UI

Unemployment Insurance and Employment Durations: Seasonal and Non-Seasonal Jobs

by David A. Green
and Timothy C. Sargent

Human Resources Development Canada
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UI Impacts on Worker Behaviour

Canada

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****This version has been reformatted 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

This paper examines whether the UI system induces significant distortionary impacts on job spell durations for seasonal and non-seasonal employment. This employment distinction is made for policy and theory reasons. Different groups of users may place different values on UI benefits if they provide temporary replacement income in an unemployment spell that is predicted perfectly in advance versus one that is unforeseen. Evidence of significant UI effects on job durations would have different policy implications for seasonal and non-seasonal jobs. From theory, one would predict a disproportionate number of employment spells will end just at the entrance requirement. Further, the probability that a spell ends in a given week will be higher for weeks following the entrance requirement relative to weeks before. Given that seasonal workers make their employment decisions within a one-year time frame, a comparison with non-seasonal UI claim patterns is useful to determine whether there are concentrations of job durations at the entrance requirement and qualification for maximum benefits, and whether the chances of job terminations increase for all weeks after the entrance requirement point is reached.

The evaluation approach is to make use of differences in the UI program across 48 UI Regions in 1989 to identify the effects of UI on job durations. We use a large sample of jobs initiated in 1989 taken from the Labour Market Activity Survey (LMAS), a large representative sample of Canadian workers. Using a hazard function model, we estimate the effects of the entrance requirement, the maximum entitlement and maximum year points which varied with the regional unemployment rate, in addition to other covariates. We find that there is no evidence of an entrance requirement effect on employment durations for seasonal jobs. However, there is strong statistical and economic significance (as many as 1 in 50) for seasonal job spells ending at the maximum year point, at which workers qualify for enough weeks of benefits to full the remainder of a 52 week period. Further, the main increase in the probability of job termination occurs after the job has lasted beyond the maximum year point. This may indicate significant tailoring of seasonal jobs to the UI system. For non-seasonal jobs, there is also statistically significant evidence of concentrations of job terminations at the entrance requirement and maximum entitlement points. However, these effects are small in absolute terms, corresponding to changes in the probability of job termination of much less than 1 per cent.

Introduction

One of the key issues in recent debates over publicly provided unemployment insurance (UI) is the question of whether the Canadian UI system is in fact insurance at all. Some researchers claim that UI in Canada is used more as a (probably highly imperfect) redistributional tool than as true insurance. (See the report of the Macdonald Commission and some of the discussion of Green and Riddell (1993a) in the volume of the *Journal of Labor Economics* in which the latter appears.) The very existence of the special fishing benefits lends credence to this point of view. To some degree, though, this is an empirical question.

One of the key issues in recent debates over publicly provided unemployment insurance (UI) is the question of whether the Canadian UI system is in fact insurance at all. Some researchers claim that UI in Canada is used more as a redistributional tool than as true insurance.

The argument that the Canadian system is not truly an insurance system hinges, in part, on claims that individuals who knew in advance that they would be unemployed or who planned their unemployment spells themselves receive benefits. For such individuals, UI benefits would not constitute insurance in the sense of allowing them to smooth their consumption across unexpected states of the world. To the extent that benefits are received by seasonal workers who would be in seasonal jobs regardless of the UI system, UI benefits may well act as a straight transfer.[1] In that case, we might want to reconsider the structure of the UI system with respect to these workers.

Of more concern, though, is the possibility that some individuals adapt their behaviour to the UI system. One way this could happen is through increased numbers of individuals choosing seasonal work patterns; arising because UI essentially subsidises seasonal work patterns. The second way this could happen is through planned termination of job spells in order to allow the worker to obtain UI spells. This latter effect implies inefficient allocations of work. In this paper we attempt to provide evidence on the latter effect by examining the impact of various parameters of the UI system on the distribution of job spell durations in 1989. We find statistically significant (though not in all cases economically significant) distortionary impacts of the UI system on job termination. Our analysis treats seasonal and non-seasonal jobs separately, both because the theoretical impact of UI on the two groups differ and because the policy analysis for the two groups should differ.

Part of our concern in this paper is to understand the predictions of basic economic theory for the impact of Canada's UI system on job spell durations. Having derived these predictions, we then proceed to search the data for evidence that they are or are not met. The theories we examine predict a disproportionate number of job terminations at the number of weeks at which an individual first qualifies for UI benefit receipt and at the number of weeks at which the individual qualifies for the maximum possible number of weeks of UI benefit receipt. Further, if individuals face a fixed horizon over which they plan their work and leisure time (as might be true of seasonal workers), theory predicts a further concentration of job terminations at the point where the individual qualifies for just enough UI benefits to last to the end of the planning horizon. Workers recognize that each added week worked past that point means one week less of UI benefit receipt before the end of the planning horizon (that is, before the start of next

1 One should note, however, that any such transfer may not be to the workers themselves if higher UI benefits lead merely to a lower compensating differential for the seasonal nature of the job.

There is no evidence of an entrance requirement effect in job durations for jobs ending in seasonal layoffs. But there is evidence of an effect at the point at which workers qualify for just enough weeks of benefits to fill the remainder of a 52 week period.

year's season in the case of seasonal workers). Because seasonal workers can be argued to face a fixed planning horizon of one year while other workers do not face any definable horizon, these latter concentrations should be evident for seasonal workers but not for others. For this reason, we divide our sample between seasonal and non-seasonal workers.

It should be noted that all of these predictions can be generated either in a straight supply side model, as is often done, or in a model which includes decisions by the firms. Use of the straight supply side models tends to put the responsibility for distortionary responses to UI on workers, but Canada's UI system creates strong incentives for firms as well. The patterns of response to the UI system we observe probably reflect both firm and worker reactions to incentives.

In our empirical work, we use data from the 1988-1990 version of the Labour Market Activity Survey (LMAS). This supplement to the Labour Force Survey is large longitudinal sample of Canadian workers includes detailed information on personal and job characteristics. We were fortunate to gain access to a version of the LMAS in which an individual's place of residence is coded as the UI region—and not just the province—in which the person lives. Since some UI parameter values are determined by unemployment rates at the UI regional level, this level of geographic coding is essential to accurately analyze the Canadian system. As argued in the theoretical section of the paper, a careful analysis also requires that data be divided between seasonal and non-seasonal jobs. Since there is no variable denoting seasonal jobs on the LMAS, we make use of recorded reasons for terminations and carry out a separate analysis for jobs ending in seasonal and non-seasonal layoffs. Quits imply a specific set of complications in this taxonomy and are discussed in the body of the paper.

Our main conclusions from this exercise are as follows. First, there is no evidence of an entrance requirement effect in job durations for jobs ending in seasonal layoffs. But there is evidence of an effect at the point at which workers qualify for just enough weeks of benefits to fill the remainder of a 52 week period—the H_{mxyr} point— and for subsequent weeks. One set of estimates suggests that almost 1 in 20 seasonal jobs ends at the H_{mxyr} point because of the incentives of the UI system. Effects of the UI system are also visible in jobs ending in non-seasonal layoffs and quits. However, the size of these effects in absolute terms (measured relative to all job starts in 1989) appears to be small.

There is evidence to support the popular notion that some people are adapting their behaviour to the UI system and not using it as insurance. However, estimates suggest that for non-seasonal workers the actual size of these effects may be small. Importantly, if we are concerned about non-insurance use of the system, there is little evidence that the entrance requirement is the parameter of the UI system on which we should focus.

This paper proceeds in five sections. The first contains a description of the Canadian UI system as it existed in 1989. This is necessary to understand the discussion of potential effects of the UI system and the description of the empirical approach and results that follow. In the second section, we present a discussion of the potential impacts of the UI system on employment durations in the context of three theoretical models. In the third section, we detail our data. The fourth section presents the empirical methodology and results, and the fifth section contains a summary and conclusions.

1. The Unemployment Insurance System in 1989

The Canadian Unemployment Insurance (UI) system in 1989 can be characterized using a few primary parameters: the benefit rate, the entrance requirement, parameters relating to the calculation of weeks of entitlement to UI benefits, and parameters relating to disqualification. Given our interest in the effects of UI on employment durations, we are primarily interested in the parameters relating to the entrance requirement and the calculation of entitlement. Several of these parameters vary with the unemployment rate in the UI region in which a recipient resides. With the exception of P.E.I., UI regions are sub-provincial geographic constructs. There were 48 such regions in 1989 and within each one the basic parameters of the UI system were constant across individuals.

To qualify to receive UI benefits, individuals had to have worked a specified minimum number of weeks in a UI-eligible job in the qualifying period. The qualifying period was defined as the 52 weeks directly preceding the filing of the claim, or the number of weeks since the start of the individual's last claim, whichever was shorter. A UI-eligible job was specified as a job of at least 15 hours per week on which the individual earned at least 20 per cent of the maximum weekly insurable earnings. Self employed work and unpaid family or volunteer work could not be used to generate eligibility. The entrance requirement for UI varied according to the unemployment rate in the individual's UI region as follows:

Regional unemployment rate (per cent)	Minimum weeks to qualify (number of weeks)
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

Special entrance requirements existed for individuals defined as either repeat users of the UI system or new entrants to the labour force. New entrants (also called re-entrants) are defined as individuals who worked fewer than 14 UI-eligible weeks in the 52 weeks prior to the qualifying period. They needed a minimum of 20 weeks of employment to qualify for benefits. Repeat users were individuals who had collected UI benefits during the qualifying period for the current claim. They faced a schedule to determine eligibility based on the unemployment rate in their UI region and the number of weeks of UI benefits they had collected in their qualifying period. The entrance requirement for repeat users varied from 10 weeks (for claimants in regions with unemployment rates of over 11.5 per cent, regardless of benefits collected) to 20 weeks (for claimants in regions with unemployment rates of less than 6 per cent who had collected 20 or more weeks of UI benefits in the qualifying period).[2]

Special entrance requirements existed for individuals defined as either repeat users of the UI system or new entrants to the labour force.

2 The full schedule is given in Dingledine (1981), Appendix 1. We use this schedule in calculating entrance requirements for individuals we identify as repeat users in our data set.

Once an individual had qualified to collect benefits, the number of weeks of entitlement were calculated according to a three-phase system.

1. Phase I, an initial benefit phase, provided one week of benefits for each week of insured employment, to a maximum of 25 benefit weeks;
2. Phase II, a 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
3. Phase III, a regional extended benefit phase, provided for two weeks of benefits for each half-a-percentage point increment in the regional unemployment rate in excess of 4.0 per cent, to a maximum of 32 weeks of benefits.

The maximum duration of benefits from all three phases was 50 weeks. Once the two-week waiting period is included, during which no benefits were paid, the maximum total length of a claim was 52 weeks. The two-week waiting period could be extended by up to six weeks for claimants who quit their previous job without just cause. (In such cases, the clock on the claim would run but no benefits were paid during this period). The actual length of the penalty for quitting was discretionary.

The benefit rate in this period was 60 per cent of average weekly earnings in the qualifying period up to a specified maximum weekly earnings level, with benefits being constant for higher earnings levels. It is important to note that benefits were based on weekly earnings, not hourly wages, and thus workers could increase their benefits by working longer hours in a week at a given wage.

Within this system, a regional unemployment rate of 11.5 per cent marks a notable cut-off. In regions with unemployment rates above this cut-off, weeks of qualification under the regional extended benefit phase were at their maximum (32), the entrance requirement was at its minimum (10), and repeat-user provisions were suspended. Thus, variations in the unemployment rate above this threshold did not cause changes in the key parameters in the system. We will refer to these regions—UI regions with unemployment rates above 11.5 per cent—as maximum entitlement regions.

2. Theoretical Models

In this section we discuss the impact of the UI system described above on employment decisions in the context of three main theoretical models: static labour supply models, search models, and implicit contracting models. A summary of the conclusions from the discussion are given at the end of the section. We begin with the static labour supply model because it provides the clearest exposition of the potential effects of UI.

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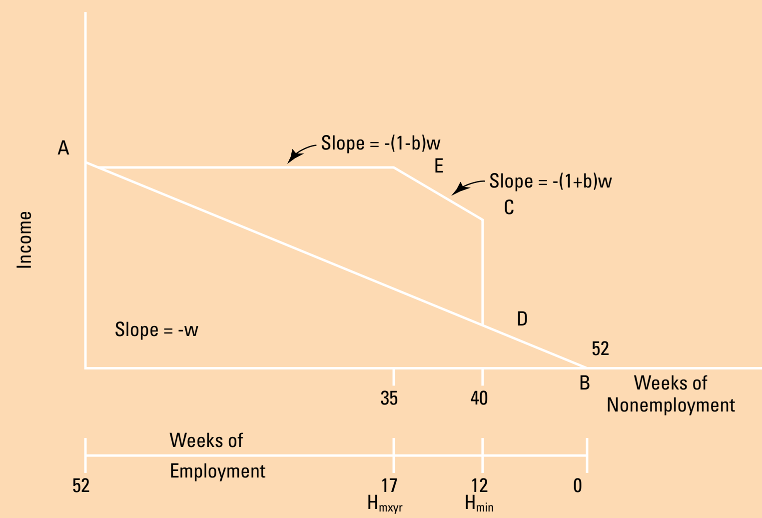
Consider an individual faced with a decision over their preferred combination of weeks of work and level of consumption in a specified period, for example, one year. In the static labour supply model, the individual makes a decision by maximizing his or her utility subject to a budget constraint. The model is static in the sense that only the specified period is under consideration: there is no reason to save, no possibility of consuming more than income in the given period, and no possibility of obtaining training in order to change outcomes in future periods. Preferences for the individual are specified as, $U(C,L)$, where C is the level of consumption of an aggregate good, L is weeks of leisure, and $U(.,.)$ is a utility function. The utility function allows the individual to rank all combinations of C and L according to the utility they bring. We assume that higher levels of both C and L lead to higher utility.

The other main component of the individual's decision is their budget constraint. This shows the set of combinations of C and L that the individual can just afford given their wage, w; their non-labour income other than UI benefits; and their potential UI benefits as determined by the parameters of the UI system. In specifying the budget constraint, the selection of the relevant decision period is critical. To our knowledge, all previous empirical studies of UI that consider the affects on the budget constraint have assumed a one-year time horizon.

To see the effects of this assumption, consider an individual who has never claimed UI benefits before but who has been in the labour force for at least the last two years (and is neither a repeat user nor a new entrant) and who lives in a UI region with an unemployment rate of 8.0 per cent. First, assume that this individual has a one-year time horizon. The relevant budget constraint is given in Figure 1, assuming that individual has no non-labour income other than (potentially) UI benefits. Note that weeks in the year can either be sold as labour or consumed as leisure; there are no part-time work weeks. Thus, if there are L weeks of leisure in the year, there are $H = 52 - L$ weeks of labour supply. In the figure, weeks of labour supply are read from right to left with $H=0$ being at $L=52$. Define H_{min} as the entrance requirement number of weeks, and AB as the budget constraint in the absence of UI. If the individual works fewer than H_{min} weeks, then his or her level of consumption just equals his or her take-home pay, which is the same as what would be received in the absence of UI. At H_{min} weeks, the individual qualifies for UI and the budget constraint jumps up by the value of the UI benefits he or she could collect at that point. In this example, $H_{min}=12$ and thus the individual is entitled to 12 weeks of benefits under Phase I of the schedule described above. Given the unemployment rate, the individual is entitled to a further 16 weeks under Phase III for a total of 28 weeks. If claimant's weekly wages are below the maximum ceiling then the jump in the budget constraint equals $0.6*w*28$. An individual who had worked exactly 12 weeks in the year, therefore, would have total income, expressed as $12*w + 0.6*w*28$ (and thus total consumption since there is no saving). This is the height of the budget constraint at 12 weeks.

Figure 1
Budget Constraint Region with an Unemployment Rate of 8%
One Year Horizon



For weeks directly following H_{min}, each added week of work yields a return of $1.6*w$, as the individual earns not only his weekly wage but entitlement to an additional week's benefits under Phase I. At 17 weeks the individual has qualified for 33 weeks of benefits. Once the two-week waiting period for UI is added in, the total of weeks worked and weeks of potential claim add up to 52 at this point. With a one-year time horizon, the individual recognizes that for every week he works beyond this point he earns an added week's pay, but can collect one week fewer in UI benefits in the time remaining to the end of his horizon. Beyond 17 weeks, therefore, the budget constraint has a slope of only $(1 - 0.6)*w$. Points beyond which workers recognize a penalty to added employment in terms of foregone UI benefits receipt in their time frame, such as the 17 week point in this example, we will call maximum year points and will label as H_{mxyr} on all figures.

Individuals choose the point on the budget constraint that provides maximum utility. Presented in a diagram, individual utility is illustrated by indifference curves, curves connecting all combinations of *C* and *L* that yield the same level of utility. Indifference curves farther from the origin correspond to higher utility levels. The solution to the problem, then, is to find the point at which the highest possible indifference curve just touches the budget constraint. Comparing the budget constraint in the absence of UI (the line *AB* in Figure 1) with the budget constraint in the presence of UI, one can see that introducing UI changes the constraint in a way that may lead individuals to select a different *C* and *L* combination as their point of highest utility. This new selection is caused by the incentives inherent in the UI system, which are captured in the shape of the budget constraint. The incentives created by this budget constraint are straightforward and have been analyzed in many papers and books. (See Mortensen (1990), Phipps (1990,1991), Moffitt and Nicholson (1982), Gunderson and Riddell (1993).)

For individuals who in the absence of UI would work fewer than H_{min} weeks or no weeks, the jump in the budget constraint at H_{min} weeks provides an incentive to increase weeks worked (or to enter the labour force) just enough to qualify for UI. On the segment from 12 to 17 weeks two offsetting effects occur. First, the increased height of the budget constraint due to the collection of UI benefits induces an income effect, pushing workers towards working fewer weeks if leisure is a normal good. In essence, workers can afford to take more time off because that time is compensated. Secondly, the increased slope of the budget constraint due to the combined effects of earning wages and increasing entitlement weeks induces a substitution effect that leads to more weeks worked. In essence, taking a week off in this range of weeks is more costly because the return to working an extra week is more than just the wage received. The net effect in this region is uncertain. Beyond 17 weeks, both the income effect of the UI benefits and the substitution effect induced by the lower net wage lead to negative effects on weeks worked. Over this latter range of weeks, the individuals view themselves as earning a lower wage on the job because they recognizes that with an extra week of work they may earn their weekly wage but are giving up one added week of UI benefits they could collect in the year.

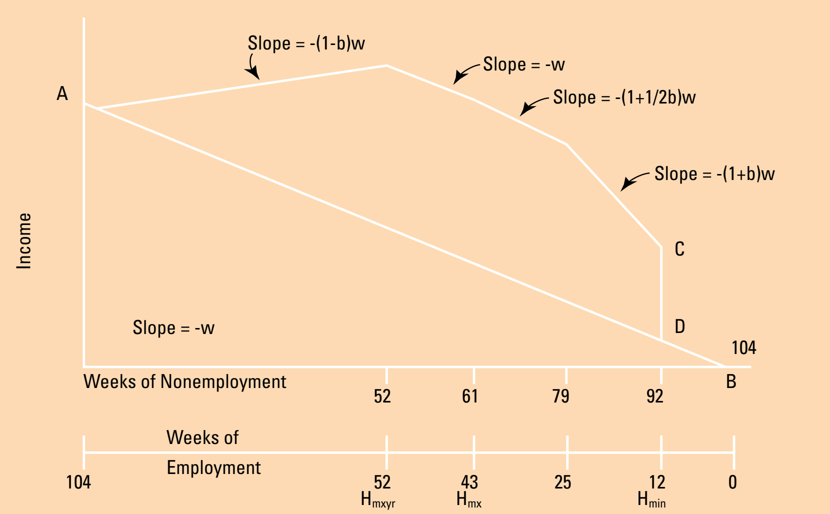
A few comments on the use of the static labour supply model are useful. First, it is important to note that this is a pure supply side model. Individuals can choose their weeks of leisure, and concurrently their weeks of labour supply, exactly and without restriction. There is no uncertainty in weeks worked and no involuntary unemployment. Secondly, it is a pure static model. Long run costs to taking extra weeks of leisure in the current period, such as erosion of skills or sending a bad message about working attitude to future employers, are not present. Thirdly, a point related to the previous two, it is not exactly clear what role UI is meant to play in this system apart from a pure transfer: there is no uncertainty over employment to require insurance and, since it is a static model, no need to help individuals facing borrowing constraints. In spite of these points, the model is still powerful and useful in summarizing the incentives inherent in the system. The danger in using this model, however, is that it would be easy to slip into viewing it as a model of the world and thus assign the worker the responsibility for labour market responses to the variables. (We return to this point later in the paper.)

But how do we move from this model to predictions for patterns of job duration? As we have seen, the incentives in the system push some individuals to lengthen their weeks worked in a year in order to qualify for UI. These individuals would work exactly H_{min} weeks (since they are trying to just qualify). Further, the income effect (the response to the fact that leisure time is now compensated) will lead some who would worked more than H_{min} weeks in the absence of UI to shorten their working year. These individuals would not choose to work fewer than H_{min} weeks, however, since they would not then qualify for benefits. Thus, a number of those who would work more than H_{min} weeks in the absence of UI will also choose exactly H_{min} weeks. For these reasons, one would expect to see a disproportionate number of jobs ending at exactly at H_{min} weeks. Further in the paper, we refer to this bunching of job durations at H_{min} weeks as a spike at that point.

In addition to this bunching at H_{min} weeks, several other predictions are possible. First, one would expect to see a low number of jobs ending just before the entrance requirement, say at H_{min}-*1* weeks. This is because as long as the job is not extremely onerous, the gain to working just one more week in the year (access to 28 weeks of UI benefits in the example above) is likely to far outweigh the disutility to the extra week of work. Secondly, the income effect discussed above implies that more jobs will be terminated at weeks above H_{min} than in the absence of UI. Thus, not only will there be a bunching of durations at H_{min}, but many longer job spells will also be shortened so that the individuals can spend part of their time in compensated leisure. Finally, both the income effects and the substitution effects mentioned earlier lead individuals who would otherwise work over 17 weeks in the above example to shorten their weeks worked. This may lead to a number of workers choosing $H=17$ and thus to a second bunching at the H_{mxyr} point. The predicted pattern of the proportion of jobs ending at each possible week in the year is given in Figure 2.

The empirical work below will be conducted in terms of hazard rates. The hazard rate at any week, is the probability that a job spell ends in at duration x given that the spell lasts at least as long as x. Thus, the hazard rate at H_{min} weeks is calculated by dividing the number of jobs that last exactly H_{min} weeks by the number of jobs that last H_{min} weeks or more. This is directly analogous to examining the simple proportion of jobs ending at each number of weeks, but provides some attractive empirical properties discussed below. The pattern in the hazard rate

Figure 2
Budget Constraint Region with an Unemployment Rate of 8%
Two Year Horizon



predicted by the static labour supply model with a one-year time horizon is essentially the same as in Figure 2.

Now consider the same worker in the same region but with a two-year time horizon. In our example we will assume that any employment occurs in one continuous spell. This rules out the possibility that the individual was using multiple UI spells to maximize UI benefit receipt. A worker who divided up employment spells in this way was in effect adopting a shorter time horizon. The budget constraint for this situation is depicted in Figure 2. As in Figure 1, there is a jump in the budget constraint at $H_{min}=12$ weeks. However, the individual no longer faced the constraint of using up eligible weeks of UI benefits within a year and thus no kink occurs at 17 weeks. Instead, the individual continued to earn their weekly wages and qualify for an extra week of benefits up to 25 weeks, the point at which Phase I of the benefits schedule ends. For weeks of work beyond 25, individuals earned their weekly wage and qualified for added weeks of benefits at the lower, Phase III rate of one week of benefits for every two weeks of work. At 43 weeks, given the regional extended benefit weeks, individuals will have qualified for the maximum of 50 weeks of benefits they can collect.

Beyond that point, the return to an added week of employment is simply w, since UI benefit income for the period can be increased no further. As in the earlier example, there is a kink in the budget constraint induced by the realization that beyond a certain number of weeks of work (52 in this case), every added week of work is one less week of benefits that can be collected in the specified time horizon. Thus, what we previously defined as H_{mxyr} equals 52 in this case.

As in the one-year time horizon case, the incentives associated with the budget constraint in Figure 2 suggest that we should see employment duration spikes at each of the kink points on the budget constraint. Thus, the hazard rate out of employment should show spikes at 12, 25, 43 and 52 weeks. The important point is that, with the exception of the entrance requirement effect, the predicted spikes in the employment hazard in this example are different from those in the previous example. Unfortunately, however, these differences arise purely from differences in assumed planning horizons: something which we cannot observe, even if they do exist and are the same for everyone.

Both the one- and two-year horizon cases just derived assume a two-week waiting period. This waiting period implies that the budget constraint facing an individual who works 51 or 52 weeks in the year (in the one-year time horizon case) is the same whether or not there is a UI system. This is true because individuals who separate from employment at 51 weeks will spend the remainder of the year in the waiting period: never actually receiving benefits before the end of the year. As mentioned above, the waiting period for individuals who quit their previous job without cause was extended by up to six weeks under the 1989 system. In the case of the maximum penalty, the spike corresponding to H_{mxyr} would occur six weeks earlier in the employment spell. Thus, quitters should be treated differently than individuals who are laid off.

The difficulty for our empirical work is that the penalty for quitting could vary from a zero- to 6-week increase in the waiting period. There is nothing in our data to allow us to even guess the length of penalty to be applied in each case.

Some individuals who, in the absence of a UI system, might decide to terminate a job at a duration less than H_{min} weeks might, with UI, decide to extend the job until they have qualified for benefits.

Moreover, the workers themselves will not know what the penalty is until they enter the UI office. Even if we did know the actual penalty to be applied, workers' actions in terminating jobs might be based on some other guessed number.

In the empirical work that follows, we treat quits separately from layoffs. Note that quitting does not affect the location of the kink associated with the entry requirement nor with the maximum entitlement point.

Search models provide an alternative viewpoint from which to examine UI effects on employment duration. These models differ from the static labour supply model in that individuals do not make decisions for a pre-specified period. Rather, they re-make their decisions each week of their life based on available information. These decisions must be re-considered each week because, in contrast to the static labour supply model, the search model includes uncertainty. In the simplest model relevant for our problem, an individual is observed after just having taken a job at weekly wage, w. In each week thereafter, the individual must decide whether to remain on that job or to become non-employed. He makes this decision by comparing his predicted utility for the rest of his life if he remains on the current job versus if he leaves this job and undertakes non-employed search for a new job. At the outset of the current job the individual's best option is to remain on the current job. We know this because he chose to take the job in the first place. In each subsequent week, however, new information about the job could be revealed that could lead the worker to revise his or her opinion of the value of the job. For example, a drop in demand for the firm's output could lead to a downward revision in the wage, causing the worker to leave the job.

A UI system with an entrance requirement can directly affect the expected duration of a job in this model. Consider an individual making a decision on whether to separate before he has reached H_{min} weeks on the job. The value of taking the non-employment option is based on consumption out of whatever savings the individual has built up and the expected value of the non-employed search (based on the individual's expectation of being able to find a better job). For the same individual making a decision in a week after H_{min}, the value of non-employment includes consumption out of UI benefits which he is now qualified to receive. Thus, with a higher value for the non-employment option after H_{min}, we would expect to see more job terminations after H_{min} weeks of duration than before.

Further, some individuals who, in the absence of a UI system, might decide to terminate a job at a duration less than H_{min} weeks might, with UI, decide to extend the job until they have qualified for benefits. Thus, as with the static labour supply model, one would predict a spike at the H_{min} point and a higher probability of a job ending after H_{min} weeks than before. For similar reasons to the static labour supply model, one would also predict a spike at H_{mx} weeks, the number of weeks of work at which an individual qualifies for their maximum possible weeks of UI receipt. What the model does not yield is the prediction of a spike at the H_{mxyr} point. Indeed, with no natural fixed decision time frame, the H_{mxyr} point does not even exist in this model.

Finally, we consider the effects of the UI system in the context of an implicit contract model. In both models above, the decision to terminate a job is entirely the worker's. A model that incorporates decisions by the firm seems more plausible, especially in the context of a system in which the reporting of the termination of a job to UI authorities is done by the firm.

Thus, consider a model in which a firm operates in a locality with a fixed pool of N possible workers. The firm operates in an uncertain environment where demand for its product could rise or fall in a given year. The average value and variation of demand is known but the exact value it will take in a given year is not. The firm draws up a contingent plan for each possible level of realized demand. The capital stock of the firm is taken as fixed but the firm can adjust to different levels of demand by adjusting the weekly wage, w, the weeks per year per worker, H, and the number of people employed, L.

As a simple starting point, we assume a production technology such that the firm is indifferent to whether it adjusts to a demand change by changing the number of weeks per worker or the number of workers. This is not a simple demand side model, however: whatever adjustments the firm makes must take into account the preferences of the N possible workers. In particular, assume that these individuals have some opportunity outside being in this firms labour pool and that the outside opportunity brings them a level of utility, U^*. For the firm to keep the workforce it needs for production, it must ensure that all of its choices permit workers to have at least U^* in utility. Thus, the workers and firms form a contract (assumed to be only implicit outside the union sector) which specifies w, H and L for each possible level of demand. The contract is set so that the firm maximizes profit subject to the workers getting at least U^* in utility. The contract is created before the level of demand is known. Once that level is known, the firm and workers just carry out the relevant part of the contract.

It is worth noting that in this standard model, if the firm is indifferent to risk while the workers are averse to risk then the employment contract may be used in part as insurance. Recall that the firm is indifferent to whether adjustments are made in terms of L or H. The workers, however, are not indifferent: adjustments in terms of L (layoffs) lead to much greater variation in consumption and thus much greater uncertainty. The workers would essentially be willing to take a wage cut in order to ensure that all adjustment is made through H. The firm, given its indifference in this regard, is willing to guarantee no layoffs in return for a wage cut.

What happens when we introduce a realistic UI system into this model? First, note that for any given wage and a fixed contract period of, say, one year, the individual workers' budget constraints will be the same as in the static labour supply model. In essence, the workers' constraints and preferences that we studied in the labour supply model will also be reflected in the contracts being discussed here. As an example, suppose that for some level of demand the firm is choosing between hiring workers for H_{min} weeks during the year versus H_{min}-1 weeks. In either case, the workers must achieve U^* level of utility. But in one case the workers get to collect UI benefits on top of their pay, and in the other case they do not.

As a result, the firm would have to pay workers hired H_{min}-1 weeks a higher weekly wage than those hired for H_{min} weeks in order to compensate for the lack of income from UI. In the example used with the static labour supply model

With UI, seasonal firms' workers receive compensation when not working, thereby reducing the extent to which seasonal firms have to pay a wage premium to attract workers.

where individuals qualify for 28 weeks of benefits as soon as they meet the entrance requirement, the wage for non-UI receiving workers would have to be substantially higher. Thus, the incentives for individual workers reflected in the budget constraints in the static model effectively become incentives for the firm. For this reason, the implicit contract model predicts the same pattern spikes as in the static model, including a spike at the relevant H_{mxyr} point for the contract period.

The implicit contract model yields several other interesting conclusions. First, as stressed by Feldstein (1976) and others, a UI system without experience rating causes non-seasonal industry firms to subsidize those in seasonal industries. In a world without UI, seasonal firms may have to pay weekly wages that are high enough to compensate their workers for the fact that they work only part of the year. If they did not do so, they would not meet the offers made by non-seasonal firms and risk losing their labour pool.[3]

With UI, seasonal firms' workers receive compensation when not working, thereby reducing the extent to which seasonal firms have to pay a wage premium to attract workers. Since seasonal firms do not have to pay extra to use the system this way, non-seasonal firms who do not make as much use of the system effectively subsidize the seasonal firms through allowing them to pay lower wages. Theoretically, then, a non-experience-rated UI system could lead to an imbalance in the industrial structure.

A second conclusion is that a UI system with an entrance requirement may also play a role in inducing layoffs by firms. The large increase in wages that is necessary if a firm offers $H<H_{min}$ implies that there is a large cost to adjusting weeks down in response to poor demand levels. The firm may then respond to poor demand by using layoffs. Since workers are risk averse, such a strategy has a cost in that firms would be obliged to offer higher wages to workers when they are employed. For some levels of risk aversion and specifications of the production function, however, it may still pay the firm to use layoffs. This is in contrast to the case without a UI system, where layoffs are not used and weeks per worker are adjusted to meet demand shocks. Thus, the UI system imposes a cost on the downward adjustment of weeks per worker at the H_{min} point which may imply that some firms respond to poor demand levels with a combination of employing some workers at H_{min} weeks and laying the rest off.

This implicit contract model is different from any other implicit contract model with UI benefits of which we are aware. In particular, other models specify the UI system as paying benefits to laid-off individuals ($H=0$ individuals) (see Feldstein (1976), Burdett and Hool (1983)). If one recognizes the qualification restrictions in the Canadian UI system, however, it becomes apparent that this is incorrect in this context. The UI system does not insure against layoffs (or rather, non-hires), but rather against variations in weeks worked per year. Thus, unlike the well known result in Feldstein (1976), UI does not encourage firms to use layoffs to adjust to variation in demand; it encourages work-sharing conditional

3 Note that it would not be necessary for a higher wage to be paid if a group of workers who have a strong penchant for leisure and would prefer to work in a seasonal industry where they can get part of the year off.

on minimum weeks of work restrictions being met. Workers are still more willing to share work because they are insured by the government, than opt for variations in weeks of work. In this model, a Canadian style UI system does not encourage layoffs; but it encourages seasonal layoffs. However, it does impose costs which may induce layoffs as well.

Theoretical Model Implications

Static Labour Supply Model

1) Spikes in the hazard rate out of employment are predicted at H_{min} (the entrance requirement number of weeks), H_{mxyr} (the number of weeks at which an individual just qualifies for enough weeks of UI receipt to cover the remainder of their decision framework period), and H_{mx} (the number of weeks at which an individual qualifies for their maximum possible number of weeks of receipt of UI benefits).

2) With the exception of H_{min}, the location of these spikes and whether they are relevant depends on the time frame being used for decision-making (for example, one year versus two years).

3) The hazard rate should be higher for all weeks after H_{min} than those before.

Search Model

1) Spikes are also predicted at the H_{min} and H_{mx} points but because the model does not contain a fixed decision period, no counterpart to the spike at H_{mxyr} is predicted.

2) Again, the hazard should be higher after than before H_{min}.

Implicit Contracting Model

1) If the contract period is the same as the decision period in the static labour supply model, then the model predicts the same spikes at exactly the same points as the labour supply model. As before, it too predicts a higher hazard after H_{min} than before.

2) The model also predicts that the UI system subsidizes seasonal firms and that it introduces costs to downward adjustment of weeks of work below H_{min} that induce firms to use more layoffs, or rather, fewer hires.

A question of immediate interest for empirical implementation is whether it is reasonable to look for the H_{mxyr} spike in the data or whether we should assume that it is an unrealistic by-product of stylized models. For many workers it would seem to be an unrealistic theoretical prediction: most individuals would likely not act as though they have only a finite period within which to consume UI benefits. However, for workers and firms in seasonal industries, this might in fact be a reasonable prediction.

Consider a group of individuals who work and qualify for UI benefits during an annual "season" that has relatively regular dates each year. Given the qualification period of 52 weeks, these individuals will recognize that they must stop drawing benefits and take a job at the start of the next season if they are to maintain a pattern of work and UI use across the years. Even the assumption in the implicit contract model of a fixed pool of workers attached to the firm seems somewhat reasonable in this case because of the association of seasonal work with small communities. In this case, it is quite striking that the need to obtain enough work to qualify for UI benefits in a year implies a work-sharing among members of the community.

Such sharing is purported to occur in resource-based towns, especially in Atlantic Canada. Thus, we believe that the predictions of the implicit contract model with a one-year contracting period are useful for examining seasonal employment. The model may also be relevant in some sectors outside the traditional seasonal-resource sector if workers in other sectors maintain an annual pattern of UI use and employment.

For most other workers, though, we believe that the simple search model is more powerful; that workers in other sectors do not have a natural planning horizon around which a model can be built but move through time in an uncertain world somewhat like that in the search model. In this case, one would not predict a spike in the hazard at H_{mxyr} for non-seasonal workers. In the empirical work that follows we differentiate between seasonal and non-seasonal workers because we believe that different models are relevant for each.

3. Data

The data used in this study are drawn from the Labour Market Activity Survey (LMAS) for 1989. This is a longitudinal survey of approximately 60,000 individuals conducted as a supplement to the Labour Force Survey.

The sample was interviewed initially during the period from January to March 1989, about their 1988 labour market activities, and then re-interviewed in the early months of 1990 and 1991 about their labour activities in the preceding calendar year. The survey asks questions about both personal characteristics—including age, education, sex, family and immigrant status—and job-related characteristics for each job held in the previous year (for up to five jobs for each individual). The job characteristics included start and end dates of jobs (specified at the weekly level), usual hourly wage, usual hours per week, firm size and union status.

The public use version of the LMAS contains place of residence coded as the province in which the respondent lives at the time of the survey. Use of this variable in studies of the impact of the UI system is misleading because key parameters of the UI system are set at the UI regional level, and provinces typically include several such regions. We were fortunate to gain access to a version of the LMAS in which residence is coded at the UI regional level.[4]

In forming our final sample, we exclude individuals who were full-time students during the sample year because UI benefits are not paid for non-employment spells that include schooling. We also exclude individuals who were over age 65 because this age group was ineligible for UI benefits in 1989. For the remaining individuals, we retain all jobs on which the individual is a paid worker and which have a weekly wage of over $15.[5] Finally, the definition of a "job" in the LMAS allows for up to five non-permanent separations in the sample year, including temporary layoffs. Because UI benefits can be collected while on temporary layoff, we have split any job spell that includes temporary layoffs into separate spells before and after the layoff. We treat all job spells as independent observations. The only exception to this is in the calculation of UI entitlement described below.

At this point, we have a sample of job spells for 1989. Our final form of data selection is to cut all spells that begin before January 1, 1989, to avoid standard length bias sampling problems (Lancaster (1990), p.95). The resulting sample of new job starts in 1989 provides generalizable results if the process generating new job spells is stationary over time.[6] Our final sample contains a total of 8,902 job spells initiated in 1989. We censor all job spells at the end of 1989 because

We censor all job spells at the end of 1989 because changes in the Variable Entrance Requirement (VER) that took place in 1990 change the nature of UI effects on employment.

4 The version of the LMAS 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 draw off the sample.

5 Self employment spells and jobs on which weekly earnings were less than $15 did not count toward generating UI eligibility.

6 The alternative approach to the length biased sampling approach is to model the non-stationary spell generating process. Our approach avoids the strong assumptions involved in this modelling by making the strong stationarity assumption.

changes in the Variable Entrance Requirement (*VER*) that took place in 1990 change the nature of UI effects on employment. (See Green and Riddell (1993b) and Baker and Rea (1993) for analyses of the 1990 changes). The LMAS over-samples individuals from rural regions, but contains weights for each individual that permit correction for this non-random selection. We use the weights in calculating all tables in this paper and in forming the log likelihood function.[7] The specification of the likelihood function is detailed in Appendix B.

For reasons discussed in the previous section, we believe seasonal and non-seasonal jobs should be analyzed separately. An optimal data set would contain a variable capturing the seasonality status of each job. No such variable exists in the LMAS. However, we do know the reason for termination for each job that ends before the end of 1990, because of the longitudinal nature of the data. One of the reasons for a job termination is “seasonal layoff.” We treat this as equivalent to a variable denoting seasonal jobs. The difficulty with assigning a seasonality label in this way is that the value of the label is revealed only at the termination of the job. Thus, we do not observe the value of the label for jobs which are still ongoing at the end of the LMAS sample period (the end of 1990). However, with the extra year of data after our sample year, this does not pose a serious problem. We simply assume that any job spell initiated in 1989 that has not been terminated at the end of 1990 is not a seasonal job. Indeed, by definition a seasonal job will not have lasted for over one year. This approach entails the disadvantage of not being able to use 1990 data on job spells, but has the advantage of presenting a more confident division of the sample based on seasonality.

This definition of seasonal jobs relies on reporting by the worker at the time of the survey. This means, in part, that we avoid having to provide an explicit definition of what constitutes a seasonal job. Such a definition might be based on observed patterns of work in the relevant occupation and industry in previous years. The danger in trying to arrive at such a definition is that the definition itself might be endogenous with respect to the measurement of the UI system: we might end up defining seasonal jobs as jobs in industries where the pattern of employment suggests an adaptation to and use of the UI system. Our approach assumes that individuals know which jobs are seasonal from the outset, with the only difficulty being that they do not reveal that knowledge until the end of the job.

We define a non-seasonal job as any job for which the recorded reasons for termination is any type of layoff other than seasonal layoffs or which is still ongoing at the end of 1990. We assume that seasonal and non-seasonal jobs are separate entities: a job that starts out as a non-seasonal job cannot be converted to a seasonal job part way through. This makes the ensuing analysis more straightforward but involves the assumption that a job that is defined as seasonal at the outset cannot end with a layoff before the end of the “season” because, for example, a plant burning down. Note that one could still expect seasonal spells which end earlier than might have been predicted at the outset—as a result of poor demand or poor supply of the key input (such as tourists or fish, for example)—

7 We first normalize the weights so that the sum of the weights across individuals equals our true sample size.

be defined as seasonal layoffs by workers. The "season" in question was simply a short one.

Jobs ending in quits pose a problem for this taxonomy. Unless we assume that seasonal workers never quit jobs, we never get to see the "seasonal" label for some seasonal jobs because the worker voluntarily ends the job before the pre-ordained layoff at the end of the season. Thus, in the general statistical sense, our data is censored by the quits process. Our estimation approach addresses this problem directly.

The final exercise in constructing our data is to generate values for UI parameters for each job. We create three dummy variables that take values of 1 in the weeks of a job spell corresponding to the H_{min}, the H_{mxyr}, and the H_{mx} points, respectively. As noted in Section 2, Theoretical Models, the H_{min} and with it the H_{mxyr} points can vary depending on the individual's use of UI in the entitlement period. Unfortunately, although the LMAS contains data on UI receipt in each non-employment spell in a sample year, there appears to be under-reporting of UI use in this data (Baker and Rea (1993), Green and Riddell (1993b)). Given calculations that well over 90 per cent of unemployed individuals collect UI benefits (Green and Riddell (1993a)), we calculate UI usage by assuming that any non-employment spell of more than two weeks in duration and that followed a UI-eligible job contained receipt of UI benefits (for jobs on which the individual was not self-employed and which meet the minimum pay and hours requirements)[8]. Using calculated spells of UI receipt based on this definition, we calculate entrance requirements under the repeat user provisions for each job. We take into account the fact that the calculated UI usage in the entitlement period can change with each week in a job spell because the entitlement period itself changes as the job progresses.[9] Similarly, the entrance requirement and other key UI parameters dictating the location of potential spikes can change during a job spell if the unemployment rate in the individual's UI region changes during the job spell.

To fully capture the incentives facing individuals, we re-calculate the values of the dummy variables H_{min}, H_{mxyr}, and H_{mx} in each week of each job spell based on regional unemployment rates obtained from Statistics Canada. Finally, there is no stipulation in the UI rules that weeks of eligibility must be generated from weeks of employment on a single job. Thus, we calculate entitlement based on weeks in (nearly) continuous employment spells, regardless of how many jobs were incorporated in those spells. Details of the way entitlement and the various UI-related variables are defined are presented in Appendix A.

8 We do not impose the requirement that the preceding job spell be at least H_{min} weeks long, because individuals may generate UI eligibility over a series of job spells.

9 Recall that the entitlement period is defined as the previous 52 weeks before the initiation of a UI claim or the number of weeks since the initiation of the previous claim, whichever is shorter.

4. Empirical Results

Our goal in the empirical work in this paper is to examine the pattern of employment durations in 1989 for evidence of the effects of the UI system.

Our goal in the empirical work in this paper is to examine the pattern of employment durations in 1989 for evidence of the effects of the UI system. We are interested in whether spikes in the proportions of jobs ending in various weeks (for example, at the H_{min} point) actually exist, the size of any such spikes that do exist, and the determinants of the size of any spikes. We will also look for evidence of predicted patterns other than spikes (such as the declining probability of leaving an employment spell as H approaches H_{min}), but we expect that these more general patterns will be harder to detect.

The key null hypothesis, then, is that there is no evidence of spikes in the density of completed employment spells at the points where theory predicts spikes should exist, and thus no evidence that UI has systematic effects on job separation decisions. The alternative hypothesis is that such spikes (and other predicted UI induced patterns) do exist.

The simplest tool for examining the impact of the UI system on employment durations is the empirical hazard rate function. Also called the Kaplan-Meier estimator, it is a non-parametric estimator of the hazard rate defined above which accounts for right-censoring of spells.[10] 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 (which is to say spells that have neither been terminated nor censored at a shorter duration). With the empirical hazard, one can inspect the shape of the employment hazard without imposing a specific functional form which might distort conclusions.

As discussed in Section 2, we believe that seasonal and non-seasonal jobs ought to be analyzed separately. In Section 1, we define seasonal jobs as all jobs ending in seasonal layoffs and non-seasonal jobs as jobs which either end in any other type of layoff or are right-censored at the end of 1990. Jobs that end in quits are viewed as jobs for which the seasonality definition is uncertain. The numbers of spells of each type are: 1,581 job spells ending in seasonal layoffs; 5,960 ending in non-seasonal layoffs; and 1,352 ending in quits.

Seasonal layoffs comprise only about 20 per cent of all layoffs, and, as we will see shortly, personal and job characteristics associated with jobs that end in quits are closer on average to characteristics of jobs that end in non-seasonal layoffs. For this reason, we assume that none of the jobs ending in quits were actually at risk to end in a seasonal layoff when we calculate the empirical hazard. Since we also assume that jobs cannot change their seasonality status mid-course, we analyse jobs ending in seasonal layoffs as a separate sample. The restrictiveness of this approach is relaxed in the duration model below.

All jobs ending other than through seasonal layoffs are grouped together under the rubric of non-seasonal jobs. However, for UI related reasons alone (see

10 In samples with fixed end dates, an employment duration, H, could be observed either because the employment spell was terminated at H, or because week H for the spell occurs in the last week of the sample period. The latter type of spells could be of true length greater than H and are termed right-censored.

Section 1 of this paper), jobs ending in quits ought to be treated separately from jobs ending in layoffs. Thus, we use an independent competing risks approach to calculating relevant failure rates for jobs ending in quits or non-seasonal layoffs.[11]

In Figure 3, we plot the empirical hazard rate for seasonal layoffs and the empirical non-seasonal layoff transition intensities for all of Canada in 1989. The differences between the failure rates for the two groups is quite striking. Failure rates for non-seasonal layoff jobs lie everywhere below those for seasonal layoff jobs and the two diverge substantially after the 10-week point. This is in line with expectations, as seasonal jobs by definition end within a year, while no such restriction is present for non-seasonal jobs. In Figure 4, we plot the empirical quits transition intensities. These have a pattern that more closely resembles the non-seasonal layoffs, though it is lower than both the other two failure rate functions.

Is there any evidence in these Figures of effects from the UI system? While there are spikes in the seasonal layoff hazard at 10 and 14 weeks (points corresponding to the entrance requirement in some regions), there is no basis for concluding that such patterns represent UI effects. Indeed, there is nothing to identify observed spikes as related to the UI system as opposed to some other unspeci-

Figure 3
Predicted Shape of the Hazard Rate Function Under the 1989 UI System

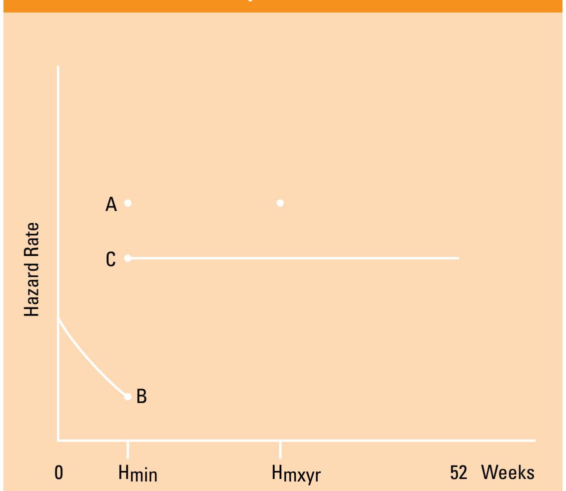
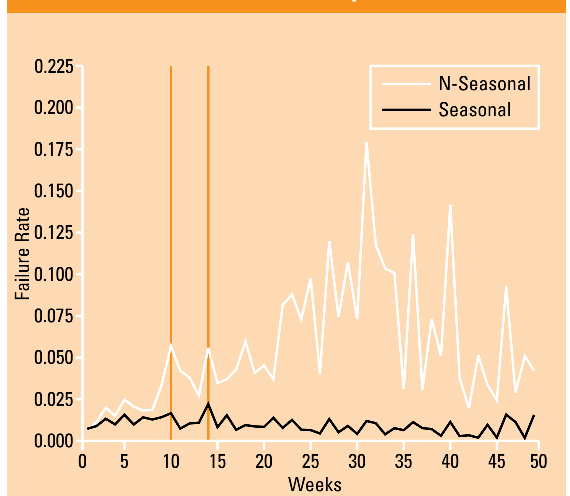


Figure 4
Empirical Failure Rates: Seasonal Vs. Non-Seasonal Layoffs, 1989



11 In a competing risks framework, processes determining non-seasonal layoffs and quits are viewed as running concurrently, with the first one to actually trigger a separation being the one we observe. Thus, if we observe a quit from a job we assume that the layoff process would have created a separation at some later date that we will never observe. In estimating the non-seasonal layoff process, we treat all quits as censored observations on that process. Given that treatment, the probability that a job ends at duration H by reason of a non-seasonal layoff conditional on the job lasting at least H weeks (called the non-seasonal layoff transition intensity) is calculated using the formula for the empirical hazard rate presented above. The sample used for that calculation consists of all non-seasonal layoffs and quits. The quits transition intensity is defined analogously, using the same sample except with jobs ending in a non-seasonal layoff treated as censored.

fied cause. The nature of the failure rate function for seasonal layoffs—high and rising rapidly to the 32nd week— also suggests that uncovering "spikes" related to the UI system will require some care.

Figure 5
Empirical Failure Rates: Quits, 1989

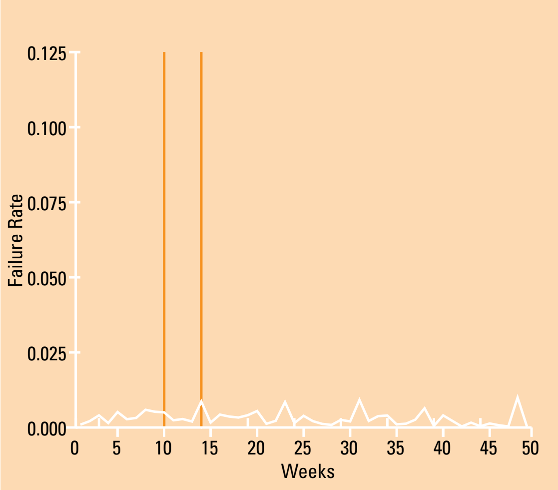
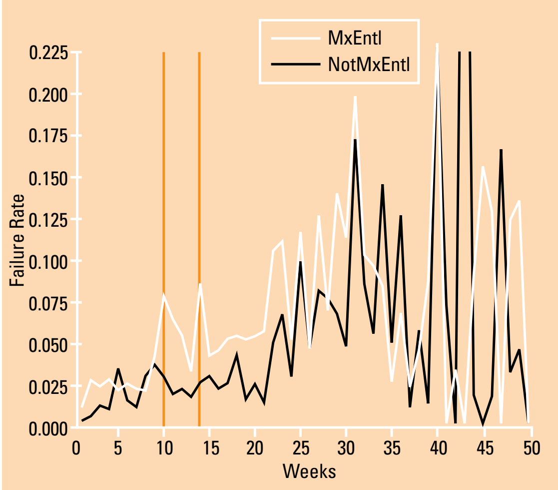


Figure 6
Empirical Failure Rates: Seasonal Layoffs, 1989
Maximum and Not Maximum Entitlement Regions



The maximum entitlement regions form a special group within the UI system. Unlike other UI regions, in these regions with unemployment rates above 11.5 per cent, the UI parameters do not vary with the regional unemployment rate (assuming that the unemployment rate does not drop below 11.5 per cent) and there are no repeat user provisions. Thus, workers in these regions face a stable and very generous UI system. This is especially true in cases where workers require some time to learn about and adapt to UI parameters.

One would expect, therefore, that the greatest reactions to the UI system would be in these maximum entitlement regions. Figure 5 provides plots of the failure rates for seasonal layoff jobs for maximum entitlement and non-maximum entitlement regions.[12] Figure 6 provides the same plots for non-seasonal layoff jobs. In Figure 5, one notes a distinct difference between the failure rates for the two regions: the failure rates for the two regions are somewhat similar up to 10 weeks, but the rates for the maximum entitlement regions are higher for weeks 10 through 24. Whether this is related to the UI system itself cannot be determined from this data, but it suggests that an assumption that the two types of regions can be treated as identical may be incorrect. In the subsequent analysis, we will provide estimates based on data for Canada as a whole and for the non-maximum entitlement regions alone to avoid errors that might arise from treating these types of regions as identical.[13]

12 Maximum entitlement regions are defined as UI regions which experienced monthly unemployment rates of over 11.5 per cent during 1989, and non-maximum entitlement regions are all other UI regions.

13 Note that we do not analyse the maximum entitlement regions on their own because there is no variation in UI parameters within them to use in identifying UI effects.

We turn next to estimating duration models, which allow estimation of failure rate patterns while controlling for observable characteristics of the individual and the job. The most important covariates in our case are the variables designed to reflect the incentives in the UI system. As we have seen, an estimation of simple empirical hazard rates alone is not sufficient to identify these incentive effects. Controlling for other covariates is also useful in ensuring that any measured effects attributed to the UI system are not in fact due to variation in other observable characteristics.

The most common duration model is the proportional hazards[14] model. It is expressed as

$$h_i(H) = h_0(H)e^{x_i(H)'\beta} \tag{1}$$

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[15] and β is a parameter vector. For different values of $x_i(H)'\beta$ the hazard function for individual i is shifted proportionally up or down relative to the baseline hazard.

Given our concern with the location of spikes in the hazard, an estimation method for the proportional hazard model which permits direct examination of the baseline hazard is preferable to ones which either impose a parametric form for the baseline hazard or eliminate it altogether. Indeed, even a flexible but smooth representation of the baseline hazard, such as one using high order polynomials in duration, may be inappropriate in this situation. It is evident from the empirical failure rate plots that the failure rate functions include many spikes. Note in particular, that there are often spikes at 14 and 25 weeks. Some of these spikes may correspond to job spells that start in the first weeks of months and end in the last weeks of months (see Green and Riddell (1993b)). There is a danger that dummy variables designed to detect UI system effects may also detect these other "calendar" effects if a smooth baseline hazard is used. For these reasons, we adopt a specification detailed in Meyer (1990). It allows us to estimate something similar to the empirical hazard function in conjunction with the elements of the β vector. With this specification, we directly estimate any "natural" spikes in the hazard and there is no danger of confusing them with UI effects.

The estimation is carried out using a maximum likelihood approach. The specification of the likelihood function is detailed in Appendix B. The main complication in the estimation is the consistent handling of spells ending in quits. As mentioned earlier, we define job spells as either seasonal or non-seasonal, based on whether or not they ended in seasonal layoff. For quits, however, we do not observe the "seasonal" status of a job. Thus, in the likelihood function we effectively calculate the probability that a quit is a seasonal versus a non-seasonal job and use that calculated value to determine whether the information from a particular job ending in a quit will be used more to calculate the seasonal or the

14 In this portion of the discussion we use the terms hazard rate and failure rate interchangeably.

15 In practice, the continuous variables in $xi(H)$ 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.

non-seasonal hazard rate.[16] Table A.1 in Appendix A presents definitions and mean sample values for the variables used in the duration models.

Central to our exposition are the variables related to incentives in the UI system: *HMIN*, which equals 1 in the week the entrance requirement is met and zero in all other weeks; *HMXYR*, which equals 1 at the H_{mxyr} point for individuals with a one-year time horizon; and *HMX*, which equals 1 at the H_{mx} point. These effects are identified in three ways. First, the values of the entrance requirement and regional extended benefits (and with them the values of H_{min}, H_{mxyr}, and H_{mx}) vary across regions with differences in the regional unemployment rates. Secondly, the values of these UI system parameters vary within UI regions as the regional unemployment rate varies within the year.[17] Thirdly, the entrance requirement for repeat users of the UI system varies across individuals according to the extent of their UI use in the previous year and the unemployment rate in their UI region.

Table A.1 in Appendix A contains a complete characterization of the relation between the UI regional unemployment rate and the values of the variable entrance requirement (*VER*) for non-repeaters and the regional extended benefits. The column on the far left shows the actual partition of the range of unemployment rate values used in calculating the regional extended benefits. In the right half of the table, we present values of relevant kinks in the static labour supply budget constraint. The kinks are presented separately for individuals with a one-year time horizon and an unlimited time horizon. For the one-year horizon case the relevant kinks are those at the H_{min} and H_{mxyr}. (In the table this is the "52 weeks used" point.) These two points coincide for the high unemployment regions. For the unlimited horizon case, the relevant kinks are those at H_{min}, H_{mx}, and at the point of transition from Phase I to Phase II entitlement calculation. We have had great difficulty identifying any effect for this latter point and have omitted it from our analysis. It is clear from Table 1 that substantial variation exists within the system.

At this point, it is worth returning to Figure 3 to define more carefully what we are trying to measure. Consider, for example, measuring the entrance requirement effect. If we include only the *HMIN* variable then for individuals in the maximum entitlement regions, where the entrance requirement is shortest, the coefficient on the *HMIN* variable will be calculated effectively by comparing their hazard rate at 10 weeks to the hazard rate of individuals in all other regions at 10 weeks. Since the individuals in other regions have not yet qualified for UI

16 In a maximum likelihood approach, we search for the set of parameters (the ß's and parameters of the baseline hazard in this case) that maximizes the probability of observing the given sample of observations on the dependent variable (in this case employment duration). The contribution to the likelihood function made by each individual observation is the probability of observing that observation's value of the dependent variable given the values of the model parameters. The total likelihood function is the product of these individual contributions. The contribution of a seasonal layoff observation is the probability that a seasonal job lasts that long and similarly for a non-seasonal layoff. A quit, however, might be used to define the parameters of the seasonal or the non-seasonal jobs. We allow it to play a role in both processes but weight its contribution in each case. Thus, the probability of the observed duration for the quit spell forms part of the seasonal job branch of the likelihood function, but is multiplied by the calculated probability that the spell is a seasonal job. See Appendix B "Construction of the Likelihood Function".

17 There is considerable variation of this form in 1989.

Table 1
Implications of Variance Entrance Requirements (VER) and Maximum Benefit Durations for Budget Constraints in a Static Labour Supply Model

			Expected Spikes (Kinks in Budget Constraints)			
UNEM RATE	MIN WEEKS (VER)	MAX PHASE III BENEFITS	ONE YEAR HORIZON	REASON	UNLIMITED HORIZON	REASON
>11.5	10	32	10	VER	10 18	VER 50 weeks UI
11.1 to 11.5	10	30	10	VER	10 20	VER 50 weeks UI
10.6 to 11.0	10	28	10 11	VER 52 weeks used	10 22	VER 50 weeks UI
10.1 to 10.5	10	26	10 12	VER 52 weeks used	10 24	VER 50 weeks UI
9.6 to 10.0	10	24	10 13	VER 52 weeks	10 25 27	VER Phase I ends 20 weeks UI
9.1 to 9.5	10	22	10 14	VER 52 weeks	10 25 31	VER Phase I ends 50 weeks UI
8.6 to 9.0	11	20	11 15	VER 52 weeks	11 25 35	VER Phase I ends 50 weeks UI
8.1 to 8.5	11	18	11 16	VER 52 weeks	11 25 39	VER Phase I ends 50 weeks UI
7.6 to 8.0	12	16	12 17	VER 52 weeks	12 25 43	VER Phase I ends 50 weeks UI
7.1 to 7.5	12	14	12 18	VER 52 weeks	12 25 47	VER Phase I ends 50 weeks UI
6.6 to 7.0	13	12	13 19	VER 52 weeks	13 25 51	VER Phase I 50 weeks UI
6.1 to 6.5	13	10	13 20	VER 52 weeks	13 25 51	VER Phase I ends Phase II ends
5.6 to 6.0	14	8	14 21	VER 52 weeks	14 25 51	VER Phase I ends Phase II ends
5.1 to 5.5	14	6	14 22	VER 52 weeks	14 25 51	VER Phase I ends Phase II ends
4.6 to 5.0	14	4	14 23	VER 52 weeks	14 25 51	VER Phase I ends Phase II ends
4.1 to 4.5	14	2	14 24	VER 52 weeks	14 25 51	VER Phase I ends Phase II ends
≤4.0	14	0	14 25	VER 52 weeks	14 25 51	VER Phase I ends Phase II ends

at the 10-week point, this measures the total height of the spike at H_{min} relative to what would exist without UI. This is essentially the height *AB*.

On the other hand, for workers in the region with the longest entrance requirement, 14 weeks in 1989, the effect of the *HMIN* variable will be calculated by comparing their hazard rate at 14 weeks to the hazard rate for workers in all other regions at 14 weeks. Since workers in all other regions will have qualified for benefits at 14 weeks, their hazard rate will be at the higher, post-qualification level. Thus, for workers in the region with the longest entrance requirement, use of the *HMIN* variable alone will produce an estimate of the height *AC*.

Since our sample includes individuals from both of these regions as well as individuals from regions with entrance requirements between the two extremes, including *HMIN* alone will produce a measure of the entrance requirement which is less than the total effect *AB*, but larger than the marginal spike effect, *AC*. To avoid this problem, we include in our estimation a variable *GEVER*, which equals 1 for all weeks H_{min} and after. This measures the impact *BC*, the extent to which the hazard is raised for all weeks after qualification. We also include variables *VERMY*, which equals 1 for weeks between H_{min} and H_{mxyr} in the seasonal process, and *VERMX*, which equals 1 for weeks between H_{min} and H_{mx} in the non-seasonal process. These latter variables allow for a more complete characterization of the patterns induced by the UI system.

Two out of the three forms of variation in the UI parameters we use to identify the effects of the UI system are caused by variations in UI region unemployment rates. This increases the probability that we might identify the effects of variations in the unemployment rate rather than effects of the UI system. To address this problem, we include as a regressor the unemployment rate for the "Economic Region" in which an individual resides.[18] "Economic Regions" are geographic constructs used by Statistics Canada which are generally smaller than UI regions. There were 70 "Economic Regions" in 1989 compared to 48 UI regions.

There is some variation, therefore, in the unemployment rate variable which will not trigger variation in UI parameters. This implies that the UI parameters and unemployment rate effects can be separately identified. Also, the nonlinearities in the schedules translating UI region unemployment rates into UI parameter values aid in the separate identification of the two. This takes into consideration the fact that UI parameters do not vary with the unemployment rate in the maximum entitlement regions. In both a simple search model and an implicit contract model, one would expect higher unemployment rates to induce fewer job separations because workers, seeing a longer time unemployed if their job ends, will have lower levels of expected wealth when non-employed. Alternatively, if higher unemployment rates reflect poor states of the world in terms of demand for the firm's product, then higher unemployment could be correlated with more job separations. The actual sign is an empirical matter.

We use several other covariates which are assumed to be related to the average wage offer an individual might expect. They include: whether the individuals are white-collar workers, their level of education, immigrant status and age; as well

18 The unemployment rate is actually a three-month moving average for the economic region, with the last month of the average being the month in which the relevant duration week occurs.

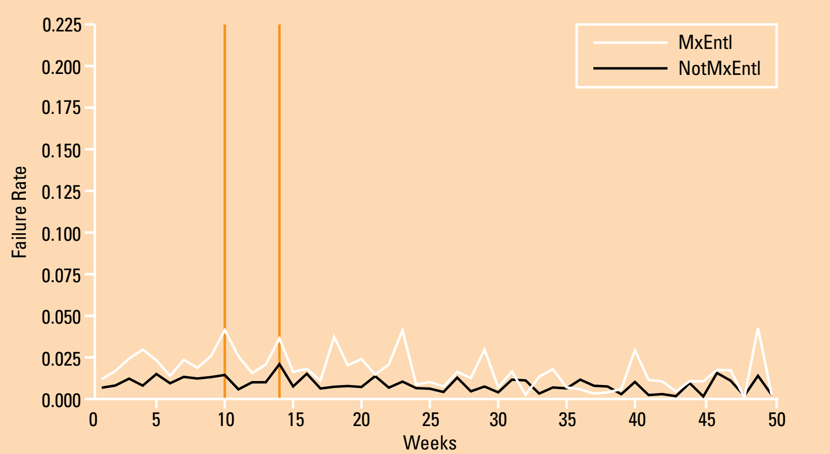
as several others related to individual preferences, such as: sex, marital status, presence of children and whether the individual is the head of the household.

We include industry dummy variables to capture differences in the production technology across industries and plant size variables on the assumption that smaller firms are themselves less stable and/or may not screen job applicants as effectively as large firms. A union status variable is included on the assumption that unions stabilize the work environment, including lengthening employment spells (or perhaps, they are able to organize more stable parts of the work force).

We also use variables that relate to weekly pay. In a search model, workers with higher wages, holding characteristics such as education constant, will be less likely to leave a job. If the wages vary across regions, perhaps due to regional differences in skill levels and mobility of workers, then controlling for the wage is an important part of isolating the effects of the UI system.

Finally, the plots of the empirical failure rates clearly indicate numerous spikes that likely are unrelated to the UI system and may complicate the search for UI effects. 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 them that way. In that case, a spike at the 23rd week, for example, may just correspond to a predominance of five-month

Figure 7
Empirical Failure Rates: Non-Seasonal Layoffs, 1989
Maximum and Not Maximum Entitlement Regions



jobs being reported in this manner. This possibility was investigated in Green and Riddell (1993b) and some support was found for the existence of these “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.

The high number of hours could exist because workers are trying to earn as much as possible during the "season". This is an especially reasonable strategy with Canada's UI system since benefits are based on the weekly wage.

Before examining the estimates from the duration model, it is worth pausing to examine the covariate means. From Table A.1 in Appendix A it is clear that the sample of jobs ending in quits and non-seasonal layoffs[19] are much more similar to each other than either one is to jobs ending in seasonal layoffs. Individuals in the seasonal sample are: more likely to be males; less likely to have a university education; more likely to be in the primary, food processing and construction industries; less likely to be immigrants; more likely to work long hours per week; more likely to be older; and more likely to live in higher unemployment economic regions. (The primary, food processing and construction industries comprise 61 per cent of the seasonal layoff sample jobs.)

These means are consistent with common views of seasonal workers. The high number of hours could exist because workers are trying to earn as much as possible during the "season". This is an especially reasonable strategy with Canada's UI system since benefits are based on the weekly wage, which can be increased by working longer hours. The variable *SAMEMP* equals 1 for jobs for which the employer employed the worker at some time in the previous year. The much higher proportion of seasonally laid-off workers who report having returned to a previous employer again fits with our notion of a seasonal job. Covariate estimates relating to the seasonal job duration process from the duration model are presented in Table 2. The first column of estimates is for the "All Canada" sample, while the second column contains estimates for the subsample of seasonal layoffs that occur in non-maximum-entitlement regions. Many of the estimates in both columns fit with predictions.

In reading these estimates, it should be noted that a positive coefficient on a variable indicates that individuals with the corresponding characteristic have a higher probability of their job ending, that is, they have shorter jobs. For example, females are more likely to have short jobs. Individuals who have completed high school or have some post-secondary education are less likely to leave a job than those who have not completed high school (the base group).[20] Jobs in virtually all other industries are likely to be shorter than jobs in the base industrial category of service and transportation, though primary sector, food processing, and public sector jobs are particularly noteworthy in this regard. Separations show no relation to plant size in either the overall or the non-maximum-entitlement region sample. The coefficient on *SAMEMP* (same employer) is negative and significant in both samples, suggesting that returning to an employer leads to longer spells. This might occur if repeated employment with the same firm means an implicit contract such as the one described earlier exists. In such a contract, firms will commit to providing enough weeks of employment to qualify workers for UI or perhaps to get them to the H_{mxyr} point.

Unions have a negative impact on the failure rate as predicted, as does age. The significant coefficients on the *AGESQ* variable suggest that the negative relationship between the failure rate and age lessens with age. *AGESQ* is defined as age squared, divided by 1000, which accounts for the size of the coefficient. Also, we subtract the mean of continuous variables such as age from the value for each

19 Note that the non-seasonal layoff sample includes jobs that are right censored at the end of 1990.

20 The insignificance of the *UNIV* variable is probably due to the very few individuals with a university education in the samples.

Table 2
Covariate Estimates from Duration Model
Seasonal Jobs, 1989

Variable	All Canada	Not-Max-Entitlement Regions
Female	0.280 (.095)*	0.228 (.108)*
Single	-0.209 (.081)*	-0.246 (.090)*
Nothead	-0.054 (.079)	-0.128 (.089)
Elem	0.236 (.092)*	0.180 (.107)+
Ps	-0.320 (.088)*	-0.306 (.095)*
Univ	0.204 (.201)	0.311 (.211)
Prim	0.563 (.127)*	0.624 (.142)*
Mfg	0.346 (.164)*	0.466 (.181)*
Food	0.701 (.187)*	0.997 (.226)*
Constr	-0.114 (.123)	-0.104 (.138)
Publc	0.627 (.213)*	0.628 (.240)*
Wcolr	-0.018 (.111)	0.026 (.128)
F2099	-0.024 (.086)	-0.024 (.095)
F10049	0.035 (.136)	0.118 (.152)
Fgt500	0.194 (.220)	0.352 (.276)
Samemp	-0.221 (.074)*	-0.335 (.087)*
Uncov	-0.393 (.080)*	-0.369 (.091)*
Presch	0.027 (.106)	-0.025 (.121)
Ochld	0.146 (.088)+	0.054 (.069)
Immig	0.088 (.116)	0.198 (.130)
Age	-1.73 (.478)*	-1.78 (.566)*
Agesq	5.124 (2.628)	5.112 (3.085)+
Wwage	-0.015 (.014)	-0.032 (.016)*
Frstwk	0.461 (.116)*	0.507 (.127)*
Ru	0.648 (1.15)	0.608 (1.541)
GEVER	0.526 (.227)*	0.550 (.254)*
HMIN	-0.100 (.422)	-0.191 (.478)
VERMY	-0.259 (.267)	-0.351 (.312)
HMXYR	0.100 (.169)	0.283 (.182)
# Observations	8968	6623
Mean Log-Likelihood Value	-2.426	-2.832

*Standard errors in parentheses. *, + Significantly different from zero at the 5% and 10% levels, respectively.*

individual observation. This is done so that the baseline hazard corresponds to an average individual in these dimensions.

For both samples—the “All Canada” sample and the subsample of seasonal layoffs that occur in non-maximum-entitlement regions—the economic region unemployment rate is positive, but not significant in both samples. The positive sign could arise either because high unemployment rates signal poor product demand states, or simply because unemployment rates rise in “out-of-season” months of the year in regions where a large proportion of workers are seasonal.

The key variables for our purposes are *HMIN*, *HMXYR*, *GEVER*, and *VERMY*. The largest effect is associated with *GEVER*. (*GEVER* equals 1 for all weeks H_{min} and after.) For both samples, there is evidence that the hazard rate increases significantly once an individual qualifies for UI. Surprisingly, the coefficient on *HMIN* is not close to significant in either sample. Note that the *HMIN* variable effect is defined as being in addition to the *GEVER* effect. Thus, these estimates indicate that the hazard does jump up at the entrance requirement point but that there is no particularly large effect exactly at this point. In the weeks between the entrance requirement and the H_{mxyr} point, the hazard is actually lower in both samples.

Finally, the coefficient on *HMXYR* is positive in both samples and close to significant in the non-maximum-entitlement sample.[21] A possible implication of the lack of a spike at H_{min}, the lower hazard rate for weeks between H_{min} and H_{mxyr}, and a possible spike at H_{mxyr} is that job spells are not constructed to last just long enough to qualify for UI, but to maximize annual consumption in combination with weeks of non-employment. This may suggest a more sophisticated use of the system than simply trying to meet the minimum entrance requirement. It is important to note, however, that when repeat use provisions are taken into account, the proportion of seasonal workers with different values of H_{min} and H_{mxyr} is not large and most of the variation that identifies the H_{min} effect comes from regions with lower unemployment rates. What we may be observing , therefore, is that the entrance effect when $H_{min} \neq H_{mxyr}$ is not significant for seasonal workers while responses are large when H_{min} and H_{mxyr} are equal. This would not suggest sophisticated use of the system as much as heavier use of the system in higher unemployment areas.

In Table 3 we present the coefficient estimates corresponding to non-seasonal layoffs. As discussed earlier, for these workers we view the relevant potential spike points as arising at H_{min} and H_{mx}, and thus include variables corresponding to these points. We also include the GEVER variable and a dummy variable corresponding to the weeks between H_{min} and H_{mx}.

21 A possible explanation for the larger effect in the non-maximum-entitlement sample is based on Figure 6. The hazard rate for maximum entitlement regions is much higher than for the rest of the economy. To the extent that in the full sample the *HMXYR* effect is calculated by comparing the hazard rate at the H_{mxyr} point for non-maximum entitlement regions to the hazard rate at the same number of weeks for the maximum entitlement regions, one could conceivably even end up with a negative coefficient on *HMXYR*. Controlling for the unemployment rate, as we do, will lessen this effect to the extent that higher hazard rates in the maximum entitlement regions are just associated with higher unemployment rates. However, we may not be able to fully control for differences between average hazard rates for maximum and not maximum entitlement regions. Removing maximum entitlement region observations from the estimation altogether may then lessen the problem of not strictly equivalent comparison groups and lead to larger estimated UI effects.

Table 3
Covariate Estimates from Duration Model
Non-Seasonal Jobs, 1989

Variable	All Canada	Not-Max-Entitlement Regions
Female	-0.073 (.034)*	-0.102 (.035)*
Single	-0.045 (.030)	-0.089 (.031)*
Nothead	-0.150 (.026)*	-0.144 (.028)*
Elem	0.023 (.041)	-0.082 (.044)+
Ps	-0.080 (.028)*	-0.073 (.029)*
Univ	-0.453 (.063)*	-0.479 (.065)*
Prim	0.211 (.058)*	0.094 (.065)
Mfg	-0.357 (.041)*	-0.357 (.042)*
Food	0.086 (.077)	-0.081 (.084)
Constr	0.102 (.036)*	0.062 (.038)+
Publc	0.339 (.071)*	0.194 (.078)*
Wcolr	-0.623 (.033)*	-0.576 (.034)*
F2099	-0.274 (.029)*	-0.271 (.030)*
F10049	-0.346 (.039)*	-0.345 (.041)*
Fgt500	-0.503 (.071)*	-0.608 (.074)*
Samemp	-0.242 (.033)*	-0.249 (.035)*
Uncov	-0.420 (.035)*	-0.353 (.037)*
Presch	-0.126 (.036)*	-0.097 (.037)*
Ochld	-0.017 (.031)	-0.017 (.032)
Immig	-0.344 (.036)*	-0.301 (.037)*
Age	-1.81 (.154)*	-1.81 (.162)*
Agesq	4.92 (.864)*	4.85 (.911)*
Wwage	0.016 (.005)*	0.016 (.005)*
Frstwk	0.838 (.039)*	0.920 (.040)*
Ru	0.855 (.449)+	0.886 (.502)+
GEVER	0.220 (.080)*	0.165 (.083)*
HMIN	0.130 (.078)+	0.158 (.079)*
VERMX	0.241 (.045)*	0.342 (.047)*
HMX	0.458 (.164)*	0.445 (.224)*
# Observations	8968	6623
Mean Log-Likelihood Value	-2.426	-2.832

*Standard errors in parentheses. *, + Significantly different from zero at the 5% and 10% levels, respectively.*

The seasonal job results indicate that the job duration hazard rate increases by 69 per cent after the job passes the UI entrance requirement length.

Several of the covariates have notably different coefficient estimates compared to the seasonal layoff sample. For example, the female variable shifts sign, implying that females exit through non-seasonal layoffs later in a job relative to males. The education and industry effects are similar to those in Table 2 apart from the fact that the *UNIV* variable is now significant and manufacturing jobs are now shown to be longer than those in the base category. The plant size variables show negative and significant effects as we originally expected, while the presence of children is now correlated with longer job spells. The immigration variable also changes sign, indicating that when considering non-seasonal layoffs, immigrants are significantly less likely to separate from a job. Interestingly, the weekly wage variable, which was insignificant in the seasonal layoffs sample, now shows a significant positive effect. This is the opposite of what a standard search model predicts. It is possible that we are indirectly capturing the effect of higher UI benefits, since these benefits are a constant proportion of the wage, up to a maximum. We investigate this possibility further below. The regional unemployment rate still has a positive coefficient which is now significant at the 10 per cent level in both samples.

The UI system variables also show quite a different pattern to comparable variables in the seasonal layoff analysis. In particular, *GEVER* is again positive and significant, though somewhat smaller than in the case of seasonal layoffs. *HMIN* is now positive and significant at the 10-per cent level in the overall sample and at the 5-per cent level in the non-maximum-entitlement sample. Similarly, the *VERMX* variable is positive and significant at 5 per cent in both samples, as is the *HMX* variable. Overall, these results suggest a more distinct pattern of spikes than in the seasonal case, with a shift up in the hazard in particular for the weeks between H_{min} and H_{mx}, and statistically significant spikes at both H_{min} and H_{mx}.

In Table 4, we present the covariate estimates for the process generating quits. The sample used in this estimation consists of all non-seasonal layoffs, job spells still ongoing at the end of 1990, and all job spells ending in quits. The estimation is done in a competing risks framework, in which the non-seasonal layoffs that are treated as censored. The coefficient estimates are often similar to those for the non-seasonal layoffs, however the UI system variables exhibit a different pattern. The estimation includes only variables for the various possible spike points. The *HMIN* and *HMXYR* variables are not significantly different from zero. In fact, the *HMXYR* coefficient is smaller than its standard error and the *HMIN* variable has a negative coefficient. The *HMX* variable coefficient, however, is positive and strongly significant. This pattern of little or no entry effect but a significant effect when the maximum possible weeks of entitlement are reached again reflects a more sophisticated use of the system than one might expect.

The seasonal job results indicate that the job duration hazard rate increases by 69 per cent after the job passes the UI entrance requirement length. Also, the results from the non-maximum-entitlement sample indicate a further 33 per cent increase at the H_{mxyr} point. The non-seasonal jobs and job ending in quits show effects of similar or larger magnitudes at predicted spike points.

Are these effects of economic interest? To help in assessing this, Table 5 presents calculations on UI effects for a worker in a region with a 9 per cent unemploy-

Table 4
Covariate Estimates from Duration Model
Quits, 1989

Variable	All Canada
Female	0.010 (.034)
Single	0.042 (.030)
Nothead	-0.025 (.027)
Hs	-0.380 (.038)*
Ps	-0.396 (.040)*
Univ	-0.834 (.077)*
Prim	0.339 (.052)*
Mfg	-0.304 (.044)*
Food	0.403 (.066)*
Constr	0.106 (.040)*
Publc	0.453 (.064)*
Wcolr	-0.510 (.034)*
F2099	-0.262 (.029)*
F10049	-0.257 (.042)*
Fgt500	-0.420 (.071)*
Samemp	-0.081 (.032)*
Uncov	-0.339 (.038)*
Presch	-0.018 (.037)
Ochld	0.058 (.032)+
Immig	-0.282 (.037)*
Age	-0.049 (.016)*
Agesq	0.015 (.008)+
Hwage	-0.257 (.028)*
Hrs	0.306 (.039)*
Frstwk	0.388 (.039)*
Ru	1.076 (.467)*
HMIN	-0.125 (.084)
HMXYR	0.066 (.076)
HMX	0.448 (.140)*
# Observations	7272
Mean Log-Likelihood Value	-1.757

*Standard errors in parentheses. *, + Significantly different from zero at the 5% and 10% levels, respectively.*

ment rate. In this region $H_{min}=11$, $H_{mxyr}=15$, and $H_{mx}=35$. The first two columns of the Table give estimates of the hazard rate for seasonal jobs at 11 weeks, between 11 and 15 weeks, and above 15 weeks. Similar hazard rates are given for non-seasonal jobs. For the seasonal job results, these hazard rates are calculated using formula 1). The baseline hazard is calculated using estimates of the seasonal job baseline hazard from the main duration model using the overall sample. The covariates are set to the average values relevant for seasonal workers, and the *ß* coefficients are the covariate effect estimates for seasonal workers presented in Table 2 for the overall sample. The non-seasonal job results are formed analogously.

Table 5
Fitted Hazard Rates
With and Without UI*

	Seasonal Jobs		
Hazard Rate at:	Without UI	With UI	Difference
H_{min} (11 weeks)	0.012	0.019	0.007
Between H_{min} and H_{mxyr} (12 to 14 weeks) average hazard rate	0.013	0.017	0.004
H_{mxyr} (15 weeks)	0.017	0.032	0.015
Greater than H_{mxyr} (16 weeks or more) average hazard rate	0.049	0.082	0.033

	Non-seasonal Jobs		
Hazard Rate at:	Without UI	With UI	Difference
H_{min} (11 weeks)	0.011	0.016	0.005
Between H_{min} and H_{mx} (12 to 34 weeks) average hazard rate	0.009	0.014	0.005
H_{mx} (35 weeks)	0.006	0.013	0.007
Greater than H_{mx} (36 weeks or more) average hazard rate	0.007	0.009	0.002

** Estimates are constructed using baseline hazard estimates and the covariate estimates reported in Tables 2 and 3. In constructing the estimates, covariates are set to their average values for the relevant group (that is, for seasonal jobs in the upper part of the table and non-seasonal jobs in the lower part). Estimates are constructed assuming an individual in a UI region with 9% unemployment. In this region, H_{min} = 11 weeks, H_{mxyr} = 15 weeks, and H_{mx} = 35 weeks.*

The fitted hazard rates in the first column are calculated with the dummy variables corresponding to UI effects set to zero. Thus, these are our best predictions of the hazard rates that would be relevant if the worker did not qualify for UI benefits. The second column gives the fitted hazard rates when the UI variables take values of 1 in the appropriate weeks (for example, the *HMIN* variable is set equal to 1 in week 11). The third column gives the difference between the two.

Table 5 indicates that, while there are more statistically significant results for the non-seasonal job UI effects, the effects for the seasonal jobs are more economically significant. Qualifying for UI raises the probability that a job ends at the

H_{min} point, or at weeks between H_{min} and H_{mxyr} for seasonal jobs, by much less than a percentage point. On the other hand, the hazard rate increases by about 1.5 per cent at the H_{mxyr} point. If we use the non-maximum-entitlement sample results, the hazard rate at 15 weeks increases from 0.035 to 0.080. This latter result indicates that nearly 1 in 20 seasonal job spells that last to the H_{mxyr} point are terminated at that point for reasons directly related to the UI system. In the weeks after the H_{mxyr} point, the average hazard rate is increased by 3.4 per cent because of UI effects.

For non-seasonal jobs, the statistically significant increases in the hazard corresponding to UI effects, in fact, lead to increases in the hazard rate of substantially less than 1 per cent. In the quits process, the failure rate in a week, H, is 57 per cent higher if that week corresponds to the H_{mx} point; a result that is statistically significant. However, since the average failure rate over relevant durations is less than 0.01 per cent, the effect seems relatively small overall.

The differences in economic significance between seasonal and non-seasonal jobs are portrayed graphically in Figures 8 and 9. They contain plots of the fitted hazard rates used in constructing Table 5 for seasonal and non-seasonal jobs with and without UI. Note that we are not claiming that the "Without UI" lines represent what the hazard would look like in the absence of a UI system, only what they would look like for an individual who for some reason failed to qualify for UI under the current system. Thus, general equilibrium adjustments in industrial structure that would surely follow a removal of the UI system are not reflected in the figures.

Figure 8
Baseline Hazards: Seasonal Layoffs, 1989
With and Without UI Effects

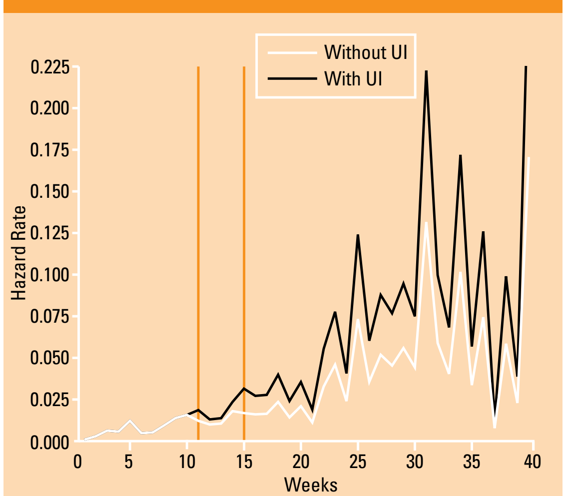
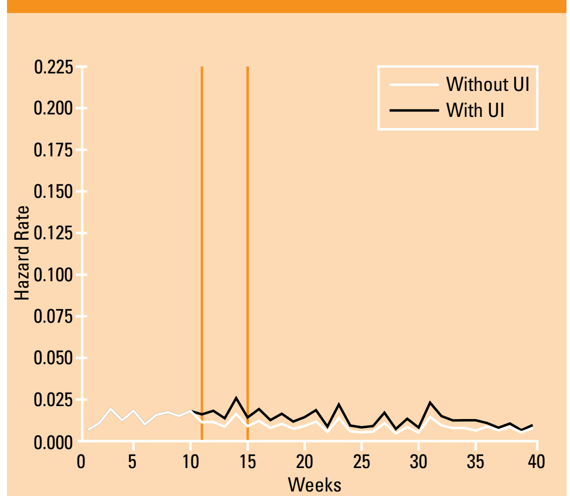


Figure 9
Baseline Hazards: Non-Seasonal Layoffs, 1989
With and Without UI Effects



The reason for differences in economic significance of UI effects for the various types of jobs stems from the proportional nature of the estimated hazard specification. One might observe apparently large percentage increases in the hazard rate attributable to the UI system, but if the hazard rates being affected are relatively small in the first place then the actual net effect may not be large.

To some extent, since seasonal jobs by definition are going to end in less than a year anyway, perhaps one should not be too concerned about these effects. However, if seasonal spells are adjusted to the UI system, then there may exist some scope for reducing the time seasonal workers spend non-employed by changing the H_{mxyr} point.

The seasonal jobs hazard rate is very high over a range of weeks following the entrance requirement. In this case, the estimated percentage effects translate into relatively large absolute effects. On the whole, therefore, the evidence points to statistically significant effects of the UI system that are small in absolute size for non-seasonal jobs and jobs that end in quits.

The lower probabilities of ending a job in the weeks before the H_{mxyr} point, the spike at H_{mxyr} (though small relative to its standard error in the overall sample), and the large increase in the hazard in week — including and after H_{mxyr} — may indicate quite a substantial tailoring of seasonal jobs to the UI system. We do not know the extent to which this tailoring involves the lengthening versus shortening of job spells. To some extent, since seasonal jobs by definition are going to end in less than a year anyway, perhaps one should not be too concerned about these effects. However, if seasonal spells are adjusted to the UI system, then there may exist some scope for reducing the time seasonal workers spend non-employed by changing the H_{mxyr} point. This could be achieved by reducing the size of regional extended benefits. As the implicit contract model suggests, the trade-off could be more individuals who do not get work at all and, ultimately, a reduction in the size of communities associated with seasonal industries. In any case the evidence here, that agents involved in seasonal jobs tailor their behaviour to the UI system, serves mainly to reinforce the notion that whatever role UI plays for these jobs, it is not simply to provide insurance.

How do these results compare to those in earlier studies? Christofides and McKenna (1993) examine job spell durations in Canada using the 1986 and 1987 waves of the LMAS. They find that job durations increase with age and education in general, are shorter for women, are longer for union jobs, and increase with firm size. Our covariate estimates conform with their findings, apart from the fact that females are estimated to have longer non-seasonal jobs once one controls for other covariates. Also, they find that average durations for jobs ending in quits are longer than for those ending in layoffs, which fits with our finding of a lower failure rate for quits. Finally, they note spikes in the probability of job termination in the 10- to 14-week range and raise the issue of whether these could be related to UI effects.

Green and Riddell (1993b) examine the effects of an accidental experiment that occurred in 1990 to identify the effects of the UI entrance requirement. In that year, because of a dispute between the House of Commons and the Senate over other issues, the variable entrance requirement schedule (see page 11, Section 1 of this paper) was not renewed. Instead, the entrance requirement became 14 weeks in all regions of the country for a 10-month period. Green and Riddell, using a similar data set and similar methodology to that used in this paper, examine the employment duration hazard for maximum entitlement regions for evidence of effects of the change in the entrance requirement in an environment in which all other parameters of the UI system are constant.

Green and Riddell find that the four-week increase in the entrance requirement that occurred in these regions for that year had significant impacts on the duration of employment spells. Their estimates indicate that the change in the entrance requirement could cause a 1.5 week increase in the average duration of employment spells and a drop in the unemployment rate by 0.3 per cent in these regions. The results in the paper by Green and Riddell are interpreted largely as being related to effects of the entrance requirement. However, in the maximum entitlement regions the entrance requirement and the H_{mxyr} point coincide. The results in this present paper indicate that what may have been estimated by Green and Riddell was an effect of moving the H_{mxyr} point. Interestingly, Green and Riddell find that the hazard rate at 10 weeks (the entrance requirement in 1989 for these regions) falls by approximately 0.02 per cent in the experiment period. This effect is in the same range as the effects of the H_{mxyr} point we estimated above for seasonal layoffs.

Baker and Rea (1993) also examine the effects of the 1990 experiment on the hazard rate out of employment. They make similar use of inter-regional differentials in the UI system. They, too, find significant effects of the change in the entrance requirement. They then calculate an H_{min} point dummy variable, treating repeat and non-repeat users the same in all regions, assigning all individuals an H_{min} value from the simple table on page 11. They find an increase in the hazard of 250 per cent at the entrance requirement point. These estimates imply much larger effects than do our estimates. In estimates in which UI eligibility is calculated differently for repeat and non-repeat users, however, they obtain estimates closer to ours.

Baker and Rea argue that the statistical evidence in their estimates point to a less accurate measurement in H_{min} when repeaters are treated differently and, therefore, place more confidence in the larger estimates. We estimated specifications in which repeaters are not given special treatment, but we obtain the opposite effects from Baker and Rea: we find that our estimates of UI effects decrease. One possible explanation for the discrepancies between the two studies is that the size of the Baker and Rea estimates are related to the fact that they use a smooth baseline hazard specification. In particular, their H_{min} dummy variable will be identified in large part by comparing the hazard rate at 14 weeks in 1990 with the baseline hazard value at 14 weeks. However, virtually all plots of employment hazard rates for Canada in 1989 show a “natural” spike in the hazard at 14 weeks. That is to say, there is a spike in the hazard at this point whether or not 14 weeks is the relevant entrance requirement for the region. By definition, a smooth baseline hazard cannot capture this “natural” spike entirely. Their H_{min} dummy variable, therefore, will pick up not only the UI entrance requirement effects in 1990, but also the “natural” spike. This could cause an over-estimate of the entrance effect. Treating repeat users of UI differently implies that the H_{min} dummy variable is not always equal to 1 at the 14-week point in 1990. This would reduce, therefore, the effect of the dummy variable capturing the 14-week “natural” spike, and could account for their finding smaller estimates when repeat users of UI are treated appropriately.

A similar issue arises with respect to the estimates in Christofides and McKenna (1994). They also obtain much larger estimates of entrance requirement effects compared to ours, and this in a somewhat similar statistical framework to ours. Contrary to our study, however, they use the Cox proportional hazards model. Their estimate of the entrance requirement effect, therefore, attributes all job separation occurring at the entrance requirement point to the UI system. If there are a large number of jobs terminated at 10 weeks in Newfoundland, where the entrance requirement is 10 weeks for everyone, then they attribute all of those terminations to an entrance requirement.[22] They do so despite the fact that large spikes at 10 weeks also occur in Ontario, where no one faces a 10-week entrance requirement, i.e. in spite of the fact that a large number of terminations appear to occur at 10 weeks even when they cannot be attributable to the UI system. Thus, their estimates form an extreme upper bound on the effects of the UI system on employment durations.

Overall, then, we believe our estimates fit with those in the two nearest papers in this area. The suggestions about the differences with respect to the Baker and Rea paper, however, are speculation: we have not attempted to estimate a specification with a smooth baseline hazard to check this conjecture.

22 The very fact that Christofides and McKenna (1994), at one point in the paper, obtain entrance requirement effect estimates for a sample made up only of jobs in Newfoundland points to the fact that they are not identifying a pure UI effect. In Newfoundland at this time, most of the UI regions were maximum entitlement regions where the entrance requirement was 10 weeks. There were no repeat user provisions to generate individual variation in the entrance requirement. Thus, with the exception of some repeat users in one UI region, there is no variation in the entrance requirement in the province to identify the entrance requirement effect separately from whatever else could cause jobs to be terminated at 10 weeks.

5. Conclusion

For seasonal layoffs... we find some evidence of a concentration of job terminations at the point at which individuals qualify for just enough benefits to fill up the rest of a year.

In this paper, we use a large sample of individual job starts in 1989 to uncover the effects of the UI system on employment durations. Evidence that such effects exist would indicate moral hazard difficulties in the use of the UI system. To the extent that such effects exist, one must question the extent to which the system is being used to insure against unexpected job terminations.

In Section 1 of the paper, we present a description of the UI system as it existed in 1989. The description suggests a potentially large amount of variation in the parameters of the system, triggered mostly by variation in regional unemployment rates. In the remainder of the paper, we make use of that variation to estimate effects of UI on job terminations. In Section 2, we discuss three theoretical models of the impact of UI on job terminations: a static labour supply model, a search model, and an implicit contract model. All three models imply that we should see a disproportionate number of job terminations at the entrance requirement point and at the point at which the individual has qualified for 50 weeks of benefits (the maximum possible). Also, there should be more job terminations in any week after the entrance requirement has been satisfied than in any week before. In addition, the static labour supply and implicit contract models with one-year time horizons imply that there should also be a disproportionate number of job terminations at the point where an individual has qualified for just enough weeks of benefits to fill the rest of the year. We conclude that an analysis should be carried out separately for seasonal and non-seasonal workers since the incentives are different for the two groups.

For reasons detailed in the paper, we cannot identify seasonal and non-seasonal jobs at the outset. We divide our sample among seasonal layoffs, non-seasonal layoffs and quits. Estimation of an econometric, capturing the effects of UI on these three processes, provides interesting results. For seasonal layoffs the probability of a job termination increases significantly in any week after the entrance requirement has been satisfied. Rather than an extra concentration of job terminations right at the entrance requirement point, however, we find that jobs are either no more likely or perhaps even somewhat less likely to end at the entrance requirement point, or in weeks just after, it relative to later weeks. In particular, we find some evidence of a concentration of job terminations at the point at which individuals qualify for just enough benefits to fill up the rest of a year. For non-seasonal layoffs we also find evidence of increased job terminations in weeks following qualification. In addition, we find significant concentrations of job terminations at the entrance requirement point and at the point where an individual has qualified for their maximum possible weeks of UI benefits. We also find evidence of a significant effect at the same point for jobs ending in quits.

All of these results are statistically significant but not all are economically significant. In particular, the concentration in at the maximum qualification point in the quits process amounts to much less than a 1 per cent increase in the number of jobs ending in a given week. Similarly, the estimated effects for non-seasonal jobs correspond to a very small proportion of non-seasonal jobs. The measured effects on seasonal job terminations, however, appear to be economically significant. By one estimate, almost 1 in 20 jobs that last to the point where the individual has qualified for enough benefits to fill up the rest of the year end at exactly that point because of incentives in the UI system.

The apparent delay of job terminations until the point of qualification for enough benefits for the remainder of the year may suggest a somewhat sophisticated use of the UI system in selecting the termination date of some jobs. In particular, seasonal layoffs and quits do not occur just when an individual qualifies for UI, but rather at later points where relevant maximum entitlement points are reached. It should be underlined that that such predictions of reaction to the UI system can be derived in models involving firm decisions as well as in models attributing the reaction to UI to workers alone. Indeed, the fact that these patterns were found in jobs ending in layoffs but not in jobs ending in quits may indicate the direct involvement of firms.[23]

To some extent, since seasonal jobs by definition are going to end in less than a year anyway, perhaps one should not be too concerned about these effects. However, if seasonal spells are adjusted to the UI system, then there may exist some scope for reducing the time seasonal workers spend non-employed by changing the H_{mxyr} point. This could be achieved by reducing the size of regional extended benefits. As the implicit contract model suggests, the trade-off could be more individuals who do not get work at all and, ultimately, a reduction in the size of communities associated with seasonal industries. In any case, the evidence here that agents involved in seasonal jobs tailor their behaviour to the UI system, serves mainly to reinforce the notion that whatever role UI plays for these jobs, it is not simply to provide insurance.

Our main conclusions from the empirical exercise in this paper are as follows. First, there is no evidence of an entrance requirement effect in job durations for jobs ending in seasonal layoffs. But there is evidence of an effect at the point at which workers just qualify for enough weeks of benefits to fill the remainder of a 52 week period, the H_{mxyr} point, and for subsequent weeks. One set of estimates suggests that almost 1 in 20 seasonal jobs end at the H_{mxyr} point because of the incentives of the UI system. Effects of the UI system are also visible in jobs ending in non-seasonal layoffs and quits. However, the size of these effects in absolute terms, measured relative to all job starts in 1989, appears to be small. There is evidence to support the popular notion that some people are adapting their behaviour to the UI system and not using it as insurance. However, estimates suggest that for non-seasonal workers the actual size of these effects may be small.

23 This suggests that the policy may have had little effect if its aim was to deny UI benefits to individuals who quit their last job without cause, and thus stop the so-called “10-40” complex or other patterns of UI use in which the aim of individuals is simply to qualify for UI benefits.

Appendix A: Sample Construction

In this appendix, we describe the construction of the UI system variables and provide means for the covariates. While the analysis in the text concerns job duration, UI eligibility is calculated using employment durations.

To understand our construction of eligibility, consider an individual with two jobs, *A* and *B*, in a year. 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. Assume, also, that the individual was non-employed before $STRT_A$. We define two variables, $ELIG_A(H)$ and $ELIG_B(H)$, as the weeks of UI-eligible work generated by the time the Hth week of Jobs *A* and *B*, respectively, are reached. In our example, by the end of the first week of Job *A* the individual has generated 1 week of UI-eligible employment ($ELIG_A(1) = 1$). Therefore, $ELIG_A(H)$ simply increases 1 for 1 with the weeks of duration of Job *A*. If $STRT_B > STOP_A + 2$ then,

$$ELIG_B(1) = 1$$

and $ELIG_B(H)$ increases with weeks of duration of Job *B*, independently of what occured in Job *A*. If $STRT_B = STOP_A + 2$ then,

$$ELIG_B(1) = STOP_A - STRT_A + 2$$

(the duration of Job *A* plus 1 week). Thus, if Job *B* starts within two weeks of Job *A* ending, then weeks of UI eligibility continue to be incremented.

If there is a non-employment gap of over two weeks between the two jobs, then it is deemed possible that the individual initiated a UI claim in the interim. In that case, UI eligibility must be recalculated. A two-week gap is allowed to account for the waiting period at the start of a UI claim. We assume that if a non-employment gap shorter than the waiting period is observed, then UI eligibility does not need to recalculated. If $STRT_B \leq STOP_A$ then,

$$ELIG_B(1) = STRT_B - STRT_A + 2$$

(the duration of Job *A* to the point when Job *B* starts plus 1 week).

Once we have defined the *ELIG(H)* function, the UI-related parameters are calculated as follows. The variable *HMIN(H)* equals 1 in the Hth week of the spell if *ELIG(H)* = the entrance requirement number of weeks for the individual. The value of this variable is recalculated on a week-by-week basis for two reasons. First, changes in the UI regional unemployment rate can trigger changes in the entrance requirement. Such changes in the entrance requirement can occur in all provinces except Newfoundland and Prince Edward Island in 1989. In some regions, the changes are quite large; for example, in Belleville/Peterborough region the non-repeater entrance requirement varied between 11 and 14 weeks in 1989.

Secondly, the entrance requirement for repeat users of UI is based on the number of weeks of UI use in the qualifying period, which is defined as the 52 weeks before a claim is filed and the number of weeks since the last claim, whichever is shorter. As a job spell progresses, this qualifying period can change with it, causing some weeks of UI use from the previous year to cease to be relevant in calculating the entrance requirement. As noted in the text, we define UI usage spells as non-employment spells that follow UI-eligible jobs.

The schedule for determining the entrance requirement for repeat UI users is as follows:

Weeks of UI in Qualifying Period	Entrance Requirement
$\leq VER$	VER
$VER + i$, $1 \leq i \leq 5$	$VER + i$
$\geq VER + 6$	$VER + 6$

where *VER* is the Variable Entrance Requirement for non-repeat UI users for the region.

The variable *HMXYR(H)* equals 1 in the Hth week of the spell if $ELIG(H) + UIQUAL(H) + 2 = 52$, where *UIQUAL(H)* is the weeks of UI receipt an individual is qualified to receive once they have worked H weeks on the current job. If the week for which $HMXYR(H) = 1$ is less than the week for which $HMIN(H) = 1$, then we set $HMXYR(H) = HMIN(H)$. This variable, too, must be recalculated week-by-week both because regional extended benefits may change with the regional unemployment rate (thus affecting *UIQUAL(H)*), and because it cannot take a value of 1 before *HMIN(H)* takes a value of 1 — and the latter variable is constantly changing.

The variable *HMX(H)* equals 1 in the Hth week of the spell if *ELIG(H)* equals H_{mx}, the point at which the individual has qualified for the maximum weeks of benefit receipt. Since UI entitlement depends on regional extended benefits, this variable too must be constantly recalculated.

Table A.1
Variable Descriptions and Means

Description		Mean Value		
		Seas. Layoff	Non-Seas Layoff	Quits
Female	=1 if female, zero else	0.320	0.453	0.530
Single	=1 if not married, zero else	0.376	0.376	0.434
Nothead	=1 if not head of household, zero else	0.540	0.480	0.591
Elem	=1 if not completed high school, zero else	0.259	0.109	0.104
Ps	=1 if some or completed post-secondary, zero else	0.201	0.318	0.292
Univ	=1 if completed university, zero else	0.023	0.086	0.072
Prim	=1 if industry is agriculture, forestry or fishing, zero else	0.259	0.053	0.034
Mfg	=1 if manufacturing, other than food proc., plus mining, zero else	0.074	0.126	0.136
Food	=1 if food and beverage industry, zero else	0.102	0.031	0.025
Constr	=1 if construction, zero else	0.250	0.151	0.085
Publc	=1 if government employee, zero else	0.051	0.046	0.040
Wcolr	=1 if occupation is manager, professional, sales or service, zero else	0.565	0.565	0.635
F2099	=1 if firm employs 20 to 99 workers at this location, zero else	0.305	0.298	0.325
F100499	=1 if firm employs 100 to 499 workers at this location, zero else	0.126	0.164	0.119
Fgt500	=1 if firm employs over 500 workers at this location	0.036	0.059	0.051
Samemp	=1 if worked for this firm in previous year, zero else	0.596	0.229	0.200
Union	=1 if job covered by a collective agreement, zero else	0.296	0.288	0.194
Presch	=1 if respondent has preschool children, zero else	0.169	0.190	0.205
Ochld	=1 if respondent has non-preschool children, zero else	0.301	0.252	0.246
Immig	=1 if an immigrant, zero else	0.100	0.154	0.172
Wwage	weekly wage (= usual hourly wage * usual weekly hours)	470.400	419.880	335.620
Age	age of individual at interview date	36.160	33.890	32.160
Ru	unemployment rate for economic region, 3 mos. moving avg.	0.105	0.082	0.082

Appendix B: Construction of the Likelihood Function

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 will last at least $H*+1$ weeks conditional on it having lasted $H*$ weeks, is given by,

$$Pr[H_i \geq H*+1/H_i \geq H*] = exp[-exp(x_i(H*)'\beta + \gamma(H*))] \quad (B1)$$

where,

$$\gamma(H*) = 1n[\int_{H}^{H*+1} h_0(u)du] \quad (B2)$$

Note that the hazard function in this model is given by 1 minus equation (3). The contribution to the log likelihood function for the ith job spell is then,

$$l_i(k_i/x_i, \gamma, \beta) = \delta_i log[1 - exp(-exp[\gamma(k_i) + x_i(k_i)'\beta])] \quad (B3)$$

$$-\sum_{H=1}^{k_i-1} exp[\gamma(H) + x_i(H)'\beta]$$

where, k_i is the observed length of the ith employment spell [24] 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, the $\gamma(H)$'s are treated 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.

If we could observe seasonality status of jobs perfectly, then a duration model such as this could be applied to samples of seasonal and non-seasonal jobs separately. The fact that in our case observation of seasonality is censored by the quits produces complications. We address these complications using the following duration model. In this model individuals are assumed to first make a decision on whether to obtain a seasonal or non-seasonal job. This decision does not depend on anticipated job lengths in the two types of jobs. Conditional on the seasonality decision, individuals then are involved in a process that determines the length of the job spell. Define a dummy variable, D_{si}. which equals 1 if the job is a seasonal job and zero otherwise. The value of D_{si} is determined by,

$$I_i = z_i\phi + \eta_i > 0 \Rightarrow D_{si} = 1 \quad (B4)$$

$$\leq 0 \Rightarrow D_{si} = 0$$

where z_i is a vector of personal and job characteristics, ϕ is a parameter vector and η_i is a mean zero disturbance term distributed independently of k_i.

For spells ending in a seasonal layoff, the contribution to the likelihood function is the joint probability that a seasonal job is chosen and that it is of observed

24 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.

length, k_i. Given the two stage decision process mentioned above, the contribution to the likelihood function is then,

$$l_i^s = l_i(k_i / x_i, \beta_s, \gamma_s, D_{si} = 1)\ \Phi\ (z_i \phi) \tag{B5}$$

where the l_i function is the same as defined in (B3), β_s is a covariate parameter vector corresponding to the seasonal layoff process, γ_s is a vector of parameters for the seasonal layoff baseline hazard, and Φ is the standard normal distribution function. Given the sequential choice assumptions, the hazard rate function is considered conditional on being observed on a seasonal job. Similarly, the contribution of a non-seasonal layoff is the probability of the individual choosing a non-seasonal job and the probability of observing a non-seasonal job of the given length, k_i,

$$l_i^{ns} = l_i(k_i / x_i, \beta_{ns}, \gamma_{ns}, D_{si} = 0)\ \Phi\ (-z_i \phi) \tag{B6}$$

where *ns* denotes non-seasonal.

Layoffs and quits are viewed as being generated in an independent competing risks framework and thus quits are treated as censored observations on the layoff processes. The difficulty is that we do not know whether a quit is from a seasonal or non-seasonal job. Thus, the contribution of a job ending in a quit in determining the parameters of the layoff processes is,

$$\begin{aligned} l_i^q = {} & l_i(k_i / x_i, \beta_{ns}, \gamma_{ns}, D_{si} = 0)\ \Phi\ (-z_i \phi) \\ & + l_i(k_i / x_i, \beta_s, \gamma_s, D_{si} = 1)\ \Phi\ (z_i \phi) \end{aligned} \tag{B7}$$

where the spell length, k_i is always treated as censored. Equation (B7) says that the contribution of a quit is the weighted average of the probability of observing the given spell length if the job were a seasonal job and the probability of observing the spell length if it were a non-seasonal job. The weights are just the calculated probabilities that the job is non-seasonal and seasonal respectively. Under the independent competing risks approach covariate and baseline parameter vectors for the quits process, *ßq* and γq, can be estimated by an application of (B3) with all layoffs, regardless of their seasonality treated as censored spells.[25]

25 This assumes that the quits process is the same in seasonal and non-seasonal jobs.

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