The Effects of the Shift to an Hours-Based Entrance Requirement

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Abstract

The new Employment Insurance (EI) system introduced by the Canadian government includes several innovations on standard unemployment insurance type systems. One of the most significant is a switch from a weeks base for calculating benefit eligibility (i.e., whether an individual can receive benefits at all) and entitlement (i.e., the amount of time an individual can potentially collect benefits once eligible) to an hours-based system. This paper examines the impact of the switch from a weeks-based to an hours-based system. We are concerned with two main questions. First, how did the switch affect eligibility and entitlement? Second, did the change affect outcomes in terms of the duration of jobs measured both in weeks and in total hours, and in terms of hours worked per week?

Our overall conclusion in terms of eligibility and entitlement is that the switch to an hours-based system had almost no impact on eligibility but did increase the number of weeks of benefits workers were entitled to, in general. A shift from a weeks-based to an hours-based system redistributes weeks of entitlement toward long hour, predominantly male, seasonal workers and away from part-year/part-time, predominantly female workers. The average wages for the groups most affected by the shift to EI are above the average wage for seasonal workers as a whole.

Our overall conclusion with respect to impacts on employment spell durations is that the shift to the hours-based system has had noticeable impacts in terms of shortening seasonal employment spells. This is plausible since seasonal workers typically work long hours per week and, as a result, can acquire given levels of EI entitlement in fewer weeks under the new system. However, this effect should be kept in perspective. Non-seasonal employment spells, which make up the great majority of all employment, show no patterns of adaptation to the UI/EI systems or changes with the introduction of the EI system.

Finally, there is virtually no change in the usual hours per week distribution between 1996 and 1997. In particular, the distribution of usual hours per week for seasonal workers in high unemployment regions, the group that appears to react most to UI incentives, does not change. Thus, while there does appear to be some reaction to the new system in terms of duration of jobs measured in weeks, there is no apparent reaction in terms of adjusting hours per week.

Introduction

The new Employment Insurance (EI) system introduced by the Canadian government includes several innovations on standard unemployment insurance type systems. One of the most significant is a switch from a weeks base for calculating benefit eligibility (i.e., whether an individual can receive benefits at all) and entitlement (i.e., the amount of time an individual can potentially collect benefits once eligible) to an hours-based system. The rationale for this change was to provide greater access to unemployment insurance (UI) for individuals working in non-standard work arrangements and to provide a clearer relationship between premiums paid to and benefits expected from the UI system (Nakamura and Diewert (1997)). Little is known about whether an hours-based system will truly be successful in meeting this objective, in large part because no other country uses an hours-based system. This paper examines the impact of the switch from a weeks-based to an hours-based system. We will be concerned with two main questions. First, how did the switch affect eligibility and entitlement? Did workers on "non-standard" jobs move from not being covered to being covered by UI? More broadly, what groups experienced changes in terms of both eligibility and weeks of benefit entitlement from this switch? Second, did the change affect outcomes in terms of the duration of jobs measured both in weeks and in total hours and in terms of hours worked per week? The switch in the system may change incentives enough to alter measurable impacts of the UI system on job durations. If so, this should be taken into account in assessing the success of the change to an hours-based system.

The paper proceeds in 5 sections. In Section 1, we discuss incentives affecting job durations under both weeks- and hours-based systems. We also discuss, briefly, earlier results in studies that examined the impact of the weeks-based entrance requirement on job durations. In Section 2, we describe the data we will use to answer the two main questions above. In Section 3, we present calculations related to eligibility and entitlement changes. We will use this in part to examine whether changes in eligibility and entitlement stemming from the switch to an hours-based system are very large, in part to investigate who experienced those changes, and in part because we need to determine whether there appears to be any substantial impact of the rule change on calculations of eligibility in order to determine whether to expect much in terms of behavioural responses. In order to investigate the nature and size of any behavioural responses, we will first examine the distribution of job durations measured in weeks and (quite briefly) the distribution of average hours worked per week to see if the switch to the new system altered incentives enough to create noticeable changes in these distributions. As a first step, we will study simple empirical hazard plots and cumulative distribution functions for evidence of such changes. However, to provide a convincing analysis we need to control for potential differences in terms of the distribution of individual worker characteristics and the state of the labour market. Thus, we also estimate more complicated duration models to permit an investigation of changes in the relevant distributions while holding these covariate effects constant. Section 4 contains the description of the methodology we use along with the results of this examination. Section 5 contains conclusions.

1. Incentives Under Hours- and Weeks-based UI Systems

1.1 Changes in the UI/EI System

We will begin with a brief description of the Unemployment Insurance (UI) system and the changes it has gone through since 1990. Given our interest in entrance requirements, we will focus on changes in rules related to eligibility and entitlement. Between November 1990 (the effective date for Bill C-21) and April 1994 (the enactment date for Bill C-17), both the entrance requirement and entitlement calculations were based on a combination of the number of UI eligible weeks worked by the applicant in the qualifying period and the regional unemployment rate. The qualifying period was the lesser of the 52 weeks directly preceding the application or the time since the end of the individual's last UI claim. The first column of Table 1 shows how the number of weeks an individual needed to work in the qualifying period in order to be eligible for benefits varied with the regional unemployment rate.[1] There is clearly a substantial range in the eligibility requirement, with individuals in very high unemployment regions needing only 10 weeks of work to qualify while individuals in regions with unemployment rates under 6 percent needed 20 weeks. In addition, there were special longer entrance requirements for individuals who were deemed new or re-entrants to the labour market by virtue of their having worked only a small number of weeks in the 52 week period preceding the qualifying period.

The entitlement calculation under Bill C-21 was broken down into three phases. Under phase 1, individuals gained entitlement to one week of benefits for each week worked, up to a maximum of 25 weeks. Under phase 2, individuals gained an additional week of benefits for every two weeks worked over and above 25 weeks, up to a maximum of 13 additional weeks. Under phase 3, the regional extended benefits phase, individuals were awarded a certain number of weeks of benefit regardless of how many weeks they worked, assuming they had worked enough weeks to satisfy the entrance requirement. The regional extended benefits were based on the unemployment rate in the individual's region, with two weeks of benefits added for each 0.5 percent increment in the unemployment rate in excess of 4 percent. The maximum weeks that could be added under phase 3 was 32. The rationale was that it would take longer to find a job in those regions and hence a longer period of benefits was required. The total entitlement for an individual could not exceed 50 weeks and was the sum of the entitlement calculated under each of the three phases.

Under Bill C-17, both the eligibility and the entitlement schedules were changed in ways that made it significantly harder to collect benefits. As shown in column 2 of Table 1, the entrance requirement was increased by up to two weeks in regions with unemployment

1 The regions used in these calculations are UI regions, which (with the exception of PEI) were sub-provincial constructs. Under Bill C-21, there were 48 such regions.

rates over 14 percent. More importantly, the functions generating entitlement were changed dramatically. First, phases 1 and 2 were reversed. After April 1994, an individual earned 1 week of benefits for every 2 weeks worked for the first 40 weeks of work in the qualifying period and then could earn 1 week of benefits for each added week of work, up to a maximum of an added 12 weeks of benefits. Regional extended benefits were maintained but now 2 weeks of benefits were added for each 1 percent increment in the unemployment rate over 4 percent, with a maximum possible 26 weeks being collectable under this component. The sum of the three components could still total at most 50 weeks of receipt. These changes drastically reduced the entitlement of many part-year workers as we will show in Section 3.

A further, large set of changes to the system were enacted under the *Employment Insurance Act*. The main change, for our purposes, was a shift from a weeks-based system to an hours-based system. At the outset, all this has meant is taking the same entrance and entitlement numbers that were relevant under Bill C-27 and multiplying them by 35, though the system opens the door for more changes in the future. Thus, column 3 of Table 1 (the EI entrance requirement) is just 35 times column 2 of Table 1 (the Bill C-27 entrance requirement). This is very important to keep in mind when we come to the discussion of changes in entitlement and eligibility. The only other change in this part of the system is that the maximum weeks of entitlement is now 45 instead of 50.

One interesting alteration in the system related to the entrance requirement is found in the calculation of the divisor used in deriving weekly benefit levels. To calculate weekly earnings, total earnings in the preceding 26 weeks are divided by the greater of: weeks actually worked and a specific divisor differing according to the unemployment rate in the region. The latter divisor is generally the entrance requirement number of hours plus 70. Consider an individual working 12 weeks at 35 hours per week in the highest unemployment rate region. This individual will just meet the entrance requirement of 420 hours but his earnings will be divided by 490 rather than 420 in calculating his weekly earnings measure used in deriving his weekly benefits. This has the effect of dividing his earnings by 14 weeks instead of by 12 weeks, establishing a lower basis for calculating weekly benefits. If, instead, the individual works two more weeks, the divisor is still 490 hours and, since his total earnings will be higher by two weeks worth of pay, the weekly earnings base calculated by the EI system will be higher. This wrinkle in the system directly discourages stopping jobs at the entrance requirement or the entrance requirement plus 1 week.

1.2 Incentives

The standard approach to discussing incentives associated with the UI entrance requirement (ER) is to make use of a static labour supply model (see, e.g., Moffitt and Nicholson (1982), Phipps (1990), Green and Riddell (1997)). In this model, individuals choose a number of weeks to work and a level of consumption in a defined period (typically a year) subject to a budget constraint. The shape of the budget constraint is affected by the attributes of the UI system and one can infer the incentives generated by the system by examining the shape of the constraint. Specifically, the budget constraint jumps up at the ER week as the individual's total non-asset income moves from equalling his labour income alone (at week ER-1) to his labour income plus his benefit receipt under UI. In the weeks directly following the ER, the budget constraint becomes steeper as each added week of work not only increases total income by the weekly wage but also by increasing the weeks of UI entitlement, and hence the total benefits received. At some point, however, the sum of weeks worked and weeks of UI entitlement will total 52. We will call this the MAXYR point since it corresponds to the point at which the individual can collect the maximum possible UI benefits in the one year period. For each week worked past the MAXYR point, the total UI receipt must decline since there is one fewer week in which to collect. The incentives, discussed extensively elsewhere (see the references above, for example), suggest a likely bunching of weeks of work at both the ER and MAXYR points.

One troubling component of this model is that its predictions depend on the time horizon used by the individual. Thus, if the individual's time horizon were two years rather than one, the MAXYR point would shift higher because it would take longer for the individual to reach the point at which he or she was running out of time to collect their benefits. This is particularly troubling because we cannot observe individual planning horizons. However, Green and Sargent (1998) argue that a one-year planning horizon is a reasonable assumption when considering seasonal workers. They look for and find evidence of a bunching of job endings at the MAXYR point among seasonal jobs in 1989. They argue, further, that a search theoretic framework may be more useful for considering outcomes for non-seasonal workers. They derive theoretical predictions from a search model that include a decline in the hazard rate as the weeks on a job approach the ER from below, a bunching of job endings at the ER point, and a higher level of job terminations after the ER has been passed than before. They also argue that comparing the number of job terminations at the ER point with the number in the weeks before the ER is not a good measure of the impact of the ER on the job duration distribution. This is so because, as seen in the search model, job terminations will be artificially low in the weeks just before the ER as terminations are "saved" for the point at which they will generate eligibility for UI. Thus, such a comparison would yield an overestimate of the actual impact of the ER. However, the extent to which job endings have been "saved" for the ER can be measured by comparing the number of job terminations at ER with the number in the following weeks.

Finally, an implicit contracting framework has also been advanced as a fruitful way to think about the impact of UI systems on employment outcomes. Such models help in making the point that, in the absence of complete experience rating, UI systems can serve to subsidize unstable work pattern industries as well as provide incentives to individual workers (see, e.g., Feldstein (1978), Anderson and Meyer (1992), Card and Levine (1994)). This framework has been suggested as particularly relevant for the Canadian situation since there is some evidence that estimated UI impacts on the job duration distribution stem at least in part from co-operation between employers and employees (e.g., Green and Riddell (1997)). In cases where the firm is assumed to view workers and weeks worked per worker as perfect substitutes in production, the firm is indifferent between adjusting to demand shocks using weeks or changes in employment. However, workers are not indifferent and, in particular, want to acquire enough weeks of work to qualify for UI. Firms will then have an incentive to follow worker preferences in terms of work schedules because workers will demand a lower wage when their preferences are met. Thus, the predictions will be the same as those under the static labour supply model that reflects worker incentives alone.

When considering the switch to an hours-based system, one possibility is that the same analytical frameworks can be used but with hours substituted for the weeks dimension. This would certainly be true if work weeks were of fixed length. However, with flexibility in setting hours per week, such a simple translation may not work well. In particular, Storer and Van Audenrode (1998) find evidence that workers and firms may have responded to the tightening of the UI system in 1994 by cutting hours worked per week in order to share out total hours in a way that generated UI eligibility for more people. If this is true, then a switch to an hours-based system will eliminate this mode of adjustment. One might then find that the number who qualify for UI will fall. As shown in Section 3, this conjecture is not borne out in the data.

A conclusion that the 1994 changes would likely lead to declines in hours worked per week and that the switch to an hours-based system would then lead to declines in eligibility seems plausible if firms and workers are focused on the ER. However, as discussed, the main impact of the 1994 changes was on entitlement, not eligibility. In a standard implicit contract model, UI benefits subsidize seasonal firms because these firms do not need to pay as large a compensating differential for their short job duration when UI benefits are available. What the 1994 changes did was reduce the amount of benefits a seasonal worker could expect to collect based on a given amount of work. This could imply a need to increase wages to keep workers in the industry. Firms might respond by reducing hours per worker in order to generate more weeks of work and hence increased UI entitlement for their workers. However, since benefits are based on weekly earnings, reducing hours would also have the effect of reducing UI based income for workers and lead to upward pressure on wages. Firms might, alternatively, keep hours of work relatively steady, reduce the number employed, and increase weeks worked among those employed to generate extra benefits for them. Which route firms adopt depends on a whole host of factors including the technology being used, how easy it is to train new workers, worker preferences concerning consumption of leisure through more hours off in a week versus more weeks off in a year, outside employment opportunities for workers, and the rate of conversion of weeks worked into UI entitlement. The result, including whether the switch to an hours-based system will lead to disentitlement, is an empirical matter.

1.3 Previous Studies

Several previous studies have examined the impact of the UI system on the distribution of job durations. Baker and Rea (1997) and Green and Riddell (1997) both examine the effects of an increase in the ER that occurred in 1990 when all regions in the country reverted to a 14 week entrance requirement because of an unrelated legislative battle between the House of Commons and Senate. Green and Riddell (1997) focus on the highest unemployment rate regions where the change caused an increase in the entrance requirement from 10 to 14 weeks. They find that the probability that a job ended in weeks 10 through 13 decreased but the probability that a job ended at the new ER (14 weeks) increased only modestly. Instead, under the new rules, more job terminations were observed in the 15 to 20 week range. This suggests that the ER does have an impact on employment outcomes but that the impact does not fit easily with a simple supply-side model in which workers unilaterally terminate jobs when they reach the ER. Baker and Rea (1997) find larger impacts directly at the ER point, perhaps because of differences in sample selection relative to Green and Riddell (1997).[2]

In a further study of the impact of the pre-1990 UI system on employment durations, Green and Sargent (1998) examine 1989 data on employment durations for evidence of UI impacts. As discussed above, they argue that breaking spells into seasonal and non-seasonal spells is important in such an examination and look for effects both at the ER and MAXYR points for seasonal workers. They find statistically significant but economically small effects in the form of an excessive bunching of job terminations at the ER for non-seasonal workers. For seasonal workers, they find that there is again an ER effect but that the MAXYR effect is larger. The effects for seasonal workers are larger than those for non-seasonal workers but are still not huge. They conclude that there is evidence of UI impacts but, apart from seasonal workers in high unemployment regions, these effects are not large. They use their estimates to generate predictions of effects from the 1994 UI changes. These predictions suggest that the decline in entitlement would cause increases in average employment durations primarily by increasing the MAXYR point for seasonal workers.

Two studies directly examine the impacts of the 1994 changes on related employment outcomes. Kuhn and Sweetman (1998) use administrative data to calculate predicted entitlement and eligibility for samples of workers and job terminations before and after the 1994 changes. They find that, using both pre- and post-1994 rules on a pre-1994 sample, the 1994 changes reduced weeks of benefit entitlement by an average of 10 weeks but that the impacts were far from uniform: the biggest reductions were in more seasonal industries and provinces. However, once behavioural adaptation to the new rules is included, by examining a post-1994 sample of workers, the losses are much more uniform: workers and firms in the high unemployment regions found ways to adjust.[3] The authors also found that impacts in terms of workers failing to gain eligibility at all were small (less than a 5 percent reduction).

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2 Christofides and McKenna (1996) also examine the impact of the ER on employment durations and find quite large effects. Unfortunately, Christofides and McKenna define their ER variable as equal to 1 for those jobs ending at the relevant ER rather than defining it as equal to 1 in the week a job passes the ER point, regardless of whether the job ends in that week. The latter is the ER variable used in all other studies of this effect. The fact that the Christofides and McKenna variable is inappropriate can be seen in the fact that they use it in their regression of log duration on various covariates. In that context, it is straightforward to see that a negative estimated coefficient on their ER variable indicates that jobs that end at the ER are shorter than the conditional average for spells that do not end at the ER. The estimated coefficient indicates nothing about how many spells ended at the ER and thus it is not possible to learn anything about the potential impact of the ER on the distribution of durations.

3 As an example, they show that in the absence of behavioural adaptation the loss in entitlement due to the1994 changes would have been 16 weeks on average in Newfoundland. In fact, the eventual loss was 7.6 weeks, which was approximately the national average after behavioural adaptation was allowed.

Finally, while the authors find evidence of excessive bunching of job terminations at the ER point, they argue that these effects are small in relation to the system as a whole: only 3.4 percent of all UI claims in the country were established with the ER number of weeks in high unemployment regions before 1994 (10 weeks), and, of course, not all those are necessarily terminations generated to adapt to the UI system. Storer and Van Audenrode (1998) use survey data linked to administrative records to investigate whether reductions in entitlement stemming from the 1994 changes caused workers to find lower paying jobs afterward. An interesting finding, mentioned above, is that hours per week on jobs appear to decline after 1994.

2. Data and Basic Results

2.1 Data

The data we use to investigate the impact of the switch to an hours-based system on employment outcomes and entitlement is the Canadian Out of Employment Panel (COEP). The COEP consists of administrative data linked to a survey. In particular, a sample of Records of Employment (ROEs) marking the end of jobs is drawn from all jobs ending in a given quarter. Workers associated with the sampled ROEs are then contacted in order to be surveyed. The survey provides extensive data on individual characteristics such as education and age, on employer characteristics such as firm size, and on household characteristics such as expenditures. This makes COEP a unique dataset that is very well suited to studying the issue at hand. Indeed, this study would be far less complete without the COEP. Before the advent of the COEP, researchers had to make do either with administrative data alone, which lack crucial information on individual characteristics such as education and family resources, or with survey data alone, the use of which necessitates considerable guessing about the exact job spells used to establish claims.[4]

At our time of study, data were available up to and including the sample of ROEs drawn in the third quarter of 1997. In order to maintain comparability, we used data corresponding to data ending in the first three quarters of 1996 and the first three quarters of 1997. Because self-employment spells and some jobs ending in quits are not eligible for UI, we restricted our attention to paid employment spells ending in layoff. In addition, following Green and Sargent (1998) we split our sample into seasonal and non-seasonal jobs. The seasonality split is created based on a question asking workers the nature of their work arrangement. We defined as seasonal those jobs listed by the worker in the survey as being of a seasonal or contract nature. All other layoffs we defined as non-seasonal.

Our examination focuses on employment duration rather than job duration since several jobs may be strung together to create Unemployment Insurance (UI) eligibility. We define an employment spell as consisting of weeks of consecutive employment. In particular, we begin with the job corresponding to the originally drawn ROE. We then consider the immediately preceding job, adding it to the ROE sample job if the end date of the preceding job is within two weeks of the start of the ROE sample job. If the preceding job is added to the ROE job for the purpose of generating an employment spell, then we move further back and consider the next previous job using the same criterion.[5] While previous jobs may not be close enough in time to count as part of the same employment spell by our criteria, they may still be used by the worker to generate UI entitlement. They are

4 We are greatly indebted both to the individuals who spent innumerable hours of work creating the COEP and to those who currently administer it (in particular, Lars Vilhuber). They have made a truly complex dataset as easy to work with as one could possibly hope.

5 If there is a gap between the preceding job and the ROE sample job of, for example, two weeks, we do not count that two week gap as part of the total length of the employment spell.

eligible to be used in this way if at least some of their associated weeks fall in the qualifying period. Recall that the qualifying period equals the lesser of 52 weeks or the time since the last UI claim ended. To establish the qualifying period we link the ROE to UI claim records included as part of the COEP. We note the length of jobs that occur in the 52 weeks preceding the start of the employment spell or the period since the end of the last UI claim, whichever is less. We use this to calculate the number of weeks of employment suitable for establishing entitlement that the individual carries into the employment spell being studied. This is particularly important for examining changes in entitlement discussed in Section 3. Our sample sizes are: 5,348 seasonal employment spells and 4,143 non-seasonal employment spells ending in the first three quarters of 1996, and 4,316 seasonal employment spells and 3,287 non-seasonal employment spells ending in the first three quarters of 1997.

3. *Impacts on Entitlement and Eligibility*

In this section we examine the impact of the switch to an hours-based system on eligibility and entitlement. The stated goal of switching to an hours-based system is to expand the coverage of the system to workers in more non-standard work arrangements and to match insurance more closely to the amount of work actually done (Nakamura and Diewert (1997)). Thus, under the old system an individual working 20 weeks at 20 hours per week generated the same weeks of entitlement as an individual working 20 weeks at 40 hours per week. Under Employment Insurance (EI), the latter individual will be counted as working twice as much for the purposes of generating weeks of entitlement. We will investigate whether the switch to an hours-based system in fact changed eligibility and weeks of entitlement and who was most affected by any changes.

Table 2 contains estimates of changes in eligibility for samples of workers under the pre-1997 Unemployment Insurance (UI) rules and the post-1997 (EI) rules. Thus, to construct the first column of the table, we applied the pre-1997 rules to the employment spell duration data for the 1996 sample of seasonal workers to calculate whether each individual appeared to be eligible to receive UI at the end of the Record of Employment (ROE) sample job under those rules.[6] We also applied the post-1997 rules to the same sample to calculate eligibility under the new rules.[7] The two sets of rules differed primarily in that the EI rules are hours-based while the UI rules are weeks-based, but they also differ in that jobs with less than 15 usual hours worked per week could not be counted toward eligibility before 1997 but could be counted toward eligibility after 1997. The first column of the table then shows the percentage of jobs in the 1996 seasonal sample that moved from being ineligible for benefits under UI to being eligible under EI (we will use the shortform, INEL→EL, to denote this group), the percentage that moved from being eligible under UI to ineligible under EI (we will use the shortform, EL→INEL, to denote this group), the percentage that were ineligible under both, and the percentage that were eligible under both. This exercise is repeated for seasonal and non-seasonal worker samples in both years. Following Kuhn and Sweetman (1998), the results using the 1996 samples could be seen as the pure impact of the rules changes while the 1997 sample results could be seen as incorporating behavioural responses to the rules changes.

Perhaps the most striking feature of the numbers in Table 2 is the small size of the EL→INEL group. Using any one of the samples, the percentage of job endings that correspond to loss of eligibility is less than 1.5 percent. For non-seasonal workers it is 1 percent or less. Thus, conjectures that taking away a firm's ability to adjust to the a

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6 In doing this, we took account of weeks of work in the qualifying period that the individual possessed at the start of the employment spell. This is described in the data section earlier.

7 In creating total hours worked, we multiplied the same weeks measure as was used for calculating eligibility under the pre-1997 rules by the usual hours worked per week reported in the survey questions related to the Record of Employment (ROE) sample job.

tightened post-1994 UI system by giving each worker fewer hours per week would have a substantial impact on eligibility proved unfounded. Given that one might expect such effects to be most prominent in seasonal jobs and, as we will show below, that seasonal jobs tend to be long-hours jobs, this may not be surprising. A switch to an hours-based system for individuals typically working over 35 hours per week actually makes qualification for UI easier because of the way the weeks-based Entrance Requirement (ER) was translated into an hours-based ER. On the other hand, the new system would have allowed approximately 5 percent of seasonal workers and 2.5 percent of non-seasonal workers to gain eligibility if it had been applied to the 1996 sample. Thus, the change in rules led to a small but statistically significant net increase in eligibility.[8]

The third and fourth columns of Table 2 give the changes in eligibility status using the 1997 data. Similar patterns to the 1996 data can be seen. However, the percentage calculated as eligible under the EI rules is actually smaller for the 1997 seasonal sample than for the 1996 seasonal sample (86.2 percent versus 88.3 percent). If the differences between columns 1 and 3 were just due to behavioural response to the rules change, one would expect a larger percentage to be eligible using the 1997 sample than was predicted using the 1996 sample. The opposite result indicates that labour market changes between the two years probably had a larger impact than any behavioural response to the Employment Insurance (EI) changes.

Who experienced eligibility changes due to the switch to the hours-based system? The sample sizes associated with the EL→INEL and INEL→EL groups are relatively small (for example only 61 non-seasonal workers in the 1997 data were in the EL→INEL group), but a few regularities are apparent. First, the EL→INEL group appears to consist largely of women in part-time/part-year jobs in the service and public sectors. In the 1997 data, 65 percent of the seasonal EL→INEL group and 85 percent of the non-seasonal EL→INEL group were women. Individuals in both the seasonal and non-seasonal EL→INEL groups reported approximately 20 usual hours per week of work on average and a little over 20 weeks of work in the previous year. They also tend to be lower wage workers, particularly among the non-seasonal workers. The INEL→EL group is also disproportionately composed of women, although not as predominantly so (approximately 54 percent of the non-seasonal INEL→EL group were women). They also tend to be part-time workers but ones with more weeks worked and more usual hours per week (27 average usual hours per week for the non-seasonal INEL→EL group) than those who lost eligibility. They are more evenly spread across sectors than the EL→INEL group. Finally, they have above average wages. Thus, both the EL→INEL and the INEL→EL groups tend to be part-time workers with the EL→INEL group having shorter work weeks and thus not having quite enough total hours to qualify. The EL→INEL group appears to be a more disadvantaged and vulnerable group and some thought should be given to addressing their loss of eligibility. None the less, this is a very small group.

Even though few workers lose eligibility with the EI rules, concerns have been raised by labour organizations that a substantial number of workers need to work many more weeks

8 These results are in line with those reported in Sweetman (1999).

to qualify. To assess this claim, Figure 1 plots the cumulative distribution function of the difference in the weeks required to attain eligibility using the 1997 rules minus the weeks needed using the 1996 rules for the 1997 sample of seasonal workers. The figure reveals that approximately three-quarters of the sample needed the same or fewer weeks to qualify under the new EI rules. However, at the same time, 11 percent of workers needed at least 10 extra weeks of work to qualify under the new rules and 6 percent needed at least 20 additional weeks relative to what was needed under the 1996 rules. Thus, while qualification is easier under EI for most workers, a substantial number of workers do need a considerable extra number of weeks to qualify in the new system. The fact that even using 1996 data, few workers are predicted to lose entitlement under EI means that most of the workers who needed extra weeks of work were already engaged in work patterns that gave them those weeks.

To examine the impact of the EI changes on entitlement, we performed a similar exercise to that used to form Table 2 but calculated total weeks of entitlement under the two systems rather than just whether individuals would be eligible or not. We calculated weeks of entitlement using the rules relevant from 1990 to 1994, those relevant from 1994 to 1996, and those relevant after 1996 (the EI rules). In Figure 2, we plot the cumulative distribution functions (CDFs) of weeks of entitlement under the three sets of rules using the 1997 non-seasonal worker sample. The left most point on each line shows the percentage who are not eligible for benefits at all. This suggests that the 1994 rule change had little impact on eligibility for this group (as indicated by Kuhn and Sweetman (1998) and predicted by Green and Sargent (1998)) but, in accord with Table 2, the EI change had the net effect of reducing the percentage who were ineligible. Much of this reduction appears to be related to the elimination of the requirement that a job must be over 15 hours per week to qualify.

In our calculations, 11 percent of all 1997 workers (i.e., seasonal and non-seasonal combined) do not work enough hours to achieve benefit eligibility under the EI rules. A recent Human Resources Development of Canada (HRDC) report (HRDC (1998)) calculates that 22 percent of the unemployed who were laid off or quit with just cause in 1997 (the closest of their groupings to our sample) did not meet entrance requirements. The difference between the two numbers could arise because the HRDC numbers are based on a survey of non-employed workers at a point in time rather than on administrative files related to job terminations, as is used here. Thus, HRDC provides a stock measure of the proportion of the unemployed who do not achieve eligibility while ours is a flow measure. Other sample selection criteria, such as the fact they examine only non-employed individuals who worked in the last 12 months, may also help account for the discrepancy.

In terms of entitlement, the largest change was generated by the 1994 changes that led to substantial reductions in entitlement. While 35 percent of the sample would have been entitled to 50 weeks of benefits under the pre-1994 rules, less than 15 percent would have been entitled to 50 weeks of benefits under the 1994-96 rules. Essentially, the whole entitlement distribution between 20 and 50 weeks shifts left by as much as 10 weeks after 1994. The EI changes actually generate slight increases in entitlement, on average, between 20 and 45 weeks. This is not surprising given that usual hours of work per week are greater than 35 for this sample.

Figure 3 repeats Figure 2 for seasonal workers. Here the rate of ineligibility is higher than for non-seasonal workers and the EI system again leads to a net improvement for this group. Again, the imposition of the 1994 changes causes a shift left in the entitlement distribution: a much larger shift than is observed for non-seasonal workers. This fits with results in Kuhn and Sweetman (1998) and Storer and Van Audenrode (1998). The switch to the EI system again increases entitlement: in this case, shifting the entitlement distribution to the right by 2 to 3 weeks. Figure 4 presents the entitlement CDFs for seasonal workers who worked between 10 and 20 weeks. This is a group for whom the losses from the 1994 changes were particularly large. Again, because this tends to be a group with high usual hours per week, EI substantially increased their eligibility and entitlement.

While the figures indicate generally small increases in entitlement due to the switch to EI, this may hide quite large individual entitlement changes. Figure 5 plots the CDF corresponding to the change in entitlement created by moving from the 1996 to the 1997 rules calculated for both the 1996 and 1997 seasonal worker samples. Focusing on the 1997 sample, the figure indicates that approximately 30 percent of the sample lost entitlement from the switch, 28 percent stayed the same, 33 percent gained between 1 and 10 weeks, and 9 percent gained more than 10 weeks. Thus, there is a balance of entitlement increases and decreases with more large increases than large decreases.

Table 3 presents sample average characteristics for groups of individuals whose entitlement: dropped by over five weeks in the switch to EI (column 1); stayed the same (column 2); and increased by over 5 weeks (column 3). The proportion of workers associated with each of the columns is .082, .28 and .16, respectively. The characteristics in the table tell a complex story. The most striking difference among the groups is in the gender composition: 73 percent of those with increased entitlement are males, compared with 50 percent of those who stay the same and only 30 percent of those who lost entitlement. Means related to other characteristics are also strikingly different, however. Among those with reduced entitlement, 20 percent have not completed high school, compared with 38 percent for those with increased entitlement. In contrast, 27 percent of those with reduced entitlements but only 11 percent of those with increased entitlement have a university degree. Hours worked per week are much shorter for those with reduced than for those with increased entitlement (20.7 versus 55.6, respectively). This is to be expected since the individuals who will have lower entitlement in a switch to an hours-based system are those with short hours per week while those who obtain more entitlement are those with quite long hours per week. The group experiencing entitlement increases is disproportionately concentrated in seasonal industries (primary, food processing and construction) and has a high proportion who have worked for their employer before. In contrast, the group with reduced entitlement has a very large concentration in the public sector and in the service sector. Finally, the mean wages for those who experience entitlement increases and reductions are well above the whole sample mean of $13.88. However, their median wages are close to the whole sample median of $12.00.

Taken as a whole, the characteristic proportions in Table 3 point to a shift in entitlement favouring low educated, male, seasonal workers and against more highly educated, female, public and service sector workers. Both the group with entitlement increases and the group with entitlement declines appear to be high wage earners; however, the fact that means and medians are so different indicates that the wage distributions for both groups are highly skewed. Thus, some in both groups are very high wage earners. The medians indicate, however, that for many of the rest, their wages are similar to those in the overall population of laid-off workers. Indeed, the group experiencing entitlement reductions are a somewhat bimodal. The top 50 percent of these individuals, arranged by wage, earn very good wages and are highly unionized workers in the public sector (likely teachers and nurses). The bottom 50 percent are much lower wage workers who are less educated, largely not unionized and work in the service sector.[9] A similar phenomenon occurs with those experiencing entitlement increases: the bottom 50 percent of wage earners in this group are even less educated than indicated Table 3. Thus, while there are clear distinctions between who experiences entitlement increases and entitlement decreases in the switch to an hours-based system, a clear conclusion on how the redistribution should be judged is difficult. The shift in systems redistributes toward low skilled, seasonal male workers, many of whom none the less earn good wages. It redistributes away from two groups of female workers: well paid, unionized, part-time public sector workers who earn quite good wages; and lower paid, non-union, part-time service sector workers who earn lower wages. Thus, it actually creates a redistribution away from one group of "non-standard" workers and toward another. Finally, it is worth re-emphasizing that whatever the composition of these groups, the group with entitlement increases is clearly larger than the group with decreases.

The other feature of Figure 5 is that it presents the change in entitlement CDFs using both the 1996 and 1997 seasonal worker samples. A difference between the CDFs for the two samples would potentially indicate behavioural adaptation to the new rules. There is, in fact, a difference in the CDFs in the 1 to 5 week reduction range. Using the 1997 sample, the probability of losing 2 to 5 weeks of eligibility in the switch to the EI system is approximately 3 percent larger than if the same probability is calculated using the 1996 sample. This would fit with a behavioural adaptation to the new system either in the form of fewer hours or fewer weeks of work for some workers. The remainder of this study attempts to uncover the size and form of any such behavioural adaptation.

9 For the lower 50 percent of those with entitlement decreases, 29 percent have not completed high school, 18 percent are unionized, 23 percent are in the public sector, and 63 percent are in the service sector.

4. Behavioural Responses

We turn now to examining potential behavioural responses to the switch to an hours-based system. We begin with an investigation of reactions in terms of job durations measured in weeks. We examine the durations in weeks because this permits an easier link to earlier papers that study the impact of entrance requirements and because, given that the new eligibility and entitlement calculations are often the same as those under the pre-Employment Insurance (EI) regulations multiplied by 35, weeks remain a natural unit of account for study. Also, we believe that measurement error will be much less in both duration measured in weeks and usual hours per week than in duration measured in hours. It is worth noting that, given the results of the previous section, one would expect to find some behavioural adaptation, although probably to a small degree. We turn, now, to investigating whether any behavioural change did, in fact, occur.

4.1 Empirical Hazards

We begin by examining empirical hazards broken down by Unemployment Insurance (UI) regions. The hazard rate at week *W*, is the probability an employment spell ends at week *W* conditional on it lasting at least as long as week *W*. The empirical hazard is calculated as the ratio of the number of employment spells terminated at week *W* divided by the number of spells at risk of ending at week *W* (i.e., the number of spells observed to last at least *W* weeks). For the purposes of this initial examination, we group regions according to their entrance requirement. In the pre-1997 period,[10] the entrance requirement was based on weeks worked in the qualifying period and the unemployment rate in the UI region. This yields the nine possible entrance requirements set out in column 2 of Table 1.[11] The data are grouped in nine entrance requirement regions corresponding to the possible requirements in Table 1. After January 1, 1997, the entrance requirement switches to an hours-based system, but one where the requirements are just the pre-1997 weeks-based requirements multiplied by 35. The hours-based entrance requirements are shown in the third column of Table 1. The number of possible entrance requirements and the regional groupings stay the same,[12] so we plot the employment duration in weeks for each of the entrance requirement (ER) regions for both pre- and post-January 1, 1997, on the same figures. Further, we break the data down according to whether the job is seasonal or non-seasonal in nature. We do this based on evidence in Green and Sargent (1998) that shows that UI policy parameter impacts are different for seasonal versus non-seasonal workers and are much larger for the former. The results are shown in Figures 6 through 23. In each case, the 1996 plot corresponds to jobs ending in the first three quarters of

10 From here on, when we refer to the pre-1997 UI system we mean the system in place just before the switch to the EI system.

11 There is also a separate entrance requirement for new labour market entrants. Our definition of the entrance requirement variable in the duration model estimates that follow takes account of this separate entrance requirement.

12 The actual UI geographical regions in each entrance requirement (ER) group do not stay the same across years. Some UI regions move across entrance requirement categories because of changes in the unemployment rate between years.

1996, and the 1997 plot corresponds to jobs ending in the first three quarters of 1997.[13] At this point, we have data only up to the third quarter of 1997. Initial plots by quarter for the pre-1997 period indicate that there is substantial seasonal variation to the hazard plots. Thus, to maintain comparability, we restrict the 1996 data to the first three quarters of the year as well. The vertical solid line marks the relevant, pre-1997 entrance requirement for the given region.

One point immediately noticeable on all 18 figures is the large number of spikes in each.[14] As discussed earlier, an ER effect should in principle appear as a spike in the hazard rate at the ER as well as a rise in the average value of the hazard after versus before the ER. The large number of natural spikes in the hazard at all different weeks makes it very difficult to establish the existence of an ER effect merely from inspecting plots of empirical hazards. However, keeping that caveat in mind, the empirical hazards for permanent employees (shown in Figures 15-23) show very little evidence of ER effects in either year. In most entrance requirement regions, the hazard is actually below average at the ER point.

For seasonal employees (Figures 6-14), the empirical hazard rates are somewhat more suggestive of ER effects. This is particularly true in the over 13 percent unemployment rate regions. Here, the hazard is substantially lower before versus after the ER and there is a spike at the ER number of weeks. Interestingly, the spike at the pre-1997 ER (12 weeks) is substantially larger in 1997 than in 1996 in the over 13 percent region. In particular, while the 1996 pattern fits predictions for a UI impact, the spike at 12 weeks is dominated by spikes at other weeks. This is not true in 1997. A more distinct UI type pattern in 1997 is not what one might predict for the impact of a switch to an hours-based system. One might reasonably have predicted a smaller ER spike in 1997, as the hours-based requirement means that the number of weeks of work required to qualify for benefits varies with the number of hours worked per week. A larger apparent spike at the pre-1997 ER in 1997 is also evident in the 10-11 percent UI region. Most other regions show limited evidence of an ER spike. Plotting empirical hazards for both years together highlights the dangers of trying to draw too many conclusions on ER effects from these simple plots. Thus, a spike in the hazard at 16 weeks in 1996 in the 9-10 percent region might be taken as evidence of an ER effect if that hazard is plotted on its own. However, the 1997 hazard shows a similar spike at the 17 weeks point, suggesting that these spikes may correspond to reported endings of four-month jobs rather than ER effects. Overall, Figures 6-23 suggest that there may be ER effects for seasonal workers in high unemployment regions but that there is little easily observable evidence for seasonal workers in other regions and for non-seasonal workers in all regions. Both the larger

13 While the 1997 system is based on reported hours worked, the ROEs still report information on duration of jobs in weeks. This latter information is used to construct the duration of employment spells measured in weeks in 1997.

14 The sample sizes associated with each ER region are as follows: for 1996 (in the order of the figures), 71, 220, 740, 870, 641, 480, 227, 411, 1696, 66, 200, 717, 995, 558, 444, 246, 234, 778; and for 1997 (in the order of the figures), 84, 475, 434, 543, 250, 182, 465, 324, 1559, 71, 384, 449, 706, 219, 158, 370, 280, 650. The small sample sizes for the under 6 percent unemployment rate regions are the source of the excessive spikeness of the hazards in Figures 1 and 10.

apparent impacts for seasonal workers and the fact that the most striking entrance requirement type pattern is in the over 13 percent region are consistent with the findings in Green and Sargent (1998). Entrance requirement effects in those regions appear to intensify rather than dissipate in 1997 relative to 1996.

4.2 Econometric Models of UI Impacts

4.2.1 Estimation Approach

Simple comparisons of empirical hazards before and after the policy change are useful for highlighting patterns and raising questions but are not a good basis for drawing conclusions on policy impacts. Differences over time could arise due to changes in labour market conditions and composition of the labour force as well as from changes in the geographic regions making up a given ER group across years. Further, what we are seeking to identify is essentially a pattern in the hazard function that includes a spike at the entrance requirement number of weeks. While there are patterns suggestive of UI impacts, most notably in the over 13 percent unemployment rate region, these patterns cannot be identified as UI effects simply by looking at the empirical hazards. Spikes could arise purely due to reporting regularities such as individuals leaving jobs at the ends of months. Similarly, those spikes could disappear in later years because of calendar changes (e.g., changes in the number of weeks included in a month) rather than because the policy has changed. In this section, we estimate duration models including a set of covariates to capture worker characteristics and the state of the labour market and another set of covariates to capture potential UI impacts. We focus on duration measured in weeks in order to provide an easier link to the earlier literature.[15]

Our estimation is based on a proportional hazards model in which,

$$1)\quad h_i(W) = h_0(W)\exp(x_i(W)'\beta)$$

where $h_i(W)$ is the hazard function for person i in week W, $h_0(W)$ is the "baseline" hazard function common to all individuals, $x_i(W)$ is a vector of observable characteristics which may vary with W and β is a parameter vector. For different values of $x_i(W)'\beta$ the hazard function for individual i is shifted proportionally up or down relative to the baseline hazard. We allow the β vector to differ across three ranges of weeks: 1-11, 12-20, and over 20. Thus, the impacts of covariates on the conditional probability of being laid-off are the same for any weeks within a range (e.g., weeks 5 and 6) but differ for any week in one range versus any week in another (e.g., weeks 5 and 12). This permits more complex differences in the distribution of weeks among workers with different characteristics, allowing more precise and reliable estimation of UI impacts.

15 Because the EI rules are based on 35-hour blocks, even under EI it is natural to study data patterns at some level of aggregation above hours. We considered the possibility of working in numbers of "standard weeks," where a standard week consists of 35 hours of work. The difficulty with that approach is that it makes links to time varying covariates such as the unemployment rate very difficult. Thus, we decided to focus our attention on weeks worked. Given that we find little behavioural response in weeks worked or in hours per week, the response in terms of total hours worked must also be small.

Given our concern with the location of spikes in the hazard, an estimation method that permits direct examination of the baseline hazard is preferable to ones that either impose a parametric form for the baseline or eliminate it altogether. We adopt a specification detailed in Meyer (1990) in which the baseline is effectively represented by a series of dummy variables corresponding to each possible duration week.

Assuming that $x_i(W)$ does not vary within a given week, Meyer (1990) shows that, with a continuous time proportional hazards model observed at discrete intervals, the appropriate log likelihood function has a contribution for the i^{th} spell given by,

$$2)\quad l_i(k_i/x_i,\gamma,\beta) = \delta_i \log[1 - \exp(-\exp[\gamma(k_i) + x_i(k_i)'\beta])] - \sum_{H=1}^{k_i-1} \exp[\gamma(H) + x_i(H)'\beta]$$

where, k_i is the observed length of the i^{th} employment spell, $\gamma(k)$ is the baseline parameter corresponding to the k^{th} week, and δ_i equals one if the spell ends before being censored. Again, this contribution is written as if there is only one β vector for simplicity of exposition, but in the actual estimation we allow the β vector to vary across the ranges listed above. Summing 2) over i yields the complete log likelihood function. We estimate this model separately for seasonal and non-seasonal workers in each year.

Our goal in estimating these models is to identify the effects of the UI system on the employment hazard. To do this, we incorporate a set of variables related to incentives associated with the entrance requirement. Primary among these is ER, which equals 1 in the week the entrance requirement is met. Thus, ER is a time varying covariate that takes a value of zero in weeks before the relevant entrance requirement week, one in the week the entrance requirement is met (regardless of whether the employment spell actually ends in that week), and zero in all following weeks. We also include dummy variables of the same form corresponding to 1-2, 3-5, 6-10 and 11 or more weeks before the entrance requirement. In addition to ER related effects, we also look for effects due to seasonal workers qualifying for enough weeks of benefits to fill out the remainder of a 52 week period. For seasonal workers who intend to return to work at the beginning of the next season, this corresponds to the point at which they are eligible for maximum possible UI benefits. If they work a week less than this, they will not be entitled to enough weeks of benefits to carry them to the next season. If they work a week more, there will not be enough time before the next season to collect their UI entitlement and hence they will receive less in total benefits. Green and Sargent (1998) find that there was more clustering of seasonal job endings at this maximum entitlement point than at the ER in 1989. To examine this effect, we include a time varying covariate that equals one in the week in which an individual qualifies for just enough benefits to fill out the remainder of a 52 week period[16] (we call this the MAXYR point) and another that equals one for weeks between the ER and the MAXYR point. Since we expect this year effect to be relevant

16 Actually, to take account of the two-week waiting period for benefit take up, we select the week in which weeks worked plus weeks of entitlement equals 50.

only for seasonal workers, we include these latter variables only in the estimations on the seasonal samples. The UI variables generate effects that are defined relative to the hazard rates for weeks beyond the entrance requirement for non-seasonal workers and for weeks beyond the MAXYR point for seasonal workers.

Returning to 1), it is useful to discuss the x vector as broken into two parts: x_{1i}, which contains time invariant covariates, such as gender; and x_{2i}, which contains time varying covariates, including the UI variables just defined. The x_{1i} component determines the height of the hazard function at all points for a person with given characteristics while the x_{2i} component and the baseline hazard together determine the way the hazard varies with W. The baseline hazard reflects basic duration dependence: the way the exit rate from a job varies with the length of the job holding constant any UI effects. The UI estimated effects reflect duration dependence that is directly related to the UI system. Our goal is to separately identify these two types of duration dependence. Note that the specification of the baseline hazard as a step function in W permits a very flexible form for basic duration dependence.[17]

The UI effects are identified primarily through variation in the value of the entrance requirement across UI regions and within regions over time with variation in regional unemployment rates. This plus similar variation in entitlement creates variation in MAXYR as well. As discussed in Green and Sargent (1998), the result of these forms of variation is substantial variation in UI system parameters. The fact that the identifying variation for ER is related to variations in UI region unemployment rates raises the possibility that we will identify the effects of variations in the unemployment rate rather than UI effects. We identify ER effects relative to unemployment effects in two ways. First, we include the provincial unemployment rate as a time varying covariate in our estimation. Since (with the exception of PEI) provinces contain several UI regions, there is variation in the unemployment rate measure that will be generated by movements in labour force conditions in UI regions other than the one in which the individual resides and hence will not trigger changes in the entrance requirement facing the individual. Second, nonlinearities in the schedules translating unemployment rates into UI parameters values aid in identification.

Table 4 contains variable definitions and means for the covariates used in our analysis. Means are presented separately by year and seasonality status. The wage variable is included to control for possible changes in the wage distribution across years. Such changes could affect firm decisions to layoff workers. The unemployment rate is included to ensure that the ER variables are not accidentally picking up differences in labour market conditions facing different individuals. While the employment spells we examine are all officially termed as layoffs, it is often difficult to separate individual from firm-initiated terminations. This is especially true with seasonal layoffs where workers knew the job would be terminated, and approximately when, when they took the job. Thus, we include

[17] One could, instead, form the baseline hazard as a polynomial in W, but such a specification would miss reporting spikes. If key UI related points occur at locations with common reporting spikes, one would then be in danger of attributing the reporting spike to UI effects.

sets of variables corresponding to individual preferences over job durations (male, maryd, cplc and spswrk), outside opportunities facing individuals (the age and education variables) and variables related to the firm and the worker-firm interaction (firm size variables, industry, empbf, union). Our overall aim is to control for determinants of job duration to ensure that changes in the job duration distribution in 1997 relative to 1996 that are actually due to changes in covariate distributions (e.g., the industrial distribution) are not mistakenly attributed to the policy change.

The covariate means presented in Table 4 indicate substantial differences between seasonal and non-seasonal workers and some important differences across years. In examining these means, it is important to keep in mind that the sample consists of employment spells observed to end in the given year. Thus, for example, university educated workers are under-represented in this sample relative to the workforce as a whole because university educated workers are less likely to lose their jobs. With that in mind, among workers ending employment spells in the first three quarters of 1996 or 1997, seasonal workers are more likely to work in high unemployment provinces, are younger, more likely to be male, more likely not to have completed high school, but also more likely to have a university education, less likely to have a spouse who works, more likely to have worked for their employer before, less likely to be unionized and more likely to be in the primary, food manufacturing, construction or public sectors. Much of this fits with a standard idea of a seasonal worker. Between 1996 and 1997, at least two significant changes occur in the sample. First, average wages of seasonal workers go down and average wages of non-seasonal workers go up, by approximately equal amounts. Second, the unemployment rate facing seasonal workers rises by half a percentage point. We will control for these types of changes in covariates using the estimates from the duration model described above.

4.2.2 Estimation Results

The results from the estimations of the duration model are presented using both tables and figures. Table 5 contains the estimates of the non-UI related coefficients and their associated standard errors for each of the three samples we examine. Again, recall that we allow the coefficients on the non-time varying covariates to differ across three ranges of weeks. In the 1 to 11 week range, the estimates indicate that males are statistically significantly more likely to face an employment spell termination than females for all groups. Education effects are in general not well defined, although those who have not completed high school are typically more likely to face a termination and those with a university degree less likely than workers who have completed high school. The firm size effect estimates become more negative between 1996 and 1997 for both seasonal and non-seasonal workers, indicating that smaller firms have become less likely to lay-off at short duration relative to firms with over 500 workers. For seasonal workers, being covered by a collective agreement makes layoffs in this short weeks range less likely but the union effect is not well defined for non-seasonal workers. In terms of industry, all industries other than the public sector are more likely to have terminations in this week range relative to service sector workers. Age has a statistically significant negative effect for all groups but it appears to be much more important for non-seasonal than seasonal workers.

Several covariate effects change in the 12-20 week range (the range that includes the ER) compared with the 1-11 week range. Higher wage workers have a statistically significant smaller probability of facing termination in the 12-20 week range, while wage had no statistically significant effect at shorter weeks. Males can no longer be said to be more likely to lose their jobs in this week range as well. The education profile of the hazard has become, if anything, flatter relative to the shortest week range. Interestingly, whether an individual worked for the employer before has no statistically significant effect at short weeks but has a large and positive effect for weeks 12-20. This may fit with a pattern in which workers and firms establish an ongoing relationship that includes a commitment on the part of the firm to ensure that workers obtain enough work to qualify for UI. The industry effects are now more of a mixed bag, although terminations are again less likely in the public sector. Age maintains its importance, becoming a much more important factor for seasonal workers in this week range in 1997 in particular.

In the 21-40 week range, there are again several differences relative to other weeks categories. The wage effect is not statistically significant at conventional significance levels for seasonal workers, as was the case for the 1-11 week range. It is still negative and significant for non-seasonal workers, however. For seasonal jobs, termination decisions seem less strongly related to education in 1997 than in 1996 in this weeks range. Whether an individual has worked for the employer before returns to statistically insignificance in this range in 1996. Thus, whether such an earlier relationship existed seems important in the 12-20 week range but not elsewhere. Being in the public sector now only has a negative and significant effect for seasonal workers.

The two time varying covariates used in the estimations are the quarter of the year and the provincial unemployment rate. As might be expected, employment spells are more likely to terminate in the third quarter, when traditional seasonal and short-term jobs are more abundant, than in the first two quarters.[18] The third quarter effects are very similar for the two years for each seasonality group. In contrast to the quarter effects, the unemployment rate effects are negative, as one would predict, and statistically significantly different from zero at the 5 percent significance level.

Table 6 contains estimates for the UI related variables for each of the groups. In creating the 1997 estimates we generated values of the unemployment insurance variables based on hours of work. In particular, for each worker we used his or her reported usual hours of work per week from the survey to calculate the week in the employment spell at which they would just qualify for EI under the hours-based system. We also calculated the week in which they would qualify for enough benefits to cover the remainder of a 52 week period (i.e., the MAXYR point). Since some workers had fewer and some more than 35 usual hours of work in a week, the newly calculated ER and MAXYR points are typically different than those that would be assigned using the 1996 weeks-based rules. The employment hazard for seasonal workers shows a statistically significant (at the

18 We tested for a difference in hazard rates between the first and second quarters as well, but did not find such a difference. We, therefore, opted for the more parsimonious specification using only a third-quarter dummy variable.

5 percent level) spike at the ER in 1996. While the estimates for 1997 do not reveal a spike relative to the hazard in weeks that follow, the estimates reveal a statistically significant (at the 5 percent level) rise in the hazard at the ER point relative to the weeks directly preceding that point. The MAXYR effect is very similar in both years and is statistically significant at least at the 10 percent level in both. In contrast, the non-seasonal hazard has a negative effect at the hazard in 1996 and one that is not significant at any conventional significance level in 1997.

To ease interpretation of the UI estimates, we use the coefficients in Table 6 to generate predicted hazard function patterns based on a flat baseline hazard. Figures 24a and 24b show effects for seasonal workers in 1996 and 1997, respectively. The plots are created as though the regional unemployment rate is over 16 percent. This implies that the ER is 12 weeks and the MAXYR point occurs at 16 weeks. The two vertical lines in each figure mark the ER and MAXYR points. The 1996 plot shows a mildly declining hazard in the weeks leading up to the ER and a spike at the ER. The hazard is lower again in weeks 13 through 15 then shows a second spike at the MAXYR point. This fits with predictions from theory, although the extent of the decline in the weeks before the ER is very small. The 1997 plot shows a much lower hazard before than after the ER point but no noticeable declining pattern in the weeks leading up to the ER. The pattern after the ER point is very similar to that in 1996 with the glaring exception of the lack of a spike at the ER point in 1997. This may be a reflection of the penalty in the form of lower calculated weekly benefits for applying for EI with the bare minimum of required hours: a new feature under EI (see the discussion in Section 1). This, plus the fact that the MAXYR spike still shows up in 1997 in spite of being at a different point under the new system than the old for most workers, indicates some relatively sophisticated adjustment to the new system. Figures 25a and 25b contain similar plots for non-seasonal workers. The 1996 plot is erratic and does not at all fit with an ER effect, especially since the ER effect itself is negative. The 1997 plot looks more like the predicted pattern.

While Figures 24 and 25 help in picturing the estimated ER related patterns, they do not aid in understanding the magnitude of the estimated effects. In Figures 26 and 27, we again use the estimated ER effects but in this case impose them upon representative baseline hazards for each group. In particular, we use the estimated baseline parameters in conjunction with the estimated coefficients presented in Table 5 to create predicted hazards (in the absence of UI effects) for a base person. The base person is a single, non-immigrant male with a high school education, working in a non-union firm employing less than 20 workers in the service sector and who has never worked for this employer before. In addition, the individual is 37.5 years old and earns $14 per hour. Finally, we assign a value of 1 for the third quarter dummy variable and set the unemployment rate to 17 percent. We construct the hazard function for this individual using each of the four sets of coefficient estimates, then use the estimated ER effects from Table 6 to add UI related impacts. From Figure 26a, one can see that the ER related spike is noticeable relative to weeks just before and just after. The spike at week 12 is approximately .013 higher than the hazard level at week 11 and .008 higher than the hazard at week 13. The latter is a measure of how many employment spells have their termination "saved up" for the entrance requirement point. It indicates that approximately 1 in 100 jobs that last to the entrance requirement end at that point because it is the entrance requirement. This effect is large enough to be worthy of notice but could not be called huge. Note that what appears to be a reporting spike at 10 weeks makes the ER effect difficult to see from a simple inspection of the hazard. The spike at 12 weeks is approximately .01 higher than the hazard rate at 12 weeks without the inclusion of the UI effects. The effect at the MAXYR point is similar in size. It is about half the size of the effect at the point where seasonal workers just qualify for enough UI benefits to cover the rest of their year as estimated in Green and Sargent (1998). The result for 1997 shows similar size reductions in the hazard in the weeks before the ER and a similar size spike in the hazard at the MAXYR point. There is an apparent spike in the hazard at the ER point but this is present in the plot generated not using the EI related variables as well and thus is not actually an ER effect. There is very little difference between the hazard with and without UI/EI effects at the entrance requirement point.

The results for non-seasonal workers presented in Figure 27, in contrast, show very little evidence of an ER related impact. The patterns portrayed in Figure 25 are now applied to a much lower baseline hazard and are somewhat difficult to observe. These results are very similar to those in Green and Sargent (1998), who conclude that any UI related impacts for non-seasonal workers are not economically substantial. Thus, any important ER effect is confined to seasonal workers.

The discussion to this point has been largely in terms of patterns within years. However, we are also interested in whether any observed patterns change after versus before the introduction of EI. In Figure 28, we plot the empirical hazards for all seasonal workers for both 1996 and 1997. The hazards are largely coincident apart from a noticeably higher hazard rate in the 12 to 15 week range in 1997. The higher hazard rate in those weeks (i.e., the higher probability of a job that lasts that long ending in the 12 to 15 week range) implies that employment spells will be, on average, shorter in 1997. Indeed, the average duration of a seasonal employment spell in our data is 32.2 weeks in 1996 and 31.2 weeks in 1997.

Before attributing the higher 1997 hazard rate in the 12 to 15 week range to the introduction of the EI system, we need to consider the possibility that differences between the two years arise from differences in the composition of workers or the levels of economic activity between the two years. To assess this, in Figure 29a we replot the With UI Effects included lines from Figures 26a and 26b on the same figure. Since the hazard rates presented in these plots are based on the same covariate vector (described earlier), any observed differences cannot arise from differences in characteristics of workers, jobs or wages, or from differences in levels of economic activity, to the extent that the latter is controlled for by our unemployment rate variable. Figure 29a also shows a higher hazard rate in the 12 to 15 week range in 1997 relative to 1996. In fact, the hazard rates are higher in 1997 over a more extended range of weeks. This translates into a shorter average duration of employment for this type of worker in 1997 (19.8 weeks) relative to 1996 (21.3 weeks). The differences between the two hazard rate functions portrayed in Figure 29a stem from two sources. First, the UI/EI response has the effect of lowering the hazard rate out of employment in the weeks leading up to the ER, and this effect is considerably larger in 1997 than 1996. This is clear in the larger negative estimates for the coefficients on the pre-ER variables in Table 6. Second, the underlying hazard (i.e., the line showing the hazard rate when all the UI/EI variables are set to zero) is higher in 1997 both before and after the ER point. This is shown in Figure 29b, where the Without UI Effects lines from Figures 26a and 26b are replotted on the same figure.

Can these effects be attributed to the introduction of the EI system? The greater reduction in the hazard leading up to the ER estimated using the 1997 sample may correspond to individuals being willing to wait for what is now a shorter ER. Since, as we shall see in the next section, seasonal workers generally have usual hours per week over 35 and the EI entrance requirement is just the number of weeks required under the old system times 35, the entrance requirement will be shorter for most seasonal workers. Thus, they and their employers may be more willing to delay termination of jobs to reach what is now a closer goal. The higher hazard rate in the 10 to 14 week range in 1997 may also reflect an adaptation to the EI system. If shorter entrance requirements cause workers and firms to agree on shorter weeks of work across jurisdictions, then one could see higher hazard rates in these weeks ranges. Note that such an effect may not be picked up in our UI/EI variable coefficient estimates if this reaction to the new system occurs all across the country because our estimates depend on cross-regional variation. Under this interpretation, the 1.5 week shorter average employment spell duration for the type of individual whose hazard rate is portrayed in Figure 29 arises largely because of the introduction of the EI system. Some caution should be exercised in interpreting the results this way, however. Our attempt to control for the overall state of the economy relies on the unemployment variable in our estimation and this variable contains only cross-sectional variation. Thus, our estimated unemployment rate effect comes mainly from comparing workers in Ontario with those in Newfoundland, for example, at a point in time. Such an effect may be a poor proxy for the effects of changes over time within a given province. To the extent we have not fully controlled for changes in the economy, they will show up in differences in the baseline hazard rates shown in Figure 29b. Also, our estimates do not explain the larger spike in the hazard at 12 weeks in the 1997 versus the 1996 plots. This could represent responses to some other large event in the economy, such as provincial government adjustments to difficulties in their economies.

To get a clearer perspective on the size of our estimated effects, we calculate the average duration of employment under alternative scenarios. In each case, we make the calculations for the same typical individual used in creating Figures 26 and 27. We provide estimates for ER = 12 regions and consider scenarios in which the entrance requirement is set at 12 weeks and an alternative one where the entrance requirement is cut to its old level of 10 weeks. All else, including the entitlement calculation and therefore the MAXYR point, is held constant between the two scenarios. An alternative interpretation of this experiment is that it shows what would happen to individuals with 42 hours of work per week in the course of a shift from UI to EI. For such an individual, the entrance requirement they face would fall from 12 weeks to 10 weeks.

The results of the exercise are presented in Table 7. Using the 1996 sample parameter estimates, the average duration of employment for seasonal workers would fall from

21.3 to 21.0 weeks. Using the 1997 parameter estimates the fall is larger (19.8 to 19.1 weeks). The latter result (a 3/4 week fall in average duration resulting from a 2 week drop in the entrance requirement) represents quite a large response. Note that, given that there is no ER spike estimated using the 1997 sample, this does not occur because of the movement of a clustering of job endings at the ER point to the left in the distribution. Rather, it occurs from the combination of the two effects visible in Figures 29a and 29b: the greater rise in the hazard from pre- to post-ER weeks in 1997 and the higher baseline hazard in 1997. Whatever the cause, the results in Table 7 reinforce earlier conclusions that there have been adjustments in employment durations to the new system. These could be the source of the patterns seen in Figure 5.

Table 8 contains the UI/EI related parameters from an alternative estimation for the 1997 seasonal workers in which ER and the related EI/UI variables are calculated using the 1996 rules.[19] We do this to see whether workers and firms appear to continue to follow old patterns at the beginning of the new system. The estimates show a spike at the ER point, as in the 1996 estimation. The spike effect is larger, relative to the hazard in ensuing weeks, than that in 1996 but the difference between the hazard rate just before the ER and at the ER is actually smaller in 1997. Indeed, the hazard before the ER using the Table 7 estimates does not fall as it approaches ER as it does in 1996. Further, there is no evidence of a MAXYR effect. Thus, apart from the ER effect, which we argue ought not to appear if agents have fully adapted to the EI system, these alternate estimates show little evidence of behavioural response to the UI/EI system.

The differences and similarities in Tables 6 and 8 potentially indicate an interesting reaction to the switch to an hours-based system. There is a measurable spike at the ER in the estimation for 1997 seasonal workers presented in Table 7 even though the no-longer-relevant weeks-based 1996 ER is being used in the estimation. The 1997 spike could arise if most jobs were approximately 35 hours in length since the hours-based ER is simply the weeks-based ER multiplied by 35. However, as we will see in the next section, most jobs do not have average hours close enough to 35 to make the effective entrance requirement stay the same. For example, over 30 percent of jobs are listed as being 40 hours per week. Even in the shortest ER region, a 40 hour a week job would mean that the individual could reach the hours-based ER in 10.5 weeks, compared with the weeks-based ER of 12 weeks. Thus, the weeks-based ER appears to be a bad proxy for the point at which most individuals meet the hours-based ER. More likely, the fact that the weeks-based ER measure is still significant in the Table 8 estimation reflects "stickiness" of work arrangements. If this is true, then the 1997 patterns should be viewed as part of an adjustment process and not as the ultimate reaction to the policy change. The fact that the MAXYR effect is economically and statistically insubstantial in Table 8 may reflect the fact that terminating jobs at the point of maximum possible benefits suggests sophisticated usage of the UI system. Workers and firms using the system in this way may be less likely to simply stick with old arrangements, especially since for most seasonal workers the

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19 These estimates come from a specification that is identical to that reported in Tables 5 and 6 apart from the use of different UI/EI related variables. The results for covariates other than the UI/EI variables are very similar to those reported in Table 5 for seasonal workers and are not included to save space.

number of weeks at which they can qualify for maximum yearly benefits is much less in 1997 than in 1996. This is true because, as we will see, seasonal jobs are more likely to have long hours per week. In contrast, the 1997 results in Table 6 do show a substantial and statistically significant MAXYR effect, fitting with the notion that those who use the system in this way did manage to adapt to the new system rather quickly.

Finally, in all the estimations above, we generate the ER and related variables based on duration in the employment spell ending with the sample ROE from the COEP. Thus, we do not make use of the information on the weeks of insurable employment the individual had at the start of the employment spell. The latter is information generated using the status vector file. It is described in more detail in Section 2 and used as part of the eligibility and entitlement calculations in Section 3. When we re-estimate the specifications detailed in Tables 5 and 6 using the insurable weeks the individual possessed at the start of the employment spell to help create ER and the related variables, we find few statistically significant or economically substantial UI/EI effects. We interpret this as evidence that firms play a role in the adaptation to the UI/EI systems suggested in Table 6. Firms may help workers in generating employment spells that last the appropriate number of weeks or hours to reach the ER and MAXYR points. It seems unlikely, however, that they would tailor their aid to the individual work histories of each employee. Consider, as an example, a worker in a region where the entrance requirement is 12 weeks. Suppose the worker has two jobs in a year, with the first job having lasted 6 weeks. The employer on the second job may make a commitment to provide enough employment for his or her employees to reach the entrance requirement. This could involve a promise of 12 weeks of employment for all employees. For the specific worker being considered, however, only 6 weeks of employment are needed to get him to the entrance requirement, and if the worker alone were trying to just qualify for UI/EI then he would stop work after 6 weeks on the second job. However, it would be costly for the firm to try to shuffle work schedules to match the requirements of each such worker and, thus, it would likely just offer 12 week contracts. Our estimates are consistent with the ER spike arising at 12 weeks rather than 6 weeks in this example, which is consistent with firm involvement in the decision.

The examination of the distribution of weeks worked in an employment spell leads to several clear conclusions. First, in the pre-1997 period when the entrance requirement is weeks-based, duration model estimates indicate a clear ER effect for seasonal workers but not for non-seasonal workers. Even for seasonal workers, however, the effect is not huge. Our estimates indicate that approximately 1 in every 100 seasonal workers whose jobs last at least as long as the ER experience a job termination at the ER point because it is the entrance requirement and they have just qualified for UI. Slightly smaller effects are found for incentives induced by seasonal workers working just enough weeks to qualify for the maximum possible benefits they can receive from UI before the start of the next season. Together, these effects imply that a two-week reduction in the entrance requirement would only reduce average employment duration by 0.3 weeks. These results are similar in form but smaller in magnitude to those found in Green and Riddell (1997) and Green and Sargent (1998). Second, in the post-1997 period, there is clear evidence of adjustment to the new system. The hazard rate rises substantially when employment spells pass the ER point and there is a spike at the MAXYR point. Since, given changes in entitlement shown in Section 3, the MAXYR point changes substantially for many workers, this suggests some quite sophisticated tailoring of employment spells to the new EI system. The fact that there still appears to be a spike in the hazard at the ER point relevant under the old system indicates that a full adjustment to the new hours-based system may not yet have taken place. As in 1996, there is little evidence of any adjustment to the system by non-seasonal workers. Third, the hazard rate is higher in the 12 to 15 week range in 1997 relative to 1996 for seasonal workers. As a result, the average duration of a seasonal employment spell is 1.5 weeks shorter in 1997 than 1996. These differences between the two years appear to be due to both larger reactions to UI/EI parameters in 1997 and a higher underlying hazard in this week range in 1997. This might be a reaction to the new EI system but more observation is required to be certain of any such conclusion.

4.3 Usual Hours Per Week

A switch to an hours-based system could conceivably show up in the distribution of usual hours worked per week as workers and employers adjust to the new form of entrance requirement. Since other EI studies focus on hours per week, we will examine the hours per week distribution only briefly to paint a more complete picture of the adjustment to the hours-based system and to allow a link between our study and those focusing on hours per week.

In Figures 30 and 31 we present the cumulative distribution functions (CDFs) for total usual hours of work. We again divide our sample into seasonal and non-seasonal workers. We examine total usual hours (equalling the sum of usual regular time hours per week and usual overtime hours per week) because we assume that workers would use all hours worked to establish UI eligibility. The plots in Figure 30 indicate that the CDFs for 1996 and 1997 for non-seasonal workers are nearly identical. At no point is the difference between the CDFs for the two years larger than 2 standard errors. Both show substantial concentrations at 40 hours per week, with over 30 percent of jobs having 40 usual hours worked per week in each year. There is also evidence of concentrations at all weeks with numbers ending in 5 and especially at 35 hours per week.

The CDF plots for seasonal workers shown in Figure 31 are also quite similar between the two years. Unlike for non-seasonal workers, however, they cannot be said to be identical. In particular, the CDF is approximately .028 higher at hour 39 in 1997 relative to 1996. This difference arises gradually between 30 and 39 hours and then dissipates gradually between 40 and 49 hours. Thus, the 1997 distribution has more mass between 30 and 39 hours and the 1996 distribution has more mass between 40 and 49 hours. All of the differences between 30 and 49 hours are greater than two standard errors in size. Differences outside the 30 to 49 hour range are small in size and not statistically significantly different from zero at conventional significance levels. This pattern might plausibly represent an adjustment to the switch to an hours-based system. Employers might reallocate hours, cutting usual hours worked per week from over 40 to under 40 in order to ensure more individuals meet the ER.

Is this difference part of an adjustment to an hours-based UI system? To investigate further, we plot the CDFs for seasonal workers in regions with unemployment rates over 13 percent, since these are the workers who show most evidence of ER effects in their weeks distributions. These plots are given in Figure 32. In contrast to what one might expect if the observed changes are due to the EI policy change, the CDFs for 1996 and 1997 for seasonal workers in these regions are virtually identical. At no point is the difference between the CDFs either substantial or statistically significant at conventional significance levels. Thus, there is no immediate evidence of a policy-related impact in these regions. In contrast, Figure 33 plots the CDFs for seasonal workers in regions with unemployment rates less than or equal to 10 percent. Here, the differences between the years are even larger than those portrayed in Figure 31, reaching nearly .05 at hours just under 40 hours per week. The differences above 30 hours per week and below 40 hours per week are greater than two standard errors. Replots of the CDFs in Figure 31 by quarter indicate that the differences arise entirely in the first quarter of 1996 versus the first quarter of 1997. This might represent a very sophisticated response to the EI system as jobs that spanned January 1, 1997, had their hours per week automatically set at 35. Thus, firms and workers could agree to cut hours per week without any effect on eligibility or entitlement. The fact that no difference in the hours distribution exists for the highest unemployment regions, however, calls such an explanation into doubt. We conclude there is little or no evidence of a significant impact of the switch to an hours-based system on the distribution of usual hours worked.

Finally, it is interesting to note the differences in the CDFs for seasonal versus non-seasonal workers in a given year. The CDFs are very similar up to approximately 35 hours per week. However, non-seasonal workers are substantially more likely to work what might be called standard full time work weeks (i.e., 35 to 50 hours) and seasonal workers are much more likely to work long weeks. The probability that a non-seasonal worker works between 35 and 50 hours is 0.63 compared with 0.54 for seasonal workers. In contrast, the probability a seasonal worker works more than 50 hours per week is 0.18 compared with 0.06 for a non-seasonal worker. This is of interest, in part, because it implies that seasonal workers may actually be better positioned to gain from an hours-based system, since more of them tend to work long hour weeks.

5. Conclusions

The conclusions from this examination can be broken down into three main areas.

1. *The new system changes both eligibility and entitlement.*

The switch to the Employment Insurance (EI) system led to a net increase in eligibility. Very small numbers of workers who would have qualified for benefits under Unemeployment Insurance (UI) are ineligible under EI. These tend to be female, part-time/part-year service workers. Since this tends to be a vulnerable group, some attention should be paid to them, even though it is a very small group.

The switch to the EI system leads to an increase in entitlement, especially for seasonal workers who work 10 to 20 week jobs (the range of weeks corresponding to entrance requirements in the old UI system). This is true because seasonal workers tend to work long hours per week and because the EI eligibility and entitlement table values (reported in hours) are just the old UI table values (reported in weeks) multiplied by 35. Thus, a given number of weeks of work generates more weeks of entitlement under EI than under the old UI system if the usual hours per week on the job is over 35.

These entitlement and eligibility effects are dwarfed by the negative eligibility and entitlement effects generated by the 1994 UI changes.

In spite of net overall gains in entitlement, both the group that experiences an increase in entitlement and the group that has a decrease in entitlement in response to the shift to EI are sizeable. Approximately 8 percent of workers have reductions of over 5 weeks of entitlement in the switch to an hours-based system versus 11 percent who have increases of more than 5 weeks.

The groups that experience increases and decreases in entitlement are quite different. Not surprisingly, given that the old weeks-based entitlement and eligibility requirements are converted to the EI hours-based requirements by multiplying by 35, the group that has declines in entitlement tends to be individuals working part-time (less than 35 but more than 15 usual hours per week) and the group that experiences increases in entitlement tends to be individuals working long hours (more than 35 hours per week). In addition, apart from the reduction in maximum entitlement from 50 to 45, a change in entitlement resulting from the system change will only be observable if the individual worked less than the amount needed to qualify for maximum benefits in 1996. Thus, both groups will tend to be part-year workers. The group experiencing increases in entitlement tend to be male seasonal workers, while the group experiencing declines in entitlement tend to be female part-year, part-time workers. Both groups have average wages above the average for seasonal workers as a whole, but both also have significant numbers of low wage workers. For the group with entitlement reductions in particular, there appear to be two groups: better paid, unionized workers in the public sector and lower paid, non-unionized workers in the service sector.

Our overall conclusion in terms of eligibility and entitlement is that the switch to an hours-based system had almost no impact on eligibility but did increase the number of weeks of benefits workers were entitled to, in general. A shift from a weeks-based to an hours-based system redistributes weeks of entitlement toward long hour, predominantly male, seasonal workers and away from part-year/part-time, predominantly female workers. The average wages for all those affected by the shift to EI are above the average wage for seasonal workers as a whole.

2. *The examination of the distribution of weeks worked in an employment spell leads to several clear conclusions.*

There is evidence in the pre-1997 data of UI effects on the weeks of work distribution but these effects are generally not large. In the pre-1997 period when the entrance requirement is weeks-based, duration model estimates indicate a clear entrance requirement (ER) effect for seasonal workers but not for non-seasonal workers. Even for seasonal workers, however, the effect is not huge. Our estimates indicate that just under 1 in every 100 seasonal workers whose jobs last at least as long as the ER experience a job termination at the ER point because it is the entrance requirement and they have just qualified for UI. Slightly smaller effects are found for incentives induced by seasonal workers working just enough weeks to qualify for the maximum possible benefits they can receive from UI before the start of the next season. Together, these effects imply that a 2-week reduction in the entrance requirement induces only a 0.3 week reduction in average employment spell duration for seasonal workers.

In the post-1997 period, there is clear evidence of adjustment to the new system. The hazard rate rises substantially when employment spells pass the ER point and there is a spike at the MAXYR point. Since the MAXYR point changes substantially for many workers under the EI versus the UI system, this suggests some quite sophisticated tailoring of employment spells to the new EI system.

Estimates using post-EI change data indicate an increase in responsiveness to entrance requirement changes among seasonal workers. A 2-week reduction in the entrance requirement generates a 3/4 of a week reduction in average employment spell durations for seasonal workers in 1997. This is quite a large reaction implying that the shift to an hours-based system has led to noteworthy reductions in job lengths among seasonal workers.

The fact that there still appears to be a spike in the 1997 hazard at the number of weeks that would have been the entrance requirement under the old system indicates that a full adjustment to the new hours-based system may not yet have taken place.

There is little evidence of any adjustment to the UI or EI systems by non-seasonal workers in either year. A 2-week reduction in the entrance requirement does not change the average duration of employment spells for non-seasonal workers.

The hazard rate is higher in the 12 to 15 week range in 1997 relative to 1996 for seasonal workers. As a result, the average duration of a seasonal employment spell is 1.5 weeks shorter in 1997 than 1996. These differences between the two years appear to be due both to larger reactions to UI/EI parameters in 1997 and a higher probability of job durations falling in this week range in 1997 even in the absence of any measured UI/EI effects. This might be a reaction to the new EI system but more observation is required to be certain of any such conclusion.

Our overall conclusion with respect to impacts on employment spell durations is that the shift to the hours-based system has had noticeable impacts in terms of shortening seasonal employment spells. This is plausible since seasonal workers typically work long hours per week and, as a result, can acquire given levels of EI entitlement in fewer weeks under the new system. However, this effect should be kept in perspective. Non-seasonal employment spells, which make up the great majority of all employment, show no patterns of adaptation to the UI/EI systems or changes with the introduction of the EI system.

3. *There is virtually no change in the usual hours per week distribution between 1996 and 1997*.

In particular, the distribution of usual hours per week for seasonal workers in high unemployment regions, the group that appears to react most to UI incentives, does not change. Thus, while there does appear to be some reaction to the new system in terms of duration of jobs measured in weeks, there is no apparent reaction in terms of adjusting hours per week.

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Appendix A: Tables

TABLE 1
Entrance Requirements

Regional Unemployment Rate	1990-94 Weeks	1994-96 Weeks	After January 1, 1997 Hours
6% and under	20	20	700
6 — 7%	19	19	665
7 — 8%	18	18	630
8 — 9%	17	17	595
9 — 10%	16	16	560
10 — 11%	15	15	525
11 — 12%	14	14	490
12 — 13%	13	13	455
13 — 14%	12	12	420
14 — 15%	11	12	420
> 15%	10	12	420

TABLE 2
Changes in Eligibility Under Pre- and Post-1997 Rules

	1996 Data (percent)		1997 Data (percent)	
	Seasonal	Non-seasonal	Seasonal	Non-seasonal
Eligible Under EI/ Not Eligible Under UI	4.5	2.3	5.2	2.6
Not Eligible Under EI/ Eligible Under UI	1.3	0.7	1.4	1.0
Not Eligible Under EI/ Not Eligible Under UI	10.5	4.2	12.4	5.3
Eligible Under EI/ Eligible Under UI	83.8	92.9	81.0	91.1

Source: Authors' calculations based on COEP data described in text.

TABLE 3 Mean Characteristics for Entitlement Decreases and Increases			
Variable	**Lost Over 5 Weeks**	**Lost Zero Weeks**	**Gained Over 5 Weeks**
Wage			
Mean	$15.30	$14.40	$15.13
Percentile			
25	$ 8.00	$ 8.00	$ 8.00
50	$11.00	$12.00	$12.00
75	$18.00	$18.00	$19.00
95	$39.00	$28.00	$27.00
Hours	20.7	36.4	55.6
Male	0.29	0.50	0.73
Maryd	0.59	0.56	0.57
Cplc	0.31	0.26	0.25
Spswrk	0.47	0.40	0.36
Age	36.2	36.6	37.2
Eshs	0.20	0.22	0.38
Sps	0.17	0.18	0.14
Collg	0.13	0.16	0.13
Univ	0.27	0.20	0.11
Flt20	0.41	0.42	0.45
F2099	0.29	0.31	0.28
F1005	0.18	0.17	0.19
Union	0.38	0.33	0.28
Empbf	0.19	0.20	0.34
Prim	0.03	0.04	0.13
Manf	0.02	0.10	0.13
Food	0.03	0.03	0.07
Constr	0.03	0.11	0.20
Public	0.43	0.31	0.13

Source: Authors' calculations based on COEP data described in text.

TABLE 4
Variable Definitions and Means

Variable	Definition	Means			
		1996		1997	
		Seasonal	Non-seasonal	Seasonal	Non-seasonal
Wage	Average hourly wage on last job in employment spell	14.07	13.78	13.88	14.07
Unemp	Provincial unemployment rate in the quarter in which the relevant week of the spell falls	0.117	0.105	0.122	0.107
Age	Age of individual in years	37.0	38.0	37.2	38.5
Male	= 1 if the individual is male	0.58	0.55	0.56	0.52
Maryd	= 1 if the individual is married (incl. common law)	0.60	0.64	0.60	0.64
Eshs	= 1 if individual's highest education is elementary or some high school (base group has completed high school)	0.28	0.24	0.29	0.24
Sps	= 1 if highest education is some but not completed post-secondary	0.16	0.17	0.17	0.17
Collg	= 1 if highest education is college certificate or diploma	0.13	0.16	0.15	0.16
Univ	= 1 if highest education is university degree	0.16	0.11	0.15	0.13
Cplc	= 1 if the family type is a couple with children	0.29	0.30	0.29	0.29
Spswrk	= 1 if the respondent's spouse was working at time of the respondent's job separation	0.40	0.46	0.41	0.46
Immig	= 1 if the respondent was not born in Canada	0.07	0.12	0.08	0.12
Empbf	= 1 if the respondent worked for the last employer earlier	0.28	0.08	0.28	0.09
flt20	= 1 if the firm employed fewer than 20 workers at the respondent's location (base group is over 500 workers)	0.43	0.43	0.46	0.44
f2099	= 1 if firm employed between 20 and 99 workers at the respondent's location	0.31	0.31	0.29	0.31
f1005	= 1 if firm employed between 100 and 500 workers at the respondent's location	0.15	0.17	0.16	0.16
union	= 1 if respondent was covered by a union agreement on the last job	0.34	0.37	0.31	0.38

TABLE 4 (continued) Variable Definitions and Means					
		Means			
		1996		1997	
Variable	**Definition**	**Seasonal**	**Non-seasonal**	**Seasonal**	**Non-seasonal**
prim	= 1 if industry of last job was agriculture, fishing or forestry (base is service industry)	0.08	0.04	0.08	0.03
manuf	= 1 if industry of last job was manufacturing other than food manufacturing	0.10	0.21	0.10	0.18
food	= 1 if industry of last job was food manufacturing	0.05	0.03	0.05	0.04
constr	= 1 if industry of last job was construction	0.15	0.09	0.14	0.09
public	= 1 if industry of last job was in the public sector (incl. health and education)	0.28	0.21	0.25	0.22

TABLE 5
Non-UI Covariate Estimates

	Seasonal		Non-seasonal	
Covariate	**1996**	**1997**	**1996**	**1997**
Effects Weeks 1-11				
Wage	0.045 (0.039)	0.038 (0.037)	0.049 (0.10)	-0.13 (0.083)
Male	0.28 (0.065)*	0.26 (0.069)*	0.24 (0.13)+	0.40 (0.13)*
Maryd	-0.30 (0.077)*	-0.025 (0.087)	-0.092 (0.15)	0.16 (0.16)
Eshs	0.13 (0.068)*	0.059 (0.075)	0.081 (0.13)	0.39 (0.15)*
Sps	0.11 (0.082)	-0.075 (0.096)	0.0011 (0.14)	0.21 (0.17)
Collg	0.045 (0.087)	-0.078 (0.096)	-0.099 (0.16)	0.18 (0.18)
Univ	-0.36 (0.10)*	-0.34 (0.12)*	-0.69 (0.25)*	-0.058 (0.24)
Cplc	0.040 (0.069)	-0.058 (0.076)	-0.012 (0.14)	-0.12 (0.15)
Spswrk	0.031 (0.068)	-0.011 (0.074)	-0.23 (0.14)+	-0.27 (0.15)+
Immig	-0.043 (0.11)	-0.13 (0.12)	0.017 (0.17)	0.096 (0.18)
Empbf	0.053 (0.057)	0.082 (0.063)	0.61 (0.15)*	0.64 (0.16)*
Flt20	0.16 (0.10)+	-0.21 (0.11)+	-0.098 (0.22)	-0.34 (0.22)
F2099	0.24 (0.10)*	-0.17 (0.11)	0.13 (0.20)	-0.29 (0.21)
F1005	0.19 (0.11)+	-0.081 (0.12)	0.076 (0.20)	-0.47 (0.22)*
Union	-0.27 (0.063)*	-0.33 (0.072)*	0.22 (0.13)+	-0.033 (0.14)
Prim	0.15 (0.096)	0.41 (0.10)*	0.23 (0.25)	0.48 (0.27)+
Manuf	0.46 (0.085)*	0.46 (0.092)*	0.29 (0.14)*	0.48 (0.15)*
Food	0.41 (0.12)*	0.26 (0.13)*	0.55 (0.24)*	0.21 (0.27)
Constr	0.51 (0.075)*	0.51 (0.085)*	0.52 (0.17)*	0.67 (0.17)*
Public	-0.27 (0.083)*	-0.43 (0.096)*	-0.42 (0.18)*	-0.59 (0.23)*
Age	-0.34 (0.26)	-0.66 (0.31)*	-3.0 (0.54)*	-2.2 (0.59)*
Effects Weeks 12-20				
Wage	-0.19 (0.051)*	-0.16 (0.058)*	-0.23 (0.10)*	-0.041 (0.082)
Male	0.086 (0.072)	0.051 (0.077)	0.049 (0.13)	0.22 (0.15)
Maryd	-0.14 (0.098)	-0.18 (0.10)+	-0.087 (0.18)	-0.13 (0.20)
Eshs	0.084 (0.088)	0.19 (0.093)	0.082 (0.16)	0.41 (0.18)*
Sps	0.44 (0.094)*	0.048 (0.11)	-0.12 (0.17)	0.32 (0.20)
Collg	0.049 (0.11)	-0.052 (0.12)	-0.19 (0.18)	0.098 (0.22)
Univ	0.22 (0.11)*	-0.072 (0.13)	-0.16 (0.22)	0.16 (0.24)
Cplc	-0.15 (0.086)+	0.052 (0.096)	0.029 (0.16)	0.17 (0.18)
Spswrk	-0.061 (0.088)	-0.10 (0.090)	-0.31 (0.16)+	-0.26 (0.17)
Immig	0.16 (0.12)	-0.061 (0.14)	0.14 (0.21)	0.019 (0.23)

TABLE 5 (continued) Non-UI Covariate Estimates				
	Seasonal		Non-seasonal	
Covariate	**1996**	**1997**	**1996**	**1997**
Empbf	0.84 (0.064)*	0.73 (0.071)*	1.2 (0.15)*	1.0 (0.18)*
Flt20	0.16 (0.12)	0.32 (0.15)	0.094 (0.24)	-0.48 (0.26)+
F2099	0.0047 (0.12)	0.13 (0.16)	-0.26 (0.24)	-0.32 (0.25)
F1005	-0.10 (0.14)	0.21 (0.16)	-0.046 (0.24)	-0.24 (0.26)
Union	-0.12 (0.082)	-0.006 (0.088)	0.21 (0.15)	-0.23 (0.16)
Prim	-0.14 (0.12)	-0.027 (0.14)	0.63 (0.25)*	0.31 (0.30)
Manuf	0.097 (0.12)	-0.25 (0.13)	0.33 (0.16)*	0.45 (0.18)*
Food	-0.37 (0.18)*	-0.09 (0.16)	-0.29 (0.40)	-0.28 (0.37)
Constr	0.21 (0.10)*	-0.004 (0.11)	0.28 (0.19)	0.044 (0.26)
Public	-0.32 (0.090)*	-0.39 (0.10)*	-0.34 (0.20)+	-0.49 (0.23)*
Age	-0.67 (0.31)*	-0.95 (0.36)*	-2.6 (0.61)*	-1.7 (0.73)*
Effects, Weeks 21-40				
Wage	0.051 (0.034)	0.038 (0.035)	-0.16 (0.058)*	-0.33 (0.084)*
Male	-0.052 (0.068)	-0.093 (0.083)	-0.040 (0.084)	0.13 (0.099)
Maryd	-0.012 (0.089)	-0.17 (0.10)	0.16 (0.12)	0.041 (0.12)
Eshs	-0.048 (0.082)	0.058 (0.094)	0.23 (0.10)*	0.23 (0.11)*
Sps	-0.19 (0.099)*	-0.015 (0.11)	0.13 (0.12)	-0.019 (0.12)
Collg	-0.20 (0.10)*	-0.047 (0.11)	0.065 (0.11)	-0.001 (0.13)
Univ	-0.20 (0.10)*	0.13 (0.11)	0.19 (0.13)	0.035 (0.16)
Cplc	0.063 (0.077)	0.13 (0.096)	-0.0079 (0.094)	-0.26 (0.11)*
Spswrk	-0.15 (0.077)*	0.0088 (0.085)	-0.14 (0.097)	-0.11 (0.11)
Immig	0.085 (0.11)	0.048 (0.12)	0.24 (0.12)*	0.14 (0.14)
Empbf	-0.19 (0.12)	1.00 (0.08)*	0.001 (0.14)	0.033 (0.16)
Flt20	0.28 (0.12)*	0.24 (0.14)+	-0.043 (0.16)	0.26 (0.19)
F2099	0.28 (0.12)*	0.20 (0.13)	-0.097 (0.16)	0.21 (0.19)
F1005	0.15 (0.13)	0.051 (0.14)	-0.070 (0.17)	0.26 (0.20)
Union	-0.090 (0.072)	-0.29 (0.090)*	0.016 (0.10)	0.13 (0.10)
Prim	0.28 (0.11)*	-0.26 (0.18)	0.34 (0.19)+	0.60 (0.19)*
Manuf	-0.037 (0.11)	0.045 (0.14)	0.016 (0.11)	0.19 (0.13)
Food	-1.0 (0.21)*	-0.20 (0.18)	-0.42 (0.26)	0.045 (0.25)
Constr	-0.052 (0.10)	-0.35 (0.15)*	0.31 (0.13)*	0.40 (0.14)*
Public	-0.22 (0.082)*	-0.038 (0.090)	0.027 (0.12)	0.25 (0.13)+
Age	-0.76 (0.30)*	-0.15 (0.38)	-1.8 (0.41)*	-0.59 (0.43)

TABLE 5 (continued) Non-UI Covariate Estimates				
	Seasonal		Non-seasonal	
Covariate	**1996**	**1997**	**1996**	**1997**
Time Varying Covariates				
Quarter 3	0.71 (0.042)*	0.78 (0.046)*	0.36 (0.065)*	0.41 (0.071)*
Unemp	0.13 (0.048)*	0.10 (0.051)*	0.076 (0.090)	0.85 (0.096)*
#Obs	5348	4316	4143	3287
Average Value Log Likelihood Function	-3.000	-3.014	-1.914	-1.919
Standard Errors in parentheses. * denotes statistically significantly different from zero, at 5 percent significance level. + denotes statistical significance at the 10 percent level.				

TABLE 6 ER Related Effects				
	Seasonal		Non-seasonal	
Covariate	**1996**	**1997**	**1996**	**1997**
Entrance Requirement	0.25 (0.13)*	0.03 (0.11)	-0.23 (0.22)	0.23 (0.22)
1 — 2 Weeks Before ER	-0.16 (0.13)	-0.29 (0.11)*	0.091 (0.17)	0.05 (0.18)
3 — 5 Weeks Before ER	-0.13 (0.14)	-0.25 (0.09)*	-0.043 (0.19)	-0.065 (0.20)
6 — 10 Weeks Before ER	-0.09 (0.15)	-0.33 (0.09)*	0.21 (0.22)	0.092 (0.23)
11 or More Weeks Before ER	-0.11 (0.18)	-0.44 (0.10)*	0.44 (0.27)	0.31 (0.28)
Max Year	0.20 (0.12)+	0.19 (0.12)+	—	—
Between ER and Max Year	0.03 (0.08)	0.06 (0.07)	—	—
Standard Errors in parentheses. * denotes statistically significantly different from zero, at 5 percent significance level. + denotes statistical significance at the 10 percent level.				

TABLE 7
Effects of Changes in the Entrance Requirement on Average Employment Spell Durations

ER = 12-Week Region		
	Seasonal Workers	
	1996	**1997**
Entrance Requirement = 12 Weeks	21.33 weeks	19.83 weeks
Entrance Requirement = 10 Weeks	20.98 weeks	19.10 weeks
	Non-seasonal Workers	
	1996	**1997**
Entrance Requirement = 12 Weeks	83.17 weeks	84.23 weeks
Entrance Requirement = 10 Weeks	83.52 weeks	84.23 weeks

TABLE 8
ER-Related Effects
1997 Seasonal Workers, Using Weeks-based UI Effects

Covariate	**Estimates**
Entrance Requirement	0.55 (0.14)*
1 — 2 Weeks Before ER	0.32 (0.14)*
3 — 5 Weeks Before ER	0.25 (0.14)+
6 — 10 Weeks Before ER	0.48 (0.16)
11 or More Weeks Before ER	0.80 (0.18)
Max Year	0.05 (0.14)
Between ER and Max Year	0.27 (0.09)*

Standard Errors in parentheses. * denotes statistically significantly different from zero at 5 percent significance level. + denotes statistical significance at the 10 percent level.

Appendix B: Figures

FIGURE 1
Cumulative Distribution Functions for Change in Weeks Needed to Qualify, 1997 Minus 1996 Rules

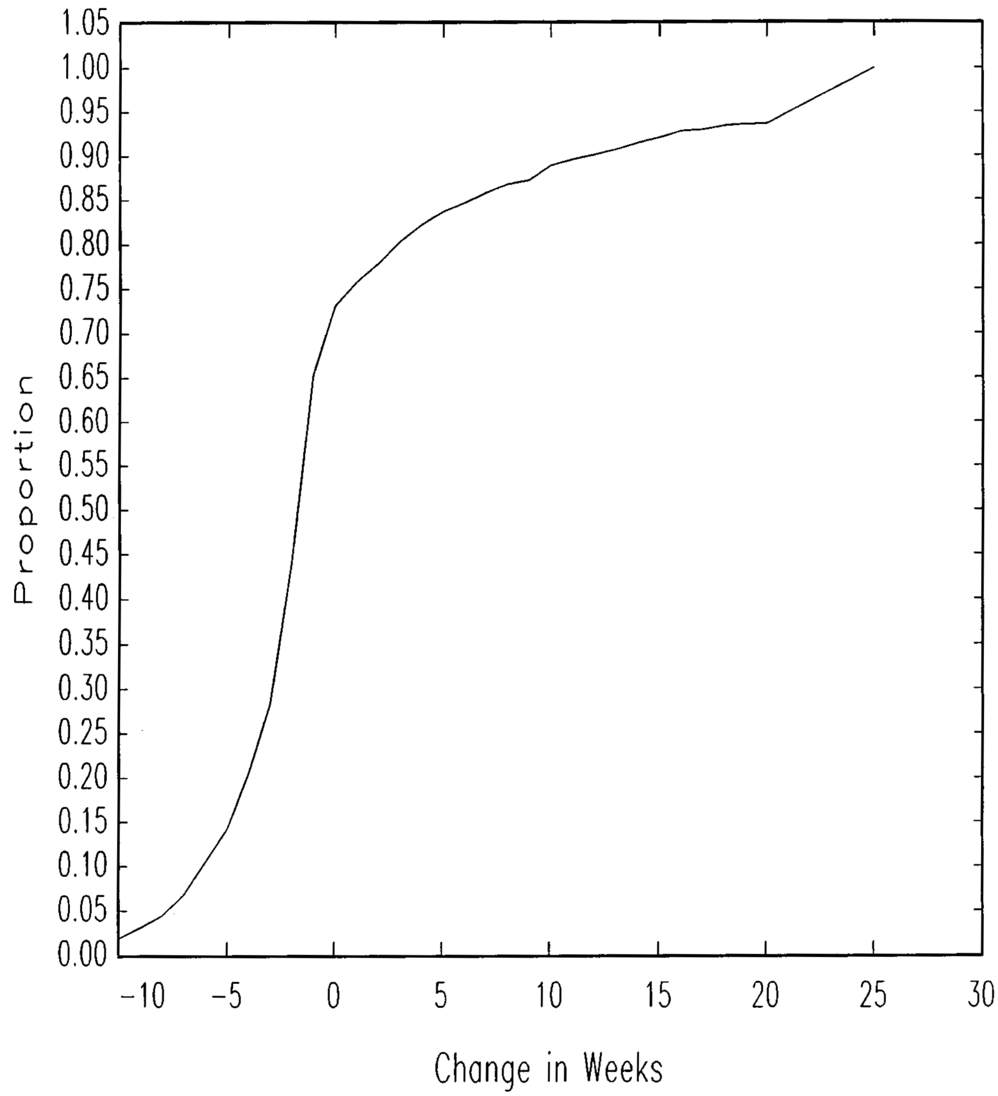

FIGURE 2
Cumulative Distribution Functions for Entitlement Under Three Sets of Rules, 1997 Non-seasonal Workers

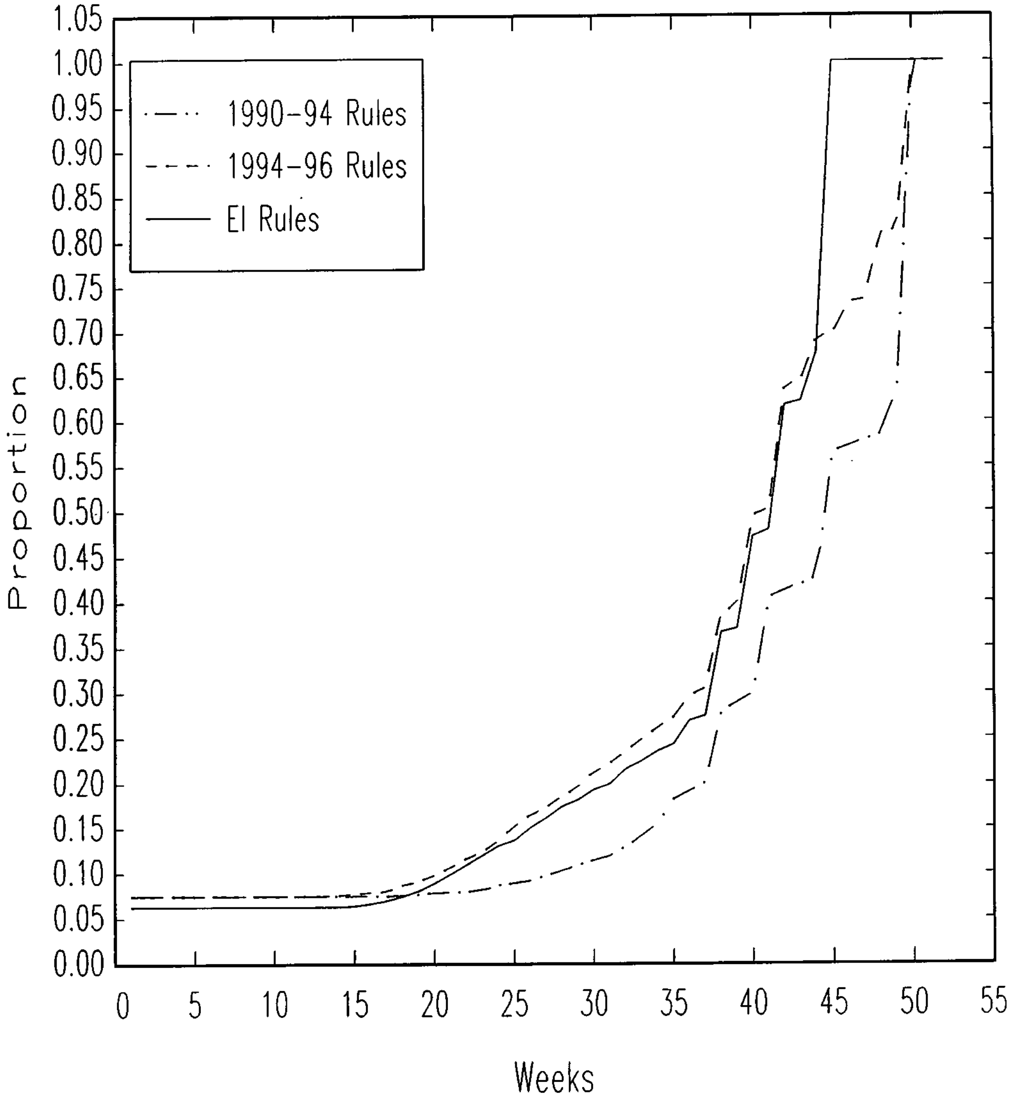

FIGURE 3
Cumulative Distribution Functions for Entitlement Under Three Sets of Rules, 1997 Seasonal Workers

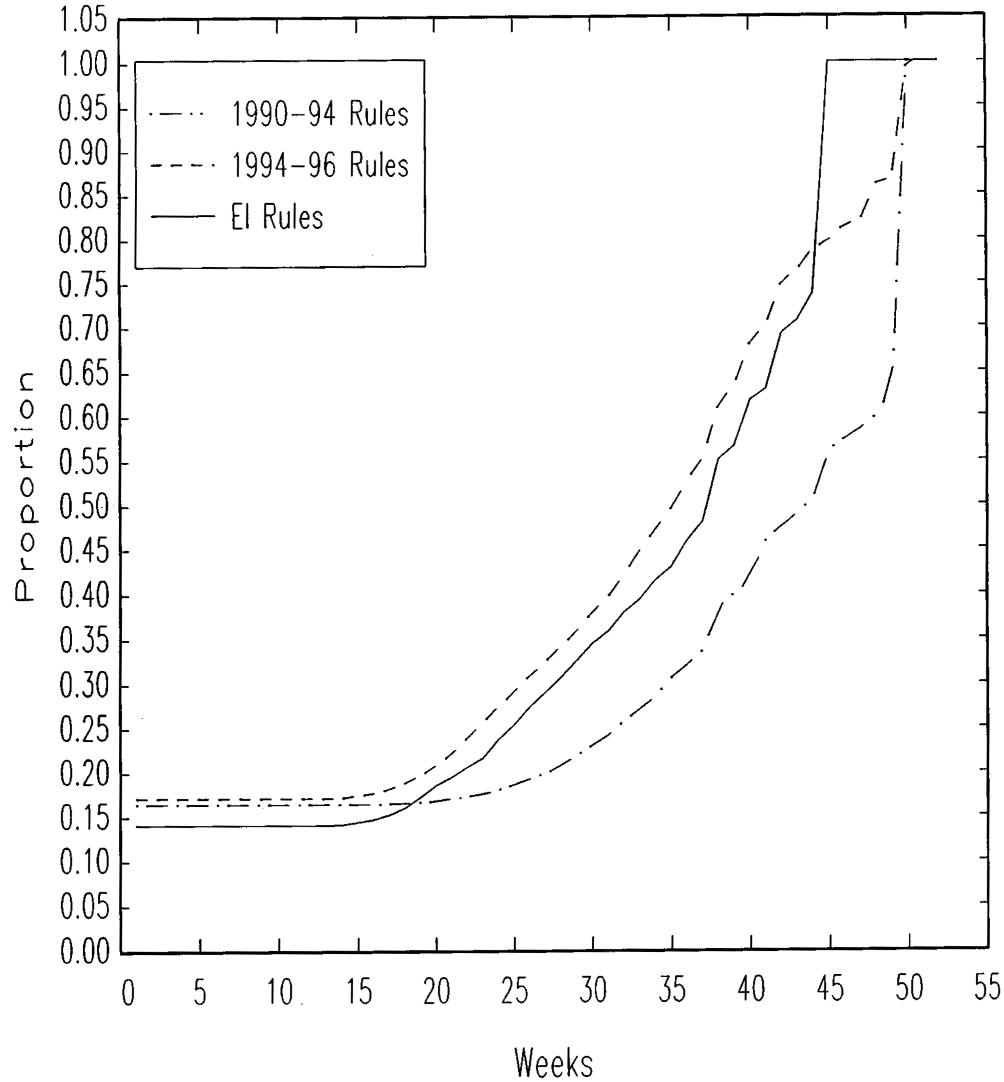

FIGURE 4
Cumulative Distribution Functions for Entitlement Under Three Sets of Rules, 1997 Seasonal Workers Who Worked 10-20 Weeks

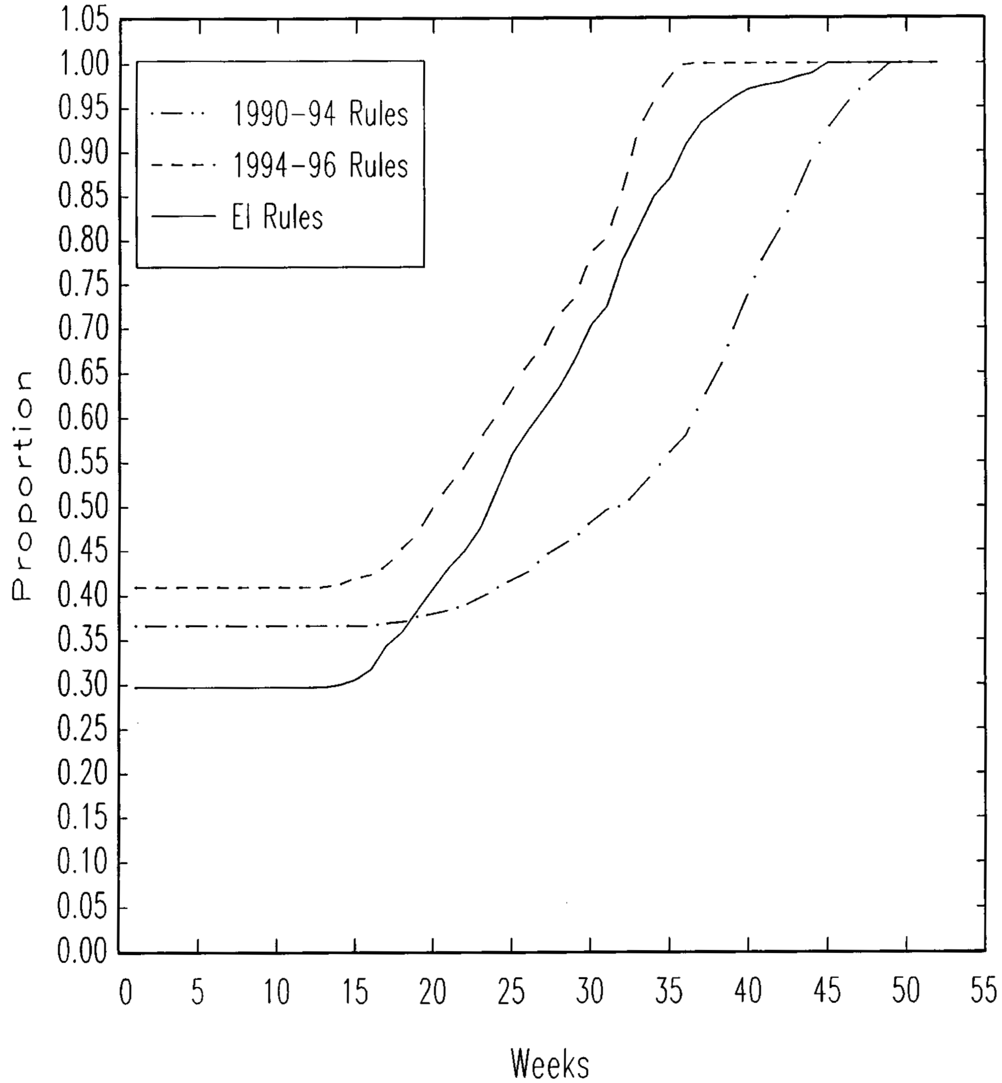

FIGURE 5
Cumulative Distribution Functions for Change in Weeks Entitlement, 1997 Minus 1996 Rules

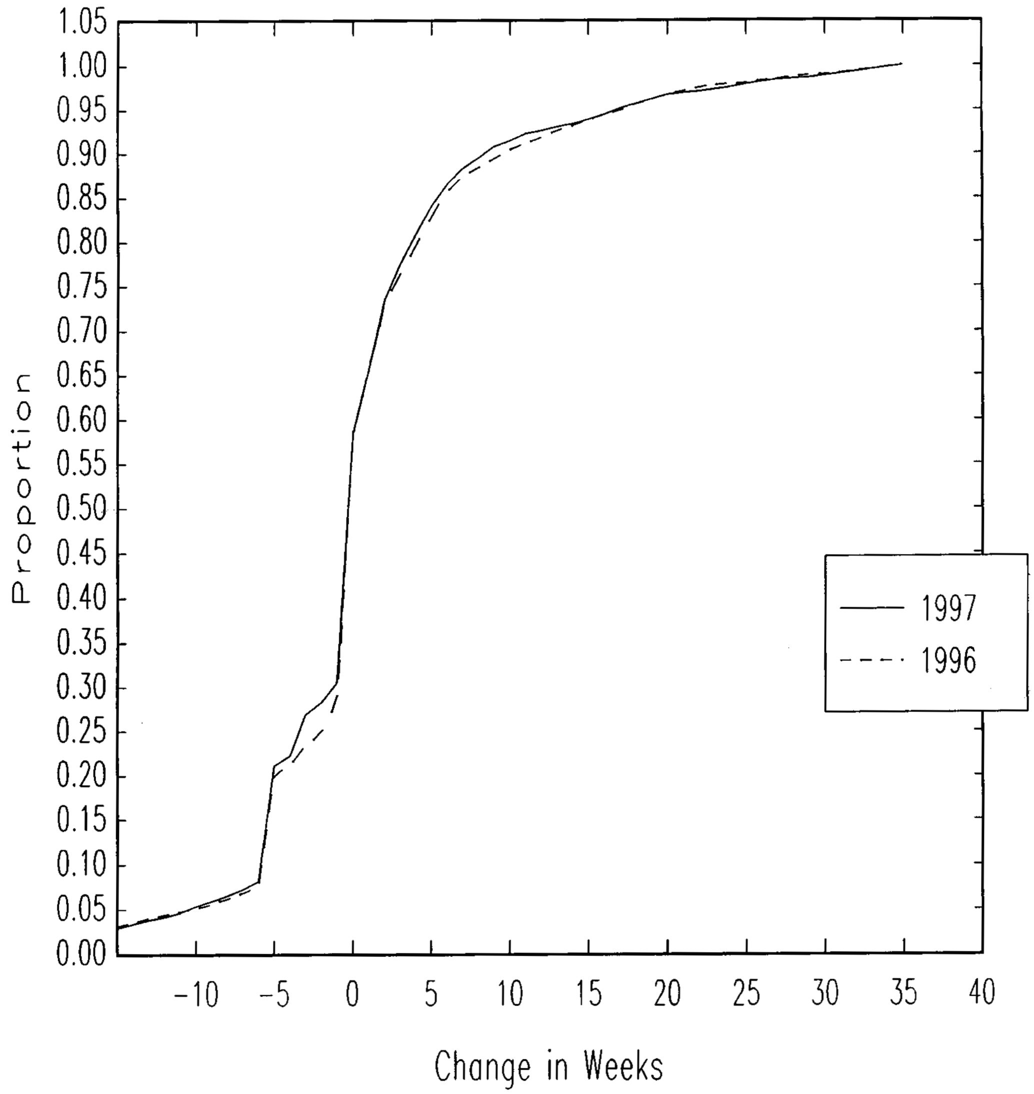

FIGURE 6
Empirical Hazard: Seasonal Workers, Unemployment Rate Under 6%, Quarters 1-3

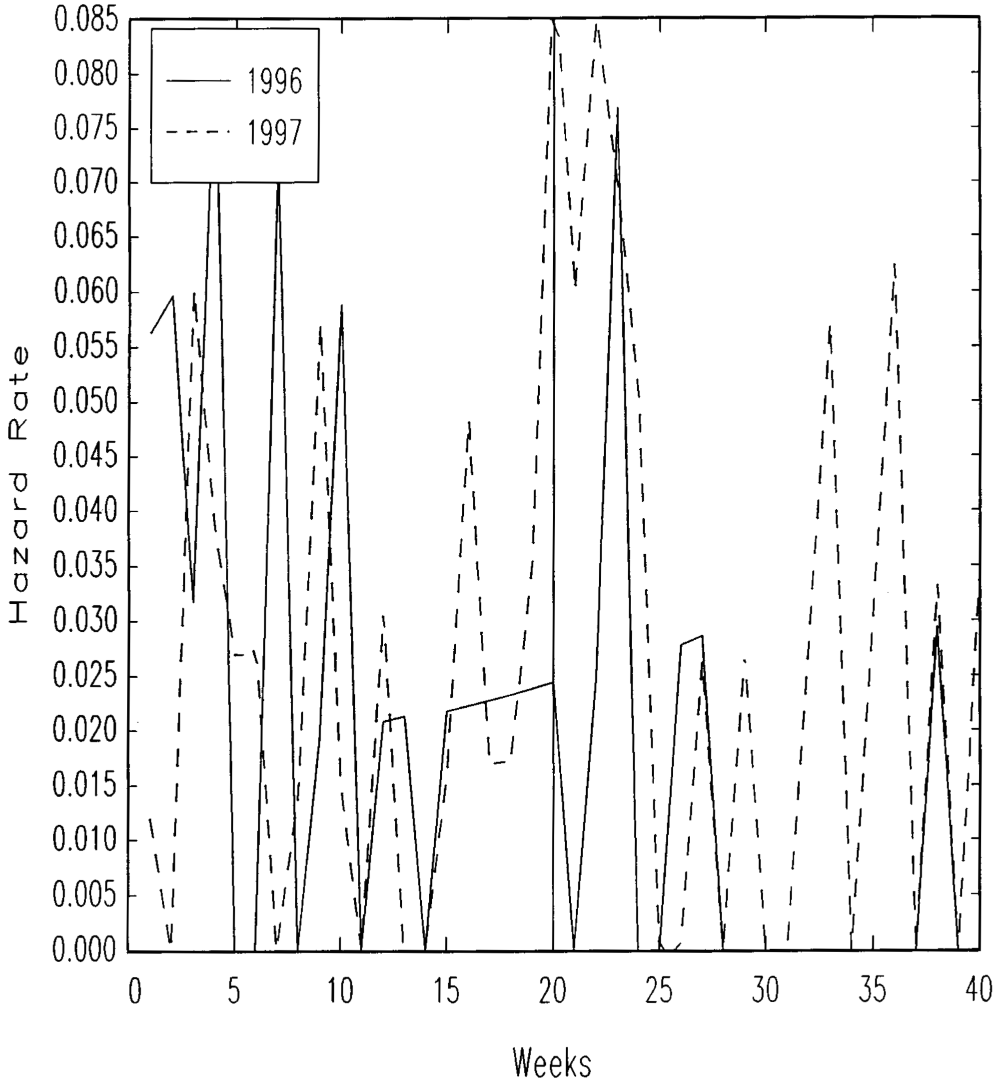

FIGURE 7
Empirical Hazard: Seasonal Workers, Unemployment Rate 6% — 7%, Quarters 1-3

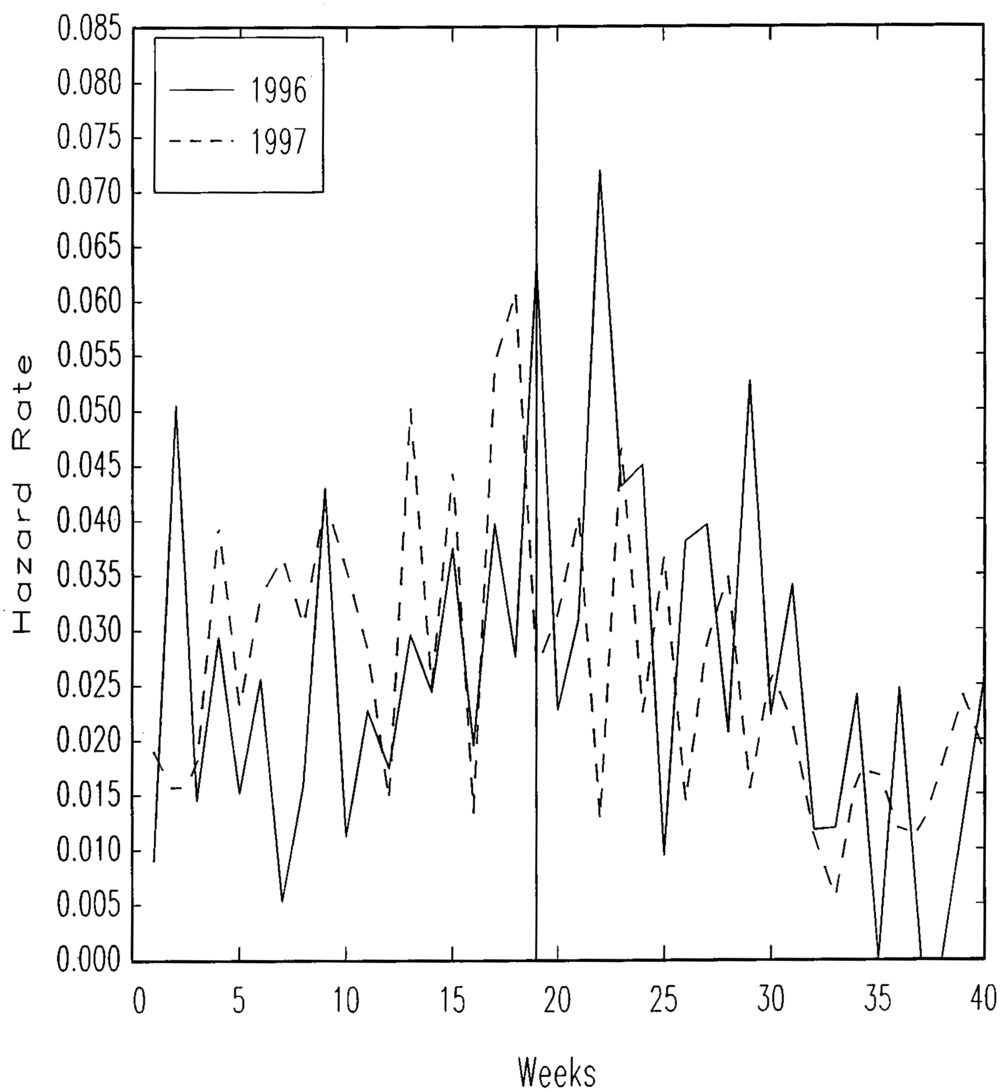

FIGURE 8
Empirical Hazard: Seasonal Workers,
Unemployment Rate 7% — 8%, Quarters 1-3

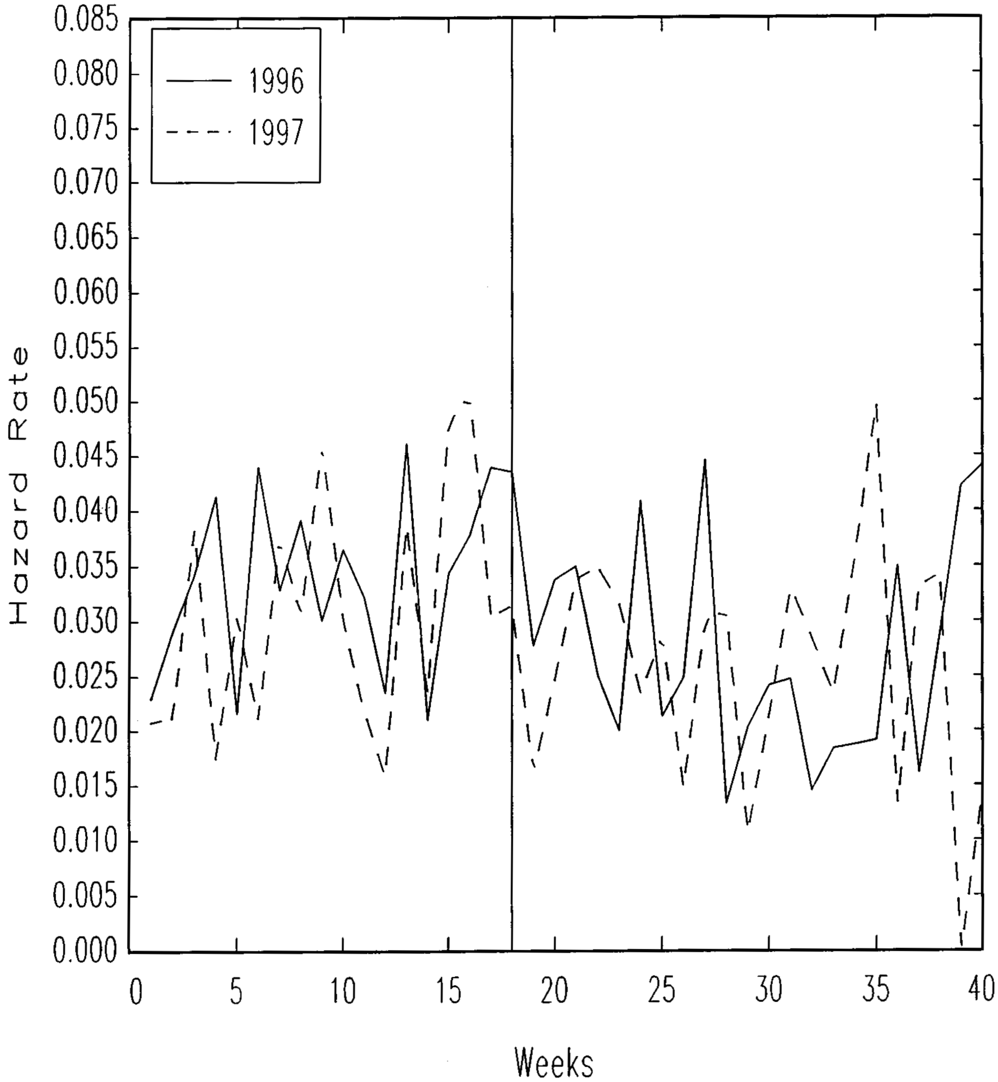

FIGURE 9
Empirical Hazard: Seasonal Workers,
Unemployment Rate 8% — 9%, Quarters 1-3

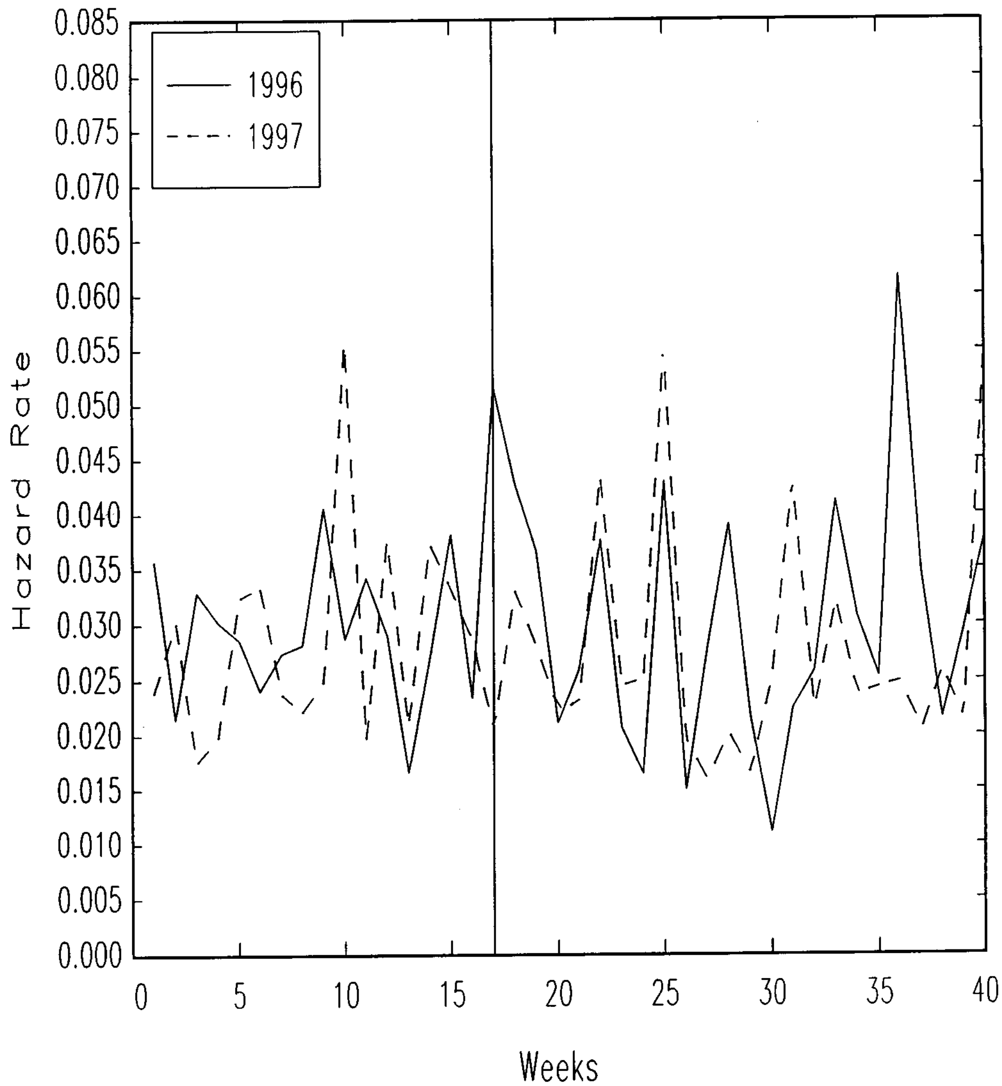

FIGURE 10
Empirical Hazard: Seasonal Workers, Unemployment Rate 9% — 10%, Quarters 1-3

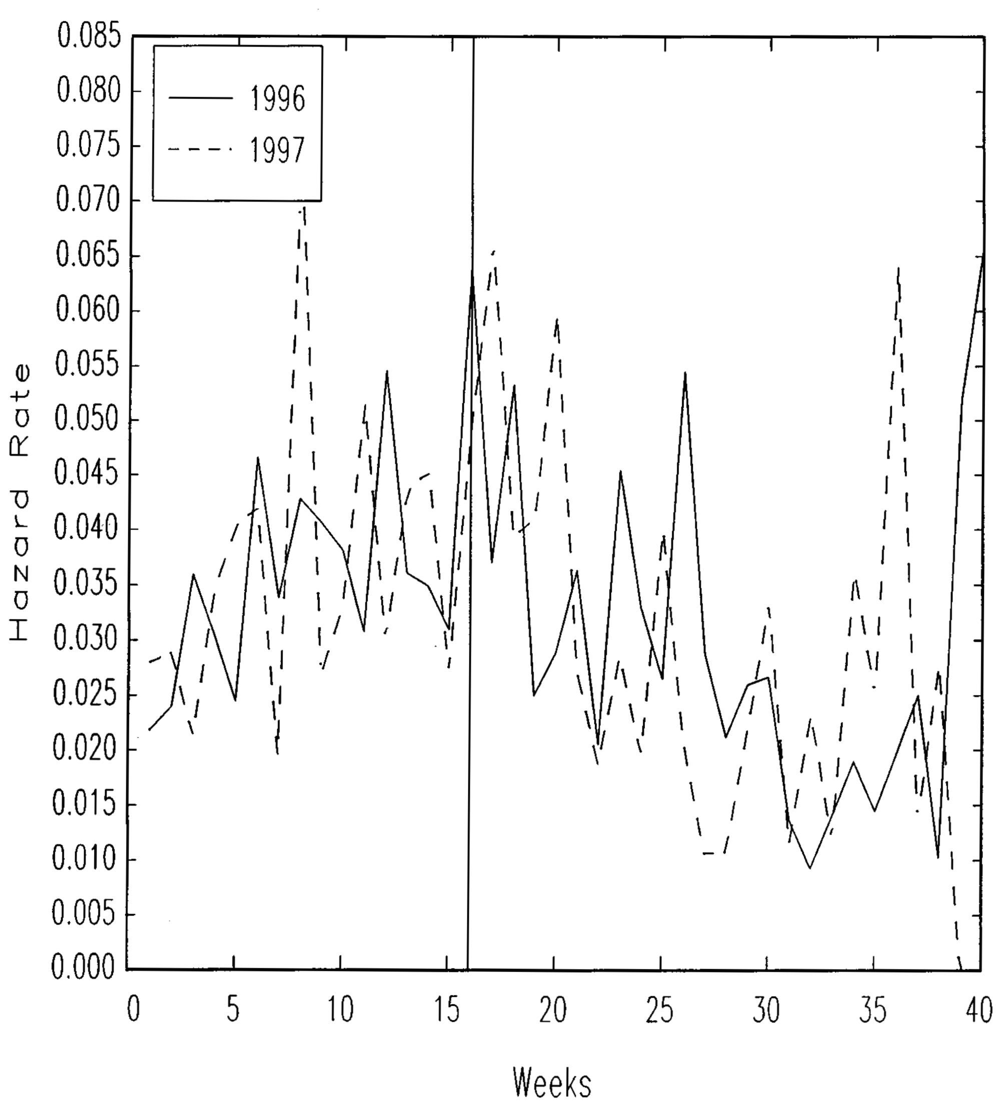

FIGURE 11
Empirical Hazard: Seasonal Workers,
Unemployment Rate 10% — 11%, Quarters 1-3

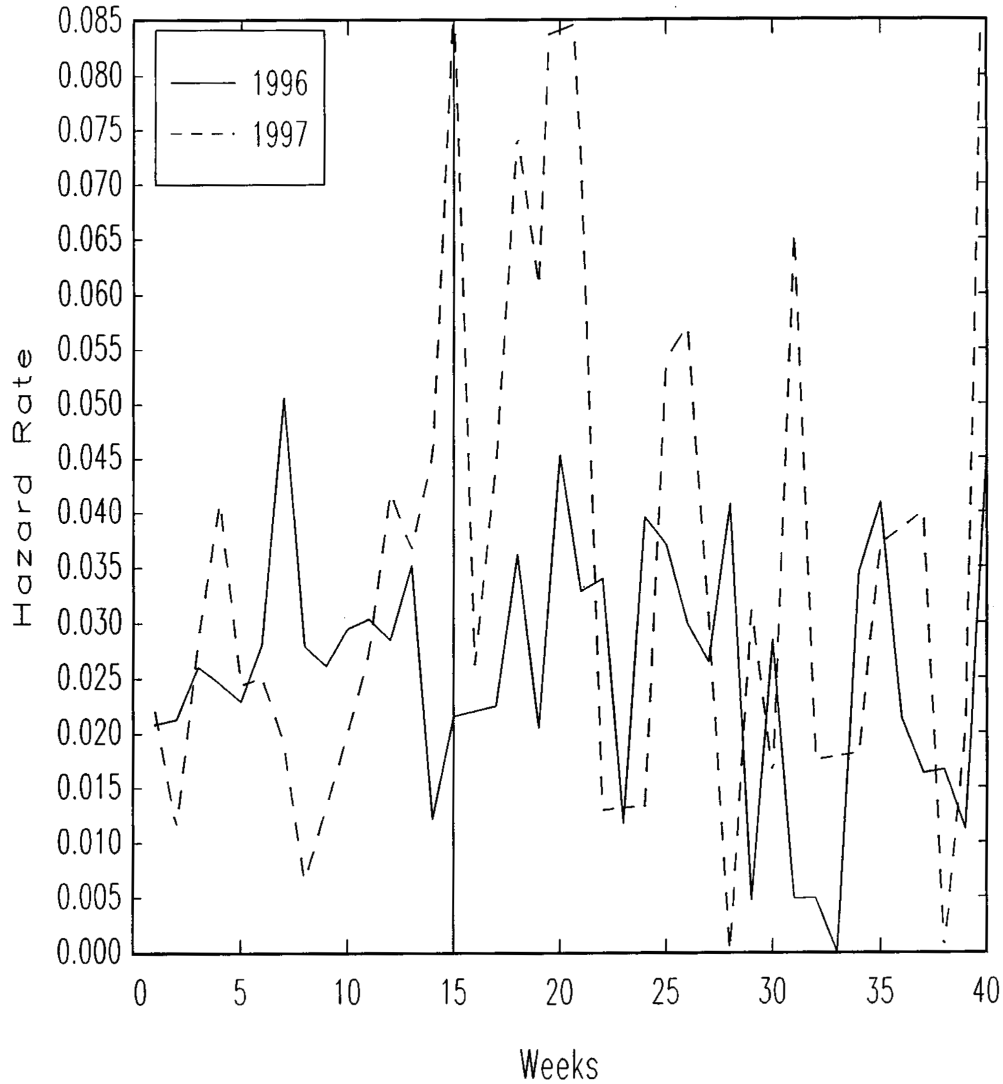

FIGURE 12
Empirical Hazard: Seasonal Workers, Unemployment Rate 11% — 12%, Quarters 1-3

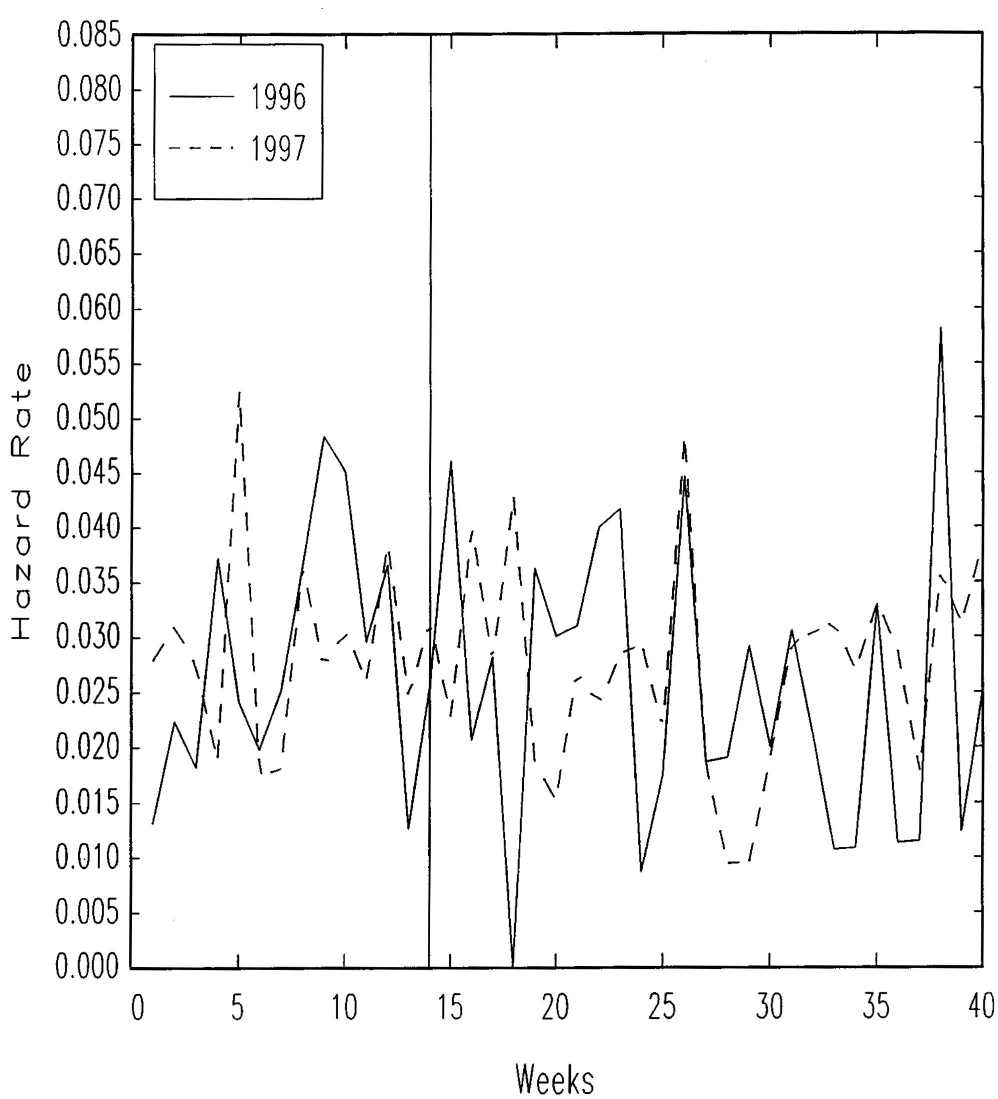

FIGURE 13
Empirical Hazard: Seasonal Workers, Unemployment Rate 12% — 13%, Quarters 1-3

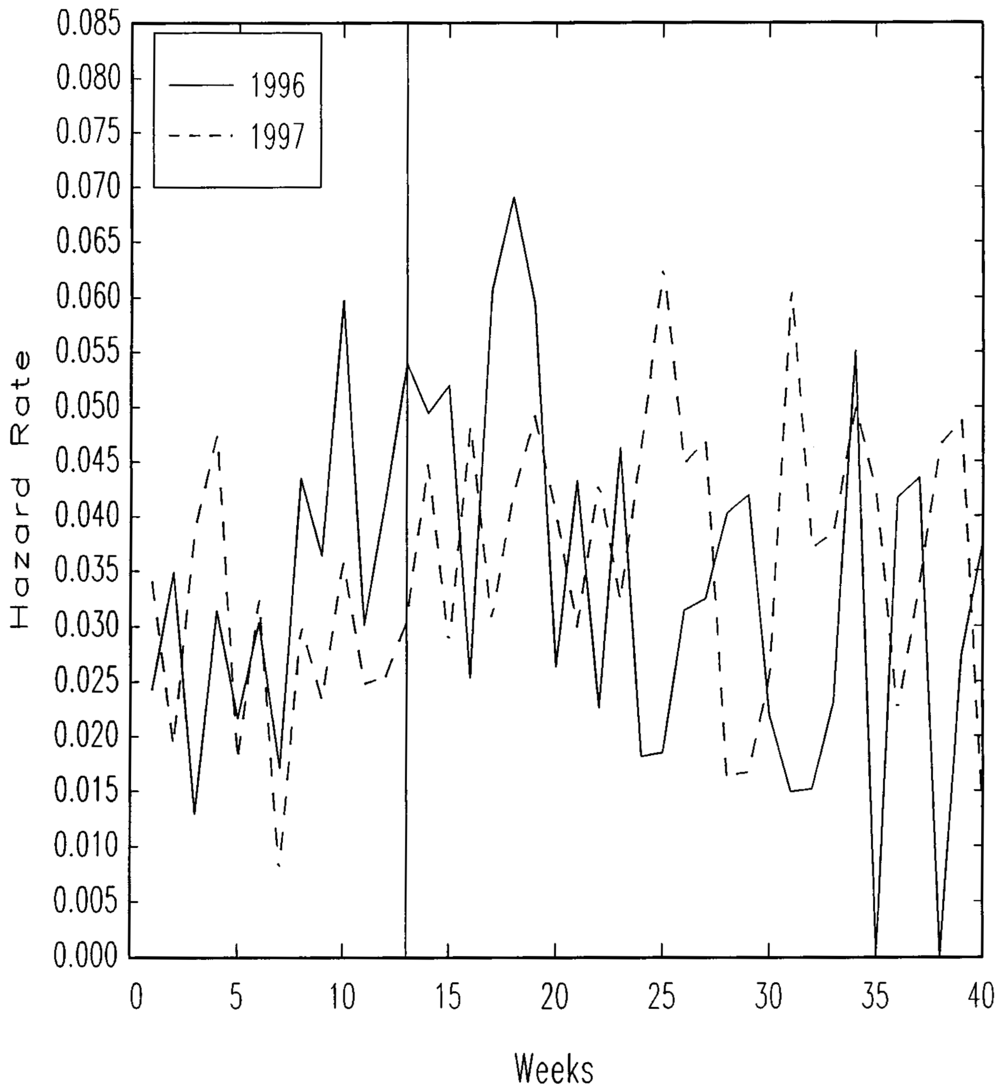

FIGURE 14
Empirical Hazard: Seasonal Workers, Unemployment Rate Over 13%, Quarters 1-3

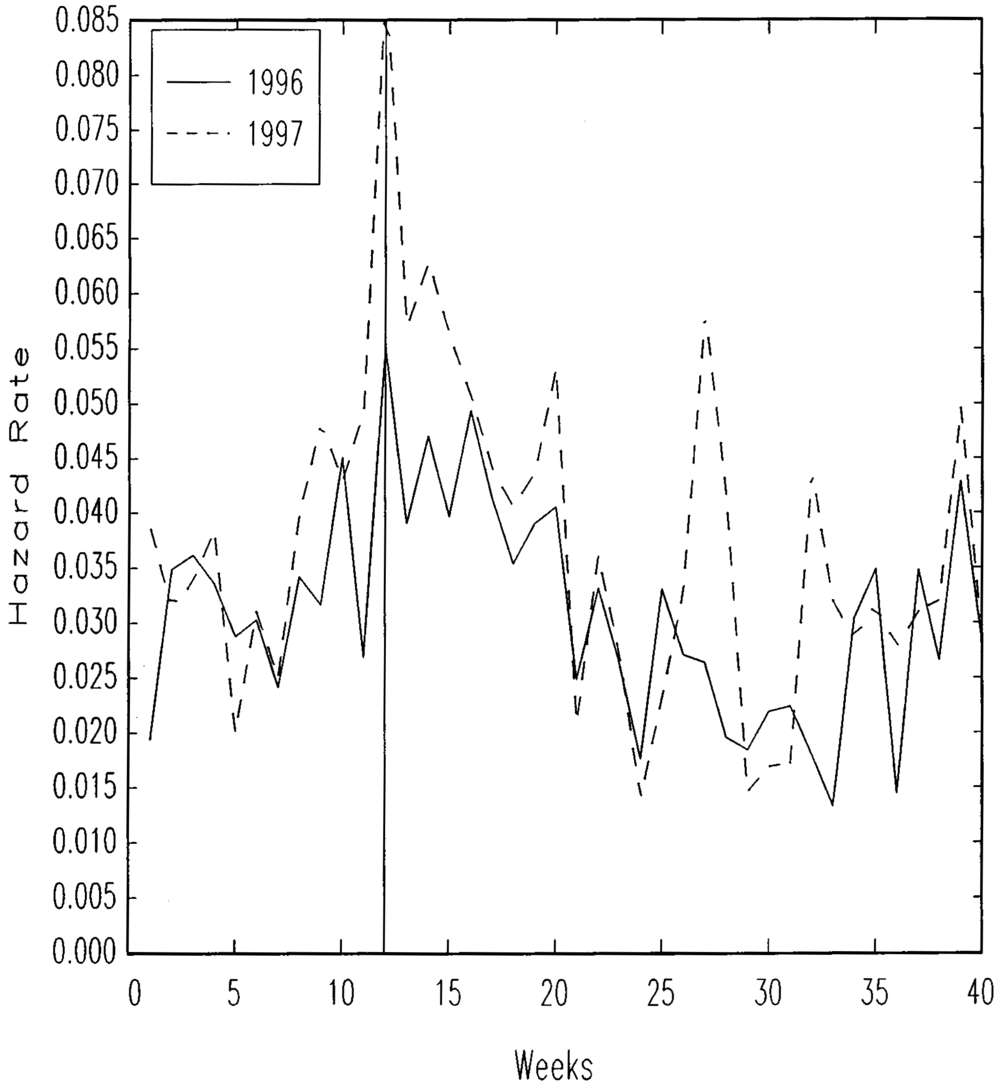

FIGURE 15
Empirical Hazard: Non-seasonal Workers, Unemployment Rate Under 6%, Quarters 1-3

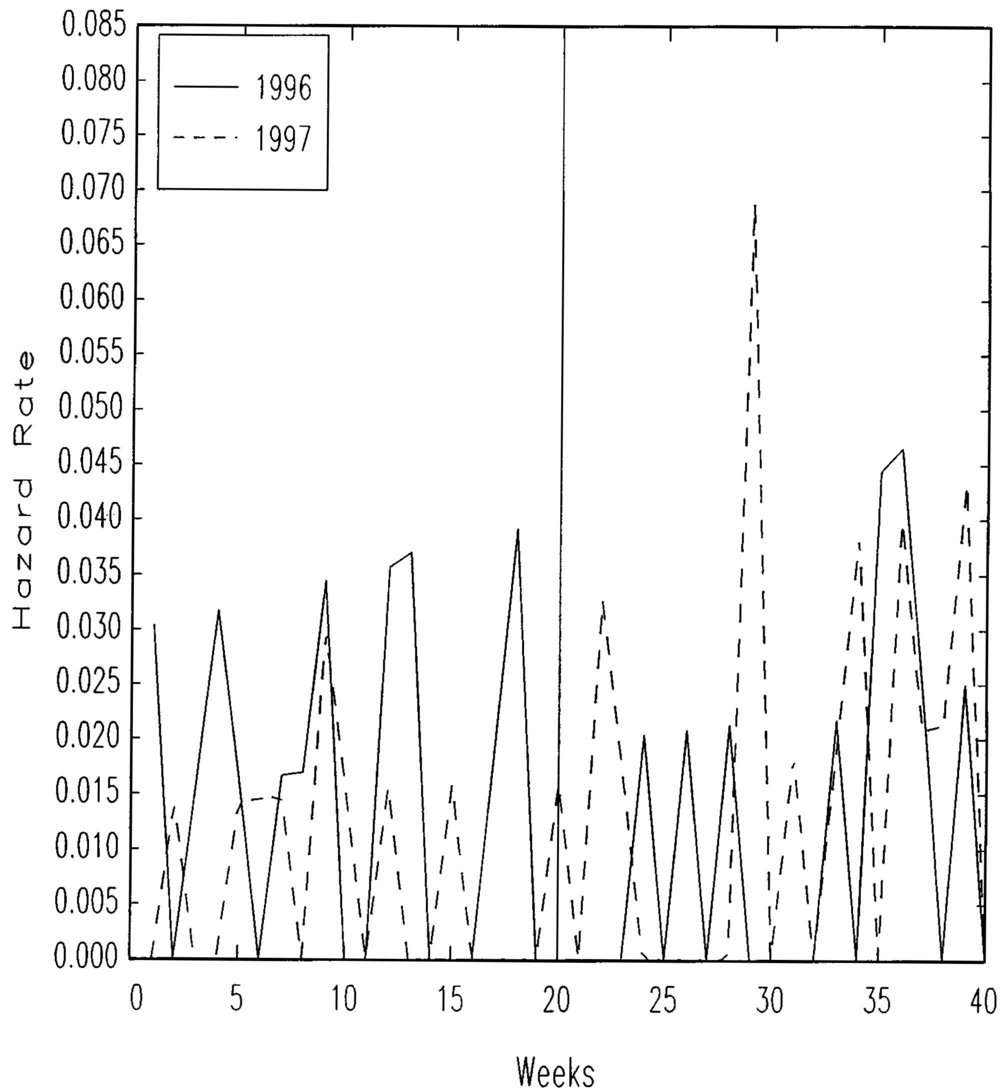

FIGURE 16
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 6% — 7%, Quarters 1-3

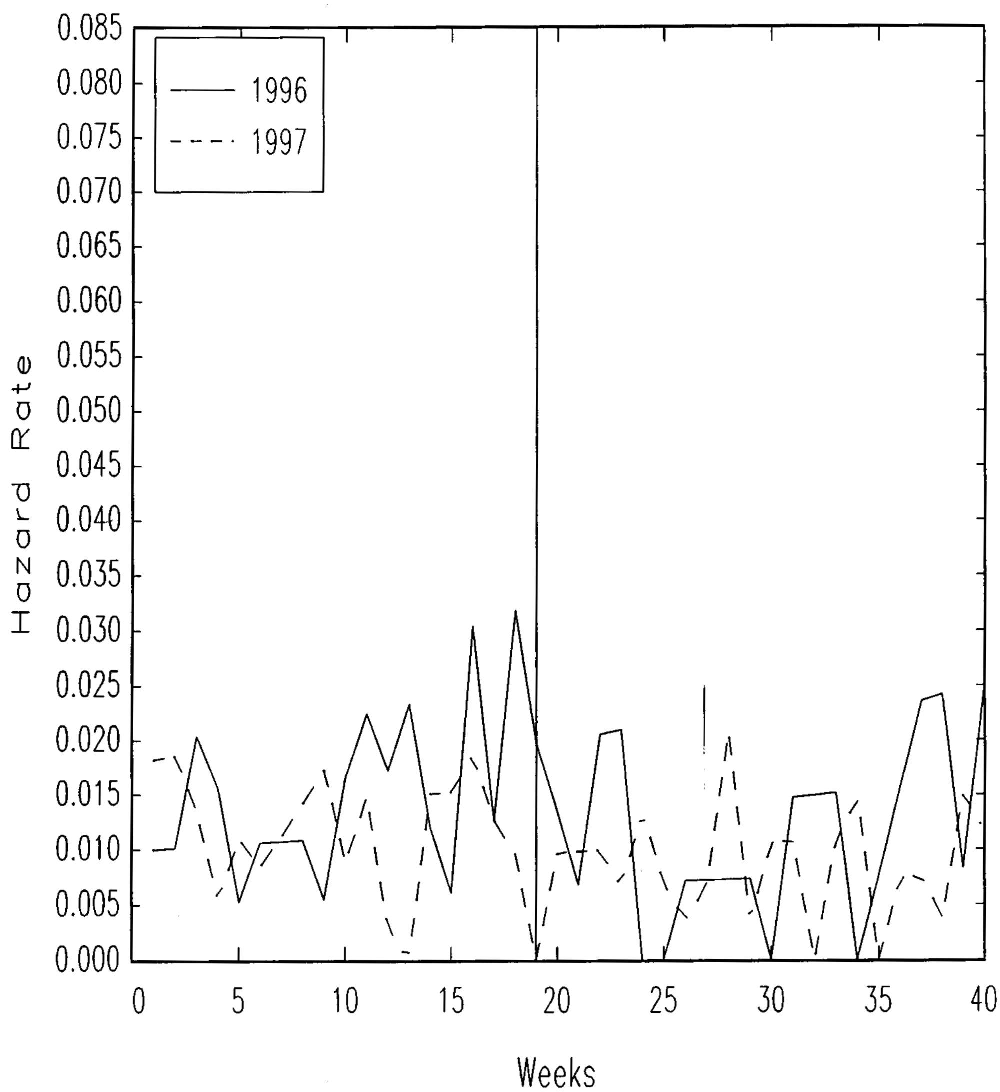

FIGURE 17
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 7% — 8%, Quarters 1-3

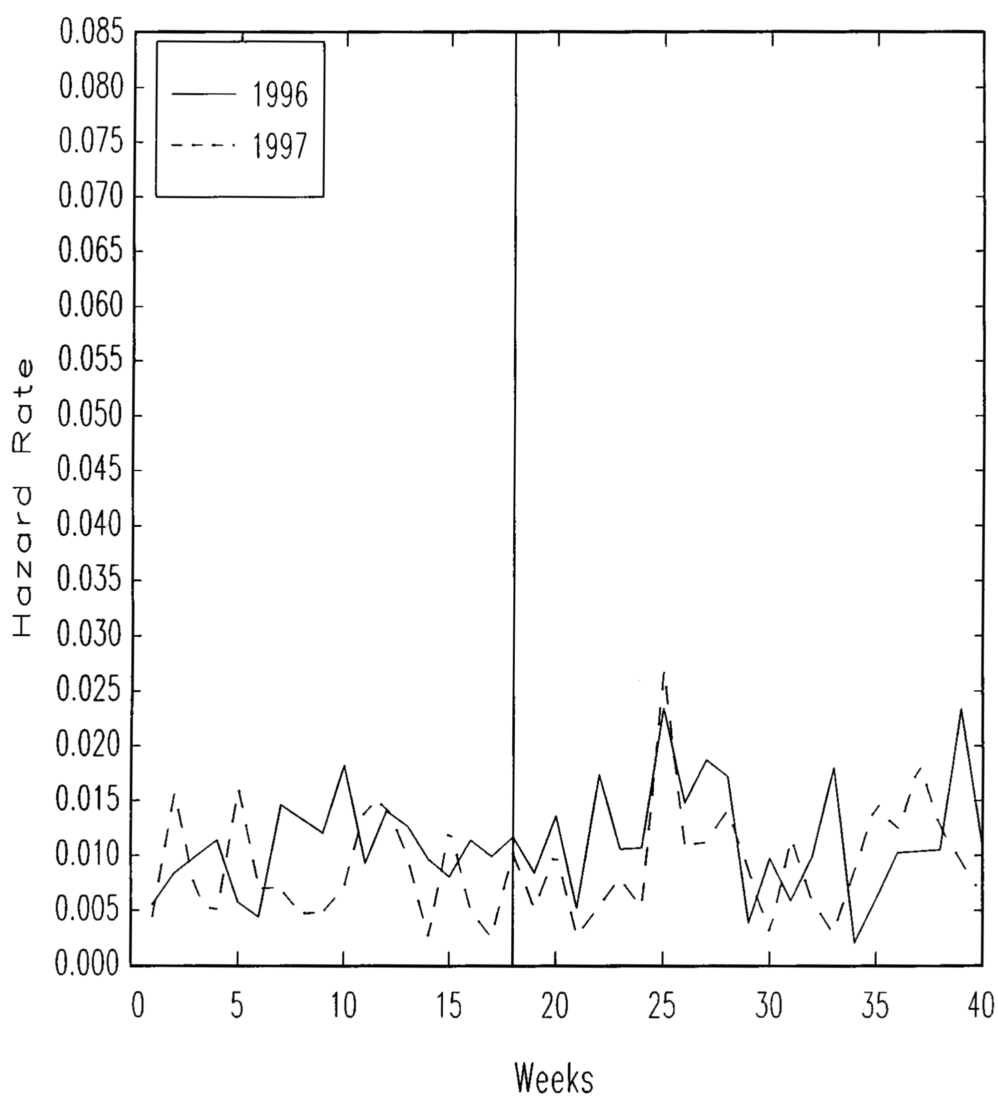

FIGURE 18
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 8% — 9%, Quarters 1-3

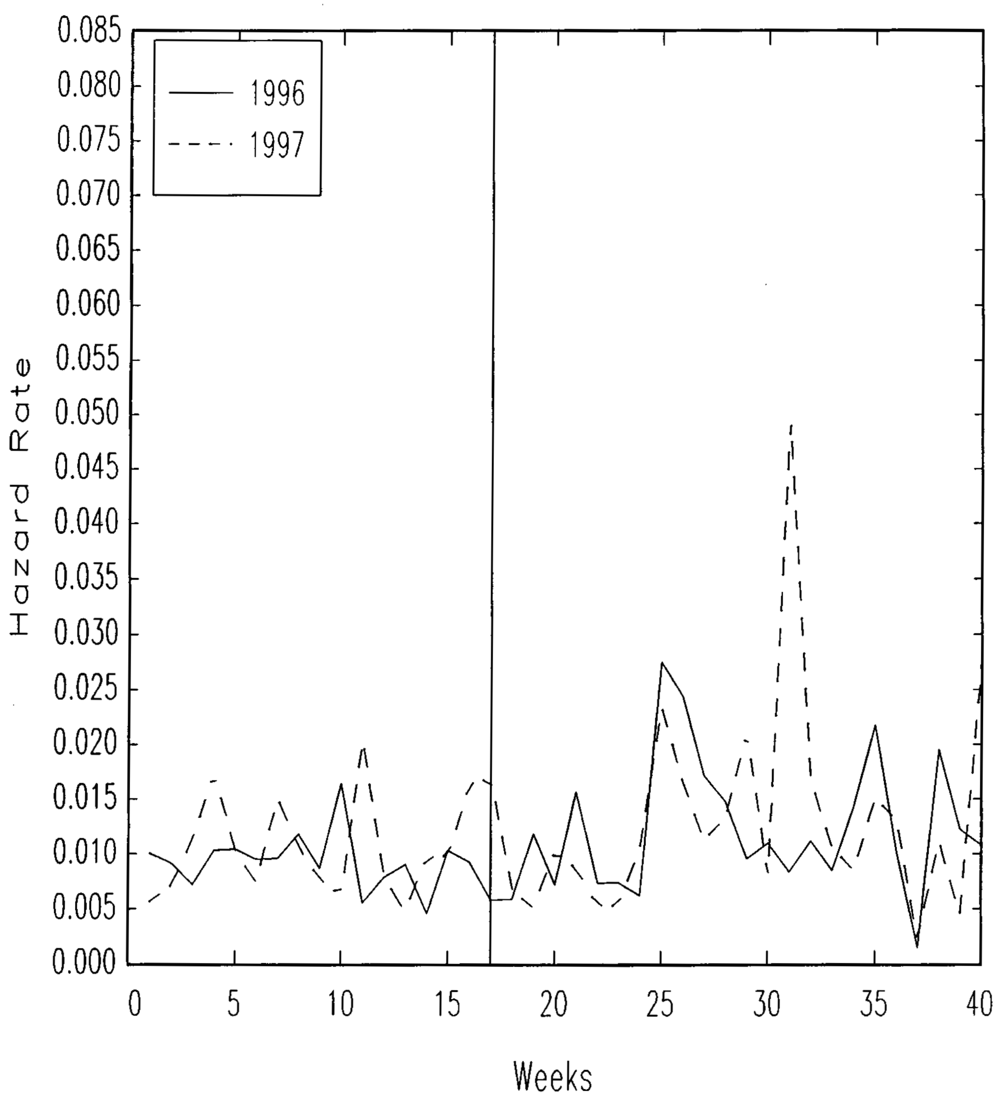

FIGURE 19
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 9% — 10%, Quarters 1-3

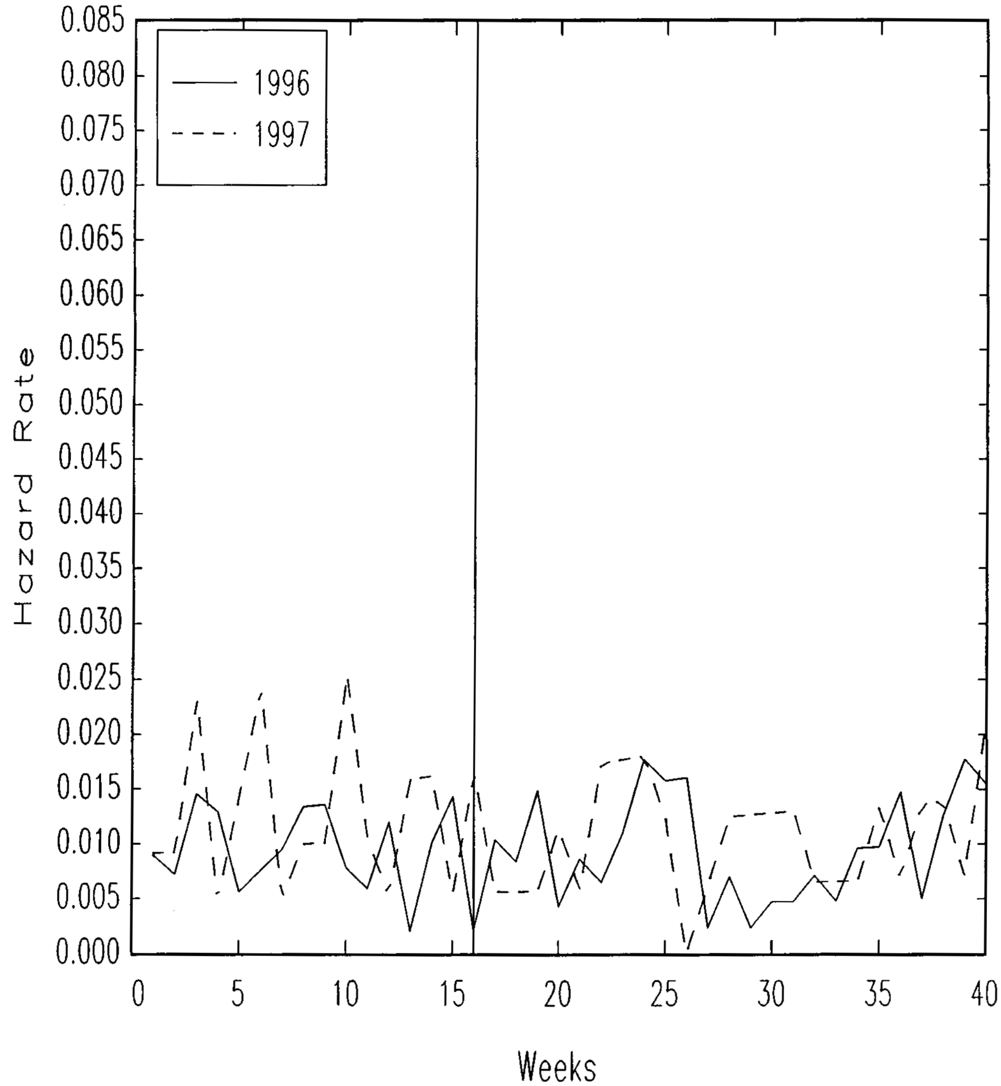

FIGURE 20
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 10% — 11%, Quarters 1-3

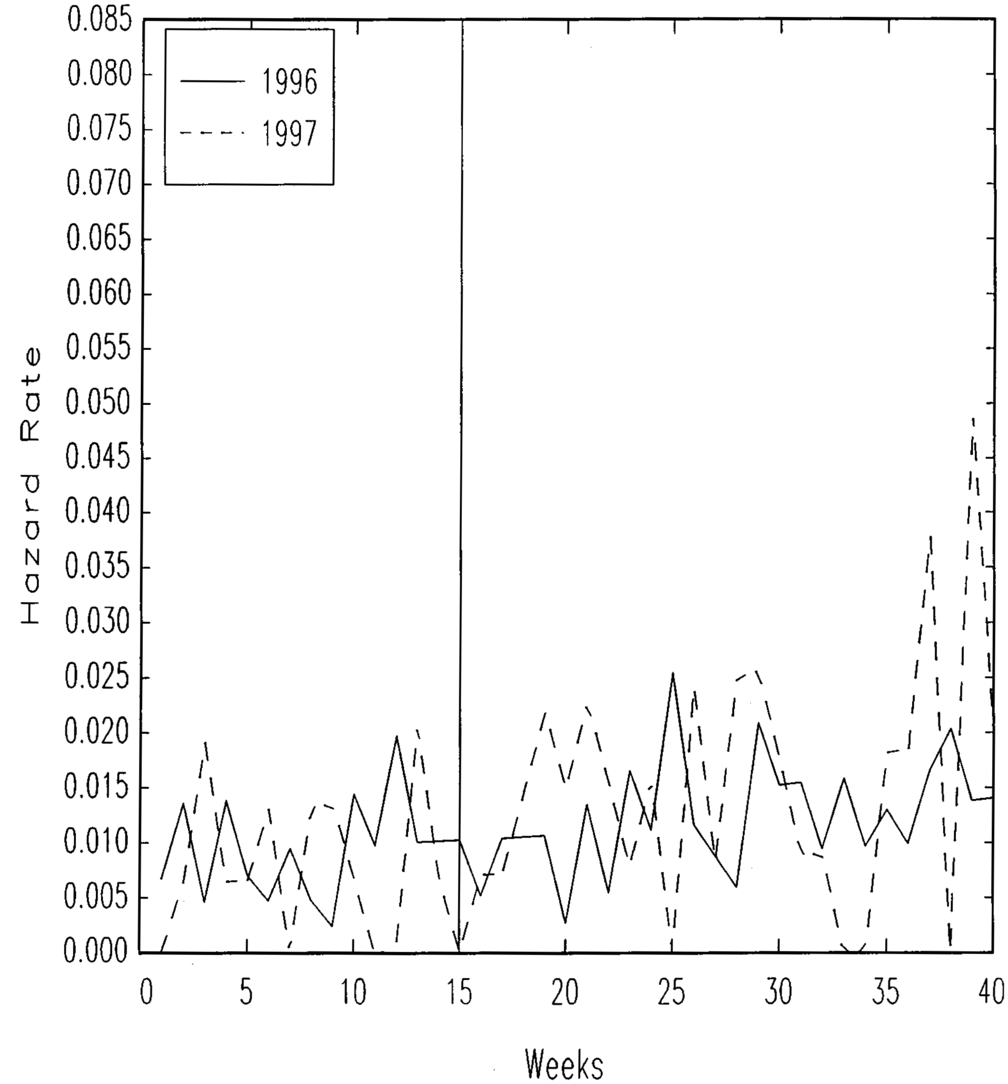

FIGURE 21
Empirical Hazard: Non-seasonal Workers,
Unemployment Rate 11% — 12%, Quarters 1-3

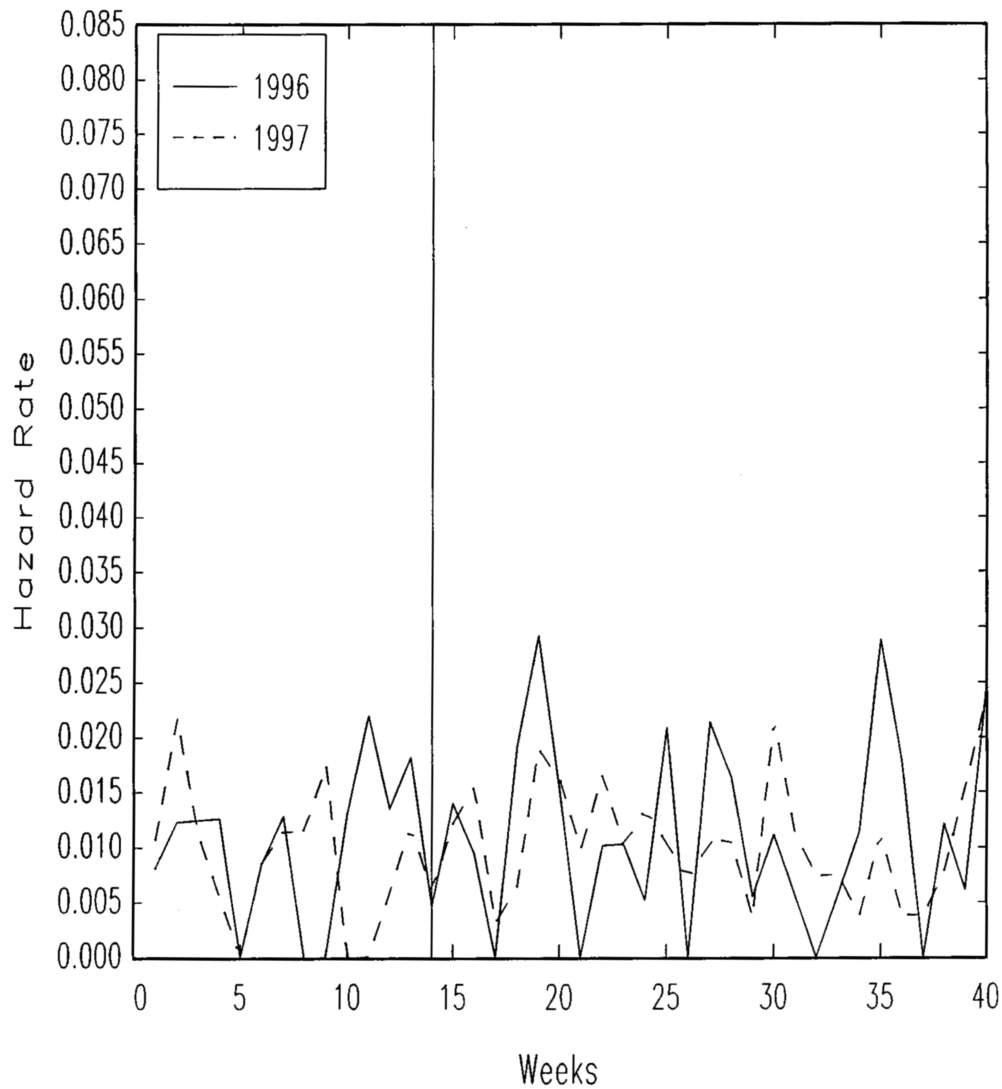

FIGURE 22
Empirical Hazard: Non-seasonal Workers, Unemployment Rate 12% — 13%, Quarters 1-3

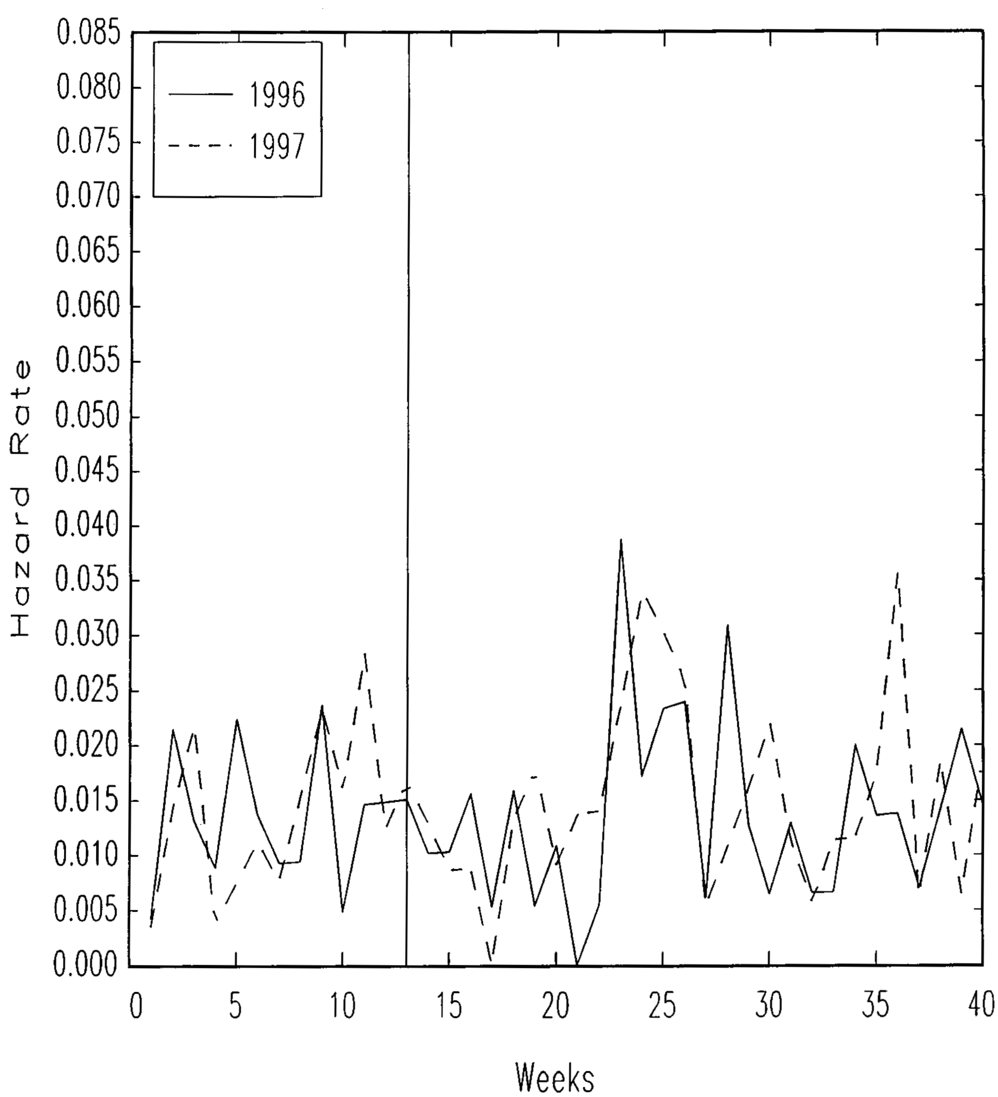

FIGURE 23
Empirical Hazard: Non-seasonal Workers, Unemployment Rate Over 13%, Quarters 1-3

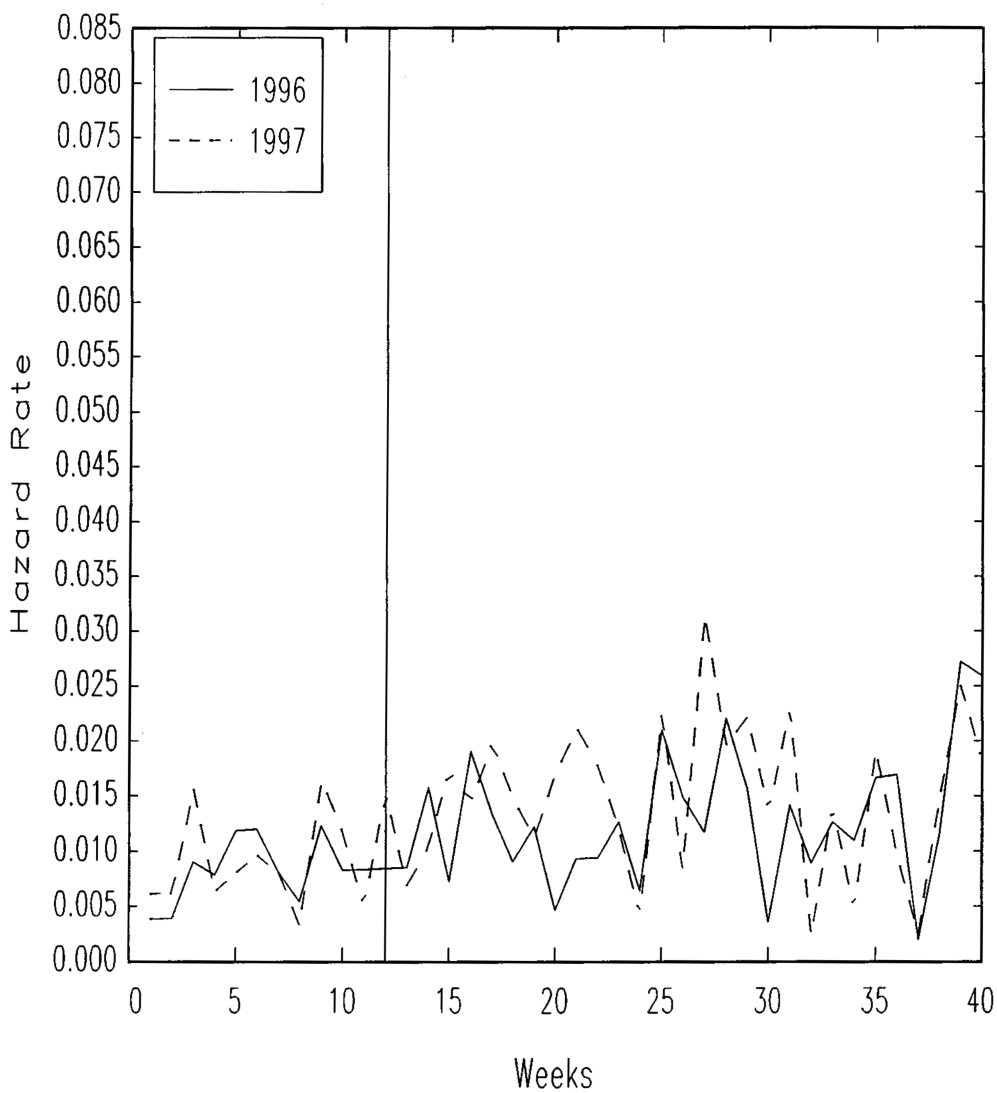

FIGURE 24A
1996 Fitted Hazard Rate Pattern:
Seasonal Jobs, Flat Baseline, Unemployment Rate Over 16% Region

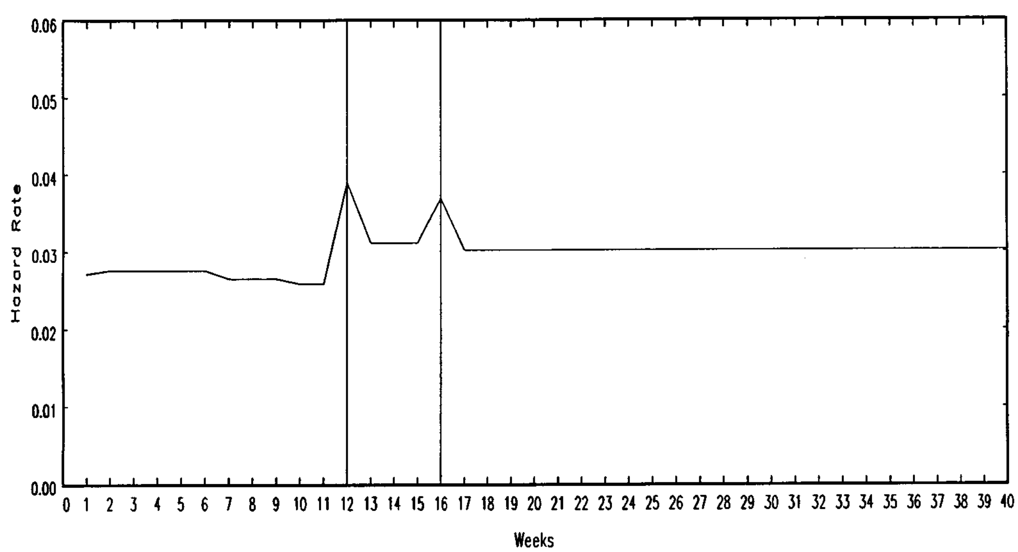

FIGURE 24B
1997 Fitted Hazard Rate Pattern:
Seasonal Jobs, Flat Baseline, Unemployment Rate Over 16% Region

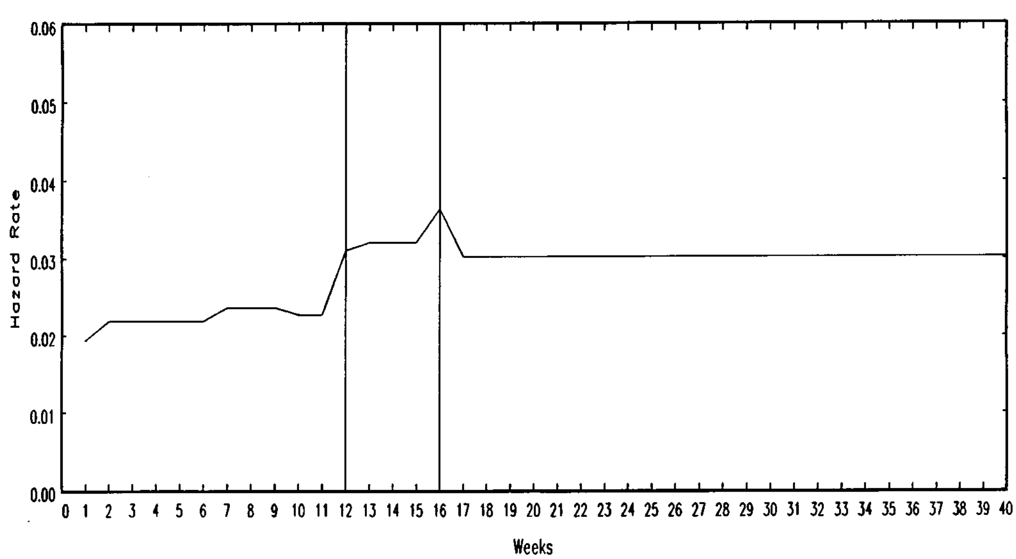

FIGURE 25A
1996 Fitted Hazard Rate Pattern:
Non-seasonal Jobs, Flat Baseline, Entrance Requirement = 12 Weeks Region

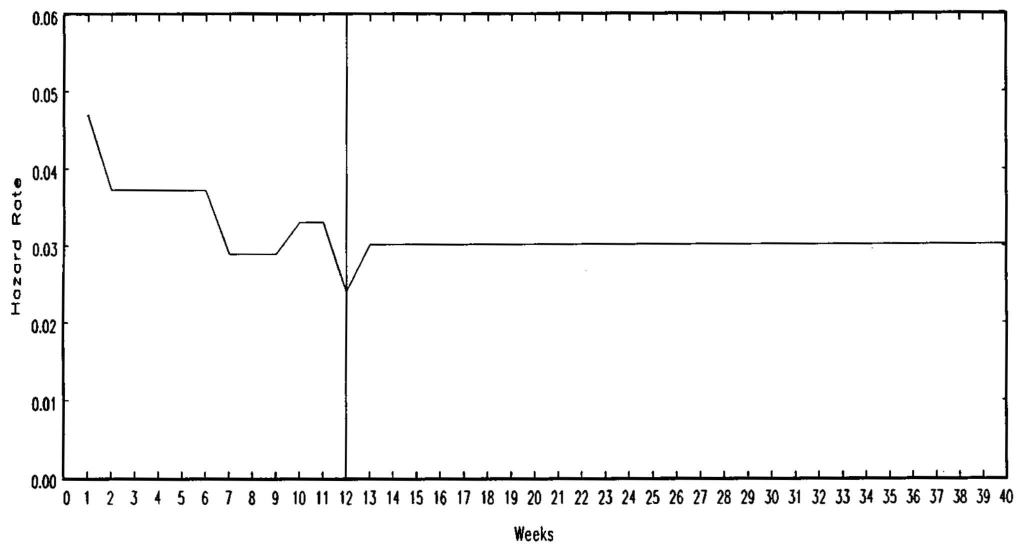

FIGURE 25B
1997 Fitted Hazard Rate Pattern:
Non-seasonal Jobs, Flat Baseline, Entrance Requirement = 12 Weeks Region

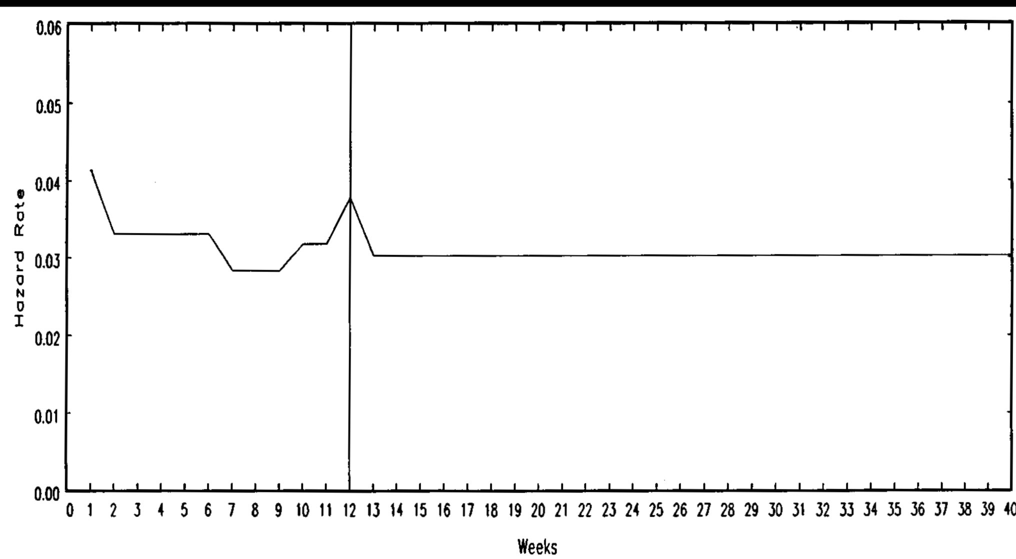

FIGURE 26A
1996 Fitted Hazard Rate Pattern:
Seasonal Jobs, Actual Baseline, Unemployment Rate Over 16% Region

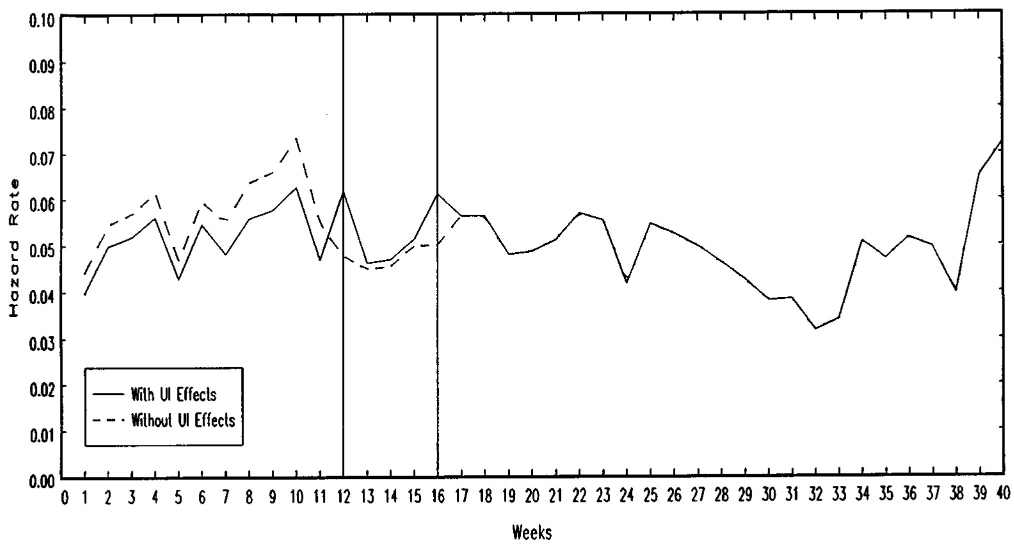

FIGURE 26B
1997 Fitted Hazard Rate Pattern:
Seasonal Jobs, Real Baseline, Unemployment Rate Over 16% Region

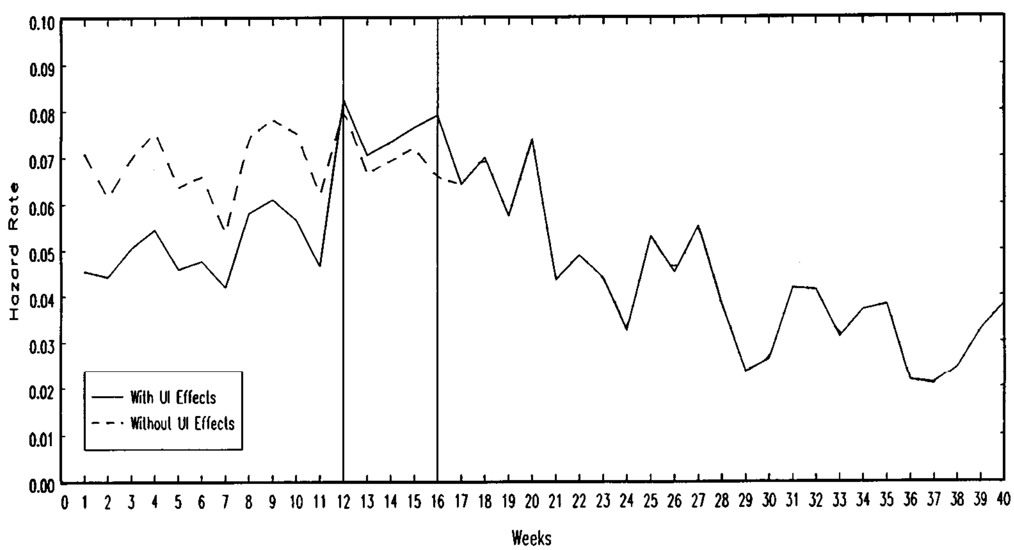

FIGURE 27A
1996 Fitted Hazard Rate Pattern:
Non-seasonal Jobs, Actual Baseline, Entrance Requirement = 12 Weeks Region

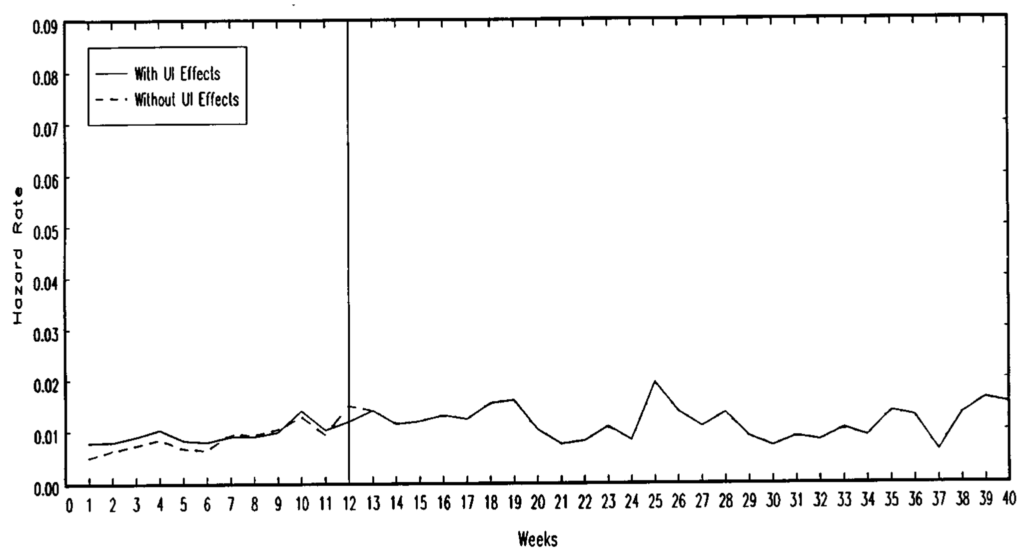

FIGURE 27B
1997 Fitted Hazard Rate Pattern:
Non-seasonal Jobs, Real Baseline, Entrance Requirement = 12 Weeks Region

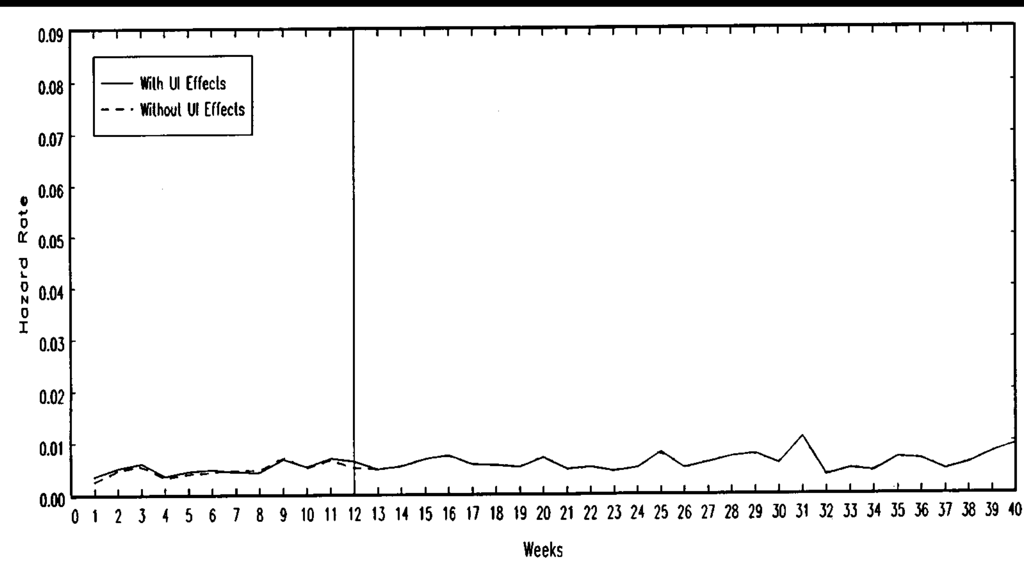

FIGURE 28
Empirical Hazard: All Seasonal Workers

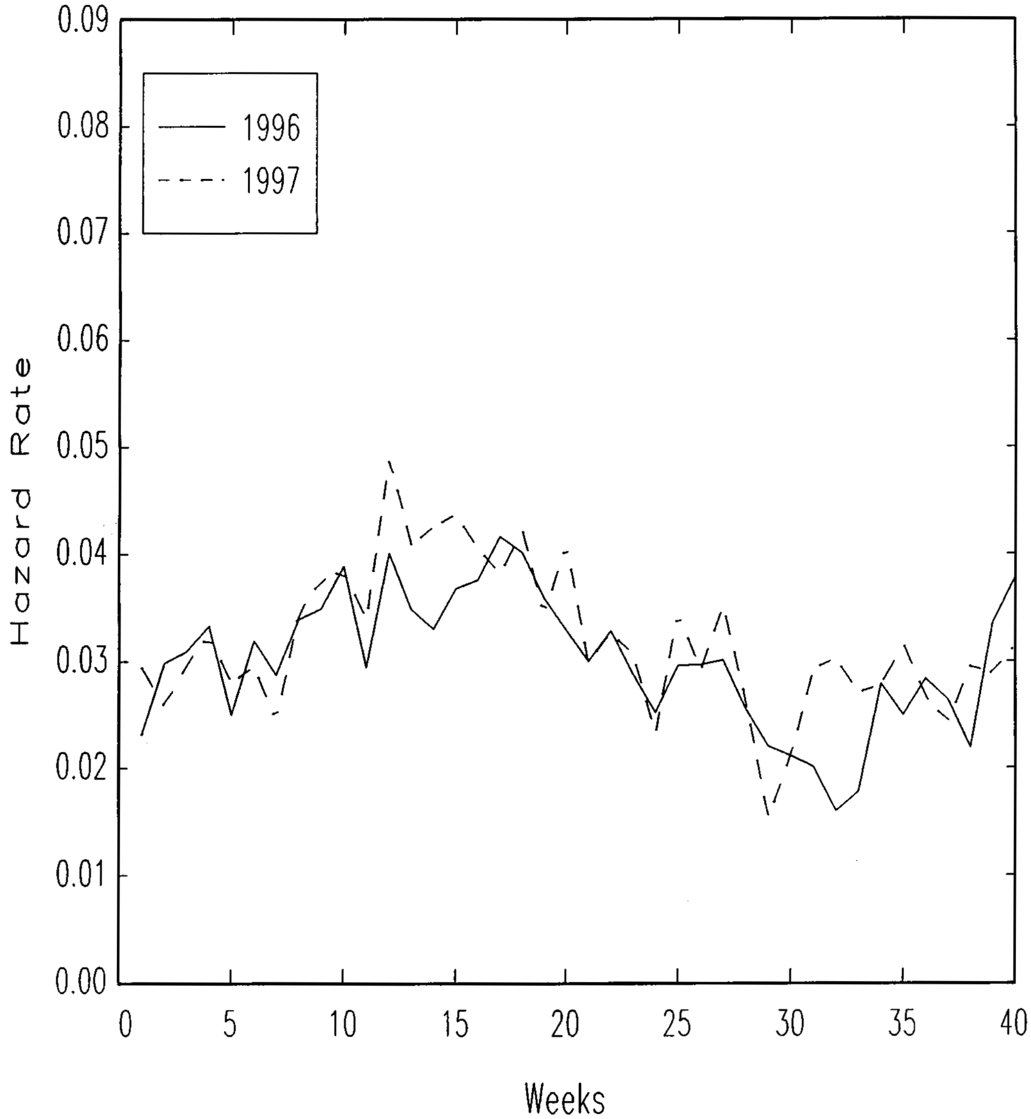

FIGURE 29A
Fitted Hazard Rate Pattern: Seasonal Jobs, Plots With UI Effects, Unemployment Rate Over 16% Region

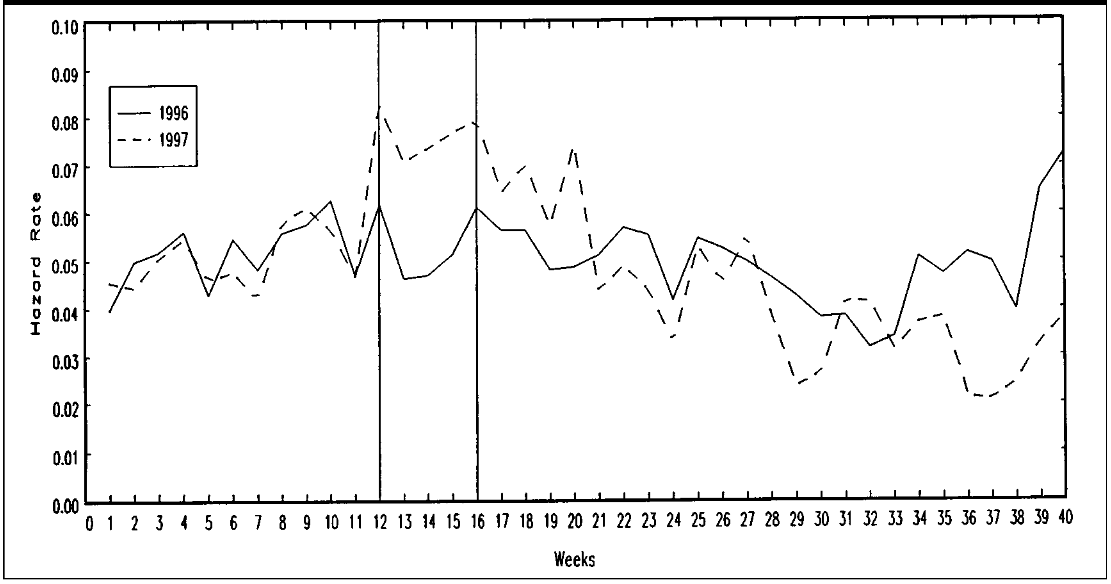

FIGURE 29B
Fitted Hazard Rate Pattern: Seasonal Jobs, Plots Without UI Effects, Unemployment Rate Over 16% Region

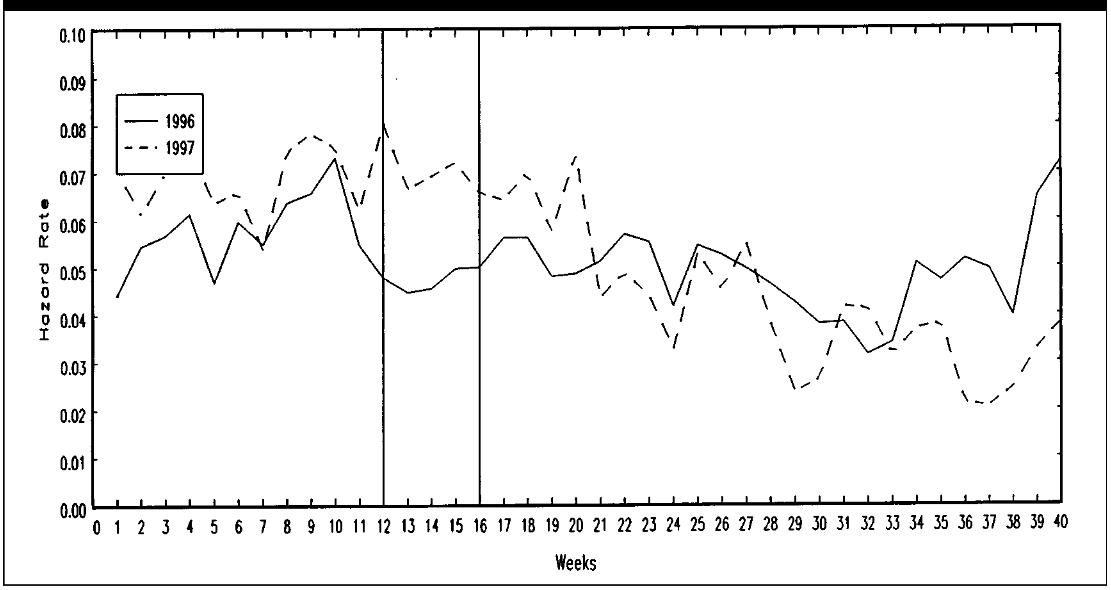

FIGURE 30
Cumulative Distribution Functions for Hours per Week, 1996 and 1997 First Three Quarters, Non-seasonal Workers

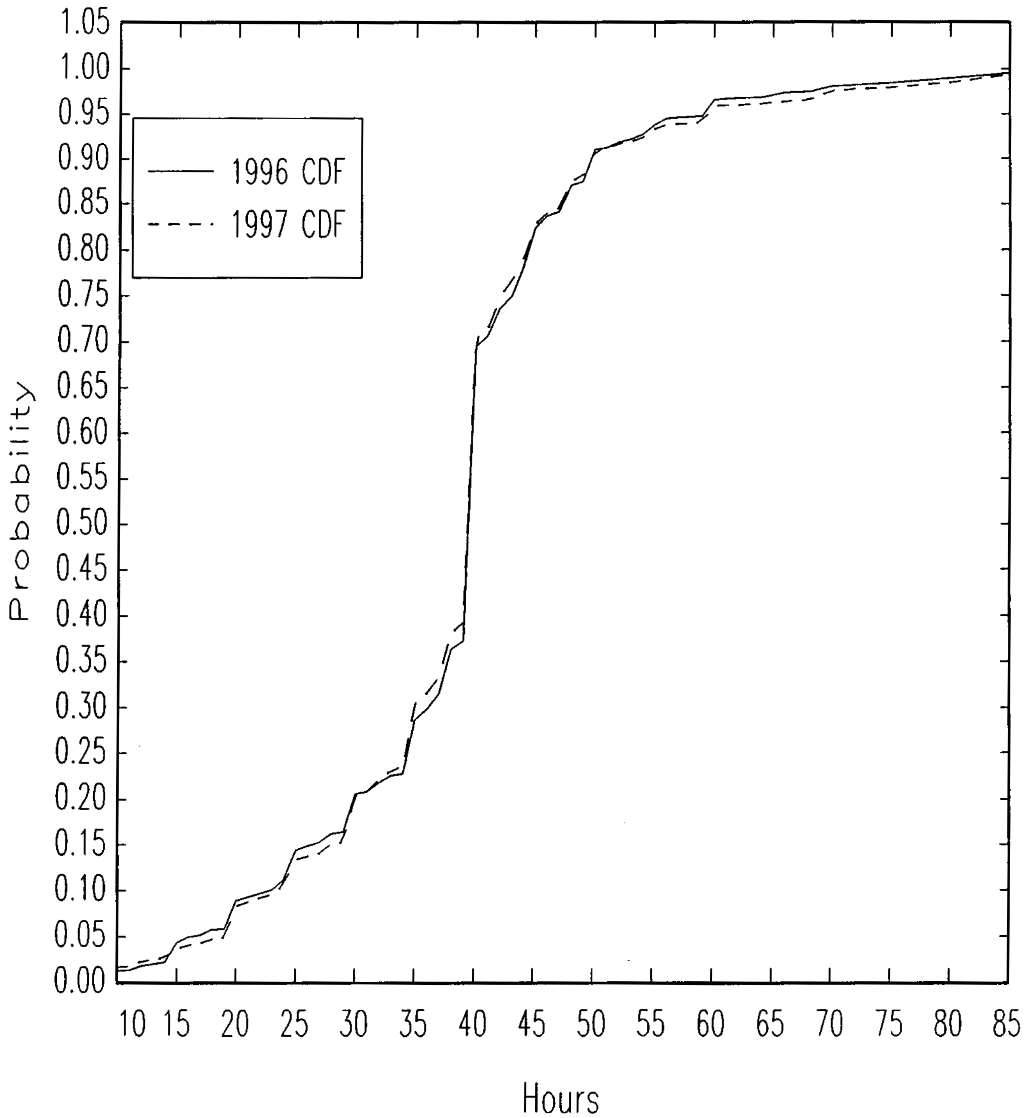

FIGURE 31
Cumulative Distribution Functions for Hours per Week, 1996 and 1997 First Three Quarters, Seasonal Workers

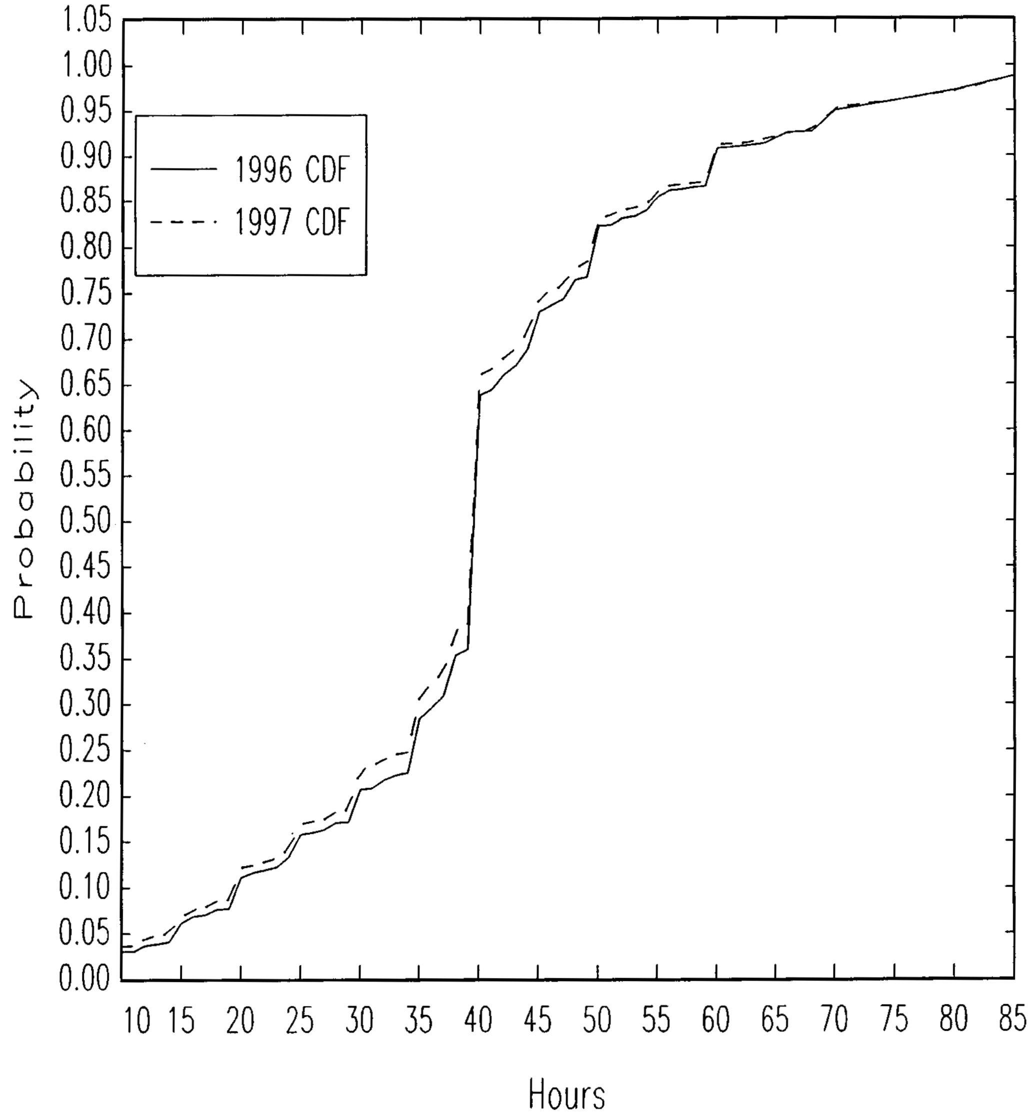

FIGURE 32
Usual Hours per Week, Cumulative Distribution Functions: Seasonal Workers in Regions with Unemployment Rate Over 13%

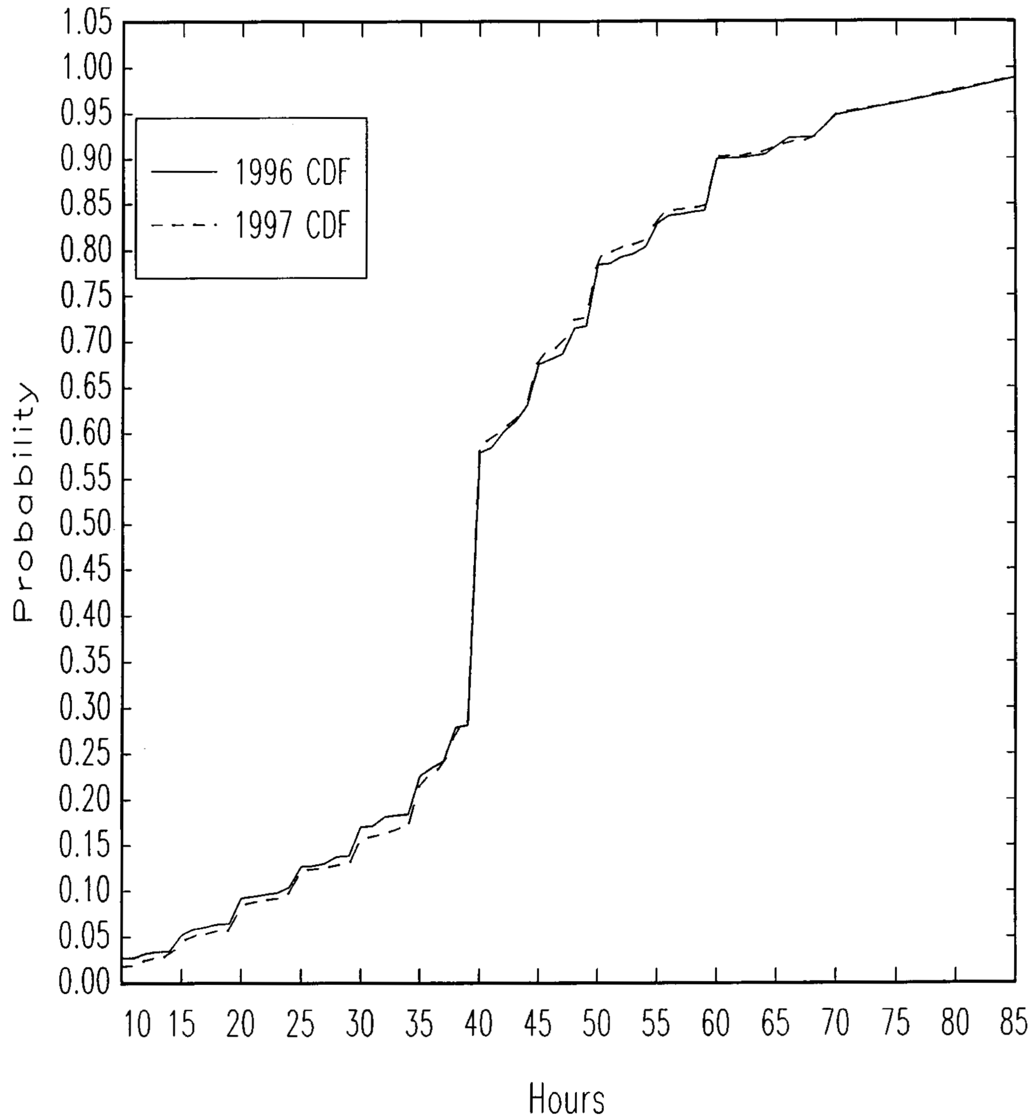

FIGURE 33
Usual Hours per Week, Cumulative Distribution Functions:
Seasonal Workers in Regions with Unemployment Rate Under 10%