment at a low wage because of the higher probability of receiving a new wage offer. Income when non-employed depends on the duration of the preceding employment spell. If the preceding spell was longer than H_{min} weeks then the individual receives a benefit, B, per period for the remainder of the non-employment spell.

In this setting, employed search can be shown to have a reservation wage property. The reservation wage, w^*, is such that in response to a productivity draw $w<w^*$, the individual will terminate the employment spell and engage in unemployed search. However, if $w \geq w^*$ the individual will continue employment for at least one more period.

10 The effects of the UI entrance requirement have not previously been analyzed in a search theoretic setting. The model is an extension of one developed by Mortensen(1990) which does not allow for UI entry requirements but that is otherwise more general. The model employs two simplifying assumptions about the UI system: individuals who qualify for UI receive a fixed benefit B per week, and benefits can be received indefinitely.

The predicted pattern of the reservation wage is shown in Figure 2(a). For $H \geq H_{min}$, w^* is independent of H. Thus, once the entrance requirement has been met the worker faces a stationary environment in which the probability of terminating the employment spell is independent of the spell's duration. However, for $H < H_{min}$, w^* is a decreasing function of H. This property arises because the closer workers are to meeting the entrance requirement, the more willing they are to remain employed—even in the presence of a poor productivity draw—in order to qualify for UI benefits once the job spell ends. The reservation wage, therefore, is predicted to decline with H until H_{min-1}, jump up at $H = H_{min,}$ and remain constant thereafter.

The predicted behaviour of the hazard rate out of employment is illustrated in Figure 2(b). The UI entrance requirement implies a falling hazard to H_{min-1} weeks of employment, a spike at H_{min} weeks, and a constant hazard thereafter.[11] This pattern is similar to that implied by the static labour supply model, with the exception that in the latter the prediction of a declining hazard as H approaches H_{min} requires specific assumptions about the distribution of preferences.

What is the impact of an increase in the entry requirement? The model supposes that the choices at various durations depend not on the value of H itself but on its

Figure 2a
Predicted Pattern for the Reservation Wage

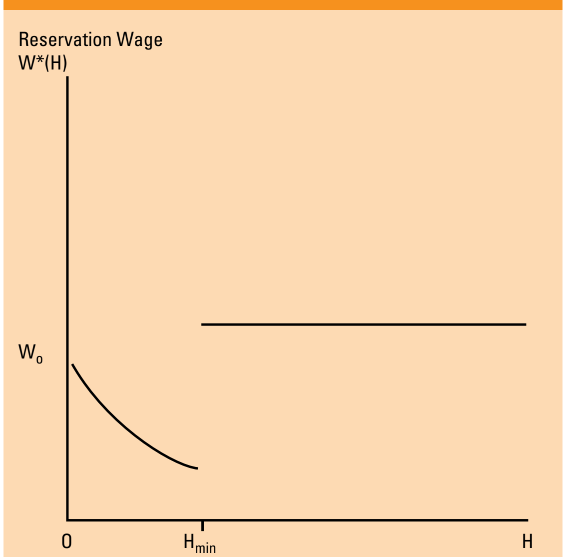
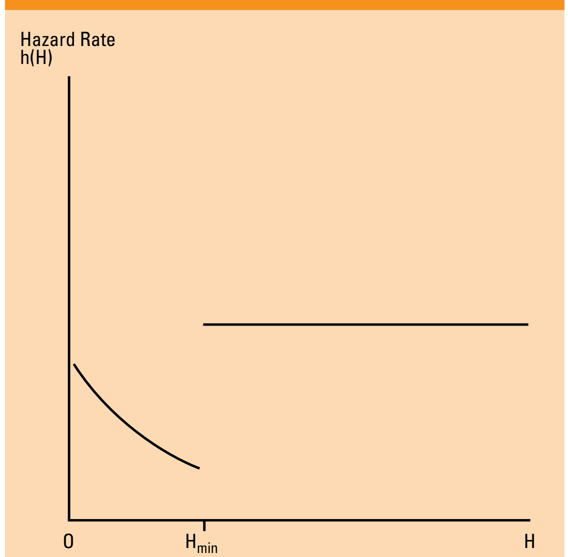


Figure 2b
Predicted Pattern for the Hazard Rate



11 The predicted spike in the hazard at $H = H_{min}$ arises from two sources. First, any new productivity draws that arrive at H_{min} will be compared with a higher reservation wage than at H_{min-1}, so that individuals receiving new draws at H_{min} will be more likely to leave employment than those who received new draws at H_{min-1}. Secondly, individuals who do not receive a new productivity draw at H_{min} but have a current wage that lies between the reservation wage at H_{min-1} and w_0 (the reservation wage for $H \geq H_{min}$) will now leave employment. The latter group have wages that were high enough for them to stay employed at H_{min-1} but not high enough for them to choose to remain employed when they can receive UI benefits while not working.

value relative to H_{min}. Thus, with a new $H_{min}' > H_{min}$, the general pattern depicted in Figure 2 will be unchanged. This is in contrast to the results from the static model in which the hazard function could change in a number of plausible ways with a lengthened qualification period, depending on the form of heterogeneity in preferences. In comparing the two models, it is interesting to note that one can generate a spike in the hazard at H_{min} either from a combination of the incentives in the UI system and individual heterogeneity (as in the static model) or purely from the incentives in the system with no heterogeneity in preferences (as in the search model). The static model might describe the choices made by seasonal workers to terminate employment spells at specific points, while the search model might reflect the choices made by workers with no such pre-set plans.

Both models discussed above are based on individual rather than employer perspectives of job terminations. While such theories may be relevant for Canada in this time period—since individuals who quit their job could qualify for UI benefits—they do not address the role of the employer in responding to the incentives created by the UI program, including the entrance requirement. In order to do so we also investigate a standard implicit contract model.

In this short run model, the employer has a fixed capital stock and a fixed pool of "permanently attached" and homogeneous workers. The firm operates with a production function, $f(L, H)$, where L is the number of workers employed, H is the weeks per year per employee, and L and H are perfect substitutes. Product demand is subject to random fluctuations. When employed, workers receive a weekly wage, w. Employed individuals also receive UI benefits from the government if they work less than the full year, with the benefits equaling bw if $H \geq H_{min}$, and zero otherwise. The employer and employees are risk neutral and risk averse, respectively, and reach a Pareto efficient contract governing the decision variables (w, L, and H).

In the absence of a UI system and with risk averse workers, the optimal contract would involve values of H and w such that there are no layoffs, where a layoff is defined as an individual in the employer's pool not working at all during the contract period (one year in our example). With a UI system with an entrance requirement, employers have a strong incentive to have $H \geq H_{min}$: for $H < H_{min}$, employers must pay a compensating differential to offset the fact that their workers do not receive UI benefits. Thus, in the optimal contract, a range of demand shocks may lead to $H = H_{min}$. In addition, the costs of downward adjustment of H below H_{min} imply that H and L are not perfect substitutes. This could lead to part of the employer's pool being laid off while the rest work at $H = H_{min}$. Finally, in the absence of experience rating, the UI system promotes the subsidization of firms that have part-year employment patterns with those that have more stable patterns. These results are in keeping with those in a study by Feldstein (1976), though in the model in that particular case, H is exogenous. Thus, this model also implies a spike in the employment hazard at H_{min} weeks, although this prediction results from incentives as much for firms as for workers.

If the entrance requirement is increased to H_{min}' a new optimal contract would hold. For some demand shocks which result in $H = H_{min}$ for all individuals in the

One model for these high-unemployment maximum entitlement regions is the "community pressure" model...

pool under the old contract we may observe layoffs for some and $H = H_{min}'$ for the remainder of the individuals in the pool under the new contract. This is in contrast to the supply side models, in which workers are free to choose $H = H_{min}'$.

These results assume a given number of permanently attached workers. It is interesting to consider how the size of the firm's labour pool is determined in the first place, and how the number of workers attached to a particular firm may adjust over time. One model for these high-unemployment maximum entitlement regions is the "community pressure" model described by the Newfoundland Royal Commission on Employment and Unemployment (1986, p. 283):

> "Since jobs are so scarce, employers come under pressure from the community to qualify as many people as possible for UI. Although not always strictly adhered to, the informal rule is that once someone qualifies, he or she should be laid off and someone else hired until that person in turn qualifies... This pattern of short periods of employment followed by long periods of unemployment...has covertly become the main form of income security in Newfoundland."

According to this "community pressure" model, the number of workers in the firm's labour pool expands until $H = H_{min}$ for each member.[12]

What do these theoretical models imply about the impact of changing the entrance requirement on aggregate labour market behaviour in these regions? All three models imply that total labour supply (aggregate weeks worked) may either increase or decrease. Increased labour supply arises because some individuals will wish to increase their weeks of work in response to the longer qualifying period. Offsetting this effect is the potential decline in labour supply associated with reduced participation (in the static model), earlier termination of employment (in the search model), or increased layoffs (in the implicit contract model). For similar reasons, total employment and unemployment may either increase or decrease. Apart from the unambiguous prediction that the labour force participation rate will not rise, the impact of the higher entrance requirement on aggregate labour market outcomes must be resolved empirically. Determining which model best describes labour market adjustment to this dimension of UI is also an empirical issue which we address below.

12 Although the implicit contract model discussed in the text assumes a homogeneous workforce, it is interesting to note that the "community pressure" version of the model predicts that even in the presence of heterogeneous preferences, all members of the firm's labour pool will work H_{min} weeks.

3. Data

The empirical analysis uses data from the 1989 and 1990 Labour Market Activity Surveys (LMAS). The LMAS surveys approximately 60,000 individuals in January through March of a given year about their labour market activities in the preceding calendar year. In addition to standard demographic information, the LMAS asks questions about each job held in the previous year (up to five jobs for each individual). The LMAS includes the start and end dates of every job spell, the usual wage and number of hours per week on the job, and the firm size, occupation, industry and union status associated with the job.[13]

The LMAS surveys approximately 60,000 individuals in January through March of a given year about their labour market activities in the preceding calendar year.

As described above, the sample is restricted to individuals residing in UI regions where the unemployment rate remained over 11.5 per cent throughout 1989 and 1990.[14] In the public use version of the LMAS, the place of residence variable is the province in which the respondent resides at the interview date. Use of this variable in studies of UI experience is misleading because provinces typically include many UI regions, each of which could face different unemployment rates and therefore different UI parameters. We were fortunate to gain access to a version of the LMAS which includes the place of residence coded as the UI region in which the individual resided at the time of the interview.[15]

In addition to the sample exclusions based on location, we remove all individuals who were full time students during the relevant sample year because they are ineligible to receive UI benefits during non-employment spells that include schooling. Similarly, we exclude individuals over the age of 65 because they were ineligible to receive UI benefits. For each individual, we retain only job spells in which the individual is a paid worker and we drop spells on which the weekly wage is less than $15 per week since those jobs do not count towards UI qualification. The definition of a “job” in the LMAS allows for up to five non-permanent separations in the sample year, including temporary layoffs. Because workers on temporary layoff can receive UI benefits, we break job spells that include temporary layoffs into separate spells before and after the layoff . Finally, because UI eligibility can be met via two or more job spells we combine job spells for an individual into employment spells in a manner described in Appendix A.

Two problems arise in using this data to estimate hazard models. First, although our data contain the start dates of employment spells that begin before as well as during our sample period, the fixed start date of the sample implies that our data face the standard length-biased sampling problem (Lancaster, 1990, p.95). To address this problem, we restrict our sample to employment spells which begin in the relevant sample year. To claim generality for the results based on such a sample, one needs

13 We do not use the 1988 wave of the LMAS because of apparent errors (described to us by sources at Statistics Canada) in which some jobs that actually started in 1987 are attributed to early 1988.

14 There are nine such regions, located in all provinces except Ontario.

15 The version we use also includes data on plant size where the individual works, which is masked on the public-use tape, and on age measured as a continuous variable. We are grateful to the Special Surveys Division of Statistics Canada for allowing us access to this data set and for access to their computer to carry out our estimation.

to assume that the process generating new employment spells is stationary over time. The alternative approach to the length biased sampling problem is to model the nonstationary nature of the spell-generating process. Our approach avoids the strong assumptions needed to do the latter modeling by making the strong assumption of stationarity across years.

Our second problem arises from potential nonstationarity in the spell generating process within a given year. Such nonstationarity within a year is reasonable in the process generating employment spells, where more spells might be generated in the spring and summer and at Christmas than at other times of year. To avoid differences between the hazards for the experimental and non-experimental periods that arise solely from seasonality, we restrict both the 1989 and the 1990 samples to end in the 46th week of the year (the 46th week in 1990 is the week just preceding that which includes November 18). Thus, spells ongoing at the 46th week in each year are censored at that point, and spells starting after the 46th week are not included.

The sample we use in what follows is composed of 2,824 employment spells that start after January 1, 1989, and before the 47th week of 1989; and 2,567 spells that start after January 1, 1990, and before the 47th week of 1990. A restricted subsample is used to examine the process by which individuals adjusted to the increase in the entrance requirement. This subsample contains individuals who had employment spells from 10 to 13 weeks and which terminated in 1989. The longitudinal nature of the LMAS allows us to examine the behaviour of these individuals in 1990.

4. Empirical Evidence

The empirical work uses a duration analysis framework which is centred on the hazard rate as defined in equation 1.[16] The main goal of the empirical work in this paper is to characterize the effect of the entrance requirement on employment durations by comparing the hazard rate functions for 1989 with those for 1990.

The main goal of the empirical work in this paper is to characterize the effect of the entrance requirement on employment durations by comparing the hazard rate functions for 1989 with those for 1990.

The simplest tool for examining the impact of the entrance requirement is the empirical hazard rate function (also called the Kaplan-Meier estimator). In samples with fixed end dates, an employment duration H could be observed either because the employment spell was terminated at H or because for this spell week H occurs in the last week of the sample period. The latter type of spell could be of true length greater than H and is termed censored. The empirical hazard rate for H equals the number of spells which actually terminate in week H expressed as a fraction of spells eligible to end in week H (that is, spells that have neither been terminated nor censored at a shorter duration). This estimator has the beneficial properties of consistently handling censored spells and of not imposing a specific functional form for the related employment duration density.

With a single year of data, or even with a sample covering several years in which the UI system does not change, examination of the empirical hazard does not permit identification of the effects of the entrance requirement. Further, since changes in the entrance requirement are generated by movements in the regional unemployment rate in typical years, comparisons of the hazard across regions does not easily allow identification of the *VER* effects separately from the effects of changes in the unemployment rate. However, the 1990 accidental experiment implies a shift in the entrance requirement that is exogenous to labour market conditions and to changes in other parameters of the UI system. If we assume that 1989 and 1990 are identical years apart from the change in H_{min}, then direct evidence of the effects of the entrance requirement exists if there is a spike in the hazard at 10 weeks in 1989 which disappears in 1990, and a spike at 14 weeks in 1990. As suggested by Figure 2, one measure of the impact of the entrance requirement is the height of the spike in the hazard function at H_{min} weeks above the hazard that is relevant for $H>H_{min}$. This measure can be seen as capturing a form of moral hazard since the spike would be composed of workers who continued their employment, or "saved" their non-employment, until they just reached H_{min} weeks.

The empirical hazard rates for 1989 and 1990 for our samples, along with associated standard errors, are presented in Table 1.[17] The hazard rates are also plotted in Figures 3 and 5. Consider first the hazard function for 1989. The theoretical analysis suggested a declining hazard leading up to the UI requirement threshold of 10 weeks, a spike in the hazard at 10 weeks, and a flat shape for the hazard thereafter.

16 See Lancaster (1990) for a complete discussion of duration models.

17 The LMAS over-samples individuals from rural regions but includes a weight for each observation which captures this non-random selection. We construct all tables in this paper using these weights. Further, each individual's contributions to the log likelihood functions are weighted and normalized so that all the weights for our sample sum to the true sample size.

Table 1
Empirical Hazard Rates
Maximum Entitlement Regions, 1989 and 1990*

Week	1989	1990
1	0.0136 (0.0022)	0.0121 (0.0021)
2	0.0203 (0.0027)	0.0184 (0.0026)
3	0.0320 (0.0035)	0.0257 (0.0031)
4	0.0195 (0.0028)	0.0204 (0.0028)
5	0.0274 (0.0034)	0.0311 (0.0036)
6	0.0199 (0.0029)	0.0187 (0.0028)
7	0.0247 (0.0033)	0.0266 (0.0034)
8	0.0219 (0.0032)	0.0223 (0.0032)
9	0.0367 (0.0042)	0.0265 (0.0035)
10	0.0555 (0.0053)	0.0375 (0.0043)
11	0.0430 (0.0049)	0.0173 (0.0030)
12	0.0314 (0.0043)	0.0325 (0.0042)
13	0.0310 (0.0044)	0.0175 (0.0032)
14	0.0624 (0.0064)	0.0710 (0.0065)
15	0.0306 (0.0046)	0.0207 (0.0036)
16	0.0256 (0.0043)	0.0531 (0.0060)
17	0.0387 (0.0055)	0.0269 (0.0044)
18	0.0427 (0.0059)	0.0341 (0.0051)
19	0.0258 (0.0047)	0.0352 (0.0053)
20	0.0309 (0.0052)	0.0417 (0.0059)
21	0.0324 (0.0055)	0.0366 (0.0058)
22	0.0385 (0.0062)	0.0360 (0.0059)
23	0.0528 (0.0075)	0.0624 (0.0080)
24	0.0236 (0.0052)	0.0277 (0.0056)
25	0.0458 (0.0076)	0.0386 (0.0069)
26	0.0170 (0.0048)	0.0203 (0.0052)
27	0.0414 (0.0077)	0.0480 (0.0083)
28	0.0274 (0.0065)	0.0190 (0.0054)
29	0.0198 (0.0058)	0.0225 (0.0062)
30	0.0239 (0.0068)	0.0077 (0.0038)
31	0.0254 (0.0072)	0.0366 (0.0086)
32	0.0410 (0.0097)	0.0060 (0.0036)
33	0.0113 (0.0054)	0.0187 (0.0067)
34	0.0209 (0.0078)	0.0175 (0.0070)
35	0.0191 (0.0079)	0.0087 (0.0051)
36	0.0061 (0.0048)	0.0324 (0.0103)
37	0.0273 (0.0106)	0.0000 (0.0000)
38	0.0025 (0.0035)	0.0206 (0.0092)
39	0.0175 (0.0101)	0.0199 (0.0099)
40	0.0203 (0.0114)	0.0224 (0.0111)
41	0.0323 (0.0153)	0.0176 (0.0105)
42	0.0254 (0.0148)	0.0123 (0.0097)

Standard errors in parenthesis. Hazard rates are 0 for weeks 43 through 46 both years.
* Define *FH* as the number of failures in week *H* and *RH* as the risk set in week *H* (the number of spells that have neither been terminated nor censored before *H*). Then, the hazard estimate is *h(H)=FH/RH* and the variance is *[h(H)(1-h(H)]/RH.* Our data is weighted so *RH* is normalized by (true sample size/weighted sample size) to calculate appropriate standard errors.

The hazard depicted in Figure 3 might be construed as fitting this general pattern, but the fit is not completely convincing. The hazard is lower for employment durations preceding 10 weeks than for the weeks directly following 10 weeks but the hazard at 9 weeks is large relative to the hazard for earlier weeks. Theory suggests that the hazard ought to be especially low at 9 weeks. There does appear to be a strong spike in the data at 10 weeks, but there are also other spikes, such as those at 14, 18 and 23 weeks, with the spike at 14 weeks dominating that at 10 weeks. The best we can say is that there is potential for the UI entrance requirement to explain part of the pattern we observe.

The hazard rate is a useful empirical tool, especially in situations in which agents update decisions each period (Kiefer, 1988). However, we are also interested in the percentage of employment spells that terminate at 10 weeks, that is, the unconditional probability of termination at 10 weeks, as this will indicate whether terminations at H_{min} are quantitatively significant relative to terminations at other points. Figure 4 presents the plot of the probability density function for completed employment spells for maximum entitlement regions for 1989.[18] Again, one can observe an apparent spike at 10 weeks, which is higher than any other in the unconditional probability density. The unconditional probability of a spell terminating at 10 weeks is 4.3 per cent. Employment spells of this length, therefore, do appear to be an important phenomenon. However, the spike at 10 weeks does not dominate that at 14 weeks by a substantial amount.

Figures 5 and 6 plot the hazard and unconditional probability density functions for 1990. As in 1989, the hazard contains several spikes, with the spike at 14 weeks becoming even more prominent. Examination of Figure 6 shows that the spike at the new 14-week entrance requirement now dominates the completed employment density.

Figure 7 contains the empirical hazards for 1989 and 1990 plotted together. The hazard rates for the two years are quite similar for durations below 10 weeks and above 20 weeks, but very different for weeks 10 through 20.[19] The large hazard rate value at 10 weeks duration in 1989 is significantly reduced in 1990 and no longer appears as a spike relative to the weeks following it. In fact, the employment durations between the old and new values of H_{min} generally have lower hazard rate values in 1990. The hazard for week 13 (the week just before the new H_{min}) suffers a particularly large drop. Further, the hazard rate at 14 weeks is larger. All of this behaviour accords well with the predictions of theoretical models incorporating a UI entrance requirement. The higher hazard rates in 1990 through much of the range between 15 and 20 weeks might arise if workers switched to jobs that ended above, but not exactly, at 14 weeks that year. An examination of the 1990 experiences of individuals who worked 10 to 13 weeks in 1989 will provide some insight on this issue. Regardless, the similarity in the hazard rates below 10 and above 20 weeks for the two years, combined with the hazard rates in the 10 to 19-week range indicating that workers were in slightly longer jobs in 1990, implies that observed differences between the two years cannot easily be attributed to recessionary effects.

18 To provide an estimate of the density of completed spells that does not suffer from biases induced by censoring, we constructed the estimates of the value of the density at each week from the empirical hazard rate estimates as, $f(H) = (1-h(1))(1-h(2))...(1-h(H-1))h(H)$.

19 The null hypothesis that the differences between the 1989 and 1990 hazard rates for weeks 1 to 9 are jointly 0 cannot be rejected at the 5-per cent level of significance. A similar test for weeks 21 to 42 also indicates no rejection at the 5-per cent level, while the same test for all weeks together soundly rejects the null at the 5-per cent level. The test statistics are distributed as $\chi^2(9)$, $\chi^2(23)$, and $\chi^2(42)$, respectively, and take values of 6.57, 32.39 and 97.72, respectively.

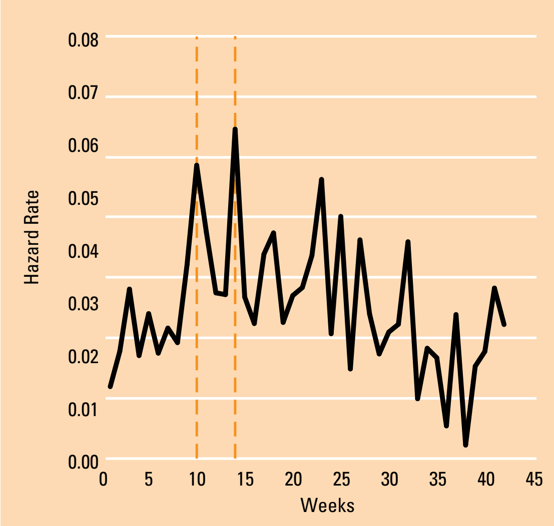


Figure 3
Empirical Hazard:
Maximum Entitlement Regions, 1989

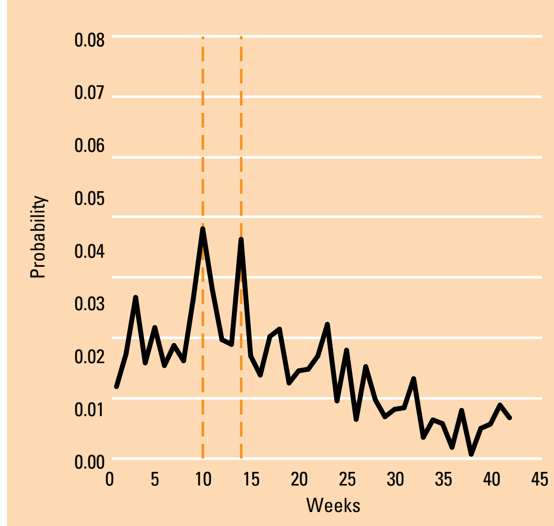


Figure 4
Empirical Density of Completed Spells:
Maximum Entitlement Regions, 1989

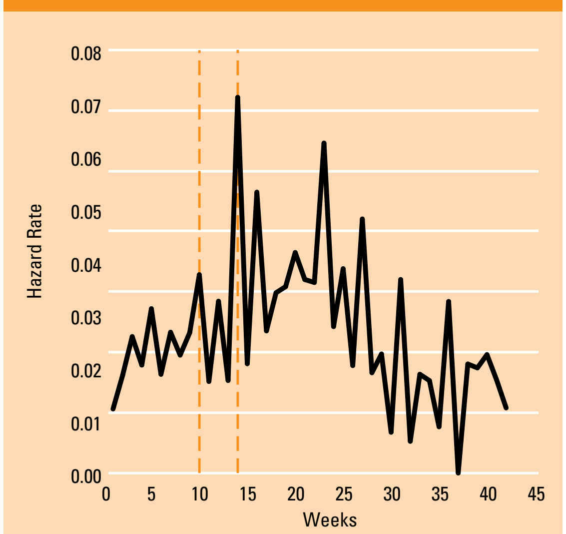


Figure 5
Empirical Hazard:
Maximum Entitlement Regions, 1990

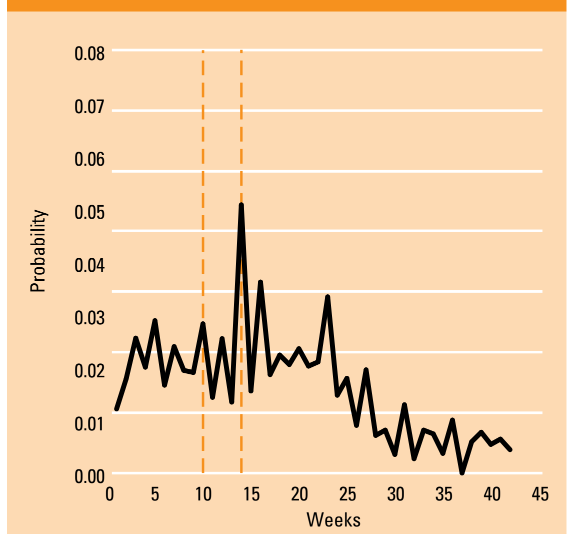
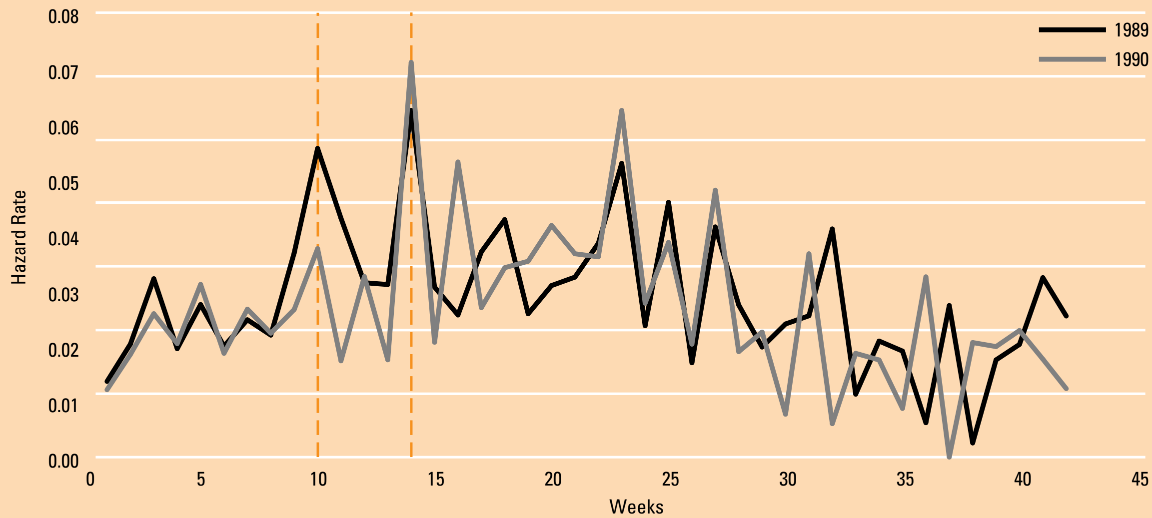


Figure 6
Empirical Density of Completed Spells:
Maximum Entitlement Regions, 1990

Figure 7
Empirical Hazards: Maximum Entitlement Regions, 1989 and 1990



It seems reasonable to claim that the spike at 10 weeks in the 1989 hazard and the relatively low values for the hazard before 10 weeks are at least partly due to the UI entrance requirement. One useful measure for determining the impact of the entrance requirement on employment durations is the height of the spike at H_{min} weeks above the hazard for $H>H_{min}$ weeks. An estimate of this measure is given by $h(14)$ in 1990 minus $h(14)$ in 1989. That difference equals 0.009, with a standard error of 0.009, under the assumption that data from the two sample years is independent. Obtaining an estimate in this way controls for anything particular about a 14-week duration that might exist regardless of whether it is the entrance requirement duration. Interpreted within a search framework, the estimate says that approximately 1 per cent of individuals who make it to H_{min} weeks do so with the intention of ending employment as soon as they qualify for UI. The estimate, however, is not significantly different from zero.

In the static labour supply model, with individuals choosing employment spell lengths with certainty, it is more natural to consider effects in terms of the unconditional probability density. The difference in the probability of a complete spell of 14 weeks in 1990 versus 1989 is 0.009 with an associated standard error of 0.006. Thus, under the static model approximately 1 per cent of all employment spells are terminated at H_{min} weeks as a result of worker preferences for working exactly H_{min} weeks.

A complete understanding of the effects of the change in the entrance requirement necessitates an examination of how the hazard changed at durations near 14 weeks as well as movements in the spike at H_{min} weeks. To facilitate this examination we present in Table 2 values for the survivor function (the probability that a spell lasts at least as long as some duration (H) for weeks 10 through

Table 2
Survivor Function Values

Week	1989	1990	Difference 1990-1989
10	0.805	0.818	0.013 (0.011)
11	0.762	0.790	0.028 (0.012)
12	0.730	0.776	0.046 (0.012)
13	0.707	0.750	0.043 (0.012)
14	0.686	0.737	0.051 (0.013)
15	0.644	0.686	0.042 (0.014)
21	0.530	0.550	0.020 (0.014)

Standard errors in parentheses.

15 and week 21 for 1989 and 1990. The survivor function value at 10 weeks is higher in 1990 than in 1989, although the difference is not significant. Increasing the entrance requirement did not result in a significant increase in exits from employment at shorter durations. At 11 weeks the survivor functions depart further and then continue to grow apart through to 14 weeks. Each of these differences is significantly greater than zero. This is in accordance with theory as the spike in the hazard at $H_{min}=10$ weeks in 1989 and the higher value of the hazard after than before H_{min} should imply a smaller probability of a spell surviving to these weeks in 1989. In fact, the probability of a spell reaching 14 weeks is over 5 percentage points higher in 1990 than in 1989, with 3.8 of the percentage points being accounted for by the relative reduction in the probability of leaving employment between 10 and 13 weeks in 1990 versus 1989. Assuming no one withdrew from the labour force in response to the increased entrance requirement, the increase in the survivor function over the 10- to 13-week range in 1990 indicates the extent to which individuals who formerly worked just enough to qualify for UI increased the length of their employment spells under the new requirement. Alternatively, if some individuals withdrew from the labour force in response to the increase in H_{min}, then the increase in the survivor function in the 10 to 13-week range could reflect the removal from the employment duration distribution of those who formerly worked this range of weeks.

At the 15-week point the survivor function values mover closer together but the difference is not eliminated. This reflects the small increase in the hazard rate at 14 weeks between the two years. By 21 weeks, the difference is once again similar to what was observed at the 10-week point and is no longer significantly different from 0. Thus, as seen in Figure 7, compared to 1989, employment spells in 1990 were less likely to end in the 10 to 13-week range and more likely to end in the 14 to 20-week range. One interpretation of these results is that, under the new requirement, individuals who had formerly been able to select 10 to 13-week (and, in particular 10-week) employment spells were forced into longer jobs that could not necessarily be terminated at exactly 14 weeks. This does not fit easily with pure supply-side models, which predict that those formerly in the 10 to 13-week range will move to exactly 14 weeks in 1990.

As a summary measure of the impact of the change in the entrance requirement, we next present estimates of the effect on the unemployment rate. We use an estimated unemployment rate formed as the average unemployment duration for these regions divided by the sum of the average unemployment duration and the average employment duration. For the average unemployment duration, we use the inverse of the average hazard rate out of non-employment for Newfoundland in 1989 as calculated from gross flows data in Jones and Riddel (1993). To construct the average employment durations for the two years, we use the estimated completed duration density up to the 20th week for each year and treat the average hazard rate for weeks 36 through 42 for 1989 as relevant for all ensuing weeks. Thus, the average employment duration is calculated as,

$$E(Emp) = \sum_{H=1}^{20} Hf(H) = [\prod_{H=1}^{20}(1-h(H))](20+\frac{1}{h_{avg}}) \qquad (2)$$

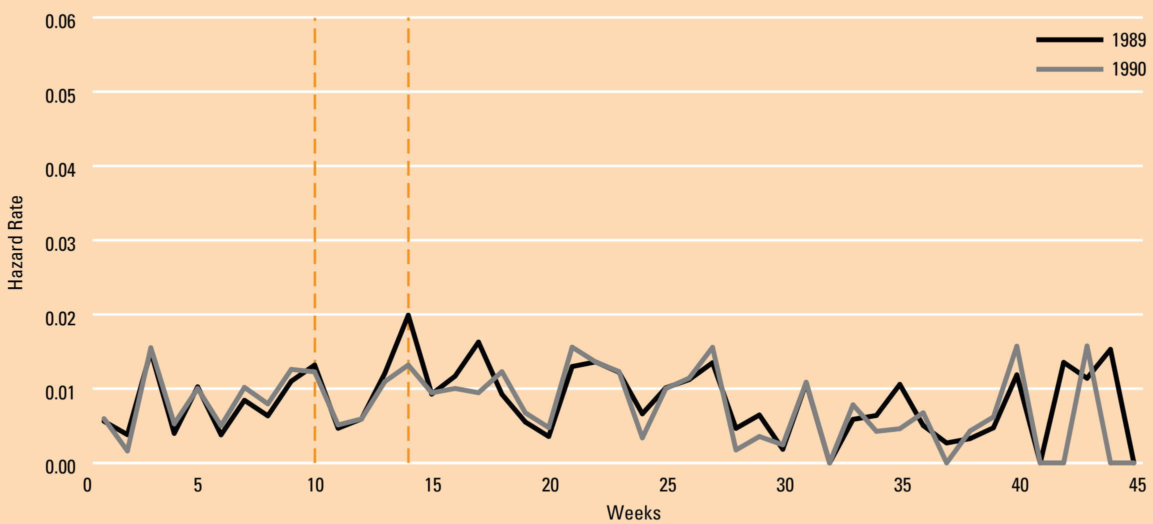
where, *f(H)* is the value of the density function for the duration of employment spells evaluated at *H*, and h_{avg} is the average hazard rate for weeks 21 through 42 in 1989. Because the hazard rates for the two years were found not to be significantly different above 20 weeks, we assume that the entrance requirement has effects up to week 20, and that the two hazards are equivalent and constant thereafter. The calculated average employment durations are 35.65 weeks in 1989 and 37.03 weeks in 1990. The calculated unemployment rate is 12.3 per cent in 1989 and 11.9 per cent in 1990. Taken literally, these calculations imply that the increase in the entrance requirement led to a one-and-a-half week increase in average employment durations and 0.4 per cent point drop in the unemployment rate in the maximum entitlement regions, impacts which are quite substantial.

In Figure 8 we plot the empirical hazard rates for 1989 and 1990 for UI regions in which the entrance requirement was either 13 or 14 weeks in the first 46 weeks of 1989. These regions provide a natural "control group" for comparisons with the experimental group, whose behaviour is shown in Figure 7.[20] Figure 8 provides only a rough benchmark because these regions do not provide as stationary an environment as the maximum entitlement regions since there are changes in other parameters of the UI system between 1989 and 1990 for most of these regions. Nonetheless, the comparison does provide a means of checking whether the patterns observed in Figure 7 can reasonably be attributed to the change in the entrance requirement.

The hazard rates for 1989 and 1990 in Figure 8 are extremely similar. There is no evidence of a sharp change in the pattern such as was observed for the maximum entitlement regions. There is a (statistically insignificant) difference at the 14-week point, but in Figure 8 it is the 1989 value that is higher at this point rather than the 1990 value, as was the case in Figure 7. Overall, Figure 8 provides no evidence to suggest that the effects observed in Figure 7 come from any source—such as a recession—that would cause between-year differences of the

20 There are seven such "control group" regions, all in Ontario and Quebec. Only two regions have VER=14 throughout 1989 and thus using this as a requirement leaves too few observations. In the samples used here, there are 1142 observations in 1989 and 848 in 1990.

Figure 8
Empirical Hazard: VER 13 or 14 Regions, 1989 and 1990



same nature in all regions. One cannot reject the null hypothesis that the differences in the hazard rates for the two years plotted in Figure 8 are jointly zero at the 5-per cent level of significance.[21,22]

As a further check on the results for the maximum entitlement regions, we also calculated the empirical hazard for the maximum entitlement regions for the last four months of 1988 and the last four months of 1989.[23] A comparison between the two years provides evidence on whether individuals appear to have altered their employment behaviour in 1989 in anticipation of the onset of the higher entrance requirement in 1990. Evidence of such anticipation would imply that the measures presented above do not fully capture the effects of the change in the entrance requirement. Plots of the hazard rates for the two years are quite similar and one cannot reject the null hypothesis that the differences in the hazard rates for the two years are jointly zero at the 5 per cent level of significance.[24] Thus, by this measure there is no evidence of anticipatory behaviour that would complicate our conclusions.

21 The test statistic is distributed as $\chi^2(42)$ and takes a value of 7.66.

22 One could form an estimate of the effects of the change in the entrance requirement in the maximum entitlement regions using the *VER=13,14* regions as a benchmark, that is, from $(h_{m,90}(14)-h_{14,90}(14))$ - $(h_{m,89}(14) - h_{14,89}(14))$, where $h_{m,90}(H)$ is the hazard rate at H weeks for the maximum entitlement regions for 1990 and $h_{14,90}(H)$ is the hazard rate at H weeks for the *VER=13,14* regions for 1990. By this measure, the switch of H_{min} to 14 weeks caused a 1.5 percentage point increase in the spike at 14 weeks in maximum entitlement regions, larger than what is reported above using $h_{m,90}(14)$-$h_{m,89}(14)$. Again, though, comparisons using the *VER=13,14* week regions are necessarily approximate for reasons specified above.

23 The employment spells were created as in the main sample but were restricted to start on or after September 1 of the given year and were censored at December 31. There are 1153 such spells in 1988 and 976 in 1989.

24 The test statistic is distributed as $\chi^2(17)$ and takes a value of 2.97.

5. Duration Model Estimates

While empirical hazard rates are useful in examining basic patterns, they do not control for the effects on employment durations of individual, job and labour market characteristics. In our case, controlling for such characteristics is particularly important since the two years being compared are unlikely to be identical.[25] The experiment year, 1990, is a year in which the Canadian economy entered a major recession; while the comparison year, 1989, is a year in which the economy reached a cyclical peak.[26]

While empirical hazard rates are useful in examining basic patterns, they do not control for the effects on employment durations of individual, job and labour market characteristics.

To control for observable characteristics, a duration model is included. The most common such model is the proportional hazards model in which;

$$h_i(H) = h_o(H)\, e^{x_i(H)'\beta} \qquad (3)$$

where $h_i(H)$ is the hazard function for person i, $h_0(H)$ is the "baseline" hazard function common to all individuals, $x_i(H)$ is a vector of observable characteristics which may vary with H[27] and *ß* is a parameter vector. For different values of $x_i(H)'ß$ the hazard function for individual i is shifted proportionally up or down relative to the baseline hazard.

Given our concern with the location and movement of spikes in the hazard, we prefer to use an estimation method for the proportional hazard model which permits direct examination of the baseline hazard over methods which either impose a parametric form for the baseline hazard or eliminate it altogether. Indeed, assuming a simple parametric form for the baseline hazard is inappropriate because theoretical discussion indicates the need for a discrete-continuous framework with different shapes for the continuous sections of the hazard before and after H_{min}. For these reasons, we adopt a specification detailed in Meyer (1990) which allows the estimation of something similar to the empirical hazard function in conjunction with the elements of the *ß* vector. Assuming that $x_i(H)$ does not vary within a given week, Meyer (1990) shows that with a proportional hazards model the probability that a spell lasts at least H^*+1 weeks conditional on it having lasted H^* weeks, is given by:

$$\Pr[H_i \geq H^*+1 / \mathrm{H}i \geq H^*] = \exp[-\exp(x_i(H^*)'\beta + \gamma(H^*))] \qquad (4)$$

where;

$$\gamma(H^*) = 1n[\int_{H^*}^{H^*+1} h_o(u)du] \qquad (5)$$

Note that the hazard function in this model is given by one minus equation (4). The corresponding log likelihood function for a sample of N employment spells is then:

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25 Estimating the relationship between employment duration and observed characteristics also provides useful information on the process determining the incidence of unemployment in the economy.

26 The 1990-1992 recession is generally dated as beginning in April 1990 in Canada.

27 In practice, the continuous variables in $x_i(H)$ are expressed as deviations from their sample means so that the baseline hazard is interpreted as corresponding to an individual with mean values for these variables.

$$\log L = \sum_{i=1}^{N} \{\delta_i \log[1-\exp(-\exp[\gamma(k_i) + x_i(k_i)'\beta])] - \sum_{H=1}^{k_i-1} \exp[\gamma(H) + x_i(H)'\beta]\} \tag{6}$$

where k_i is the observed length of the ith employment spell[28] and δ_i equals one if the spell terminates before being censored and zero if the spell is censored. In maximizing the log likelihood function, we treat the $\gamma(H)$'s as parameters to be estimated. These parameters form the basis for estimates of the value of the baseline hazard at each duration. With them one can examine the hazard for potential spikes and patterns of interest after controlling for the effects of covariates, and without having to impose a specific functional form.

In addition to studying the location and movement of spikes in the baseline hazard, we would like to know whether employment spells ending at exactly H_{min} weeks are observationally different from other spells. Information is needed on the particular nature of spells ending at H_{min} weeks in order to understand the impact of altering the entrance requirement. Thus, we also carry out estimations in which the ß vector is allowed to take different values at 10 weeks in 1989 and 14 weeks in 1990. The log likelihood function in this case is:

$$\log L = \sum_{i=1}^{N} \{\delta_i \log[1-\exp(-\exp[\gamma(k_i) + (1-\theta_{89i}-\theta_{90i}) x_i(k_i)'\beta_0 + \theta_{89i} x_i(k_i)'\beta_1 + \theta_{90i} x_i(k_i)'\beta_2])] - \sum_{H=1}^{k_i-1} \exp[\gamma(H) + (1-\theta_{89i}-\theta_{90i}) x_i(H)'\beta_0 + \theta_{89i}(H)'\beta_1 + \theta_{90i} x(H)'\beta_2]\} \tag{7}$$

where θ_{89i} equals 1 in the 10th week of a spell if the spell occurs in 1989 and zero otherwise; θ_{90i} equals 1 in the 14th week of a spell if the spell occurs in 1990 and zero otherwise; and $ß_0, ß_1, ß_2$ are parameter vectors. Finally, to estimate equations 6 and 7, the data for 1989 and 1990 are pooled and the $\gamma(H)$ parameters are allowed to differ between the two years to account for the effects of the higher entrance requirement in 1990.[29]

Table A1 in the appendix presents definitions and mean sample values for the variables used in the duration models. The covariate that reflects the most direct attempt to control for differences between the two years in economic activity is the regional unemployment rate. This is a three-month moving average for the economic region in which the individual resides, with the last month of the average being the month in which the relevant duration week occurs.[30] This regional unemployment rate is the only time-varying covariate in our estimation.[31] Note that in our sample, a variation in this unemployment rate does not trigger any

28 This equals the actual length of the spell if the spell is terminated during the sample period and equals the observed length of the spell up to the time it is censored if the spell is censored.

29 We also estimated model (6) allowing for unobserved heterogeneity. This was done by multiplying the right -hand side of (3) by eq, where q is a heterogeneity parameter. Individual contributions are integrated over q where q is assumed distributed as a gamma distribution. The results did not differ significantly from the model without the heterogeneity correction, especially with respect to H_{min} effects, and so are not reported.

30 Economic Regions are geographical constructs used by Statistics Canada that are generally smaller than UI regions.

31 However, changes in other covariates within an employment spell are allowed at job change points.

changes in the UI system. In the context of the model of employed- versus unemployed-search, one would expect local unemployment rates to have a negative effect on the hazard out of employment because workers, seeing fewer job opportunities, will face lower levels of expected wealth when non-employed. If the regional unemployment rate is a proxy for the state of product demand, the implicit contract model would have the opposite implication.

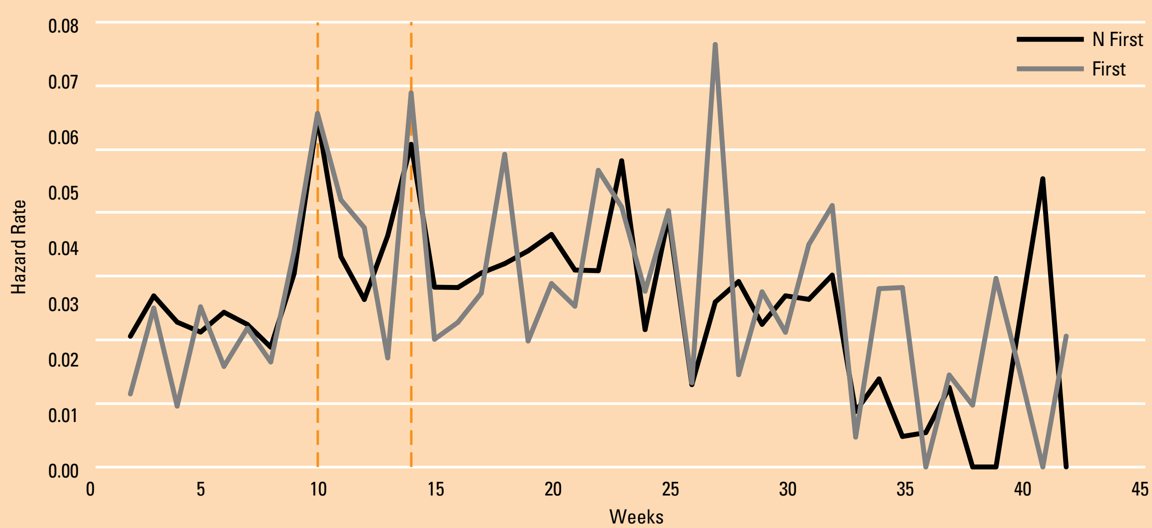
We also control for other covariates in order to account for more complicated reactions to the recession in terms of changes in the composition of the workforce. The covariates we use include several which are assumed to be related to employment stability and the position of the wage offer distribution, such as education, age and white collar status. Increases in all of these variables are expected to increase the opportunity cost of non-employment spells and thus lead to longer employment spells. There are also variables relating to preferences, such as sex, marital status and whether the individual is the head of the household. Typically, one would expect females, single people and those who are not the head of the household to have a lower stake in employment stability and therefore shorter employment spells. Further, we include variables relating to job characteristics. Industry dummy variables are included to capture anticipated differences between seasonal and non-seasonal industries and other industrial differences. Plant size is included on the basis that smaller firms may themselves be less stable and/or may not screen job applicants as effectively as large firms. Union status is included because unions may stabilize the work environment, including lengthening employment spells (or they may organize more stable parts of the work force).

We also use variables that relate to weekly pay. In a search model, workers with higher wage draws, holding constant characteristics that shift the mean of the wage distribution, will be less likely to leave a job. If the recession or other forces changed the wage distribution then controlling for the wage may be important. Hourly pay and usual weekly hours are included separately rather than being combined to form weekly pay because usual hours may enter preferences in their own right, with long hours more likely to induce separations.[32] Under job-matching models, the wage variable will be endogenous. We justify treating the wage as exogenous in spells censored at 46 weeks on the basis that in a typical job very little selection based on match quality will have occurred by this point.

Finally, it is interesting to note that there are spikes in the empirical hazards at 14, 18, 23 and 27 weeks in 1989; and at 23 and 27 weeks in 1990. One explanation for these spikes is that jobs may be likely to start in the first week of a month and end in the last week of a month—or at least that survey respondents may be likely to report job spells that way. In that case, the spikes at 14, 18, 23 and 27 week may simply correspond to a predominance of jobs lasting from 3 to 6 month being reported in this manner. To examine this possibility, in Figure 9

32 Alternatively, seasonal jobs often involve long hours and for that reason one might expect to see a higher hazard for higher hours per week. Specifications were also estimated including only the weekly wage. The weekly wage variable was insignificant and other estimates were essentially unchanged from those reported.

Figure 9
Empirical Hazard: Maximum Entitlement Regions By Start Week, 1989



we present plots of the 1989 empirical hazards for all employment spells that start on the first week of a month and for spells that do not begin in the first week of a month.[33] The difference between the two lines is striking. The "first week start" hazard still shows a spike at 10 weeks but also has large spikes at regular intervals thereafter that are much stronger than the spikes at the same points in Figure 3. In contrast, the "not first week start" hazard shows a smaller spike at 14 weeks and virtually no spike at 18 or 27 weeks. This suggests that the spikes at those points in Figure 3 might in fact be what are called calendar effects. To capture this effect in the duration model we use a variable, *FRSTWK*, which takes a value of 1 for weeks which are the last week of a month if the particular spell being examined started on the first week of a month, and zero otherwise.

Results from estimation of the simple proportional hazards model captured in the likelihood function (6) are presented in Tables 3a and 3b.[34] Table 3a contains the estimates of the coefficients on the covariates for the 1989 and 1990 sample, while Table 3b contains the estimates of the baseline hazard for each year. The covariate estimates indicate a plausible pattern in which individuals with low skills and low labour force attachment are more likely at any point to leave employment.[35] Females and single individuals are more likely to end an employment spell than are men and married individuals. The probability of failing out is significantly lower for those with some post-secondary or university education relative to those who have completed high school and decreases with each increase in education level. The age variable is insignificant at standard signifi-

33 Forty per cent of the spells in our sample begin in the first week of a month.
34 In the following estimations, the first partial derivatives are formed analytically. The estimated standard errors were constructed from the matrix of outer products of the first partial derivatives.
35 Many of our covariate estimates are in line with results in Christofides and McKenna (1993).

Table 3a
Duration Model Estimates: Covariates
Basic Proportional Hazards Model, 1989 and 1990

Variable	Estimates
Female	0.166 (0.048)*
Single	0.073 (0.039)+
Nothead	0.038 (0.040)
Elementry	0.031 (0.045)
Ps	-0.155 (0.055)*
Univ	-0.448 (0.135)*
Prim	0.533 (0.056)*
Mfg	0.098 (0.066)
Food	0.658 (0.078)*
Constr	0.366 (0.059)*
Publc	0.438 (0.080)*
Wcolr	-0.343 (0.053)*
P2099	-0.118 (0.041)*
P100499	-0.241 (0.062)*
Pgt500	-0.426 (0.123)*
Samemp	-0.058 (0.035)+
Uncov	-0.357 (0.041)*
Age	-0.031 (0.020)
Age2	0.000256 (0.000118)*
Hwage	-0.096 (0.038)*
Hrs	0.194 (0.036)*
Ru	0.733 (0.415)+
Frstwk	0.359 (0.054)*

Average Value of Log Likelihood -1.914
Standard errors in parentheses.
*Significantly from 0 at the 5 % level.
+Significantly different from 0 at the 10 % level.

cance levels, but together with the age squared variable indicates that increases in age may reduce the probability of failing out, albeit at a decreasing rate.

The impacts of the job-related variables are also reasonable. Employment spells in the primary and food processing industries (composed mainly of fish processing in the regions we examine) are significantly more likely to be terminated than are those in the base group (the service and transportation sector). Construction workers are also more likely to fail out of employment while manufacturing workers do not show significantly different employment experience compared to the base group. Compared to blue collar workers, white collar workers are significantly less likely to leave employment. As predicted, the larger the plant at which individuals, work the lower the probability of leaving employment. Also as expected, the union variable indicates that those covered by a collective bargaining agreement are less likely to leave employment.

The wage variable coefficient in our estimates is negative, as one would predict from a basic search model, and is significant, which is often not the case in empirical models of employment duration (Devine and Kiefer, 1990). The usual hours-per-week variable has a significantly positive effect on the hazard, which could result from long-hours jobs being unpleasant situations which workers seek

Table 3b
Duration Model Estimates:
Baseline Hazard: Basic Proportional Hazards Model, 1989 and 1990

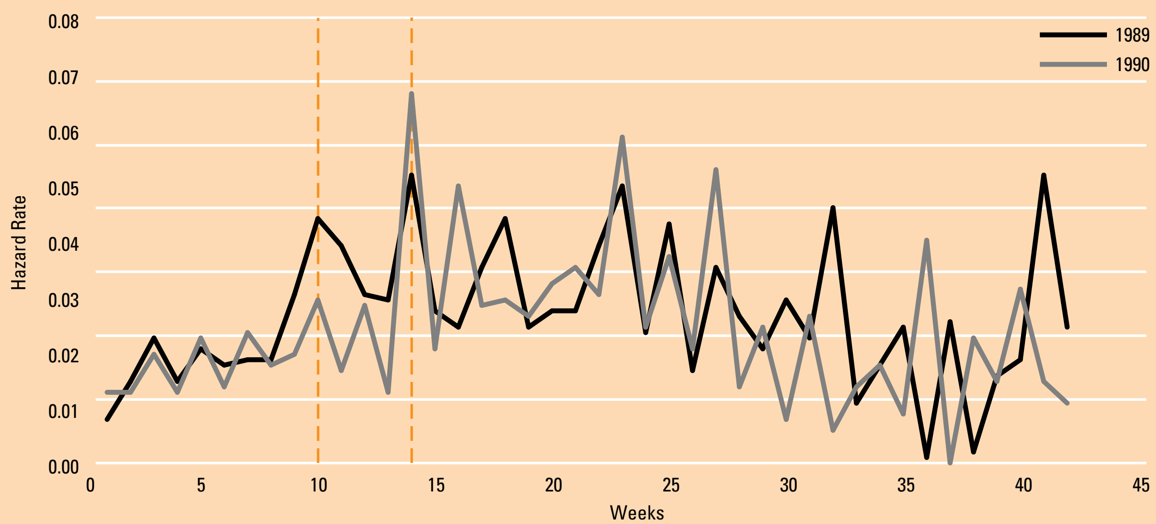
Week	1989	1990
1	0.008 (0.001)	0.013 (0.002)
2	0.015 (0.002)	0.013 (0.002)
3	0.023 (0.003)	0.020 (0.002)
4	0.015 (0.002)	0.013 (0.002)
5	0.021 (0.003)	0.023 (0.003)
6	0.018 (0.003)	0.014 (0.002)
7	0.019 (0.003)	0.024 (0.003)
8	0.019 (0.003)	0.018 (0.003)
9	0.031 (0.004)	0.020 (0.003)
10	0.045 (0.005)	0.030 (0.004)
11	0.040 (0.004)	0.017 (0.003)
12	0.031 (0.004)	0.029 (0.004)
13	0.030 (0.004)	0.013 (0.003)
14	0.053 (0.006)	0.068 (0.007)
15	0.028 (0.004)	0.021 (0.004)
16	0.025 (0.004)	0.051 (0.005)
17	0.036 (0.005)	0.029 (0.005)
18	0.045 (0.005)	0.030 (0.005)
19	0.025 (0.005)	0.027 (0.004)
20	0.028 (0.005)	0.033 (0.003)
21	0.028 (0.004)	0.036 (0.007)
22	0.040 (0.006)	0.031 (0.005)
23	0.051 (0.007)	0.060 (0.008)
24	0.024 (0.005)	0.025 (0.005)
25	0.044 (0.007)	0.038 (0.007)
26	0.017 (0.004)	0.021 (0.005)
27	0.036 (0.007)	0.054 (0.010)
28	0.027 (0.005)	0.014 (0.003)
29	0.021 (0.006)	0.025 (0.007)
30	0.030 (0.008)	0.008 (0.003)
31	0.023 (0.007)	0.027 (0.005)
32	0.047 (0.010)	0.006 (0.003)
33	0.011 (0.004)	0.014 (0.005)
34	0.018 (0.007)	0.018 (0.006)
35	0.025 (0.010)	0.009 (0.006)
36	0.001 (0.001)	0.041 (0.011)
37	0.026 (0.008)	0.000 (0.000)
38	0.002 (0.005)	0.023 (0.010)
39	0.016 (0.008)	0.015 (0.007)
40	0.019 (0.012)	0.032 (0.016)
41	0.053 (0.025)	0.015 (0.011)
42	0.025 (0.009)	0.011 (0.010)

* Standard errors in parentheses.

to leave. The positive sign could also come from the tendency of seasonal jobs to be long hours jobs. The regional unemployment rate has a positive effect, which is not what is predicted in a simple search model. The positive sign is consistent with the implicit contract model in that higher regional unemployment indicates reduced product demand and increased layoffs.

Table 3b contains estimates of the baseline hazard for the two sample years, while Figure 10 shows values of the fitted hazard when covariates are set at their 1989 sample average values.[36] The patterns in Figure 10 are much the same as

Figure 10
Baseline Hazard: Maximum Entitlement Regions, 1989 and 1990



those seen for the empirical hazard. In 1989 there is a spike at 10 weeks, but spikes are also evident at 14, 18 and 23 weeks. In 1990, the hazard at week 10 is not noticeably high relative to adjacent weeks but very large spikes exist at 14 and 23 weeks. The *FRSTWK* variable, which has a positive and significant coefficient in the estimation, has apparently not eliminated the "calendar effect" spikes seen in Figure 9. The difference in the hazard values at 14 weeks for the fitted baseline hazard is 0.015 with a standard error of 0.009, indicating that more individuals than implied by the empirical hazard save their non-employment for H_{min} weeks. Values for the survivor function for key employment durations are presented in Table 4. The pattern is very similar to that seen in Table 2: the survivor functions grow apart between 10 and 14 weeks, such that employment spells are 5 per cent more likely to last to at least 14 weeks in 1990 than in 1989, with 4.3 percentage points of that difference arising from differences in weeks 10 through 13. Again, this evidence suggests a significant moral hazard effect. The main difference between the survivor functions based on the empirical and fitted baseline hazards is the larger difference in survival function values at 21 weeks in the latter case. The baseline hazard does not suggest as strongly that employment spells that formerly would have ended in the 10 to 13-week range ended in the 14 to 20-week range in 1990, but rather that there are differences between the two years at even longer durations.

36 The numbers in this table appear large relative to the empirical hazard estimates in Table 1. However, while the pattern of the baseline hazards is relevant for all employment spells under the proportional hazards model, the height of the hazard will depend on the base group in the covariate definitions. Thus, this hazard corresponds to a married, male blue-collar worker of average age, facing average wages, hours and unemployment rate, and who is not unionized and works in a plant with under 20 workers in the service or transportation sector. Such an individual would be expected to have a relatively high probability of leaving employment.

Table 4
Survivor Function Values Fitted Baseline Hazard

Week	1989	1990	Difference 1990-1989
10	0.847	0.856	0.009 (0.010)
11	0.810	0.831	0.021 (0.011)
12	0.778	0.817	0.039 (0.011)
13	0.754	0.794	0.040 (0.012)
14	0.732	0.784	0.052 (0.012)
15	0.694	0.732	0.038 (0.013)
21	0.577	0.606	0.029 (0.014)

Standard errors in parentheses.

There is little difference between the results of calculations of mean employment durations and unemployment rates using the fitted baseline hazard in the formulas described earlier and those based on the empirical hazard. The mean employment duration is 38.4 weeks in 1989 and 39.9 weeks in 1990, a difference of 1.5 weeks. The calculated unemployment rate is 11.5 per cent in 1989 and 11.1 per cent in 1990, again implying a reduction of 0.4 per cent due to the increase in the entrance requirement. The evidence indicates that increasing the entrance requirement led to significant changes in employment behaviour.

6. Adjustment to the New Entrance Requirement

Perhaps the most striking finding is that spells ending in quits show little or no evidence of entrance requirement effects.

The purpose of this section is to examine the mechanisms by which the apparent adjustments to the change in the entrance requirement take place. We begin by examining the hazard plots for employment spells that end in quits as opposed to layoffs. Figures 11 and 12 provide these plots for 1989 and 1990, respectively.[37] The plots are derived using a competing risks framework where for the layoff "hazard", spells ending in quits are treated as right-censored. In both years, the hazard rates for spells ending in quits are substantially lower than for those ending in layoffs, at least for durations below 30 weeks. Perhaps the most striking finding is that spells ending in quits show little or no evidence of entrance requirement effects. Rather, the spells ending in layoffs display the patterns observed in Figure 7. Indeed, those patterns appear even stronger when layoffs alone are examined. These results are more consistent with the implicit contract model than with the search model, in which job terminations would be employee-initiated. The dramatic difference between the two figures may seem surprising given that under the Canadian UI system during this period, one could quit a job and still qualify for benefits. However, the potential penalty for quitting a job was an extension of up to six weeks of the initial waiting period during which no benefits are paid. This penalty, together with the absence of experience rating, may have been sufficient to induce firms to initiate separations. Indeed, implicit in the "community pressure" model is the notion that the employer is best placed to ensure that as many members as possible qualify for UI. The employer's role is particularly important if some employees would prefer to work more weeks than the minimum entrance requirement.[38]

In Table 5 we present the covariate estimates from the duration model associated with the likelihood function specified in equation 7 above. In this estimation the estimated coefficients on the covariates are permitted to take different values at the 10-week point in 1989 (column 2) and the 14 week point in 1990 (column 3). A comparison of these estimates with the covariate coefficient values relevant for the rest of the range of weeks (column 1) provides information on the characteristics of individuals who worked just enough to qualify for UI in each year. The column 1 estimates of the covariate values for the main range of weeks are similar to the estimates in Table 3a, though effects such as education and age are stronger. The key difference is the change in the unemployment rate effect from positive to significantly negative. These suggest that the earlier positive coefficient may have been capturing a combination of a positive effect at the spike points and negative effects everywhere else. A more positive unemployment rate effect at the entrance requirement is consistent with the implicit contract model where poor demand shocks will cause more firms to bunch their workers at $H = H_{min}$.

37 The distinction between quits and layoffs is based on response to the question as to why the individual separated from the last job in an employment spell.

38 It is interesting to consider this evidence with the results in Katz and Meyer (1990), which show an increase in the hazard out of unemployment at the entitlement exhaustion point for the U.S. even for layoffs with recall.

Figure 11
Empirical Hazard: Quits Vs. Layoffs, 1989

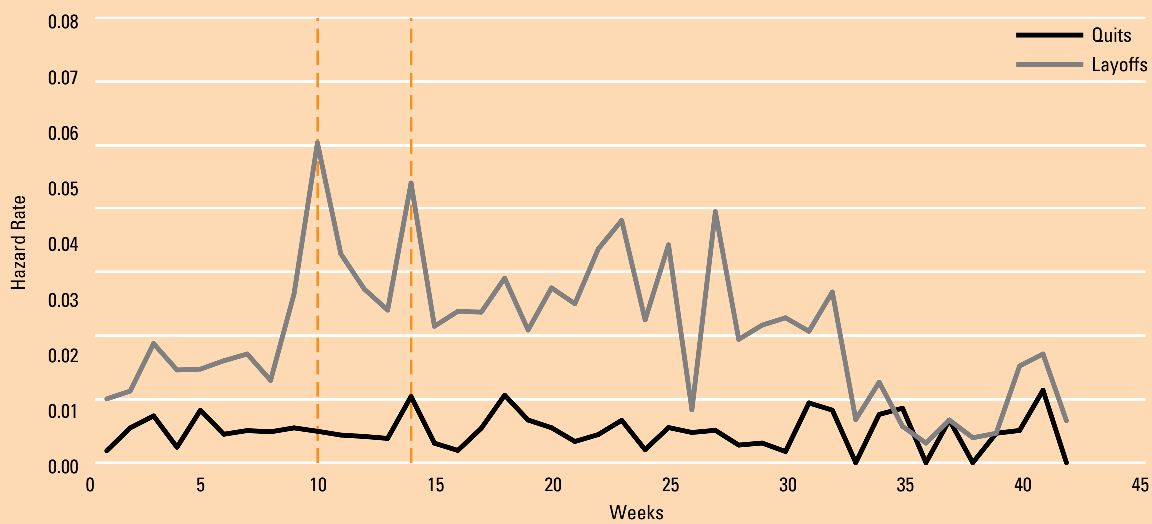
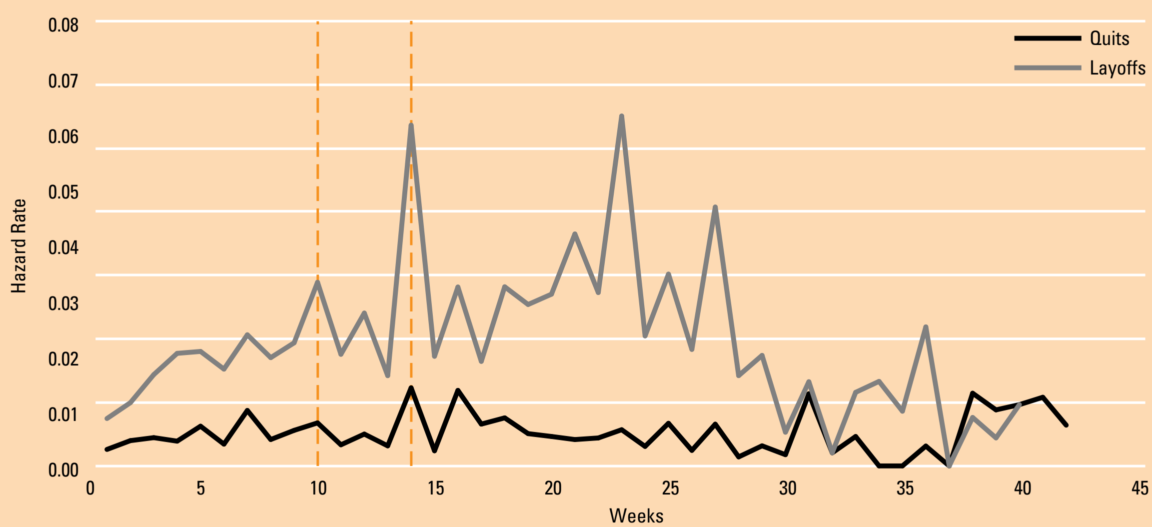


Figure 12
Empirical Hazard: Quits Vs. Layoffs, 1990



In general, the estimated coefficients for the 10-week point in 1989 fit with expectations about individuals who might work just enough to qualify for UI. Individuals with post-secondary or greater education are significantly less likely to leave an employment spell at this point than those who have completed high school. In fact, our sample contained so few university educated individuals sep-

Table 5
Duration Model Estimates: Covariates Spike Interactive Proportional Hazards Model, 1989 and 1990

	Main Range	1989, 10 Week	1990, 14 Week
Variable	Estimates	Estimates	Estimates
Female	0.126 (0.049)*	0.559 (0.333)+	0.076 (0.285)
Single	0.057 (0.039)	0.132 (0.232)	0.226 (0.252)
Nothead	0.067 (0.041)+	-0.276 (0.303)	-0.018 (0.245)
Elem	0.148 (0.047)*	0.155 (0.259)	0.492 (0.268)*
	-0.297 (0.057)*	-0.732 (0.322)*	-1.092 (0.384)
Univ	-0.695 (0.136)*	-	-
Prim	0.411 (0.057)*	1.163 (0.365)*	0.728 (0.360)*
Mfg	0.041 (0.068)	-1.323 (1.054)	0.317 (0.359)
Food	0.588 (0.080)*	1.480 (0.496)*	0.428 (0.555)
Constr	0.326 (0.060)*	0.263 (0.437)	0.727 (0.336)*
Publc	0.337 (0.083)*	1.088 (0.361)*	0.183 (0.570)
Wcolr	-0.408 (0.054)*	0.065 (0.343)	-0.357 (0.343)
P2099	-0.123 (0.042)*	-0.395 (0.278)	-0.101 (0.241)
P100499	-0.179 (0.063)*	-1.885 (0.658)*	0.137 (0.434)
Pgt500	-0.437 (0.126)*	-	-
Samemp	-0.123 (0.036)*	0.245 (0.265)	-0.315 (0.240)
Uncov	-0.295 (0.043)*	-1.058 (0.333)*	-0.179 (0.289)
Age	-0.043 (0.020)*	-0.089 (0.124)	0.100 (0.117)
Age2	0.000169 (0.000121)	0.000383 (0.000758)	0.000321 (0.000691)
Hwage	-0.099 (0.039)*	0.177 (0.230)	-0.376 (0.316)
Hrs	0.217 (0.037)*	-0.160 (0.233)	-0.197 (0.330)
Ru	-1.321 (0.438)*	1.583 (2.899)	2.252 (2.983)
Frstwk	0.567 (0.059)*	0.259 (0.227)	0.428 (0.224)+

Average Value of Log Likelihood -1.914
Standard errors in parentheses.
*Significantly from 0 at the 5 % level.
+Significantly different from 0 at the 10 % level.
For the spike point estimates there are too few university educated individuals to permit estimation of the university coefficient. University educated individuals are placed in the PS category. Similarly Pgt500 is combined with p100499.

arating from employment at this point that we had to combine that group with individuals with lower levels of post-secondary education for purposes of estimation. Perhaps not surprisingly, individuals in primary and food processing industries show a much stronger propensity to fail out at this point than at other weeks. Interestingly, construction does not seem to be an industry that generates employment spells of exactly H_{min} weeks. Firm size is another strong indicator for leaving employment at H_{min} weeks, with individuals working for larger firms being much less likely to fail out at this point. This result could arise if the more personal relations in small firms lead to collusion between the firm and worker in the timing of employment spells.[39] Alternatively, small firms may be more likely to be based in the region and to be more susceptible to community pressure. The *SAMEMP* variable, which takes a value of one for individuals who have worked

39 One way such collusion could take place is through labeling of separations. Workers may find it easier to get a small firm to label a separation as a layoff, thereby avoiding the penalty associated with quitting.

for the current employer in a previous year, is an attempt to capture implicit contracts between workers and firms on recurring work patterns, but the evidence is at best weak that individuals who have worked for a firm before are any more likely to leave a job at 10 weeks. On the other hand, unionized workers are much less likely to fail out at 10 weeks. This could reflect union organization of more stable jobs or more formal work arrangements making firm-worker collusion more difficult. Finally, the regional unemployment rate does not have a significant impact on the likelihood of a separation occurring at 10 weeks.

With a few exceptions, the estimates for coefficients at 14 weeks in 1990 are more poorly defined than their counterparts in columns 1 and 2 in Table 5. These estimates are also closer in both magnitude and sign to the estimates for the main range of weeks than to the estimates for the 10-week point in 1989. Given the relatively small increase in the spike in the hazard at 14 weeks between 1989 and 1990, the estimates in column 3 may not accurately reflect the characteristics of individuals who plan on working exactly H_{min} weeks. As discussed earlier, individuals may not be able to quickly target their employment spells to exactly H_{min} weeks when the entrance requirement is changed. In that case, the estimates in column 2 deserve more attention. Those estimates suggest that individuals working just the minimum number of weeks tend to be disproportionately poorly educated, non-unionized, and in the primary and related processing industries.

We next examine individuals who worked just long enough in 1989 to qualify for UI benefits in order to see how they adjusted to the increase in the entrance requirement. In this analysis we allow for the possibility that some individuals meet the entrance requirement by working several jobs separated by brief non-employed searches. Thus, we create a variable, *TOTEND*, which equals the total duration of job spells not separated by periods of UI receipt for an individual.[40] We select a sample of individuals who have employment spells in 1989 such that $10 \leq TOTEND \leq 13$, and whose last employment spell in 1989 ends before November 18. The latter restriction is imposed to make examination of work experience in the experimental period in 1990 more straightforward. Given our goal of selecting a sample of minimum employment UI users, length biased sampling issues are not a concern. In drawing this sample we include employment spells that begin in 1988, providing they end in 1989. Once this sample is selected, we compare their labour force experience in the first 46 weeks of 1989 and 1990.

Table 6 provides mean characteristics for 1989 and 1990 for the following: all workers in Canada who were not full-time students and were under age 65 in the given year; all workers living in the maximum entitlement regions who were not full-time students and who were under age 65; and the sample of individuals who worked for from 10 to 13 weeks in 1989. The characteristics recorded for the 10 to 13 weeks sample accord well with the pattern for the estimated covariates at the 10-week spike in 1989. These individuals are substantially less educated

40 The UI receipt variable on the LMAS appears to under-report receipt of UI benefits. Given ratios of weeks of UI receipt to weeks of unemployment for Canada as a whole of over 0.9 for the late 1980s (Card and Riddell, 1993), we define a period of UI receipt as any non-employment spell preceded by a job which creates UI eligibility (that is, a job with the minimum required wage, number of hours per week, and length longer, held by an individual who would be covered).

Table 6
10 to 13 Week Workers

	1989			1990		
	Comparison Groups		10-13 Week	Comparison Groups		10-13 Week
Variables	All Canada	Maximum Entitlement Regions	Workers	All Canada	Maximum Entitlement Regions	Workers
% Female	46.1	44.4	46.3	46.1	43.9	46.3
% Married	74.6	75.9	71.3	74.8	76.2	71.3
% Head	55.3	53.0	42.5	55.3	53.4	42.5
Education (%)						
Elementary	9.7	14.6	27.5	7.8	11.8	27.5
High School	45.1	45.6	57.5	46.3	47.4	57.5
Post Sec.	32.4	31.1	13.8	33.1	32.0	13.8
University	12.8	8.7	1.3	12.8	8.8	1.3
Plant Size (%)						
1-20	40.7	47.4	70.6	40.0	47.5	69.4
21-99	30.8	28.1	21.3	31.1	28.5	17.5
100-499	19.7	16.8	5.0	19.0	15.9	11.9
500 +	9.8	7.7	3.1	10.0	8.2	1.3
% Union	40.4	41.3	20.6	40.5	42.0	23.1
Industry (%)						
Agriculture	2.0	2.3	5.6	2.3	2.4	3.1
Forests	1.3	3.3	6.3	1.1	2.9	5.6
Fishing	0.9	2.7	21.3	0.8	2.2	13.1
Food & Bev.	3.9	6.8	14.4	3.8	7.2	15.6
Other Manu	16.4	13.4	4.3	15.6	12.9	6.8
Constr.	6.7	8.2	8.8	6.4	7.9	11.9
Transport.	8.2	7.2	3.1	8.6	8.4	5.0
Sales	16.4	15.1	6.3	16.1	14.5	11.9
Finance	4.7	3.3	1.3	4.8	2.9	0.6
Edn. & Health	17.4	16.7	5.6	18.1	17.7	5.6
Service	13.8	12.6	17.5	13.7	12.4	13.8
Publc. Admin.	8.4	8.5	5.6	8.3	8.3	6.2
Hourly Pay						
Mean	12.5	11.5	8.5	13.1	12.2	8.9
Std Dev	6.9	6.6	3.7	6.9	6.6	4.8
Weekly Hours						
Mean	38.8	40.8	47.7	38.5	40.6	46.1
Std Dev	13.6	14.7	13.6	13.1	14.4	14.8
% Part Time	16.5	14.9	3.8	16.4	15.0	7.5
% UI Eligibl	93.4	94.7	98.8	93.9	95.3	98.1
% Rec'd UI	22.6	38.5	94.4	22.7	37.2	93.8
Number of Jobs						
1	79.3	79.7	71.8	84.3	83.9	72.5
2	16.1	16.2	25.2	12.6	12.6	21.9
3+	4.6	4.2	3.1	3.2	3.5	5.6
Job Leavers: Reasons for Leaving						
Quit	50.6	32.1	6.9	49.4	34.2	13.2
Seasonal Job	22.4	37.0	58.1	22.1	35.4	53.7
Other Layoff	26.9	30.9	34.3	28.4	30.4	32.4
Sample Size	31,975	6,379	160	32,287	6,550	160

Notes: All job related variables refer to the last job held in the sample period. Union status refers to individuals covered by a collective bargaining agreement. The other Manufacturing industry category includes mining. The Transportation industry category includes communications and utilities. The Weekly Hours variable is usual hour per week on the job. A worker is classified as UI Eligible if he or she worked more than the minimum required number of hours per week and earned more than the minimum insurable weekly earnings in his/her last job in the sample period.

than the average worker in their regions and, especially, when compared to Canadian workers overall. They are much more likely to work for establishments of under 20 people and are about half as likely to be unionized as other Canadians. They are much more concentrated in seasonal industries with 33.2 per cent of them being in agriculture, forestry or fishing compared to 9.3 per cent for the maximum entitlement regions as a whole and 4.2 per cent for Canada. They are also disproportionately found in the food and beverage industry and the service industry and are significantly under-represented in other manufacturing, sales, and education and health. The seasonal nature of their jobs is reinforced by their reason for leaving the last job in the year (conditional on having left a job in the sample year).

Only 6.9 per cent of the 10 to 13-weeks workers quit their job as compared to 32.1 per cent for the maximum entitlement regions overall and 50.6 per cent for Canada as a whole. This accords with the evidence presented earlier on the quits vs. layoffs breakdown and with the idea that if one is going to work just enough weeks to qualify for UI it is best to do so in a job from which a layoff is expected in order to avoid the penalties associated with quitting. Fully 58 per cent of the 10 to 13-weeks sample gave as their reason for leaving their last job that the job had ended because of its seasonal nature. This compares to 22 per cent for Canada as a whole. The 10 to 13-weeks workers also take much lower paying jobs and tend to work much longer hours when they do work. The long hours are consistent with the structure of UI benefits which are calculated based on weekly earnings. Workers trying to generate high UI benefits—especially those with little education facing a low hourly market wage—can best do so by working long hours. Finally, the 10 to13-weeks workers are much more likely to have two jobs in the year than other workers. Overall, one forms a picture of individuals with low education who work for low pay and long hours in predominantly seasonal industries.

How did these workers adjust to the increase in the entrance requirement? First, almost all of them continue to work in 1990; only 3 out of 170 individuals who worked between 10 and 13 weeks in 1989 did not work in 1990.[41] The great majority of them manage to increase their weeks of work beyond the new entrance requirement. Of those who do not have a job spell ongoing at the end of the year (85 per cent of the original sample), 9.7 per cent work between 1 and 9 weeks in 1990, 18.4 per cent work between 10 and 13 weeks as they had the year before, 18.4 per cent work exactly 14 weeks, 36.9 per cent work between 15 and 20 weeks, and the remaining 16.6 per cent work 21 weeks or more. These results accord with the evidence from the hazard rate analysis that few individuals who had worked from 10 to 13 weeks in 1989 continued to do so and that many moved not just to jobs that ended exactly at the new H_{min} but in the 15 to 20-week range as well.

41 The characteristic means reported in Table 6 are for the 167 individuals who are observed to work in both years.

The 1990 job characteristics for the sample of individuals who worked for 10 to 13 weeks in 1989 indicate offsetting patterns of response. On the one hand, these workers show dramatic movements out of the primary and service sectors. The proportion in fishing alone falls from 21 per cent in 1989 to 12.6 per cent in 1990. These movements are offset by increases in a variety of industries including Food and Beverage and Other Manufacturing. Further, there is a slight increase in their likelihood of being unionized and a fall from 58.1 per cent in 1989 to 51.8 per cent in 1990 in the percentage who say their last job terminated because of its seasonal nature. Thus, there appears to be a movement towards more stable job patterns. On the other hand, the probability that these individuals work part time rises. The percentage who leave jobs through quitting also increases with over 5 per cent of the sample claiming they quit because of low pay in 1990, an answer that is not given by anyone in 1989. The percentage holding two jobs in the year decreases but this is offset both by increases in the percentage holding exactly one job and the percentage holding three or more. Similarly, a lower proportion of individuals work for establishments of under 100 workers but the proportion working for the largest category of establishments also drops. Thus, there is evidence that some workers ended up in less stable work situations. Some of the hours of work changes could be associated with the onset of the recession; however, the decline in hours of work in these regions is much larger than that experienced in Canada as a whole.

In an attempt to understand the role of firm-worker interaction in adjusting to the new requirement, we also selected a subsample of the 10-13 weeks workers who worked for exactly one employer in 1989 and who returned to that employer in 1990. Because there are only 68 such individuals, we do not report complete tabulations; however, the breakdown of weeks worked in 1990 is instructive. Of the portion of this subsample who do not have a right-censored spell at the end of 1990 (48 individuals), 4 per cent worked less than 10 weeks in 1990 for the employer they worked for in 1989, 27 per cent worked for between 10 and 13 weeks in 1990, 19 per cent worked 14 weeks, 50 per cent worked for 15 to 20 weeks, and the rest worked for 21 or more weeks in 1990. Thus, the results are mixed. Returning to the same employer made it more likely that an arrange-

Table 7
Aggregate Labour Market Behavior in the Maximum Entitlement Regions and Canada as a Whole, 1989 and 1990

	Maximum Entitlement Regions			Canada		
	1989	1990	% change	1989	1990	% change
Employment (000s)	861.0	869.0	+0.9	12,486.0	12,572.0	+0.7
Employment/population ratio	53.0	53.4	+0.8	62.0	61.5	-0.8
Unemployment (000s)	125.0	124.0	-0.8	1,018.0	1,109.0	+8.9
Unemployment rate	12.7	12.5	-1.6	7.5	8.1	+8.0
Labour force participation rat	60.6	61.0	+0.7	67.0	67.0	0.0

Notes: 1. Data for the maximum entitlement regions are based on Statistics Canada data for economic regions and an algorithm for mapping from economic regions to UI regions. This algorithm involves some approximations.
2. Sources for the data on economic regions are CANSIM monthly series and Labour Force Annual Averages 1990 and 1991 (Ottawa: Statistics Canada) and calculations by the authors.

ment could be made yielding 14 to 20 weeks or work (though no more likely than for the sample as a whole that exactly 14 weeks would be worked) but also more likely that one would only be offered employment for the same 10 to 13-week spell as in 1989.

Overall, an examination of the subsample of individuals who worked just enough to qualify for benefits in 1989 reveals that a significant number managed to increase their weeks of work to the 14 to 20-week range in 1990. They appear to have done so largely by shifting toward more stable industries. Some of the adjustment also took the form of a decline in hours of work per week. Finally, it is noteworthy that virtually the same proportion of the sample received UI payments in 1990 as in 1989.

Our final piece of evidence is presented in Table 7. It compares the aggregate labour market behaviour in the maximum entitlement regions to that for Canada as a whole. Between 1989 and 1990, the employment/population ratio increased in the maximum entitlement regions but declined in Canada as a whole. The increase in total employment was also somewhat greater in the maximum entitlement regions. The behaviour of unemployment is dramatically different. While unemployment increased by about nine percent in Canada as a whole, unemployment declined by almost one percent in the maximum entitlement regions. The 0.2 per cent decline in the unemployment rate in these regions (versus an increase of 0.6 per cent for Canada as a whole) is consistent with the estimates of the impact of the change in the entrance requirement made earlier in the paper. Generally speaking, these aggregate statistics, while admittedly crude, do support the conclusions of the micro-analysis that the change in the entrance requirement resulted in increased employment, reduced unemployment and little (if any) change in labour force participation.

7. Conclusion

In this paper, we examined the impact of the UI entrance requirement on the distribution of employment durations by studying an accidental experiment that took place in Canada's UI system in 1990. In that year, as a result of a dispute between Canada's Senate and House of Commons over issues other than the entrance requirement, the variable entrance requirement was suspended. In some regions this meant an increase from 10 to 14 in the number of weeks of employment needed to qualify for UI. We focused our attention on the differences between 1989 and 1990 in the hazard rate out of employment for the "maximum entitlement regions", those in which there was an increase in the entrance requirement from 10 to 14 weeks and no changes in other features of the UI system.

We look for evidence of the effects of the entrance requirement change using both simple empirical hazard (Kaplan-Meier) estimates as well as duration model estimates that control for the influences of covariates. In particular, we attempt to identify the effects of the entrance requirement by comparing the hazard rates out of employment for the "experimental" period (most of 1990) with those for the equivalent period in the previous year. Both with and without controls for covariate effects, we find that there is a substantial decline in the hazard in the 10 to 13-week range between 1989 and 1990, with a particularly noticeable decline at 10 weeks. These changes are offset by an increase in the spike in the hazard at 14 weeks and by general increases in the hazard between 14 and 20 weeks. Plots of the empirical hazard for "control group" regions—those in which the entrance requirement changed very little or not at all between the two years—show no significant difference between 1989 and 1990. This evidence strongly suggests that the patterns observed in the maximum entitlement regions can indeed be attributed to the increase in the entrance requirement in those regions.

As further evidence, we plotted the empirical hazards for spells ending in quits versus layoffs for the two years. There appears to be no entrance requirement effect on spells that end with quits but there is a strong effect for spells ending in layoffs. We also examined a sample of individuals who worked for between 10 and 13 weeks in employment spells ending in 1989, a group of individuals who may have been targeting their weeks of employment to just qualify for benefits. These individuals were found to be disproportionately less well-educated, non-unionized workers in small establishments in the primary and service sector. They also tended to be paid lower wages but to work longer hours per week when they did work. Following these same individuals into 1990, one finds that most are able to increase their weeks worked sufficiently to qualify for UI and that this was done at least in part by moving to jobs in different sectors.

In general, these results appear to be more consistent with an implicit contract model (or at least a model that reflects direct worker-firm interactions) than standard supply-side models that are often used to examine UI effects. The fact that the entrance requirement effects showed up for layoffs but not for quits fits this pattern. One might argue that with a UI penalty for quitting and no cost to a firm relabelling a quit as a layoff, such a pattern could arise if workers were to make

Canada's UI system provides strong incentives for individuals in high unemployment regions to work the minimum number of weeks required to qualify for UI.

small side payments to firms to relabel quits as layoffs. In that case, though, one should not expect to see any recorded quits in Canada, something which Table 6 illustrates is far from being the case.

The pattern of adjustment to the higher entrance requirement also accords more with contracting models. Workers who formerly had employment spells of between 10 and 13 weeks were generally able to find enough work to meet the new requirements. However, under the new requirements, these workers tended to record employment spells which ended at longer durations rather than at exactly 14 weeks. This is contrary to labour supply model predictions that these workers will work exactly 14 weeks in the new regime. Under a contracting model, on the other hand, increasing the entrance requirement may lead to layoffs of workers from firms that formerly provided exactly the entrance requirement weeks of work. The laid-off workers may then be forced to find work in industries where technological constraints prevent employment from being set exactly at the entrance requirement. The fact that one main response of 10-13 weeks workers to the increased requirement was to switch industries fits this explanation. That these switches were generally out of seasonal industries toward more stable employment pattern industries is indirect evidence of the subsidization provided by Canada's UI system to less stable industries.

Finally, that higher unemployment rates tended to raise the hazard out of employment more at the entrance requirement than at other spell lengths, while a weak result in our paper, fits with the community-pressure version of the contract model. In poor demand states of the world, firms would be under pressure not to let anyone work extra weeks, employing more workers at exactly the entrance requirement to ensure that everyone obtains UI. This would cause a rise in the hazard at the entrance requirement with a rise in unemployment, a result that is more difficult to reconcile with search or static labour supply models. Overall, models that directly incorporate firm actions seem better able to predict the response to this particular UI parameter. At the very least, supply side models tend to attribute too much of the responses to UI parameters to the actions of workers alone.

Canada's UI system provides strong incentives for individuals in high unemployment regions to work the minimum number of weeks required to qualify for UI. This situation arises from the combination of a short entrance requirement and long maximum-benefit duration. The evidence presented in this paper indicates that workers and firms in these regions adjust their employment patterns and practices in response to these incentives.

Appendix A: Sample Construction

In this appendix, we describe the construction of the employment duration variable used in our analysis as well as detailing sample exclusions and providing the descriptions and means for covariates. To understand our construction of employment spells from job spells, consider an individual with two jobs, *A* and *B*, in a year and call the start weeks for jobs *A* and *B*, $STRT_A$ and $STRT_B$, respectively, and the stop weeks $STOP_A$ and $STOP_B$, respectively. If $STRT_B \leq STOP_A+1$ then jobs *A* and *B* are combined into one job with start week $STRT_A$ and stop week $STOP_B$. If $STAT_B$ is greater than $STAT_A$ and $STOP_B$ is less than $STOP_A$ then job *B* is termed a moonlighting job and is dropped. Moonlighting jobs are dropped because these spells do not add weeks that contribute to UI eligibility. Covariates

Table A.1
Variable Descriptions and Means

Variable Name	Description	Mean Value 1989	Mean Value 1990
Female	=1 if female, 0 otherwise	0.393	0.395
Single	=1 if not married, 0 otherwise	0.342	0.348
Nothead	=1 if not head of household, 0 otherwise	0.527	0.524
Elem	=1 if not completed high school, 0 otherwise	0.197	0.165
Ps	=1 if some or completed post-secondary, 0 otherwise	0.250	0.234
Univ	=1 if completed university, 0 otherwise	0.038	0.035
Prim	=1 if industry is agriculture, forestry or fishing, 0 otherwise	0.187	0.160
Mfg	=1 if manufacturing, other than food proc., plus mining, 0 otherwise	0.066	0.080
Food	=1 if food and beverage industry, 0 otherwise	0.086	0.094
Constr	=1 if construction, 0 otherwise	0.158	0.161
Publc	=1 if government employee, 0 otherwise	0.063	0.058
Wcolr	=1 if occupation is manager, professional, sales or service 0 otherwise	0.422	0.418
F2099	=1 if firm employs 20 to 99 workers at this location, 0 otherwise	0.256	0.276
F100499	=1 if firm employs 100 to 499 workers at this location, 0 otherwise	0.115	0.115
Fgt500	=1 if firm employs over 500 workers at this location, 0 otherwise	0.029	0.031
Samemp	=1 if worked for this firm in previous year, 0 otherwise	0.403	0.441
Union	=1 if covered by a collective bargaining agreement on the job, 0 otherwise	0.267	0.258
Wage	usual hourly wage	10.00	10.13
Whrs	usual weekly hours	44.30	43.33
Age	age of individual at interview date	34.04	34.43
Ru	unemployment rate for UI region, expressed as a 3-month moving average		

related to the job, such as plant size, could differ between job *A* and job *B* and the probability of leaving employment in any week should be related to the covariates relevant to the job held in that week, rather than some average over the employment spell. In forming the employment spell from *A* and *B*, we keep the job related covariates from the two jobs and the start week for job *B*. In the estimation, these covariates are treated as time-varying covariates that switch from their job *A* values to their job *B* values at the point at which job *B* starts.

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