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EMPLOYMENT SPELLS AND UNEMPLOYMENT INSURANCE ELIGIBILITY REQUIREMENTS

by

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Abstract

In this paper we examine whether the requirements which workers must satisfy to qualify for UI benefits in any succeeding period of joblessness affect the duration of employment spells. This behavioral consequence of a UI system has been neglected in empirical research, which has instead focused on the effects of UI parameters on the actions of the unemployed. The effect is identified by a unique change in the eligibility requirements of the Canadian UI system in 1990, which is orthogonal to changes in the economic environment. We find a significant increase in the employment hazard in the week that an individual satisfies the eligibility requirement in many regions of the country. In the spirit of Feldstein's (1976) study of temporary layoffs, the results provide new evidence of the impact of UI system parameters on the actions of employers and workers.

JEL: J2

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The design of unemployment insurance (UI) programs has important implications for individuals' labor supply decisions. The effects of program parameters on the actions of the *unemployed* is well documented in the literature (e.g., Solon 1985, Ham and Rea 1987, Meyer 1990). Another part of the story, which has received much less attention, is how UI regulations affect the *employment* relationship.

Most UI systems require individuals to work some minimum period of time to qualify for benefits. For example, in the U.S. most states demand that individuals work a certain number of weeks or calender quarters, and/or earn a specified amount of wages, over the "base period" prior to a claim. In Canada, claimants must accumulate a minimum number of insured weeks of employment during a "qualifying period", which is the 52 week period prior to the claim. These eligibility restrictions are intended to measure workers' attachment to the labor market, and thus target UI benefits to those in the labor force. We examine how these requirements might affect individuals' decisions to leave employment, or alternatively the length of employment contracts which employers' make available.

The influence of UI system parameters on the employment relationship has been the subject of some previous research. For example, Brechling (1981) and Topel (1984) investigate Feldstein's (1976) conjecture that the imperfect experience rating of UI employer premiums increases the incidence of temporary layoffs.[1] We also note that UI parameters are sometimes offered as an explanation of differences in unemployment rates across countries.[2] The eligibility requirements that are examined here have recently played a role in Canada–U.S. comparisons.[3]

The application is to the parameters of the Canadian UI system. Until recently,

[1] Hamermesh (1979) examines the entitlement effect of UI benefits, while Deere (1991) examines the effect of experience rating on inter-industry employment. Solon (1984) examines the effects of benefit disqualification for those who leave their jobs voluntarily on job quitting behavior.

[2] This hypothesis has been offered primarily in comparisons of American and European unemployment. See for example Burda (1988).

[3] See Card and Riddell (1992).

individuals in Canada had to accumulate between 10 and 14 "insured weeks" of employment in the year preceding a UI claim to qualify for benefits.[4] The precise number of weeks, called the Variable Entrance Requirement (VER), varied across 48 economic regions within the country according to the local unemployment rate.[5] At the beginning of 1990, the legislation enabling this feature of the system expired, and new legislation was not passed until the end of the year. This lead to an 11 month period in which the VER's were temporarily suspended. As a result, the entrance requirement was set to 14 weeks in all regions of the country, regardless of economic conditions, for the duration of the suspension. We examine the extent to which this increase in the entrance requirement (for most regions), which is orthogonal to changes in the economic environment, affected the weekly exit rate from employment in the initial stages of a spell.[6] Our empirical method is to look for "spikes" in the employment hazard in the week that individuals qualify for UI benefits.[7] We use the variation in the entrance requirement resulting from the suspension of VER's to identify this behavioral effect. The information in Figure 1 leads one to suspect that there is some behavioral effect. For the province of Newfoundland, we graph the distribution of UI claimants by the number of insured weeks of employment accumu-

[4] These rules were modified in November 1990 with the passage of Bill C-21.

[5] Different requirements applied to individuals who had claims in the past 52 weeks ("repeaters"), and to people who had not worked in the previous two years ("new entrants"). In general, the minimum employment period for these groups was longer.

[6] Christofides and McKenna (1993) examine the correlation between provincial average VER's and the employment hazard in the period prior to 1988. Our analysis differs from their's in that we are able to identify the actual VER's individuals face at the regional level (48 regions versus 10 provinces), and we use the "natural experiment" provided by the suspension of the VER's in 1990 to identify their effect on the employment hazard. Regional identifiers are not available in the public use versions of the *Labor Market Activity Survey.* Green and Riddell (1993) examine the effects of the suspension of the VER's in 1990 for a subset of 9 regions. These are areas of chronic unemployment, and therefore their results may not generalize to other regions of the country. Finally, a more dramatic change in entrance requirements took place in 1971 when the required weeks of work fell from 30 to 8. Data on employment duration were not available at that time.

[7] This analysis is in the same spirit as Munts (1970). He identifies spikes in the earnings distribution of workers on partial UI benefits which correspond to points in the benefit schedule where the marginal tax rate on earnings changes discretely. See also Moffitt's (1985) and Katz and Meyer's (1990) studies of spikes in the unemployment hazard which are correlated with the potential duration of UI benefits.

lated prior to the claim, in 1989 and 1990. In 1989 workers must have accumulated at least 10 insured weeks to qualify for UI benefits, while in 1990 the eligibility requirement was 14 weeks. We observe large spikes in the distribution of claimants in each year at the point of the minimum requirement. Also, the effects of the change in the minimum requirement between the two years is reflected in the large spike at 14 weeks in 1990, which has no precedence in the previous year.[8] We seek similar effects among the larger sample of workers who completed employment spells in these two years.

Our results indicate a significant increase in the employment hazard in the week in which individuals qualify for UI benefits. Because the employment hazard displays spikes for a number of reasons beyond UI eligibility, the estimates exhibit some sensitivity to the specification of the baseline hazard. For an economically diverse subset of provinces, however, the results are robust to such experimentation. Furthermore, the estimates display little sensitivity to the specification of observed and unobserved heterogeneity. We conclude, therefore, that employers and employees in many parts of the country recognize and respond to the subsidy to short-term employment, this feature of the UI system represents.

In the next section we provide a brief theoretical motivation for the effects of UI eligibility requirements on employment tenure. The following section contains a description of the data. We pay special attention to the methods used to identify the week in which individuals become eligible for UI benefits. In the next two sections we outline the empirical framework for the analysis and present the main results of the study. In the final section we provide an interpretation of the results and some suggestions for future research.

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[8] The data is for claims initiated in the period January through November in each year. Since the eligibility requirement is 14 weeks in 1990, we do not observe potential claimants who accumulated fewer than 14 weeks in this year.

Theoretical Motivation

A static model of labor supply provides strong intuition for the effects of UI eligibility requirements on employment tenure. The suitability of this type of model is enhanced by the fact that we expect people with loose attachment to the labor force to be most affected by this parameter of the UI system.

There are two dimensions of the UI system which must be considered in our analysis. The first is the number of weeks one must work to qualify for some "initial" benefit entitlement, which "usually" varies by region. The second is how this entitlement changes with additional weeks of work beyond the regional minimum. An individual qualifies for an initial entitlement of one week of benefits for each week of insured employment (up to a maximum of 25), given they have satisfied the VER. Further weeks of benefit entitlement, "labor force extended benefits", up to a total of 13 are awarded for every 2 weeks of work beyond 25. Finally, there is an additional entitlement, called "regionally extended benefits", which is determined by the local unemployment rate and is available to all who satisfy the VER.[9] The maximum benefit entitlement from these various "phases" is 50 weeks in a 52 week benefit period. The minimum employment requirement will provide close to the maximum benefit entitlement in high unemployment regions.

Rea (1977) presents a model of labor supply which captures the effects of UI on employment and unemployment, assuming that unemployment is voluntary and agents have limited time horizons.[10] In Figures 2 and 3, the worker allocates the weeks in the period between work and leisure along the budget constraint ABCD to maximize utility. Eligibility for UI after a number of weeks of employment leads to a convex kink in the budget constraint, reflecting the increase in income as a result

[9] This entitlement ranges from 0 weeks of benefits in regions where the unemployment rate is 4 % or lower, to 32 weeks in regions where the unemployment rate exceeds 11.5 %.

[10] The specific conclusions of this model are sensitive to the length of the decision period, two years in the following example, but the general insights are valid for other decision periods.

of benefit receipt. In Figure 2, the incentives for a worker in "high" unemployment region (i.e., unemployment rate > 11.5 per cent) are outlined.[11] After 10 weeks work and a 2 week waiting period, the worker is eligible for 10 weeks of initial benefits and 32 weeks of regionally extended benefits, at 60% of his or her weekly wage.[12] Therefore, the return for the 10th week of work is $(W + (42 \cdot 0.6 \cdot W))$ (the shift in the budget constraint from B to E). Additional work to a total of 18 weeks yields further entitlement. At 18 weeks the worker qualifies for 50 weeks of benefits which is the maximum. This leads to the second kink in the budget constraint at G. Finally, a worker who chooses to work more than 52 weeks would face a 60 per cent tax rate on earnings, because each additional week of work would eliminate 1 week of benefits. This final kink at H is a result of the 2 year period of the analysis.

The impact of an increase in the eligibility requirement to 14 weeks in this region is shown by shift in the initial kink from E to F. This shift provides a strong incentive for workers with between 10 and 13 weeks employment to gain the additional weeks necessary to qualify for UI. For example in moving from 10 to 14 weeks, the worker receives her weekly age, $(4 \cdot W)$, plus the value of the UI benefits, $(0.6 \cdot 46 \cdot W)$, a figure over thirty one times the actual weekly earnings.

The analysis of a low unemployment region (unemployment rate ≤ 6.0 per cent) is presented in Figure 3. After 14 weeks of employment, the worker is eligible for 14 weeks of initial benefits. This leads to an initial kink in the budget constraint at 14 weeks (the shift from B to E). Additional work to a total of 51 weeks brings a longer potential duration of benefits. Note that between 25 and 51 weeks of work, benefits are accumulated at a slower rate under the labor force extended benefits

[11] The eligibility requirements for the period 1980 through 1989 were: 14 weeks if the unemployment rate ≤ 6.0; 13 weeks if $6.0 <$ the unemployment rate ≤ 7.0; 12 weeks if $7.0 <$ the unemployment rate ≤ 8.0; 11 weeks if $8.0 <$ the unemployment rate ≤ 9.0; and 10 weeks if the unemployment rate > 9.0.

[12] This is a simplification. Weekly UI benefits are capped at roughly 60 per cent of the economy average weekly wage. Also, UI benefits are taxable in Canada.

phase (segment EF versus FG). After 51 weeks of work benefit entitlement equals 38 weeks, the maximum possible in this region given economic conditions. At 64 weeks, further work eliminates weeks of potential benefit receipt due to the 2 year period of analysis. Finally, the increase in eligibility requirements in 1990 has no effect in this region as the 14 week requirement is already in effect.

The kinks in the budget constraints in Figures 2 and 3 suggest we should observe a "bunching" of employment spells at the point where eligibility is attained; that is, the expectation of a spike in the employment hazard in this week. Given these incentive effects are derived in a static model, we might expect to find the strongest empirical evidence of their existence in areas of high employment instability. Using the provincial unemployment rate as a proxy for such instability identifies the Atlantic provinces and certain regions of Quebec and British Columbia as areas of particular importance. To the extent that UI induces participation from those who would otherwise be out of the labor force, however, the spikes in the employment hazard could be evident in all parts of the country.

Many dynamic models of employment tenure are based on the job-training or job-matching hypotheses. In the job- training or human capital model workers accumulate human capital with tenure. Wages are likely to increase with tenure, whether training is specific or general. In a simple version of the model quit rates fall with tenure because quasi rents from investment in specific human capital are being paid. The job matching model suggests that workers and firms are engaged in a process of matching worker qualifications with the job. Separations occur when the match is found to be poor or when better matches appear. Mortensen (1988) has combined the two hypotheses into one comprehensive model of employment duration. Predictions concerning transitions from employment to unemployment differ according to whether the job-training or job-matching model is assumed. The job-training hypothesis predicts that the hazard increases with tenure given the current wage. The job matching hypothesis may lead to the opposite result. In the limit, the employment hazard decreases with tenure given a current wage greater than the value of unemployment. These conclusions are based on the assumption of constant unemployment insurance benefits. The increase in benefits in the week of eligibility will raise the probability of an employment separation under either approach.

The Data

Our analysis is based on data from the *Labor Market Activity Survey* (LMAS). This is a panel data set covering a probability sample of individuals over the period 1988 through 1990. In the initial year of the survey, information is collected through a supplement to the regular *Labor Force Survey* (LFS), which is comparable to the U.S. *Current Population Survey.* In subsequent years, respondents are re-contacted and interviewed about their labor market activities over the intervening period of time. This survey provides a wealth of information on individuals' socio-economic characteristics. Important for our purposes here is that it also contains weekly information on the duration of their periods of employment. The sample design provides us with a stratified sample of all employment spells which were initiated in Canada during the period 1988 through 1990. Our sample includes all individuals between the ages of 16 and 64 (excluding the self-employed) who initiated employment spells in the periods January through June, 1989 and 1990.[13]

We conduct most of our analysis at the provincial level. Due to a high degree of decentralization, Canada's provinces possess a fair bit of discretion over economic matters and are the largest employers in each province. Also, income supplement programmes, which are an alternative source of support for marginal labor force par-

[13] We exclude respondents with no provincial identifier, and those for whom the regional and provincial identifiers did not match at the time the spell commenced. These exclusion restrictions lead us to discard 159 spells which represented 1.4 per cent of the total sample. See the discussion below.

ticipants, are a provincial and municipal concern. Finally, many job creation programs, which are potentially the source of short-term jobs that provide UI eligibility, are initiated at the provincial level.[14]

We should be careful to note that the unit of analysis in the study is the employment spell, not the length of time in a given job. By this we mean any continuous period of employment which is bordered by time spent unemployed or out of the labor force.[15] This choice is deliberate, as the duration of employment is the appropriate unit of analysis for the unemployment insurance program. UI entrance requirements do not distinguish between jobs, but rather focus on total weeks of employment within the qualifying period.

The eligibility requirements increased to 14 weeks of employment in all regions of Canada over an approximately 11 month period between January and November 1990. Our strategy is to compare employment hazard rates for spells which were initiated in the period January through June 1990, with those which began in the corresponding period in the previous year (January through June 1989). The use of fresh spells eliminates problems of left-censoring. The choice of spells starting through June guarantees that all of those initiated in 1990 will potentially be affected by the suspension of the VER's before their renewal in November 1990. Because the increase in eligibility requirements affected all parts of the country, and very few regions had a 14 week requirement prior to this event, there is no "natural" control group for unobserved time effects which may affect the employment hazard. We attempt to minimize the influence of this source of bias by using the variation in the eligibility requirement across comparable periods of a year (i.e., the same months). We believe that the following evidence is strong enough to allay any

[14] There is an incentive for provincial governments to provide these sorts of jobs since they offer a type of longer term income maintenance which is federally funded as an alternative to welfare which is funded at the provincial level.

[15] Many individuals will hold a number of jobs within a single employment spell.

doubts that particular events in 1989 or 1990 lead to the observed spike in the employment hazard in the week of UI eligibility.

Identifying the Week of UI Eligibility

There are three potential sources of variation in UI entrance requirements in our data over the period. The first is the "across time" variation, induced by the country-wide suspension of the VER's in 1990. As noted above, usually eligibility requirements vary across 48 economic regions in the country based on local economic conditions. Therefore, there is also "across region" variation for the spells which start between January and June 1989. Finally, the entrance requirement will also vary across individuals. People enter employment spells with different histories, and therefore different amounts of insured weeks accumulated from past spells. For example, if the employment requirement were 10 weeks, and an individual started employment with 4 insured weeks from a previous spell, he or she would become eligible to collect UI benefits in the 6th week of the employment spell. Also, individuals who have made a benefit claim during the qualifying period (repeaters), or who have not worked in the previous 2 years (new entrants) face an entrance requirement in excess of the local VER. Therefore, a final source of variation in this parameter is "across individuals".

We are able to exploit these three sources of variation to differing degrees in our data set. The most precisely identified are the across time and across region variation, requiring only information on the period in which the spell started, and each individual's UI region. Our data set includes monthly regional unemployment rates. These allow us to determine the entrance requirements and maximum duration of benefits by UI region for each month in 1989. Determination of a claimant's region is not necessarily a straightforward exercise, however, because individuals may choose to work in one region of the country but establish a subsequent UI claim in another.[16] We locate respondents by their reported UI region of permanent residence as of January 1989. Since there is only variation in the entrance requirements across regions in 1989, this method of locating respondents should not be too severely affected by subsequent migration.[17]

To identify the individual specific variation, we need information on any employment spells, or UI claims, an individual may have had in the qualification period preceding a given spell. If the person is a "new entrant", this qualification period is two years rather than one. Each person accumulates insured employment starting at the end of any preceding benefit claim. A person who has made a claim during their qualification period is identified as a "repeater", and faces an eligibility requirement which is generally longer than the standard for their region.[18]

The LMAS data provide only partial information on each individual's history. First, we would need data from 1987 to identify new entrants in the 1989 sample, and therefore are forced to assume that respondents are not new entrants. Second, the LMAS does not identify the exact weeks in which UI benefits were received, but rather whether benefits were received sometime during a year.[19] While this information allows us to identify most repeaters, we are unable to determine the temporal ordering of any UI claims and employment spells in a given year. The latter information is important for determining the number of insured weeks an individual has accumulated at the start of the employment spell which we observe.

[16] The incentive for this action is that a longer benefit entitlement for a given number of insured weeks may be available in another region.

[17] The duration of benefits was still affected by regional unemployment in 1990. There are a small number of cases in which the regional identification is not consistent with the reported provincial identifier at the time the spell is initiated. We delete these observations from our sample. To the best of our knowledge this exclusion has no material effects on the results.

[18] Repeaters face the same entrance requirements as "regular" claimants if the local unemployment rate is 11.6 per cent or higher.

[19] The number of weeks of UI receipt (but not the precise timing of those weeks) is known for claims that are made during temporary absences from an ongoing job (e.g., layoffs). Similar information for benefit receipt between jobs, however, is not available. Therefore the summary question on UI receipt over the year provides the only information which is consistent across all types of claims.

We take two approaches to constructing an individual's history from these data. In the first approach ("Duration"), we assume that all individuals enter employment spells with no accumulated insured weeks.[20] Therefore, we identify the week in which an individual qualifies for UI benefits by comparing the duration of their spell to the local entrance requirement. Under this method the only individual specific variation in the entrance requirements is due to repeater status. Clearly the bias in this procedure is that we may underestimate the number of insured weeks individuals have at the start of their spells.

In the second approach ("Insured Weeks"), we assume respondents receive benefits in each of the weeks they were not working, if they report any UI income during the year. While this assumption allows us to determine the number of insured weeks with which individuals enter their employment spells, it measures the variation across individuals with error. This error arises because we may erroneously assign benefit receipt to weeks in which none were received, and because UI benefit receipt is systematically underreported in the LMAS. Discussions with officials at Statistics Canada suggest the degree of underreporting may be as high as 20 per cent. The first source of error may lead to an underestimate of the number of insured weeks accumulated at the start of a new employment spell, for individuals who reported UI income in the previous year. The second source of error may lead to overestimates of accumulated insured weeks within the rest of the sample. In the following analysis we note the effects these two methods of identifying individual specific entrance requirements have on our results.

[20] Recall that employment spells are preceded by a period of non-employment. Therefore we assume that all individuals either made UI claims in this period or are entering or returning to the labor market after a sufficient period of time such that previous employment does not contribute to satisfying the VER. While evidence on a decline in the UI take up rate (Burtless, 1983) would make such an assumption troublesome for the U.S., there are no studies, we are aware of, documenting similar trends for Canada. Also, eligibility rates for UI among the unemployed are higher in Canada than the U.S., and there is evidence that many individuals classified as out of the labor force receive UI benefits in Canada. See Card and Riddell (1992). Finally, this assumption makes some compensation for the underreporting of UI benefit receipt in the LMAS.

Characteristics of the Sample

In the first two columns of Table 1 we report the number of spells in our sample, by province and year. The variation in the number of spells by province reflects their relative populations, the sampling frame of the survey and local economic conditions. Some smaller provinces are oversampled to obtain more reliable statistics. Also, provinces with less employment stability are likely to have more spells initiated over the period. Therefore, while Newfoundland has a population of less than 20 per cent of British Columbia's, its higher sample weight, and perhaps higher unemployment rate, lead to a greater number of spells from this region.

In the next two columns of Table 1, we report the percentage of spells which are censored in each year. The cross year variation results from our sample selection criteria. All spells in progress as of November 10, 1990 are censored, which is approximately when legislation renewing the VER's was finally passed into law. Spells starting in 1989 can be up to 96 weeks long before censoring, while those starting in 1990 can be as long as 43 weeks. Therefore, a higher proportion of spells initiated in 1990 will be incomplete. The cross province variation, however, provides an interesting view of the labor market in each area. It is quite evident that the percentage of censored spells is lower in certain provinces, in particular those in the Atlantic region (Newfoundland, Prince Edward Island, Nova Scotia and New Brunswick). This indicates the prevalence of short employment spells in these areas, which in turn is consistent with a relative lack of employment stability. Therefore, this information identifies the Atlantic region as an area in which the spikes in the employment hazard may be largest.

In the next column we report the number of individuals in the sample. In each province some individuals have more than one employment spell over the sample period. The proportion tends to be higher, however, in the Atlantic region, again suggesting relatively higher employment instability.

In the final three columns of Table 1 we report some indicators of the state of the provincial labor markets. First, there are large difference in the provincial unemployment rates. For example, the rate in Newfoundland was generally three times that in Ontario. Together with the information on mean hourly earnings, these statistics highlight the differences across provinces. The Atlantic provinces have higher unemployment rates and lower average earnings. As indicators of loose labor markets, this information is consistent with the previous evidence of employment instability in this region.

Empirical Framework

We model employment exits as a discrete process, in which the weekly hazard takes the logistic functional form. This model has been used previously by, among others, Nickell (1979) and Ham and Rea (1987) to model unemployment duration, and Gunderson and Melino (1990) to model the duration of strikes. The conditional probability (hazard) that an employment spell ends in week t, is

$$h_i(t) = \frac{1}{1 + \exp[-(\theta + \beta' X_i(t))]} \tag{1}$$

where θ is a constant, $X_i(t)$ is a vector of time varying explanatory variables for the i^{th} spell and β is the corresponding vector of parameters.

Using (1), the likelihood function for a sample of complete spells C, of lengths t_i^* and incomplete (censored) spells $\bar{C}$ of lengths $\hat{t}_i$ is

$$L = \prod_{i \in C} f_i(t_i^*) \cdot \prod_{i \in \bar{C}} [1 - F_i(\hat{t}_i)], \tag{2}$$

where

$$f_i(t_i^*) = [\prod_{t=1}^{t_i^*-1} (1 - h_i(t))] h_i(t_i^*), \tag{3}$$

and $F_i(\cdot)$ is the cumulative distribution function. In (3), the probability of observing a completed spell of length t_i^*, is just the probability of departure in that week, multiplied by the probability that the individual does not exit in any of the preceding weeks.

Our explanatory variables include controls for three age categories, five education categories, gender, marital status, school attendance, past UI receipt, hourly earnings,[21] the provincial unemployment rate, completed duration and a year effect. We enter duration into the hazard using alternatively a polynomial (quartic) and a step function. The justification for these different specifications of the "baseline hazard" is discussed below.

We capture the effects of qualification for UI benefit receipt with three dummy variables. The first, ELIG1, is a variable which takes the value 1 in the week that an individual becomes eligible to claim UI benefits, and the value 0 in all other weeks. This variable will pick up any spike in the employment hazard in the week of eligibility. A second variable ELIG2 takes the value 1 in the period between the initial week of eligibility and the week in which the maximum benefit entitlement is attained (50 weeks of benefits); weeks in which additional employment brings further benefit entitlement. This period corresponds to, for example, the segment EG on the budget constraint in Figure 3. As noted above, in high unemployment regions this period is quite short because the sum of the base entitlement and regionally extended benefits can be as high as 42 weeks. Finally, we also include a dummy variable ELIG3 which takes the value 1 in the weeks in which the individual qualifies for the maximum benefit entitlement. This variable should capture some of the more permanent effects of UI eligibility on employment duration. On the other hand, this variable may just serve as an additional control for duration, particularly if the

[21] A possible objection is that hourly earnings are endogenous. Our estimates of the eligibility parameters are not substantively affected if we delete this variable from the specification.

effects of UI eligibility on the employment hazard are confined to the first weeks within the state. A full description of all the variables used in the analysis is presented in the Appendix.

Results

The ELIG1 variable is constructed to capture spikes in the employment hazard which are correlated with the week in which individuals qualify to make UI claims. This correlation is critical to identification in this study, since raw employment hazards typically have spikes at numerous points for various reasons. These can arise due to digit preference,[22] calender effects[23] or local employment initiatives which provide a relatively large number of jobs of a fixed duration. Our approach attempts to determine the extent to which any spikes correspond to the UI entrance requirements which individuals face.

We present some evidence on this point in Figures 4 and 5. To conserve space we plot the empirical hazards for just two of the ten provinces: Newfoundland and Ontario. By measure of the unemployment rate, these regions had the weakest and strongest labor markets, respectively, over the sample period. We also confine the figures to the first 24 weeks of employment. This is the period in which the eligibility requirements should have their greatest effect. The statistics underlying these plots are recorded in Table A1 in the Appendix

The hazards, by year, for Newfoundland are presented in Figure 4. They display fairly regular spikes which roughly correspond with integer multiples of two weeks. These are likely partly the result of response error, as individuals round off the reports of their employment duration to the nearest even number. Also, the

[22] The tendency of individuals to report the lengths of their employment spells rounded off to the nearest even number, or to a multiple of one month.

[23] By this we mean an observed tendency for spells to begin or end at the beginning or end of a calender month.

spikes at approximate "month ends" appear a little larger, suggesting an additional source of aggregation.

Newfoundland is one of the few provinces in which the VER in all of its regions over 1989 was 10 weeks. Therefore the suspension of the VER's in 1990, led to a province- wide increase in the eligibility requirement to 14 weeks. A rough indication of the effect of this change on people's behavior would be a distinct drop in the hazard at 10 weeks in 1990 and an increase in the hazard at 14 weeks.[24] The decline at 10 weeks is quite evident as the 1990 hazard is just 66 per cent of its 1989 level. The increase in the 14 week hazard is more moderate, however, at about 11 per cent. To control for temporal changes in the absolute level of the hazards we can compare the ratio of the 10 week to 14 week hazard in each year. In 1989 it is 1.21, while in 1990 it is 0.72. It appears the 10 week spells were relatively less common after the entrance requirement was raised to 14 weeks.

The empirical hazards for Ontario are plotted in Figure 5. In 1989 the VER in this province ranged between 10 and 14 weeks and therefore there is no particular spike of interest in the hazard for this year. In 1990, the entrance requirement was 14 weeks province wide. The hazard at 14 weeks falls moderately between 1989 and 1990, from 0.0713 to 0.0669. What is striking, however, is the relative decline in the magnitude of the hazards at earlier weeks. For example, the ratio of the hazards at 10 and 14 weeks is 0.99 in 1989 and 0.69 in 1990. The corresponding ratios for the 12 week hazard are 0.65 and 0.54. Again there appears to be a relative decline in the importance of employment spells less than 14 weeks in 1990.

Our alternative specifications of the control for completed duration separate the eligibility spikes from spikes due to digit preference and calendar effects in different ways. The quartic specification enforces some smoothness on the hazard, identify-

[24] This inference will not be exact as people with accumulated insured weeks from previous spells will become eligible at other points in time.

ing the eligibility effect by systematic deviation from the base that is correlated with the variation in the eligibility parameters.[25] The step function includes dummy variables for weeks 2 through 14, 15-16, 17-18, 19-20, 21-25, 26-30, 31-40, 41-50, 51-60 and 61+. Therefore, this specification can easily accommodate the various spikes evident in Figures 4 and 5, as well as their common movement over time through the year effect. The risk in this approach, however, is that because the primary effect of interest is also a spike, we may over-parameterize the model.

In Table 2 we present the estimates of the eligibility parameters when duration is used to determine UI eligibility. The first three columns of results are for a specification in which we condition only on a year indicator and the quartic in duration without other control variables (labelled "No Controls"). The ELIG1 parameter is both positive and significant for most provinces. The one exception to this conclusion is Manitoba, for which the estimate is positive, relatively small and imprecisely estimated.

The estimates of the other eligibility parameters are generally positive, although often imprecisely estimated. The ELIG2 parameter tends to be larger, and the inference more precise, in central and western provinces which had relatively lower unemployment rates over the period. The estimates for the Atlantic provinces are generally much smaller and have larger standard errors. As noted above the ELIG2 parameter will not be well identified in these areas, because individuals will initially qualify for UI with close to the maximum benefit entitlement. Furthermore, the ambiguous sign for ELIG2 is not surprising given that there are competing incentives for workers with this duration. Entitlement to UI should increase the hazard but additional work generates greater duration of benefits.

The results for ELIG3 are harder to interpret as this variable may be providing

[25] A quartic is the highest order polynomial which can be supported for all provinces, and is used to provide a consistent basis of comparison.

an additional control for duration. For 7 of the 10 provinces the estimates are larger than those for ELIG2. We might expect this result, as once initial eligibility has been established there is an incentive to remain employed until benefit entitlement has reached the maximum. More generally, however, the inference on this effect is quite imprecise.

The next columns of Table 2 contain estimates of the eligibility parameters which are net of observable differences across individuals (labelled "With Controls").[26] The controls for observable heterogeneity slightly alter the magnitude of some of the ELIG1 estimates, but the effect is quite marginal for all provinces. The effects on the estimates of ELIG2 and ELIG3 also vary by province, although again they do not result in any substantive change in inference.

In Table 3 we present a second set of estimates in which we use our measure of insured weeks to determine the week of eligibility within each spell. Again a set of results without controls is presented in the first three columns. The estimates of ELIG1 are generally smaller than those in Table 2, typically by between one-third and one-half. They do indicate a significant rise in the hazard in the week of eligibility, however, except in Manitoba for which the inference is robustly insignificant across the two sets of results. The estimates of the other parameters are also typically smaller than in the previous results, and in more cases insignificant. Also, the point estimates are negative in a larger number of cases.

The results net of observable heterogeneity are presented in the final three columns. The week of eligibility effects are generally smaller in this specification, although not enough to affect inference in most cases. The effects on the other two eligibility parameters are more substantial. The point estimates of ELIG2 and ELIG3 are now typically negative, although in most cases insignificant.

[26] The parameter estimates for the various control variables are available from the authors on request.

Tables 4 and 5 contain a corresponding set of results using the step function specification. Using duration to determine eligibility (Table 4) there are both increases and decreases in the parameter estimates across provinces, although the general pattern suggests smaller effects in the week of eligibility. Furthermore, the ELIG1 parameter is now significant for only half the provinces. This group includes the Atlantic provinces where we expect large eligibility effects, but also Quebec and Saskatchewan. Also, the estimates of ELIG2 are again larger and more likely significant for the prairie provinces. Finally, the difference between the results with and without controls mirrors that seen in Table 2.

The results using insured weeks to determine eligibility (Table 5) bear little resemblance to those in Table 3. Most of the parameter estimates are small and insignificant at conventional levels. Furthermore, most of the point estimates of ELIG2 and ELIG3 are negative. There is no evidence in these results that UI eligibility requirements have any effect on the employment hazard.

Comparing the results using duration to determine UI eligibility across the quartic and step function specifications, a conservative summary is that there is a significant increase in the employment hazard in the week of UI eligibility for 5 of the 10 provinces. These include areas of typically high unemployment, but also Quebec and Saskatchewan. There also appears to be significant effects in the weeks between initial eligibility and maximum benefit entitlement in some of the western provinces.

The change in the results across the two methods of determining UI eligibility (duration versus insured weeks), however, lends some uncertainty to this inference. The estimates using our measure of insured weeks are not robust to the change in specification of the baseline hazard. The duration approach clearly overstates the week of eligibility in some spells. In the insured weeks approach we may also overestimate the week of eligibility when we allocate UI benefit receipt to a preceding spell of non-employment in which no UI claim was made. In other cases, however, we will underestimate the week of eligibility due to the systematic under-reporting of UI receipt within the sample.

If the two sources of bias in the insured weeks approach lead us to misallocate the week of eligibility in a higher proportion of spells, we would expect smaller estimates of the ELIG1 parameter. This is consistent with the results in Tables 3 through 6. Also, more misallocation will reduce the correlation between exits in response to the eligibility requirements and eligibility parameters. In the quartic specification, misallocation of the week of eligibility may simply result in the eligibility parameters capturing spikes in the hazard that are due to other sources. This, however, should be less of a problem in the step function specification. As noted above, digit preference and calender effects which are relatively stable over time will be accommodated by the duration dummy variables.

Comparing the changes in the duration and insured weeks estimates across the two specifications lends some support to this interpretation. The results using duration to determine eligibility are smaller using the step function, while the insured weeks results are obliterated. This is what we would expect, assuming that the insured weeks method led to greater misallocation of the week of eligibility. If this assumption were true, the insured weeks estimates of the eligibility parameters in the quartic specification are more likely to be capturing the effects of digit preference. Therefore, we would expect these estimates to go to zero once we adopt the more flexible step function formulation of the baseline hazard. This is exactly what we observe comparing the results in Tables 3 and 5.

Clearly neither method of identifying the week of eligibility is free of measurement error. Nevertheless, we believe that the systematic underreporting of UI receipt in the LMAS, as well as the greater sensitivity of the insured weeks results to the specification of the baseline hazard, suggest that the duration method leads to more reliable inference.[27]

To calibrate the implications of the estimated eligibility effects, we have run some simulations of the employment hazard for Newfoundland and Ontario, using the duration method to identify the week of eligibility and the parameter estimates which are net of the controls for observable heterogeneity.[28] For Newfoundland we compare the employment hazard assuming a 10 week eligibility requirement, to the result if none of the eligibility requirements were in effect. The spike in the employment hazard in the week of eligibility is 2.3 times the base level using the estimates from the quartic specification (i.e., ELIG1 is 0.905 with a standard error of 0.128). Using the estimates from the step function specification the multiplier on the hazard in the week of eligibility is 1.43 (i.e., ELIG1 is 0.378 with a standard error of 0.171). For Ontario we assume a 14 week eligibility requirement. Using the estimates from the quartic specification the rise in the hazard in the week of eligibility is 2.1 times the base. The estimates from the step function specification indicate an increase of 1.27, although this increase is not statistically significant at conventional levels (see Table 4).

A final concern is the issue of unobserved heterogeneity. One indication that this issue is of limited relevance here is that additional controls for observed heterogeneity have relatively little influence on the estimates. In the specification of the hazard model which we use, the addition of controls for unobserved differences across individuals is fairly straightforward using a semi-parametric procedure suggested by Nickell (1979) and Heckman and Singer (1984). The approach is to allow the constant term in (1), θ, to vary across individuals following a discrete distribution with points of support $\theta_1, \theta_2, ..., \theta_K$ and associated probabilities $P_1, P_2, ..., P_K$

[27] The previous studies (cited above) of UI eligibility effects do not, to our knowledge, address this issue, and use the duration method to determine the week of eligibility.

[28] The simulations are constructed evaluating the weekly hazard at the means of the demographic and economic control variables.

(where $P_K = 1 - \sum_{k=1}^{K-1} P_k$). Given that we observe more than one spell for some individuals, this specification seems more appropriate than a simpler formulation which allows θ_k to vary across spells. The potential effect of the heterogeneity terms is to allow variation in the mean hazard. Since the eligibility spikes are measured relative to this mean, their magnitude could be affected by this innovation.

Suppose we observe n_{ch} complete spells and $n_{\bar{c}h}$ censored spells for individual h. Then the contribution to the likelihood function for this individual is

$$L_h = \sum_{k=1}^{K} P_k \left((\prod_{i=1}^{n_{ch}} f_i(t_i^*; \theta_k)) \cdot (\prod_{j=1}^{n_{\bar{c}h}} [1 - F_j(t_j^*; \theta_k)]) \right), \tag{4}$$

and where $f_i(t_i^*; \theta_k)$ is defined through appropriate modification of (3).

Given the large number of equations which potentially need to be re-estimated, we first examine the results of an information matrix test for model specification suggested by White (1982) and Chesher (1984). This test compares the alternative estimates of the information matrix provided by the Hessian and the outer product of the score vector. Under the null hypothesis of a correct specification, the difference between these two estimates should converge to zero.[29]

Our specific interest is in the possible mis-specification of the constant term, which would provide some evidence of the presence of unobserved heterogeneity. Accordingly we focus on the components of the information matrix corresponding to this term. Ham and Melino (1987) show that this test of the heterogeneity parameters is simplified by singularities in the information matrix. As a result, the test statistic has degrees of freedom equal to the number of points of support in the distribution of θ_k. Therefore the initial test for unobserved heterogeneity in the specification with a single constant term, θ_1, has one degree of freedom.

The test statistics, and their associated probability values, are reported in Table 6. We perform the test for the various specifications including the controls for

[29] In implementing test, we use the sample variation in the test statistic to estimate its standard error.

observed heterogeneity. Overall the results indicate that unobserved heterogeneity is not a problem. At conventional levels of significance, we can reject the null hypothesis in a few cases (most notably New Brunswick and Quebec), but the evidence is not overwhelming.

To confirm this inference, we have re-estimated our model allowing for additional points of support. The application is to New Brunswick, which the results in Table 6 suggest is the province for which unobserved heterogeneity may have the greatest impact. While we do find an additional point of support, its associated probability is very small. Furthermore, the estimates of the eligibility parameters are unaffected by its inclusion. Finally, when we recalculate the Information matrix test we cannot reject the hypothesis that there are no further points of support.[30]

Interpretation and Conclusions

The effect which we observe may in part be due to the prominence of the UI system in Canada and Canadians' high degree of familiarity with the program. Besides serving as insurance against involuntary job loss, over the sample period UI benefits were available to people who quit their jobs,[31] to those in special job training and women on maternity leaves.[32] A high proportion of Canadians have some involvement with the program. Card and Riddell (1992) estimate that the ratio of UI recipients to the unemployed was consistently above 80 per cent throughout the 1980's. The corresponding ratio for the U.S. was 25 to 35 per cent. Baker and Benjamin (1993) provide evidence that in 1990, 30 per cent of individuals and 25 per

[30] For example, using duration to determine eligibility and the quartic specification, the second point of support is 0.211 with a standard error of 0.986. Its associated probability is 0.006. Finally the Information matrix test ($\chi^2(2)$) for a third point of support is 1.007 with a p-value of 0.316.

[31] This feature of the UI system was changed in 1993.

[32] As a consequence some have argued that the system is more akin to an income supplement programme than a insurance system as such. See Commission of Inquiry on Unemployment Insurance (1986).

cent of families in Canada reported receiving some UI income.

The results should extend to other countries where workers are well informed about UI. We expect that awareness of UI programs will be highest in industries or regions with employment instability, because frequent spells of unemployment disseminate information about the UI system throughout the relevant population.

Eligibility requirements may both lower the labor supply of those with loose attachments to the labor force and induce participation of nonparticipants. Which effect dominates remains an important question for future research. Likewise, the relative roles of demand and supply side factors, leading to the disproportionate number of employment spells satisfying the minimum eligibility requirement, is yet unknown. What is clear from our results is that employers and employees are jointly taking advantage of the subsidy to short-term employment.

More generally, our evidence points to an additional behavioral consequence of the UI system. The increase in the employment hazard at the time of eligibility is substantial in many parts of Canada. Together with studies such as Feldstein (1976), the results provide further instances where UI parameters have implications beyond their previously documented impact on the search activities of the unemployed.

Appendix

Variable Definitions

1. ELIG1: A dummy variable which takes the value one in the week in which an individual satisfies the local UI eligibility requirement.

2. ELIG2: A dummy variable which takes the value one in weeks that the individual has satisfied the local UI eligibility requirement, but has not reached the maximum benefit entitlement.

3. ELIG3: A dummy variable which takes the value one in weeks that the individual has satisfied the local UI eligibility requirement and has reached the maximum benefit entitlement.

4. Duration: The number of weeks the spell has been in progress.

5. Year: A dummy variable which takes the value one in weeks during the year 1990.

6. UR: The monthly unemployment rate in the respondent's province of residence.

7. Hourly Earnings: The respondent's average hourly earnings in a given year. All values are converted to 1989 dollars using the Consumer Price Index.

8. Age: Dummy variables are coded according to the following categories:

   a) 14 to 24 years of age in 1988,

   b) 25 to 44 years of age in 1988.

   The excluded group is respondents 45 to 64 years of age.

9. Education: Dummy variables are coded according to the following categories:

   a) High School Graduate,

   b) Post Secondary Education,

c) Trades Certificate or Diploma,

b) University Degree.

The excluded group in the analysis is respondents who have not completed high school.

10. Gender: A dummy variable for females.

11. Marital Status: A dummy variable which takes the value one if the respondent is married.

12. School Attendance: A dummy variable which takes the value one if the respondent attended school in the year of the current week.

13. Past UI Receipt: A dummy variable which takes the value one if the respondent received UI income in the year preceding the spell.

Table A1: Empirical Hazards: Newfoundland and Ontario

	Newfoundland				Ontario			
	1989		1990		1989		1990	
Week	Hazard	Std. Err.	Hazard	Std. Err.	Hazard	Std. Err.	Hazard	Std. Err.
1	0.021	0.005	0.012	0.004	0.013	0.004	0.014	0.004
2	0.012	0.004	0.020	0.005	0.013	0.004	0.012	0.004
3	0.017	0.004	0.022	0.005	0.018	0.005	0.005	0.003
4	0.011	0.004	0.023	0.005	0.011	0.004	0.009	0.004
5	0.016	0.004	0.021	0.005	0.017	0.005	0.016	0.005
6	0.027	0.006	0.030	0.006	0.014	0.004	0.012	0.004
7	0.015	0.004	0.035	0.007	0.010	0.003	0.019	0.005
8	0.039	0.007	0.034	0.007	0.014	0.004	0.011	0.004
9	0.033	0.006	0.013	0.004	0.020	0.005	0.019	0.005
10	0.088	0.011	0.058	0.009	0.070	0.010	0.046	0.008
11	0.047	0.008	0.031	0.007	0.017	0.005	0.018	0.005
12	0.052	0.009	0.027	0.007	0.046	0.008	0.036	0.008
13	0.039	0.008	0.032	0.007	0.017	0.005	0.018	0.005
14	0.073	0.011	0.081	0.012	0.071	0.010	0.067	0.011
15	0.024	0.007	0.049	0.010	0.028	0.007	0.025	0.007
16	0.040	0.008	0.041	0.009	0.037	0.008	0.055	0.010
17	0.064	0.011	0.047	0.010	0.042	0.008	0.021	0.006
18	0.075	0.012	0.067	0.012	0.066	0.011	0.067	0.012
19	0.043	0.010	0.077	0.013	0.026	0.007	0.038	0.009
20	0.057	0.011	0.050	0.011	0.027	0.007	0.015	0.006
21	0.058	0.012	0.049	0.012	0.014	0.005	0.009	0.005
22	0.077	0.014	0.038	0.011	0.026	0.007	0.026	0.008
23	0.060	0.013	0.081	0.016	0.022	0.007	0.033	0.009
24	0.043	0.011	0.044	0.013	0.010	0.005	0.008	0.005

Source: 1988-1990 LMAS Longitudinal Files.

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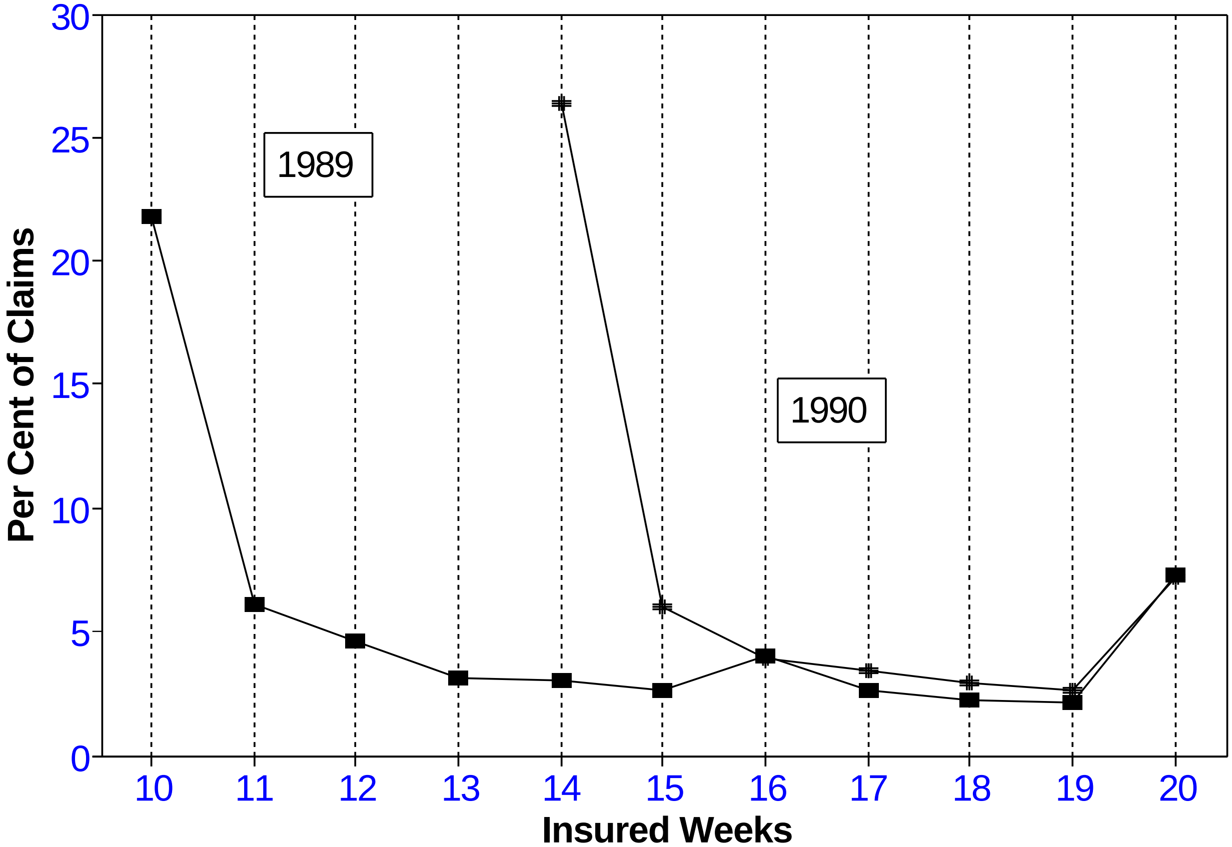
Figure 1: Distribution of Insured Weeks
January-November Claims in Newfoundland



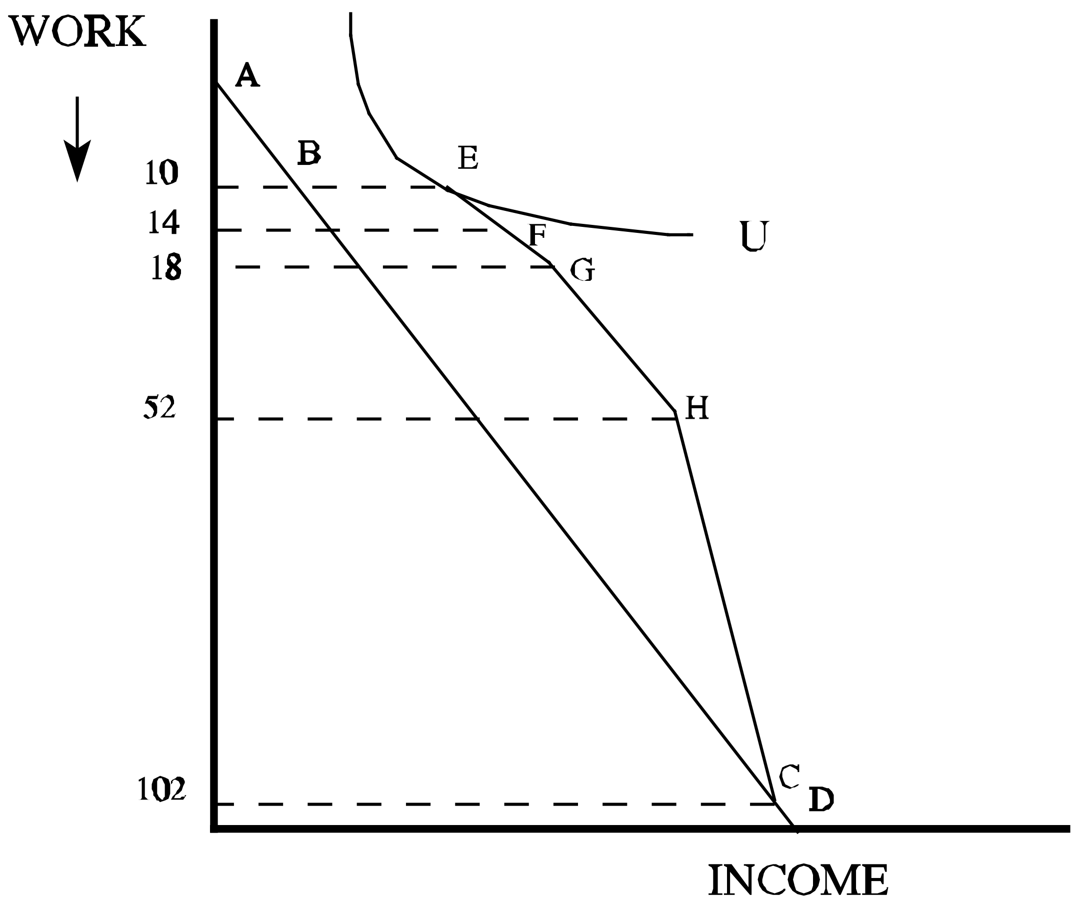


Figure 2

Unemployment Insurance in a High Unemployment Region

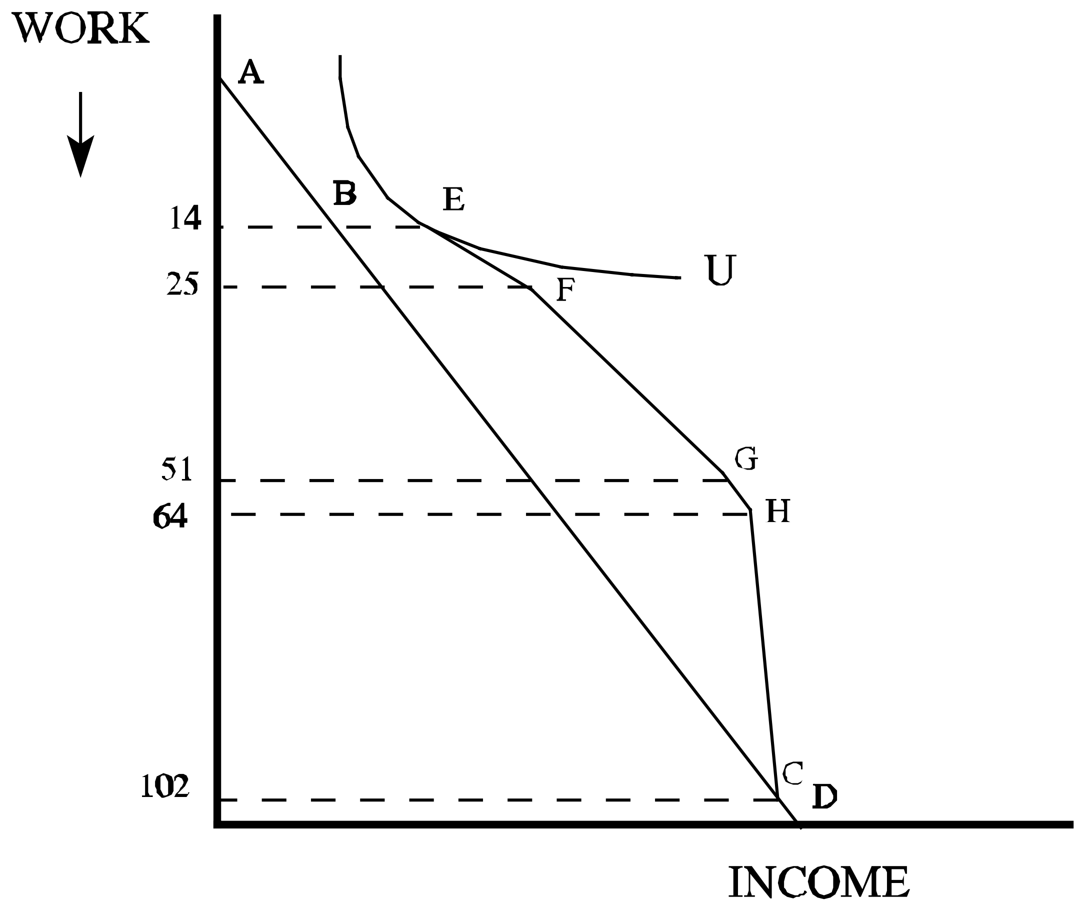


Figure 3

Unemployment Insurance in a Low Unemployment Region

Figure 4: Empirical Employment Hazard
Newfoundland (LMAS Data)

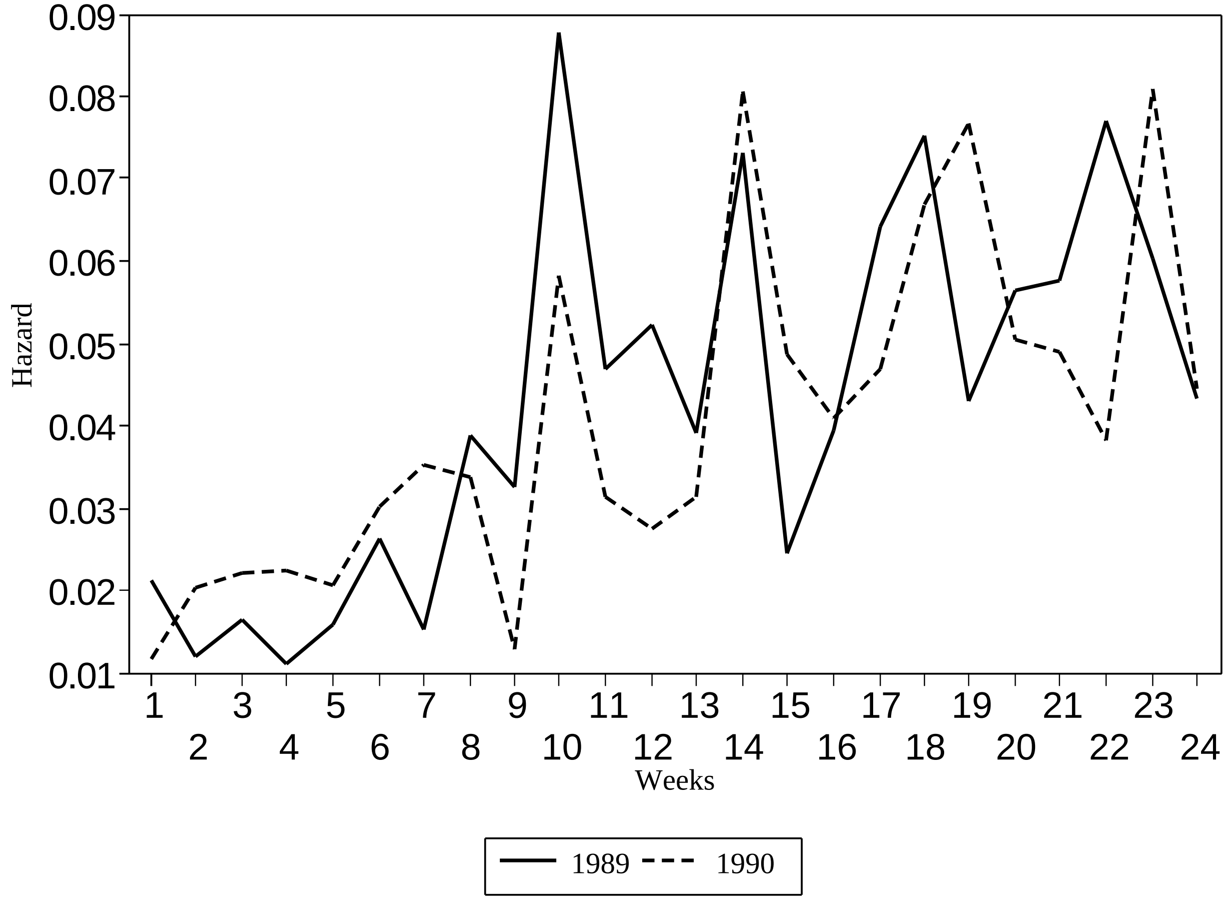


Figure 5: Empirical Employment Hazard
Ontario (LMAS Data)

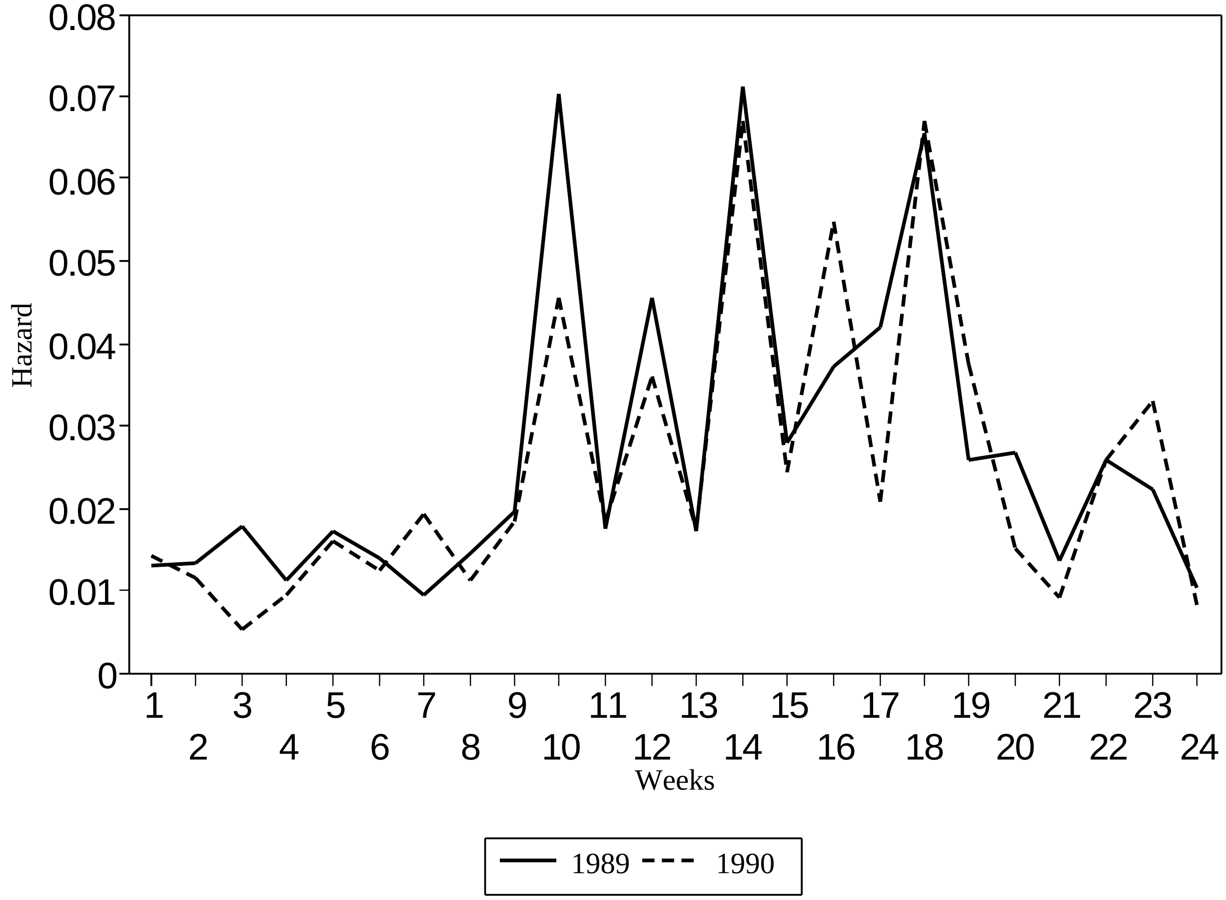


Table 1: Summary Statistics

	Number of Spells		Per Cent Censored		Number of Individuals	Per Cent Female	Unemployment Rate		Hourly Earnings
	1989	1990	1989	1990			1989	1990	
British Columbia	614	513	24.8	46.0	841	48.1	9.1	8.3	10.75
Alberta	723	538	24.3	47.2	943	41.8	7.2	7.0	10.15
Saskatchewan	452	338	25.9	49.7	574	45.2	7.5	7.0	9.42
Manitoba	376	303	22.6	43.6	487	44.9	7.6	7.2	9.28
Ontario	923	782	29.0	49.0	1304	46.4	5.1	6.3	10.44
Quebec	946	829	21.7	42.9	1258	44.5	9.3	10.1	10.27
New Brunswick	822	696	13.9	38.8	999	40.7	12.5	12.1	8.61
Nova Scotia	270	206	23.7	45.1	344	40.8	9.9	10.5	8.96
Prince Edward Island	301	235	13.6	31.1	347	48.2	14.2	15.1	8.48
Newfoundland	948	850	10.0	31.2	1178	45.2	15.8	17.1	8.16

1/ Source: 1988-1990 LMAS Longitudinal Files.
2/ Per cent Female and Hourly Earnings are calculated using weighted data. Hourly Earnings (1989 dollars) are an average over the reported wages in the year in which individuals' spells ended.

Table 2: Eligibility Parameter Estimates Using Duration to Determine Eligibility (Quartic)

	No Controls			With Controls		
	ELIG1	ELIG2	ELIG3	ELIG1	ELIG2	ELIG3
British Columbia	0.790*	0.291	0.277	0.781*	0.286	0.272
	(0.184)	(0.159)	(0.199)	(0.186)	(0.160)	(0.202)
Alberta	0.727*	0.319*	0.363	0.681*	0.290*	0.592*
	(0.170)	(0.143)	(0.190)	(0.171)	(0.144)	(0.192)
Saskatchewan	0.867*	0.646*	0.835*	0.869*	0.648*	0.871*
	(0.229)	(0.185)	(0.298)	(0.231)	(0.189)	(0.302)
Manitoba	0.311	0.423*	0.365	0.248	0.367*	0.328
	(0.269)	(0.181)	(0.301)	(0.271)	(0.186)	(0.310)
Ontario	0.808*	0.510*	0.719*	0.772*	0.485*	0.687*
	(0.149)	(0.125)	(0.265)	(0.150)	(0.127)	(0.263)
Quebec	0.792*	-0.010	0.237	0.792*	0.004	0.252
	(0.134)	(0.118)	(0.148)	(0.135)	(0.120)	(0.149)
New Brunswick	0.789*	-0.027	0.330*	0.794*	-0.026	0.291
	(0.140)	(0.126)	(0.165)	(0.141)	(0.127)	(0.166)
Nova Scotia	1.025*	0.234	0.533	1.029*	0.236	0.469
	(0.267)	(0.247)	(0.288)	(0.268)	(0.249)	(0.292)
Prince Edward Island	0.824*	-0.190	0.158	0.835*	-0.173	0.181
	(0.223)	(0.227)	(0.298)	(0.224)	(0.227)	(0.300)
Newfoundland	0.922*	0.252*	0.648*	0.905*	0.261*	0.626*
	(0.128)	(0.120)	(0.160)	(0.128)	(0.120)	(0.161)

1/ * indicates significant at the 5% level.
2/ Source: 1988-1990 LMAS Longitudinal Files.
3/ "No Controls" includes a quartic in duration and a year indicator.
4/ "With Controls" includes a quartic in duration, a year indicator plus controls for the provincial unemployment rate, hourly earnings, age, education, gender, marital status, school attendence and past UI receipt.
5/ Standard errors in parentheses.

Table 3: Eligibility Parameters Using Insured Weeks to Determine Eligibility
(Quartic)

	No Controls			With Controls		
	ELIG1	ELIG2	ELIG3	ELIG1	ELIG2	ELIG3
British Columbia	0.587*	0.071	0.025	0.455*	-0.118	-0.112
	(0.203)	(0.124)	(0.125)	(0.205)	(0.129)	(0.132)
Alberta	0.593*	-0.159	-0.209	0.452*	-0.376*	-0.366*
	(0.178)	(0.105)	(0.127)	(0.180)	(0.112)	(0.135)
Saskatchewan	0.530*	0.104	-0.272	0.424	-0.043	-0.351
	(0.249)	(0.140)	(0.173)	(0.250)	(0.150)	(0.186)
Manitoba	0.355	0.416*	0.080	0.212	0.252	-0.081
	(0.295)	(0.140)	(0.187)	(0.297)	(0.150)	(0.202)
Ontario	0.395*	-0.023	-0.425*	0.179	-0.266*	-0.504
	(0.171)	(0.088)	(0.145)	(0.173)	(0.094)	(0.152)
Quebec	0.575*	-0.094	-0.151	0.482*	-0.200*	-0.199
	(0.141)	(0.100)	(0.104)	(0.142)	(0.102)	(0.109)
New Brunswick	0.502*	-0.187	-0.094	0.454*	-0.231*	-0.122
	(0.147)	(0.106)	(0.110)	(0.149)	(0.110)	(0.117)
Nova Scotia	0.665*	-0.263	-0.176	0.501	-0.409*	-0.373
	(0.262)	(0.198)	(0.209)	(0.264)	(0.204)	(0.216)
Prince Edward Island	0.590*	-0.292	-0.477*	0.579*	-0.256	-0.406
	(0.214)	(0.201)	(0.228)	(0.218)	(0.205)	(0.242)
Newfoundland	0.651*	0.002	0.137	0.490*	-0.134	-0.058
	(0.122)	(0.104)	(0.119)	(0.123)	(0.104)	(0.122)

1/ * indicates significant at the 5% level.
2/ Source: 1988-1990 LMAS Longitudinal Files.
3/ "No Controls" includes a quartic in duration and a year indicator.
4/ "With Controls" includes a quartic in duration, a year indicator, plus controls for the provincial unemployment rate, hourly earnings, age, education, gender, marital status, school attendence, and past UI receipt.
5/ Standard errors in parentheses.

Table 4: Eligibility Parameter Estimates Using Duration to Determine Eligibility
(Step Function)

	No Controls			With Controls		
	ELIG1	ELIG2	ELIG3	ELIG1	ELIG2	ELIG3
British Columbia	0.117 (0.212)	0.136 (0.191)	0.210 (0.230)	0.098 (0.215)	0.116 (0.197)	0.188 (0.235)
Alberta	0.276 (0.191)	0.179 (0.179)	0.539* (0.224)	0.220 (0.193)	0.073 (0.184)	0.414 (0.229)
Saskatchewan	0.721* (0.258)	0.603* (0.220)	0.703* (0.313)	0.743* (0.260)	0.630* (0.226)	0.766* (0.318)
Manitoba	0.538 (0.290)	0.816* (0.228)	0.727* (0.337)	0.487 (0.294)	0.745* (0.237)	0.670 (0.347)
Ontario	0.404 (0.211)	0.893* (0.191)	0.994* (0.311)	0.251 (0.218)	0.728* (0.199)	0.778* (0.317)
Quebec	0.534* (0.160)	0.257 (0.139)	0.513* (0.167)	0.538* (0.162)	0.276 (0.142)	0.537* (0.170)
New Brunswick	0.433* (0.175)	0.199 (0.158)	0.875* (0.219)	0.452* (0.176)	0.230 (0.158)	0.882* (0.220)
Nova Scotia	0.778* (0.326)	0.499 (0.309)	0.974* (0.366)	0.794* (0.329)	0.532 (0.313)	0.960* (0.371)
Prince Edward Island	0.295 (0.295)	-0.104 (0.276)	1.752* (0.563)	0.303 (0.296)	-0.077 (0.278)	1.796* (0.565)
Newfoundland	0.431* (0.170)	0.300 (0.165)	0.617* (0.244)	0.378* (0.171)	0.274 (0.165)	0.587* (0.245)

1/ * indicates significant at the 5% level.
2/ Source: 1988-1990 LMAS Longitudinal Files.
3/ "No Controls" includes a step function in duration and a year indicator.
4/ "With Controls" includes a step function in duration, a year indicator plus controls for the provincial unemployment rate, hourly earnings, age, education, gender, marital status, school attendence and past UI receipt.
5/ The step function in duration includes dummy variables for weeks 2 through 14 as well as for weeks 15-16, 17-18, 19-20, 21-25, 26-30, 31-40, 41-50, 51-60 and 61+.
6/ Standard errors in parentheses.

Table 5: Eligibility Parameters Using Insured Weeks to Determine Eligibility
(Step Function)

	No Controls			With Controls		
	ELIG1	ELIG2	ELIG3	ELIG1	ELIG2	ELIG3
British Columbia	0.105	-0.011	-0.019	-0.001	-0.215	-0.167
	(0.217)	(0.129)	(0.127)	(0.217)	(0.135)	(0.136)
Alberta	0.277	-0.209	-0.272*	0.156	-0.460	-0.459*
	(0.186)	(0.108)	(0.130)	(0.187)	(0.116)	(0.140)
Saskatchewan	0.427	0.074	-0.317	0.337	-0.075	-0.393*
	(0.256)	(0.144)	(0.175)	(0.257)	(0.155)	(0.189)
Manitoba	0.475	0.492*	0.119	0.339	0.322	-0.037
	(0.301)	(0.142)	(0.189)	(0.304)	(0.154)	(0.206)
Ontario	0.049	-0.028	-0.482*	0.156	-0.309*	-0.594*
	(0.184)	(0.090)	(0.150)	(0.185)	(0.098)	(0.159)
Quebec	0.315*	0.022	-0.076	0.234	-0.094	-0.120
	(0.153)	(0.104)	(0.105)	(0.152)	(0.107)	(0.112)
New Brunswick	0.126	-0.149	-0.037	0.094	-0.180	-0.026
	(0.165)	(0.114)	(0.114)	(0.164)	(0.118)	(0.123)
Nova Scotia	0.357	-0.279	-0.148	0.197	-0.436*	-0.344
	(0.293)	(0.210)	(0.221)	(0.290)	(0.218)	(0.230)
Prince Edward Island	0.145	-0.242	-0.457	0.133	-0.192	-0.358
	(0.263)	(0.236)	(0.253)	(0.262)	(0.239)	(0.270)
Newfoundland	0.203	0.011	-0.067	0.005	-0.213	-0.331*
	(0.152)	(0.129)	(0.143)	(0.148)	(0.128)	(0.148)

1/ * indicates significant at the 5% level.
2/ Source: 1988-1990 LMAS Longitudinal Files.
3/ "No Conrtols" includes a step function in duration and a year indicator.
4/ "With Controls" includes a step function in duration, a year indicator, plus controls for the provincial unemployment rate, hourly earnings, age, education, gender, marital status, school attendence, and past UI receipt.
5/ The step function in duration includes dummy variables for weeks 2 through 14 as well as for weeks 15-16,17-18, 19-20, 21-25, 26-30, 31-40, 41-50, 51-60 and 61+.
6/ Standard errors in parentheses.

Table 6: Information Matrix Tests for Unobserved Heterogeneity

	Quartic		Step Function	
	Eligibility Determined by:		Eligibility Determined by:	
	Duration	Insured Weeks	Duration	Insured Weeks
British Columbia	1.868 (0.172)	1.925 (0.165)	2.017 (0.156)	2.257 (0.133)
Alberta	0.671 (0.796)	0.939 (0.333)	0.546 (0.460)	0.964 (0.326)
Saskatchewan	3.591 (0.058)	3.132 (0.077)	3.274 (0.070)	2.867 (0.090)
Manitoba	5.232 (0.022)	4.877 (0.027)	5.231 (0.022)	4.794 (0.029)
Ontario	3.507 (0.061)	2.484 (0.115)	2.973 (0.085)	2.744 (0.098)
Quebec	7.638 (0.006)	7.862 (0.005)	7.595 (0.006)	7.565 (0.006)
New Brunswick	8.372 (0.004)	8.510 (0.004)	7.422 (0.006)	8.298 (0.004)
Nova Scotia	0.019 (0.890)	0.083 (0.774)	0.084 (0.772)	0.087 (0.769)
Prince Edward Island	4.064 (0.044)	2.745 (0.098)	3.900 (0.048)	2.949 (0.086)
Newfoundland	5.108 (0.024)	4.176 (0.041)	5.106 (0.024)	5.890 (0.015)

1/ All test statistics are distributed $\chi^2(1)$.
2/ The test is calculated for "With Controls" specification.
3/ Probability values in parentheses.