Eligibility for UI Benefits, Take-up of Benefits and the Financial Liability of the UI Account

Final Report

Prepared for:

Human Resources Development Canada

May 1998

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SP-AH049E-05-98
(également disponible en français)

Acknowledgements

The authors wish to thank Anne Routhier and Ging Wong for many helpful comments at several stages of this research. Anne Routhier also provided invaluable help with the data.

Preface

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

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

As part of this process, the Department commissioned five formal evaluation studies on how Canadians adjusted to the 1994 UI reforms. These studies were performed by external academic subject-matter experts. Each evaluation represents a stand alone analysis of a specific topic.

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Executive Summary

This study evaluates the impact of Bill C-17 upon unemployment insurance (UI) eligibility status of job losers and the probability that eligible persons actually claim these benefits. Increases in weeks needed for eligibility were introduced in C-17 and these had the potential to significantly reduce the probability that many job losers were eligible for benefits. The impact of these changes to eligibility on benefit receipt rates might be further amplified if they coincide with reductions in claim rates for eligible persons. These effects, combined with reductions in the length of benefit entitlement periods under C-17, have the potential to significantly alter the financial obligations of the UI program. This study conducts a statistical analysis of the effects of C-17 upon eligibility for benefits and claim rates and then applies the results of these analyses to estimate the dollar values of changes to the liability of the UI account.

The possibility of reductions in benefit claim rates due to C-17 is suggested by recent academic studies in Canada and the United States that have found that a significant proportion of the unemployed do not claim the benefits to which they are entitled. These studies often find that a reduction in the generosity of UI benefits reduces the likelihood that persons will find it in their interest to claim them. In the United States, for example, there is evidence that UI benefits were claimed less often as these benefits became subject to income taxation during the 1980s. Inasmuch as C-17 reduced certain aspects of UI generosity, we might expect to see a similar fall in claims among those eligible for benefits as a result of C-17. A complete analysis of the impact of C-17 upon UI pay-outs needs to consider this claim-rate effect.

To quantify these effects of C-17, this study considers a group of job losers, many, but not all, of whom are eligible for benefits. For this group, each element of the following sequence of possible outcomes is examined:

1.) Establish Eligibility → ***2.) Decide to Claim Benefits*** → ***3.) Exhaust Benefits***

To investigate the effect of C-17 upon eligibility for benefits, the first link in this sequence, statistical methods are use to determine how eligibility status has changed under C-17 for persons with given observable characteristics (young women in Nova Scotia, for example). Next, for those persons eligible for benefits, possible changes in the propensity to actually claim benefits due to C-17 are quantified. Here, we examine the so-called 'take-up rate', which is defined as:

$$\frac{\textit{number of jobless persons who claim UI benefits}}{\textit{number of jobless persons who are eligible for UI benefits}}$$

Given that C-17 affected the generosity of unemployment insurance benefits in several dimensions, the C-17 reform had a clear possibility to change this take-up rate.

Any changes in eligibility for benefits and take-up rates would be of interest to policy makers for several reasons. First, any large drop in take up of benefits would imply a significant fall in the perceived value of these benefits. This would suggest that the adequacy of the income support provided by UI was jeopardized by C-17 changes. If take-up rates do not change, on the other hand, this would suggest that benefits continued to be viewed as a valuable source of income support.

A second source of interest in the impact of C-17 on eligibility and take-up of UI is that changes in the rate at which benefits are received have an impact on the financial state of the UI account. There is a fairly clear and immediate effect of changes in eligibility rates upon the requirement to pay out benefits. Similarly, if take-up probabilities rise or fall in any significant way the financial obligation of the UI fund will experience a corresponding increase or decrease. A traditional method of evaluating the effect of a UI reform would involve calculating changes to benefit entitlements of eligible workers. This method is incomplete, however, if there are also changes in take up behaviour. This study addresses this issue and provides measures of the impact of C-17 on the UI account that quantify the impact upon take-up probabilities and convert this into dollar amounts. Finally, given that C-17 may have changed the length of time during which benefits can be received, this study attaches dollar amounts to these reductions for persons who exhaust benefits under C-17 benefit entitlements but not pre-C-17 rules.

The quantitative analysis produced by this study is facilitated through the use of data from surveys conducted by Human Resources Development Canada in 1993 and 1995. These surveys, known as the Canadian Out of Employment Panel (COEP) surveys, were based on samples of persons known to have separated from an employer due to the filing of a Record of Employment. Random samples of persons separating from a job were contacted and questioned regarding their personal and job characteristics. They were re-contacted at later dates in order to determine the status of their searches for new jobs. In addition to a wealth of information provided directly by the surveys, links to HRDC administrative data provided an even more detailed description of the persons in the samples.

The mathematical relationships uncovered in this analysis of the COEP data are used to conduct "what if?" analyses of policy issues. We use our statistical relationships to calculate 'predicted' eligibility rates, claim rates and weeks of benefit eligibility for the 1993 and 1995 COEP samples. We do this using the actual UI rules in force at the time and also for the rules at the time of the other

sample. The benefit of our statistical analysis is that it can tell us what eligibility rates would have been if, say, the C-17 policy rules had actually applied to 1993 COEP characteristics and behaviour or if 1993 rules had applied to the 1995 COEP sample.

Given these 'what if' results, we can then see how payments to the 1995 COEP sample would have changed if the 1993 pre-C-17 rules had still been in effect. This analysis separates the effects of reductions in eligibility probabilities, benefit replacement rates, reductions in weeks of benefit entitlement and changes in the take up rate. Taken together, these results show how much money C-17 saved the UI account as well as why these savings occurred.

The first set of results presented in the report simply look at average take-up rates in the 1993 and 1995 samples. There is evidence of modest declines in administrative data take-up rates between 1993 and 1995 (from 82.6 percent to 80.7 percent). These differences are small, however, and it is actually true that the fraction of persons who reported that they applied for UI was higher in 1995 (67.9 percent) than in 1993 (65.6 percent). More demanding eligibility criteria in 1995 could explain why this did not translate into higher administrative-data take-up rates.

The study next goes beyond raw data analysis to conduct a statistical analysis of the probability that persons with given demographic factors (young women in Nova Scotia, for example) are actually eligible for UI benefits. There are changes in these relationships but few generalizable trends are observed. An interesting finding is that once we control for the regional unemployment rate, workers leaving seasonal jobs have lower probabilities of being eligible for benefits than do non-seasonal workers. This suggests that C-17 may have reduced eligibility for low-attachment workers. Also, the regional unemployment rate has a positive effect in 1995 while it was negative in 1993. This 1995 result is what we would expect given that variable entrance requirements make qualifying for benefits easier in high unemployment areas. The 1993 result may have captured a reduced availability of jobs needed to generate insurable weeks.

Probabilities of claiming benefits were also linked to observable variables in the same way as eligibility probabilities. Prior to C-17, take-up rates were lower for seasonal workers than for non-seasonal workers while the opposite became true afterward. The reason for this change is not immediately apparent. It is important to recognize, though, that weeks of benefit entitlement are entered separately in the analysis so that reductions in entitlements that were more significant for seasonal versus non-seasonal workers would be captured by the entitlement variable directly.

We next used these statistical analyses to see how average eligibility status, take-up probabilities and benefit entitlement weeks would have been different under the pre- and post-C-17 rules. This was done for the characteristics and statistical behavioural rules of both the 1993 and 1995 COEP for the entire COEP samples. This yields what is perhaps the most striking results of this study: C-17 seemed to have had a relatively small impact upon the probability that an unemployed persons was eligible for some benefits but large effects upon numbers of weeks of benefit entitlement.

These results can be illustrated by two graphs. First, Figure 1 (see page20), which shows most UI claimants had at least the minimum number of weeks needed to qualify in their region both before and after C-17. This, combined with a slight rise in claim rates, shows why C-17 had little impact upon eligibility. There was however, a much larger effect upon the number of weeks of benefits to which eligible persons were entitled. As Figure 2 shows (see page 21), most persons in the 1993 COEP sample received the maximum benefit entitlement while the distribution of benefit entitlements lengths in the 1995 COEP was spread out much more over the range of possible numbers of weeks. This was the prime way that C-17 impacted upon the persons in the COEP sample.

A useful summary of these changes in eligibility, take-up and benefit weeks and amounts is obtained by attaching dollar amounts to C-17 reforms. Using the 1995 COEP sample as a reference group, it was estimated that C-17 reduced the maximum potential benefit pay-out (the pay-out if all eligible persons collected all the benefits to which they were entitled for the maximum period of time) by 28 percent. Most of this change was due to a drop in maximum weeks of benefit entitlement (61 percent of the total reduction) and cycling behaviour (31 percent) rather than reduced eligibility (0.2 percent) or lower benefit replacement rates (7 percent). The slight drop in take-up rates and the fact that most spells end before benefits are exhausted mean that the actual saving realized was more in the order of 21 percent.

The policy conclusions derived from this study are as follows. First, C-17 does not seem to have lowered the attractiveness of UI benefits enough to provoke a substantial fall in claim probabilities. Second, C-17 had a relatively minor effect on eligibility for benefits but a large effect on the length of benefit entitlement periods. This could be viewed as a positive outcome since most unemployed persons still had access to some benefits, although the maximum benefit entitlement fell for lower-attachment workers. On the other hand, shorter benefits under C-17 may have meant that some unemployed persons eventually encountered significant financial difficulties. Finally, regardless of how we interpret the effect on the unemployed, changes in benefit duration under C-17 resulted in significant savings for the UI account.

1. Introduction

This study evaluates the impact of Bill C-17 upon unemployment insurance (UI) eligibility status of job losers and the probability that eligible persons actually claim these benefits. Increases in weeks needed for eligibility were introduced in C-17 and these have the potential to significantly reduce the probability that a typical job loser is eligible for benefits. The impact of these changes to eligibility on benefit receipt rates might be further amplified if they coincide with reductions in claim rates for eligible persons. These effects, combined with reductions in the length of benefit entitlement periods under C-17 have the potential to significantly alter the financial obligations of the UI program. This study conducts a statistical analysis of the effects of C-17 upon eligibility for benefits and claim rates and then applies the results of these analyses to estimate the dollar values of changes to the liability of the UI account.

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The possibility of reductions in benefit claim rates due to C-17 is suggested by recent academic studies in Canada and the United States that have found that a significant proportion of the unemployed do not claim the benefits to which they are entitled. For Canada, studies such as Storer and Van Audenrode (1995) place this figure at between 10 and 15 per cent of the eligible unemployed. Slightly higher percentages are obtained in studies for the US such as McCall (1995).

A common finding in these studies is that a reduction in the generosity of UI benefits lowers the likelihood that persons will find it in their interest to claim them. For example, in a seminal analysis of U.S. data, Blank and Card (1991) find evidence that UI benefits were claimed less often as these benefits became subject to income taxation during the 1980s. Inasmuch as C-17 reduced certain aspects of UI generosity, we might expect to see a similar fall in claims among those eligible for benefits as a result of C-17. A complete analysis of the impact of C-17 upon UI pay-outs needs to consider the claim rate effect.

To quantify these effects of C-17, this study considers a group of job losers, many, but not all, of whom are eligible for benefits. For this group, each element of the following sequence of possible outcomes is examined:

1.) Establish Eligibility → 2.) Decide to Claim Benefits → 3.) Exhaust Benefits

To investigate the effect of C-17 upon eligibility for benefits, the first link in this sequence, statistical methods are use to determine how eligibility status has changed under C-17 for persons with given observable characteristics (young women in Nova Scotia, for example). Next, for those persons eligible for

benefits, possible changes in the propensity to actually claim benefits due to C-17 are quantified. Here, we examine the so-called 'take-up rate', which is defined as:

$$\frac{\textit{number of jobless persons who claim UI benefits}}{\textit{number of jobless persons who are eligible for UI benefits}}$$

The take-up behaviour of the unemployed has two major consequences from the specific viewpoint of analyzing Bill C-17. First, the decision to take up unemployment insurance (UI) benefits is necessarily the outcome of a comparison of the benefits and costs of claiming. Legislation such as C-17 had in some cases a major impact upon the benefits of claiming and it will be interesting to know how this has affected the take-up decision. Any changes in take-up behaviour will have direct implications for the perceived adequacy of UI benefits, one of the preoccupations of the current review of C-17. Another aspect of C-17 under review is its impact upon the financial state of the UI account. Changes in take-up behaviour have implications for this question because the relationship between UI revenues and expenses will change if eligible unemployed persons become more or less likely to claim their benefit entitlement.

The quantitative analysis produced by this study is facilitated through the use of data from surveys conducted by Human Resources Development Canada in 1993 and 1995. These surveys, known as the Canadian Out of Employment Panel (COEP) surveys, were based on samples of persons known to have separated from an employer due to the filing of a Record of Employment. Random samples of persons separating from a job were contacted and questioned regarding their personal and job characteristics. They were re-contacted at later dates in order to determine the status of their searches for new jobs. In addition to a wealth of information provided directly by the surveys, links to HRDC administrative data provided an even more detailed description of the persons in the samples.

The mathematical relationships uncovered in this analysis of the COEP data are used to conduct "what if?" analyses of policy issues. We use our statistical relationships to calculate 'predicted' eligibility rates, claim rates and weeks of benefit eligibility for the 1993 and 1995 COEP samples. We do this using the actual UI rules in force at the time and also for the rules at the time of the other sample. The benefit of our statistical analysis is that it can tell us what eligibility rates would have been if, say, the C-17 policy rules had actually applied to 1993 COEP characteristics and behaviour or if 1993 rules had applied to the 1995 COEP sample.

Given these 'what if' results, we can then see how payments to the 1995 COEP sample would have changed if the 1993 pre-C-17 rules had still been in effect. This analysis separates the effects of reductions in eligibility probabilities, benefit replacement rates, reductions in weeks of benefit entitlement and changes in the take-up rate. Taken together, these results show how much money C-17 saved the UI account as well as why these savings occurred.

The study is structured as follows. We begin by presenting the economic framework for the analysis and then discussing the statistical methods used to uncover behavioural relationships from the data. We next discuss how these statistical results can be used to address the policy questions of interest. The features of the data are then described and statistical results are presented. We discuss two types of 'what if' experiments conducted using what we have learned from the statistical analysis: the effects of C-17 on eligibility, claim rates and benefit weeks followed by calculations of how C-17 affected the financial obligation of the UI account. A brief section of conclusions ends the report.

2. Economic Framework for the Study

In a simple framework where unemployed persons can quantify the costs of claiming UI benefits, a claim will be filed if the expected present value of the net income from UI benefits exceeds the lump-sum cost of filing for benefits. If benefits are b_i for each week of the benefit entitlement period and a corresponding per-week cost of c_i is incurred, then benefits are claimed if:

In a simple framework where unemployed persons can quantify the costs of claiming UI benefits, a claim will be filed if the expected present value of the net income from UI benefits exceeds the lump-sum cost of filing for benefits.

$$\sum_{i=1}^{d} \beta^i \ (b_i - c_i) > C_0 \qquad (1)$$

Here, C_0 is a fixed cost of filing incurred at the beginning of the period. The upper limit of the summation operator, d, is the *lesser* of either the expected duration of unemployment in weeks or the maximum benefit entitlement. Both of these periods are measured in weeks. The discount factor β is used to convert future benefits into present values. In Canada, $b_i = 0$ for $i=1,2$ due to a mandatory two-week waiting period. The *expected* duration of unemployment is relevant here because persons are required to anticipate how long they expect to receive UI benefits. Persons who anticipate a short period of joblessness will find it less beneficial to file a claim than someone who feels that they are likely to face a long wait before a new job arrives. The fixed-cost variable C_0 captures both the time taken to submit an application form and any other lump-sum costs. The recurring costs c_i capture the per-week cost of continuing a UI claim.

Several predictions flow from this model. As was also shown by Anderson and Meyer (1994), one implication is that take-up rates should be positively related to the expected duration of unemployment since d rises with unemployment duration until the point of benefit exhaustion. For persons who expect to have a duration of two weeks or less, there is no value to filing for benefits due to the two-week waiting period. After this point, the value of claiming is an increasing function of the expected duration of benefit receipt.

Higher UI benefit replacement rates should increase take-up rates because they increase the value of claiming. Also, other things being equal, persons having previously earned higher wages should have a higher probability of claiming benefits because b_i is generally an increasing function of past wages, although this may not hold if wages function as a proxy for other factors such as unobserved heterogeneity. While the theoretical effect of the cost variables C_0 and c_i on the take-up decision is clear, it is far from obvious how these costs should be measured in practice. It may be possible to associate certain demographic factors such as rural versus urban residence with these costs.

Unfortunately, a large aspect of these costs is non-pecuniary in nature and this prevents a precise determination of the costs of filing a UI claim. In the 1995 COEP, and to a lesser extent the 1993 data, a series of questions are asked designed to obtain some measure of these non-pecuniary costs. For example, persons who did not claim UI benefits are asked why they did not do so. Among the possible codes are “don’t believe in U.I” and “too much trouble”. Similarly, persons who delayed filing may state that they “didn’t get around to it”. Persons whose responses receive these codes indicate that for reasons of political belief or motivation they did not wish to claim benefits, at least not immediately.

3. Statistical Methodology for Data Analysis

As explained above, the definition of the take-up rate adopted in this study is the following:

$$\frac{\textit{number of jobless persons who claim UI benefits}}{\textit{number of jobless persons who are eligible for UI benefits}} \qquad \textbf{(2)}$$

This type of measure has been analyzed in studies by McCall (1995) and Storer and Van Audenrode (1995). It is somewhat different from the definition used by Blank and Card (1991) which looks at the fraction of eligible insured weeks for which UI benefits were actually received. Given that one of the goals of this study is to see how C-17 has changed the propensity of individuals to claim benefits, the individual "spell-based" definition in (2) above is appropriate.

Initially, the econometric analysis looks at the impact of Bill C-17 on UI take-up behaviour. Using COEP data, we determine the UI eligibility status of each job loser. In a first step, we use pre-C-17 COEP data to model the determinants of the decision to claim benefits prior to the changes. This is done using limited dependent variable estimation methods to estimate the probability that an eligible unemployed individual will actually claim benefits. In this approach, the probability of claiming benefits is modelled with the following equation:

$$\textit{Pr (claim UI benefits)} = F(X\Gamma) \qquad \textbf{(3)}$$

The matrix X is a group of explanatory variables and the parameter vector Γ contains the response of the take-up probability to changes in these variables. Among the X variables, particular attention will be paid to regional/sectoral determinants of take-up rates. Individual differences in take-up behaviour will be captured by the parameters Γ of these variables among the X matrix. For the function $F(X\Gamma)$, either a logistic function or the standard normal distribution function can be used, giving rise to logit and probit specifications respectively. The sensitivity of results to this choice will be examined.

While we will also look at the impact of UI parameters on take-up by estimating the equation above with both pre- and post-C-17 data, in a first step we estimate the impact of Bill C-17 on take-up rates assuming no behavioural change is induced by the bill. In other words, in this first analysis it is assumed that the Γ coefficients capturing take-up responses are unaffected by C-17 and it is only values of the X variables (such as insurable weeks) that change as a consequence of C-17. This is done by

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applying parameters calculated with pre-C-17 COEP data to changes in the X matrix induced by C-17.

In a second step, we use post-C-17 COEP data to re-estimate determinants of take-up behaviour so that the parameters may themselves change due to C-17. This will allow us to estimate the effect of some specific changes in UI parameters introduced by C-17. Modification of behavioural responses might be due to a change in attitude with respect to UI, or by a perception that it could be more difficult to re-qualify in the future, or again, by changing the pool of eligible persons. With this in mind, the impact of replacement rates (particularly differences between individuals receiving 60 percent versus those receiving 55 percent or less) and the impact of the duration of benefits on the decision to claim is studied. Formal tests of behavioural changes (individual or regional, for example) before and after C-17 are conducted through the use of econometric tests for the constancy of certain elements of the Γ vector[1].

The estimation and testing outlined above will enable us to estimate whether Bill C-17 has significantly changed take-up behaviour, either indirectly, through changes in replacement rates or weeks of benefit entitlement (captured by X), or directly, by inducing behavioural changes in the take-up decision (captured by Γ). This part of the project allows us to infer how perceptions regarding the adequacy of UI benefits were affected by C-17 and to estimate the impact on the UI account of C-17 induced changes in take-up behaviour.

The remainder of the statistical analysis examines in detail the economic motivation behind the take-up decision, a question related to the econometric evaluation of the effects of UI on unemployment durations in general. Many eligible non-claimants have short spells of unemployment. Thus far, no evidence exists showing whether this correlation is due to pure luck (people are lucky finding a new job quickly and don't have time to claim) or to conscious decisions (people don't want to claim and search intensively as a consequence). The lack of good instruments that could permit the unravelling of this simultaneity between the decision to claim and the duration of the spell has made it impossible to resolve this question. We try to remedy this by using the information provided in the 1995 COEP regarding the timing of claims. Several avenues can be explored. First, a simple analysis of the timing of claim is conducted to determine whether people always claim immediately after losing their job or tend to wait before claiming. Second, we consider the motivation for the unemployed to delay (or not) their claim.

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[1] We use likelihood ratio tests which look at the decrease in the ability of the model to explain the data when certain algebraic restrictions are added. Statistical methods are used to determine whether the losses in explanatory power associated with these restrictions are large enough to be 'significant'. If losses are not significant, we cannot reject the restrictions.

Finally, we fit a proportional hazard model of the time elapsed between the beginning of a spell of eligible unemployment and the moment of claiming, allowing for censoring when a job is found before the end of the benefit entitlement period or a claim still has not been filed at the end of the observation period. In this specification, the probability of claiming benefits (the "hazard rate") for an individual with characteristics X who has been eligible to claim for d periods is determined by the equation:

$$\lambda\,(d,X_i) = \lambda_0(d)\,\exp(X_i\delta) \qquad \textbf{(4)}$$

Here $\lambda_0(d)$ is the so-called baseline hazard which is common to all individuals i but may vary with the elapsed duration of eligible unemployment d. The advantage of this proportional hazard model of time-to-claim, relative to alternate models such as the "accelerated failure" model, is that there is a non-parametric baseline hazard which is not itself estimated. This is done through the use of a partial likelihood method in which a likelihood function to be maximized is independent of the baseline hazard. There is some suggestion in the literature (see Meyer (1990), for example) that the δ parameters obtained with this method are relatively unaffected by the presence of unobserved individual-level heterogeneity.

Use of this specification allows us to identify unemployed persons with exceptionally long or short delays in their claim timing (by looking at the difference between the actual time of the claim and the integrated hazard obtained from the hazard rate estimation). We then determine whether these unemployed differ significantly from other claimants and non-claimants with regard to unemployment duration and new wages.

For the analysis of the probability that persons are eligible, the same methodology described above for the probability of claiming benefits is applied. This is done by replacing the benefit claim probability in equation (3) by a benefit eligibility probability. We avoid turning our equation into a policy rule by using personal characteristics in the X matrix. As for take-up of benefits, we examine how C-17 affected the eligibility probability of different groups.

4. How to Link Statistical Results to Policy Questions

While of interest in themselves, the primary purpose of the econometric analysis outlined above is to contribute to the resolution of the following two policy-related questions:

(1) How has Bill C-17 changed UI take-up behaviour and what do these changes imply about the perceived adequacy of UI in Canada?

(2) What is the dollar-value of changes to the UI account implied by C-17 induced changes to UI take-up rates?

These two questions will be discussed in turn.

First, the logit and probit results for the probability of claiming UI under the pre- and post-C-17 regimes are used to examine how perceptions of the value and adequacy of UI benefits has changed as a result of C-17. It is possible that persons with a given number of insurable weeks and regional unemployment rate would have filed for benefits under the regime in place prior to C-17 but not file given the post C-17 program. This is because the length of the benefit entitlement period fell in many cases so that the value of claiming also fell. The value of claiming was also affected by a drop in the replacement rate from 57 to 55 percent. In terms of the theoretical model presented above, persons with a certain range of fixed costs C_0 might determine that after C-17 the return to filing no longer exceeds the cost. The logit regression coefficient allows us to determine the extent to which persons have become less likely to claim and this gives an indication of how much the perceived value of UI benefits has fallen.

It is worth noting that this analysis can only be done if the effect of C-17 on UI take-up is separated into an eligibility/entitlement effect due to changes in the calculation of benefit entitlement and a behavioural effect of a change in take-up probabilities given the number of insurable weeks. Only the behaviour of individuals can inform us regarding the perceived value of UI benefits. Mechanical effects due to the change in the rule for determining eligibility must therefore be eliminated from the calculation and the econometric methodology outlined above achieves this separation.

For the second question, the decomposition of the effects of C-17 is less crucial. Here, we seek to quantify the reduction in terms of benefits paid due to C-17. There is the clear possibility of savings due to reduced eligibility, lower replacement rates and the shortening of benefit entitlement periods for

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certain groups. This is not the entire effect, however, since take-up rates may have fallen under C-17 so that costs for a given distribution of eligibility in the population may have also fallen. To evaluate this take-up effect, it is necessary to translate changes in take-up behaviour into dollar amounts.

This dollar amount is calculated by using the following methodology. First, the characteristics of the 1995 COEP are used to calculate weekly benefit amounts and maximum weeks of benefit entitlements under the post-C-17 rules. This amount is called the "maximum potential pay-out". In other words, this is the potential dollar amount if all eligible persons claimed their maximum benefit entitlement. Next, two smaller amounts are calculated: "maximum claimed pay-out" and "actual dollars paid". The maximum claimed pay-out is less than than the maximum potential pay-out because some eligible persons will not claim (e.g. the take-up rate for benefits is less than one). Finally, actual dollars paid is less than the maximum claimed pay-out because many unemployment spells end prior to the end of the maximum entitlement.

The following table summarizes the definitions of these three measures of potential benefits pay-outs:

Term	Payment Conditions
Maximum Potential Pay-out	All eligible persons use maximum benefit entitlements
Maximum Claimed Pay-out	All claimants use maximum benefit entitlements
Actual Dollars Paid	Actual (not maximum) benefit payments

These three measures of potential and actual dollar pay-outs are re-calculated using the pre-C-17 rules. In other words, a "what-if"experiment is conducted in which we see how much more or less would have been paid to the unemployed in 1995 if pre-C-17 rules had still been in effect. By looking at changes in the maximum potential pay-out, the maximum claimed pay-out and the actual dollars paid, we are able to separate effects of changes in eligibility and benefit entitlement rules, take-up of benefits and the relationship between benefit entitlement weeks and unemployment duration.

5. *An Overview of the Data Used*

The primary sources of data for use in this study are the 1993 and 1995 Canadian Out of Employment Panel data-sets ... supplemented with administrative data.

The primary sources of data for use in this study are the 1993 and 1995 Canadian Out of Employment Panel data-sets. In both of these panel studies, information about UI claims is available. While the 1995 study provides more information about take-up of UI benefits there is nevertheless an extensive set of common information available in the two samples. Given that the 1993 COEP covers the period before Bill C-17 and the 1995 COEP samples unemployment spells beginning after C-17 came into force, these data are ideally suited to the purpose of this study.

The sampling methods of both COEP data sets are the same: Record of Employment information is used to identify persons leaving a job. These persons are then surveyed roughly half a year and one year after job loss. In both samples, two cohorts were used in order to have some control for the effects of seasonality. Cohort One was essentially composed of persons losing jobs in February or early March of 1993 or 1995 while Cohort Two lost jobs through late April to early June of the respective year. It is worth noting that these dates will not permit us to see persons leaving seasonal jobs in industries such as fishing or forestry since these jobs are likely to end during the summer or early fall. Retail sector seasonal employment will be much more likely to be captured by these sampling dates.

For each cohort, follow-up waves of questions were asked some time after the moment of job loss. The 1993 COEP had three waves of interviews at averages of 23 weeks, 38 weeks and 58 weeks. Just two waves were used in the 1995 COEP and the average elapsed time before each wave was 31.5 weeks for wave one and 57 weeks. The 1995 second wave and the 1993 third wave were thus at roughly the same point. The 1995 first wave fell roughly mid-way between 1993 first and second waves. For this study, the precise timing of the interviews is not of great importance because we are studying the take up of UI claims after job loss. If there were large differences in elapsed times between the 1993 and 1995 COEP samples we might worry about differential recall bias but the timing is so similar here that this should not be a problem.

The COEP data are supplemented with administrative data. It is possible to link unemployment spells to unemployment insurance administrative records and this gives a raw measure of the take-up rate defined as in equation (2) above. However, by definition, administrative data alone do not allow us to determine which individuals applied for benefits but were turned down nor why persons not filing a claim chose not to do so. Furthermore, the

administrative data do not indicate how long an individual waited before claiming benefits. This type of information is available in the COEP samples and it is exploited in the econometric analyses outlined above in order to further our understanding of the nature of eligible persons not claiming benefits.

The method used to study take-up of benefits was the following. First, persons who say that they voluntarily left their jobs were excluded from the sample. This choice largely reflects the special nature of this group and their sometimes ambiguous eligibility status under UI law. Next, for persons who claimed benefits we used administrative file information to determine whether they were deemed to be eligible for benefits and, if so, the length of their benefit entitlement period. For non-claimants it was necessary to impute eligibility for benefits by combining information on insurable weeks, the UI region of residence and the appropriate UI law (pre- or post-C-17). We examined the accuracy of these imputed values by comparing them in cases where both types of information were available. We found that our method, while imperfect, did perform reasonably well.

Special attention was given to persons who claimed benefits but who might have an on-going UI claim. For these persons, the appropriate measure of benefit entitlement is the number of weeks remaining at the time of job loss and not at the beginning of earlier claim. We were able to calculate the correct number of weeks for this group[2]. Replacement rates were either 55 percent or 60 percent according to family situation and income. For claimants, information was provided regarding the actual replacement rate. For non-claimants, the 55 percent default replacement rate was used as the potential replacement rate since there was not enough information about family composition and income to determine whether the higher 60 percent rate should have been applied.

An important feature of these data required further attention. The original 1993 COEP sample was designed so as to deliberately over-represent person claiming UI benefits. This was done because claimants were the group of interest. It is thus necessary to use weights provided by HRDC in order to ensure comparability of the two samples with regard to probabilities of claiming[3]. Analysis of the take-up definition (2) makes it clear that, without adjustment for weights, raw take-up rates will be biased upward in the 1993 COEP data.

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[2] For these persons, we computed the remaining number of weeks on the ongoing claim (i.e. the maximum number of weeks of benefit entitlement minus the time elapsed between the start of the benefit period and the Record of Employment date).

[3] Over-sampling of certain types of job loss was also done in the 1993 COEP and this also needs to be controlled for when comparing the 1993 and 1995 data.

6. Statistical Results

A. Sample characteristics

We begin this section by analyzing the characteristics of the persons found in the 1993 and 1995 COEP samples. This is done in Table 1. Of note for this study is the fact that 68 percent of the 1995 sample applied for UI benefits while this proportion was slightly lower at 66 percent in the 1993 data. While this raw figure does not take into account changes in the composition of the unemployed, it is nevertheless a small surprise. If reductions in the generosity of UI had made applying less attractive then we would have expected this figure to fall rather than rise.

A fall in reported take-up rates is apparent in both the weighted and un-weighted figures. The drop is much smaller, however, in the weighted data. Similar patterns are observed for take-up rates derived from administrative records ...

One advantage of the COEP data-sets is that information on UI take-up rates is available from two sources. While the act of claiming benefits is a verifiable act, the issue of eligibility for benefits is subject to some interpretation. The actual eligibility status of an applicant as determined by HRDC may in some cases differ from that perceived by the applicant. It is thus possible to determine two measures of the pool of eligible persons based on either self-reported eligibility from the survey or actual eligibility status as determined by HRDC administrative records[4]. In this survey, both methods are used since each has its value. Take-up rates constructed using survey data provide a better measure of the behaviour of persons who perceive that they are eligible for benefits while administrative rates may give better information regarding potential pay-outs from the UI account.

The all-sample average figures for take-up rates (rather than application rates) are more in keeping with the hypothesis that eligible unemployed workers may have judged that it was less profitable to claim UI with the 1995 versus the 1993 rules. A fall in reported take-up rates is apparent in both the weighted (80 percent to 75 percent) and un-weighted (89 percent to 76 percent) figures. The drop is much smaller, however, in the weighted data. Similar patterns are observed for take-up rates derived from administrative records rather than survey responses.

Some interesting facts regarding the period of time elapsed before a claim is filed are apparent. Looking at weighted survey responses, 94 percent of 1995 respondents recall having claimed within 2 weeks or less of losing a job while this figure was only 86 percent in 1993. In the administrative records, the

[4] When determining eligibility using administrative information, a person's eligibility is assessed based only upon documented employment history. Documentation can come from either an accepted claim or insurable weeks indicated on a record of employment.

corresponding figures are 77 percent versus 64 percent. Survey data may show more persons filing within two weeks or less because survey respondents might have felt that a short delay would somehow look better or because delays get shortened when remembered. Whatever the explanation, while the levels of these percentages differ between survey responses and administrative data, the two sources corroborate one another regarding differences in time-to-claim in the two years.

The greater time taken before filing a claim in 1993 versus 1995 may explain why application rates were higher in 1995. If theunemployed waited longer prior to applying in 1993 it is possible that some persons did not apply because they found a job before claiming. Of course, this raises the question of why these persons waited so long before claiming in 1993 and why they were faster to claim in 1995.

Table 1
Descriptive Statistics

	Coep 93		Coep 95	
Survey Data	**Not Weighted**	**Weighted**	**Not Weighted**	**Weighted**
Applied for UI	.790	.656	.694	.679
Will apply				
(% of those who didn't apply)	.131	.112	.085	.079
Reason for Not Applying				
Not eligible	.446	442	.162	.162
No separation papers	.005	.004	.001	.001
Found a job	.287	.301	.038	.039
Don't want UI	.068	.075	.015	.017
Already on UI	.085	.075	.023	.021
Take-up Rate	.888	.800	.761	.746
Weeks Before Applying				
0	.508	.523	.708	.703
1	.253	.242	.164	.165
2	.093	.093	.072	.076
3	.042	.041	.027	.029
4	.043	.036	.007	.008
Administrative Data				
Take-up Rate	.907	.826	.841	.807
Weeks Before Applying				
0	.429	.493	.708	.700
1	.157	.094	.043	.045
2	.091	.054	.028	.029
3	.057	.034	.017	.018
4	.034	.020	.014	.014

The potential for differences to be observed between survey responses and administrative records prompt us to examine Table 2. The first panel of this table shows that there are inconsistencies between 1995 COEP survey respondents' recollection of having applied for UI and administrative records.

For example, of 1,820 persons who claim to have not applied for UI, 566 (31 percent) do show up as having a claim in the administrative data. This could indicate that persons wish to hide the fact that they claimed UI, perhaps because of stigma effects or possible fear that some dishonesty on their part may be detected.

Some variance in estimated versus actual UI eligibility is also observed. Of 991 persons classified as not eligible in the UI files, 790 (80 percent) appeared to be eligible based on their characteristics and the UI rules in place under C-17. Of the 5,080 eligible persons, only 74 (1.5 percent) were incorrectly classed as ineligible on the basis of survey data. By far, the most serious type of misclassification based on survey data was the characterization of the ineligible as eligible rather than the opposite error.

A final interesting comparison is of the time-lag between layoff and the filing of a UI claim. As mentioned above, survey data do show more persons filing in two weeks or less than is found in administrative data. Table 2 reveals that this is mainly because some persons who claim to have filed quickly have delays of over four weeks according to administrative data sources. For example, 25 percent of persons who report having waited a week are listed as having waited over four weeks in administrative data.

Table 2
Comparing Claim Information in Survey and Administrative Data

These tables compare the UI situation as declared by respondents to the COEP 95 survey with their UI situation estimated from administrative records

Existence of a Claim:	**Administrative Data**	
Survey Data	**Claim Identified**	
Did Apply for UI	**No**	**Yes**
No	1254	566
Yes	586	3665

Evaluation of their eligibility:	**Administrative Data**	
Survey Data	**Identified Status**	
Estimated Status	**Not Eligible**	**Eligible**
Not Eligible	201	74
Eligible, did not claim	483	1102
Eligible, did claim	307	3904

Time between layoff and UI claim (in Weeks)

	Administrative Data					
Survey Data						
Weeks:	**0**	**1**	**2**	**3**	**4**	**>4**
0	1598	71	49	35	31	319
1	319	27	7	10	4	120
2	142	22	15	5	5	32
3	45	13	8	1	1	12
4	6	1	4	1	0	9
>4	304	14	12	5	5	156

There are also examples of faster claim behaviour in administrative date for some individuals. For example, of the group of people who, according to administrative records, waited less than a week after being laid-off before claiming benefits, 12.6 percent reported that they had waited more than four weeks. Only 66 percent of administrative data "rapid claimers" self-report that they waited less than a week before claiming. While it could be true that some UI claimants feel that they should give the impression that they only applied for UI after determining that they were not going to find a job quickly, the behaviour of these persons is unlike the self-reported fast claimers who show up as slow claimers in administrative data. These finding suggest that different types of recall and reporting errors quite likely coexist for different persons.

These differences observed between survey and administrative data reveal interesting facts about individual responses to surveys but do not jeopardize the usefulness of these surveys for the analysis contained in this report. For persons who claim benefits, information from UI files is used to determine eligibility for benefits and weeks of benefit entitlement. Eligibility for persons who do not claim benefits is determined from record of employment information regarding insurable weeks combined with regional unemployment rates and the UI rules. These same factors determine benefit entitlement. Responses from the survey data are used to provide the demographic information used in the statistical analysis of the report. While self-reported UI claim status and weeks worked may be subject to reporting errors, we do not use this information in the statistical analysis of the report and the information that we do use (such as education levels, province, sex, etc.) is less likely to be reported with error.

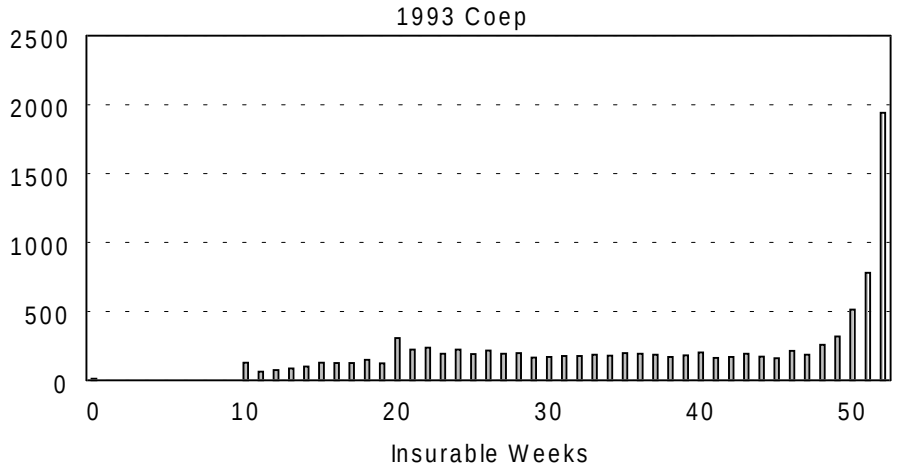
To end this section, it is useful to examine Figures One and Two which show how our measures of benefit entitlements are constructed. Figure One looks at the distribution of numbers of insurable weeks for persons who claimed benefits in both the 1993 and 1995 COEP samples. Insurable weeks are calculated from status vector information. These distributions look very similar, although one notable exception is that the lower limit of the distribution is ten weeks in 1993 versus twelve weeks in 1995. This change reflects changes in eligibility requirements for benefits in C-17.

The similarity of these two graphs is in stark contrast to the differences observed between the two panels of Figure Two that translate these insurable weeks distributions into weeks of maximum benefit entitlement under the two UI regimes. Under the pre-C-17 rules, most persons are entitled to 50 weeks of benefits. Once the C-17 reforms are in place, there is a slightly larger fraction entitled to zero weeks of benefits. Most persons continue to be eligible for some benefits but the number of weeks of benefit entitlement is severely reduced for most groups. The shape of the graph changes completely under C-17 as

many more persons have fewer than 50 weeks of entitlement. This points to a relatively minor effect of C-17 upon eligibility for benefits but a large effect upon the number of weeks of benefit entitlement. This will be examined more formally using statistical methods in the following sections of the report.

This points to a relatively minor effect of C-17 upon eligibility for benefits but a large effect upon the number of weeks of benefit entitlement.

Figure 1
Evolution of Insurable Weeks



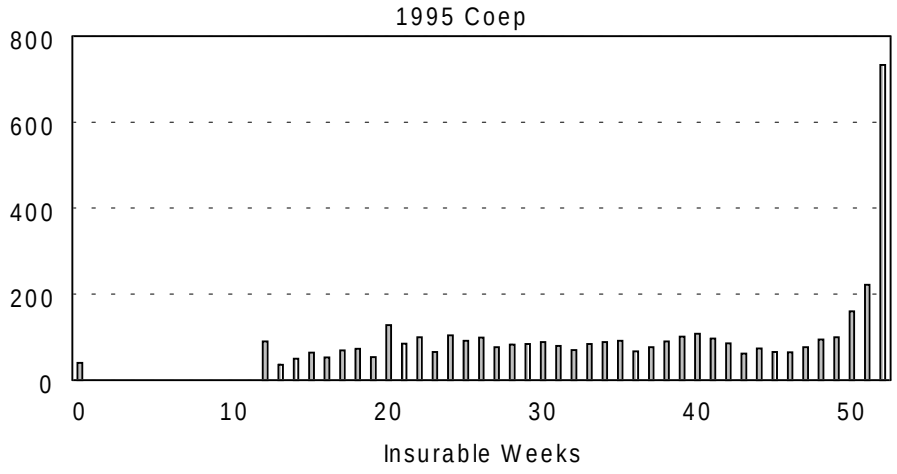
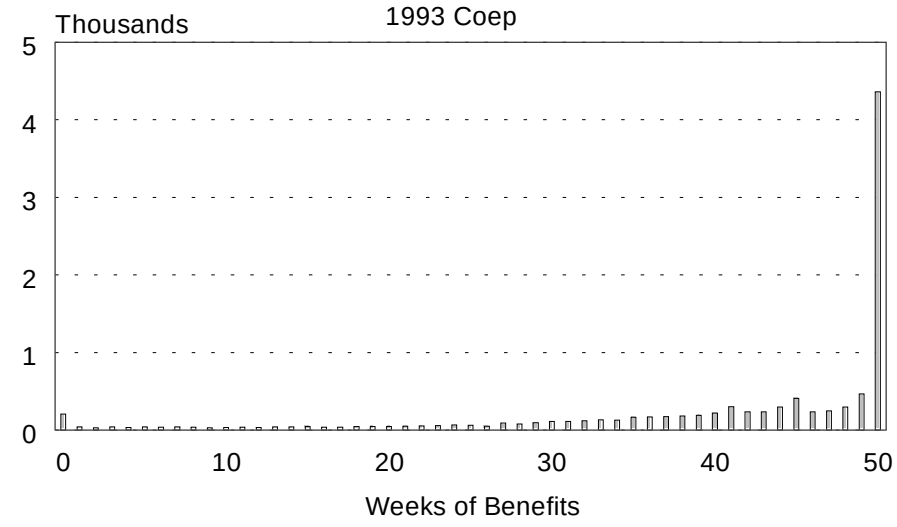
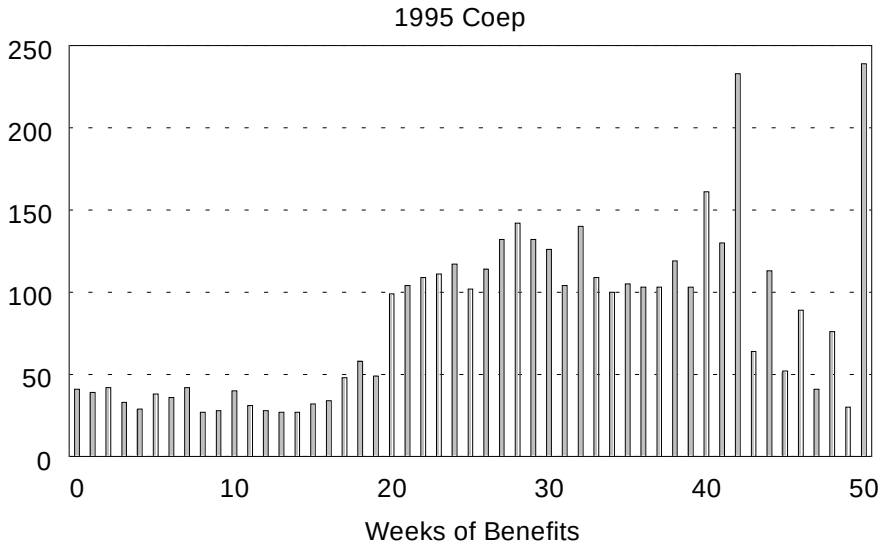


Figure 2
Weeks of Benefits





B. Determinants of the probability of being eligible for UI benefits

One important impact of C-17 on the unemployed is that it change eligibility requirements for workers with relatively low numbers of insurable weeks. As a result, we can expect to find that persons with a given set of characteristics were, on average, less likely to be eligible in 1995 than they would have been under the 1993 rules. The graphs shown in the previous sample suggest that C-17 had a greater impact upon benefit entitlement than upon eligibility and the goal of this section is to analyze the determinants of eligibility using formal statistical methods.

To quantify the effect of C-17 upon eligibility for benefits, we examined the relationship between observable demographic characteristics and the probability of eligibility for benefits using the 1993 and 1995 COEP samples periods. The characteristics examined were socio-demographic in nature rather than determinant of eligibility as specified in UI legislation. This is because we sought to determine how C-17 affected the probability of UI eligibility given that a person was, for example, a high-school educated young woman in Nova Scotia. These results thus combine both changes in eligibility criteria and the ability of various demographic groups to meet these criteria.

These comparisons are done in three steps. First, only personal characteristics such as age, marital status, minority or disabled status, sex, education, language of the interview and the province of residence were used to explain the probability of being eligible for UI. Next, only job-related characteristics such as the wage lost, union status of the lost job, seasonality of the lost job and information about notice and recall are included. By examining person-related and job-related characteristics in a sequential fashion we are able to examine factors that predict UI eligibility status when restricted groups of factors are examined.

In a final step, these two types of information are combined and the regional unemployment rate is added. The regional unemployment rate controls for the variable entry requirement of the UI system so that our provincial variables are not capturing this aspect of the eligibility criteria. Tests of constant coefficients before and after C-17 reject the hypothesis of no change for each of the equations. This means that C-17 did change the probability of being eligible conditional on observable socio-demographic factors.

In the first two columns of Table 3, changes in coefficients between the 1993 and 1995 samples are observed for the variables capturing disabled status, the language of the interview and some of the province of residence variables. With job-related variables only, some large changes in coefficients between

1993 and 1995 are noted for variables such as unionization of the lost job and the seasonal nature of the job. When only job-specific information is used, otherwise identical persons losing a seasonal job are less likely to be eligible for benefits than those losing a non-seasonal job but this effect is greater in 1993 than in 1995. The second part of this result may seem surprising at first since a consequence of C-17 was that eligibility was tightened most significantly for persons losing a job with relatively low tenure and seasonal jobs are typically associated with low tenure and weak attachment to the labour force. This apparent paradox is resolved once additional information is added to the eligibility equation.

Combining job-specific and person-specific variables and adding regional unemployment rates makes the seasonal variable more negative for the 1995 sample but less negative for the 1993 sample, as we would expect. This change likely reflects the importance of the regional unemployment rate variable. This variable had a negative impact on eligibility in 1993 but a positive effect in 1995. A positive effect of the regional unemployment rate variable seems sensible given that fewer insurable weeks are needed to be eligible in high unemployment regions. The negative impact prior to C-17 is more difficult to interpret but may be due to other impacts of regional unemployment such as demand-side constraints upon weeks worked. Under C-17 these effects may have become relatively less important.

Overall, there is evidence that C-17 did reduce UI eligibility status for certain groups such as persons losing seasonal jobs. Controlling for seasonal status, persons losing jobs in high unemployment regions and provinces were relatively more likely to be eligible under C-17. The estimated eligibility effects of this section are quantified further when the impact of C-17 on the financial liability of the UI account is examined in section 8 of this report.

C. Determinants of the probability of claiming benefits

Once the effects of C-17 on eligibility have been examined, it is possible to see how take-up behaviour changed conditional on eligibility. This is examined in Table 4 which repeats the sequence of columns from Table 3 and adds a fourth set of variables designed to measure the attractiveness of UI: weeks of maximum benefit entitlement (both alone and squared to capture non-linear effects) and the ratio of UI benefits to lost wages (the UI replacement rate).

Statistical tests of constant coefficients were again conducted for the four types of equation specifications. Results of these tests are provided in Table 5 which shows that we rejected the formal hypothesis that coefficients were the

Table 3
The Probability of Being Eligible for UI Benefits
Probit analysis of the probability of being eligible for UI Benefits
(standard errors in parentheses)

	Coep 93	Coep 95	Coep 93	Coep 95	Coep 93	Coep 95
Age	.007	.005			.007	.004
	(.002)	(.002)			(.002)	(.002)
Married	.125	.128			.154	.116
	(.049)	(.043)			(.048)	(.045)
Minority	.037	.046			.004	.042
	(.063)	(.051)			(.062)	(.054)
Disabled	.262	-.278			.270	-.264
	(.218)	(.073)			(.235)	(.078)
Sex	.068	.050			.069	.080
	(.046)	(.041)			(.046)	(.044)
Interview in English	.077	-.055			.078	-.020
	(.105)	(.091)			(.105)	(.097)
Schooling: *						
Elementary or less	-.258	.106			.336	.212
	(.115)	(.108)			(.128)	(.117)
Some Secondary	-.176	-.045			-.233	.020
	(.085)	(.069)			(.097)	(.074)
High School Degree	-.018	.088			-.028	.091
	(.082)	(.066)			(.094)	(.069)
Some College	-.015	-.038			-.099	-.084
	(.104)	(.092)			(.114)	(.096)
College Degree	-.093	.082			-.099	.060
	(.099)	(.078)			(.112)	(.083)
Some University	-.176	-.342			-.190	-.294
	(.118)	(.089)			(.132)	(.095)
Other Training	-.095	.017			-.072	-.009
	(.135)	(.125)			(.125)	(.131)
Province: **						
Newfoundland	-1.034	.284			-.862	-.117
	(.135)	(.160)			(.154)	(.182)
PEI	-.790	.433			-.710	.198
	(.211)	(.371)			(.224)	(.372)
Nova Scotia	-.225	.182			-.150	.143
	(.142)	(.137)			(.156)	(.148)
New Brunswick	-.544	.346			-.564	.368
	(.118)	(.133)			(.126)	(.144)
Ontario	-.049	-.147			-.131	-.025
	(.099)	(.093)			(.108)	(.100)
Manitoba	-.139	-.525			-.219	-.428
	(.153)	(.140)			(.167)	(.148)
Saskatchewan	.039	-.296			-.076	-.103
	(.174)	(.148)			(.192)	(.160)
Alberta	-.122	-.164			-.247	.007
	(.116)	(.110)			(.128)	(.118)
British Columbia	-.222	-.181			-.262	-.065
	(.111)	(.108)			(.123)	(.115)
N.W.T. and Yukon	.133	-.239			.397	-.935
	(.438)	(.391)			(.459)	(.423)
Wage Lost			.009	-.001	.005	-.000
			(.002)	(.001)	(.003)	(.001)
Job Lost Unionized			.295	.093	.152	.036
			(.051)	(.043)	(.056)	(.047)
Received Notice			.163	.024	.123	.007
			(.050)	(.049)	(.056)	(.052)
Had recall Date			.361	.111	.280	.144
			(.082)	(.049)	(.089)	(.053)
Job Lost Seasonal			-.393	-.272	-.294	-.355
			(.046)	(.042)	(.055)	(.045)
Regional Unemployment Rate					-.034	.053
					(.008)	(.009)
Pseudo R-Squared	.038	.029	.025	.010	.067	.048
Number of Observations	9,774	6,043	11,109	5,969	8,977	5,800

* Missing: University degree

** Missing: Quebec

same in 1993 and 1995 for all but one of the cases examined. Non-rejection happened for the specification with just personal variables and where the intercepts vary between 1993 and 1995 but all other coefficients remained the same. In other words, when only personal characteristics are included it is not possible to reject the hypothesis that there was simply a constant shift of the take-up probability. This would mean that the change in take-up probability could not be attributed to changes in behaviour of any identifiable group but rather to the same change regardless of demographic characteristics.

It is not useful to compare in detail the coefficients of the first two columns of Table 4 because we cannot reject the hypothesis that the coefficients only differ due to random sampling error. The third and fourth columns show large changes between 1993 and 1995 for almost all job-related variables. The fifth and sixth columns of this table look at UI program variables and is thus important for the goals of this study. The negative effect of the squared benefit weeks variable means that while claim probabilities initially increase with the number of weeks of benefits this happens at a progressively slower rate as the number of weeks increases. In fact, extrapolating this theoretical relationship even gives a point at which this relationship becomes negative so that the graph of the weeks/probability of claiming relationship has an inverted U-shape. In 1993, the highest point on this graph (just before the slope turns negative) would be at 80 weeks while this happens at just 45 weeks in 1995[5]. Although these numbers of weeks are not themselves relevant because they involve considerable extrapolation, the figures do imply that the impact of benefit weeks on claim probabilities declines faster in 1995.

When all variables are included in the take-up rate equation, the C-17 changes to the effect of firm variables is still apparent. It continues to be the case that take-up rates are now higher for those losing seasonal jobs whereas before C-17 they were lower. This is a puzzling result that is nevertheless very significant because it suggests that, other things equal, C-17 increased the likelihood that UI-eligible low-attachment seasonal workers actually claimed UI benefits. The reason for this finding is not immediately apparent. It is important to recognize, though, that weeks of benefit entitlement are entered separately in the analysis so that reductions in entitlements that were more significant for seasonal versus non-seasonal workers would be captured by the entitlement variable directly.

[5] These figures are obtained by noting that in a quadratic relationship of the form $y = bx - cx^2$, the maximum value of y occurs for $x = b/2c$.

Table 4
The Probability of Claiming UI Benefits
Probit analysis of the probability of claiming UI Benefits when eligible
(standard errors in parentheses)

	Coep 93	Coep 95	Coep 93	Coep 95	Coep 93	Coep 95	Coep 93	Coep 95
Age	.009	.006					.009	.005
	(.002)	(.002)					(.002)	(.002)
Married	.035	-.044					.060	-.066
	(.038)	(.043)					(.046)	(.044)
Minority	-.141	.006					-.002	-.004
	(.048)	(.050)					(.057)	(.051)
Disabled	.118	.067					.053	.048
	(.160)	(.085)					(.184)	(.086)
Sex	-.110	-.203					-.087	-.217
	(.036)	(.041)					(.044)	(.043)
Interview in English	-.064	-.129					-.138	-.143
	(.089)	(.088)					(.105)	(.090)
Schooling: *								
Elementary or less	.104	.117					-.147	.114
	(.102)	(.102)					(.123)	(.106)
Some Secondary	.125	.284					-.051	.257
	(.069)	(.070)					(.087)	(.072)
High School Degree	.097	.103					-.059	.087
	(.063)	(.064)					(.081)	(.066)
Some College	.063	-.021					-.027	-.039
	(.083)	(.092)					(.102)	(.094)
College Degree	.138	.112					.022	.107
	(.080)	(.076)					(.101)	(.078)
Some University	.012	-.069					-.009	-.070
	(.099)	(.097)					(.125)	(.099)
Other Training	.238	.166					.129	.194
	(.095)	(.125)					(.115)	(.128)
Province: **								
Newfoundland	-.065	.213					.014	232
	(.164)	(.147)					(.192)	(.165)
PEI	.029	-.330					-.126	-.337
	(.272)	(.263)					(.293)	(.267)
Nova Scotia	.078	.064					.181	.079
	(.136)	(.127)					(.158)	(.132)
New Brunswick	-.073	.091					-.083	.123
	(.127)	(.112)					(.142)	(.117)
Ontario	-.152	-.157					-.003	-.121
	(.088)	(.091)					(.104)	(.095)
Manitoba	-.205	-.220					.063	-.236
	(.131)	(.158)					(.160)	(.163)
Saskatchewan	-.198	-.406					-.053	-.350
	(.138)	(.151)					(.163)	(.157)
Alberta	-.314	-.143					-.143	-.117
	(.100)	(.108)					(.119)	(.113)
British Columbia	-.136	-.130					.051	-.071
	(.100)	(.107)					(.119)	(.110)
N.W.T. and Yukon	-.028	-.932					.007	-.811
	(.327)	(.386)					(.353)	(.416)
Wage Lost			-.007	-.003			-.004	-.002
			(.002)	(.001)			(.002)	(.001)
Job Lost Unionized			.433	.104			-.008	.113
			(.040)	(.043)			(.049)	(.045)
Received Notice			.500	-.026			.190	-.025
			(.045)	(.048)			(.054)	(.049)
Had recall Date			.614	.010			.405	-.027
			(.078)	(.048)			(.090)	(.050)
Job Lost Seasonal			-.093	.138			-.143	.141
			(.043)	(.044)			(.055)	(.048)
Benefit Entitlement					.048	.027	.040	.029
(Weeks)					(.0043	(.007)	(.005)	(.007)
Benefit Entitlement					-.0003	-.0003	-.0002	-.0003
Weeks Squared					(.000)	(.000)	(.000)	(.000)
UI Rate (1)					.174	.124	.148	.107
					(.011)	(.051)	(.014)	(.053)
Regional Unemployment					-.008	.020	-.011	-.003
Rate					(.005)	(.006)	(.008)	(.008)
Pseudo R-Squared	.017	.024	.045	.004	.076	.009	.093	.032
Number of Observations	9,169	5,074	10,319	5,013	10,134	5,059	7,949	4,957

* Missing: University degree
** Missing: Quebec

(1) UI Replacement Rate (excluding any cap on earnings). When workers did not applied, the default replacement rate at the time of the job loss was used.

Table 5
Formal Test of Behavioural Change Between COEP 93 and COEP 95
(Standard Errors in Parentheses)

	Equation		
Eligibility Equation:	**1**	**2**	**3**
No Restrictions:			
Log. Likelihood	-4767.1	-5434.2	-4153.7
Same Coefficient, Different Int.			
Value of Coep95 Dummy	-.500 (.031)	-.516 (.028)	-.504 (.036)
Log. Likelihood	-4851.7	-5448.7	-4276.4
Prob > Chi2	0	.001	0
Same Coefficient, Same Int.			
Log. Likelihood	-4981.3	-5619.4	-4376.2
Prob > Chi2	0	0	0

	Equation			
Take-up Equation:	**1**	**2**	**3**	**4**
No Restrictions:	-5766.8	-7377.8	-6444.6	-4727.0
Log. Likelihood				
Same Coefficient, Different Int.				
Value of Coep95 Dummy	-.466 (.029)	-.227 (.026)	.304 (.042)	.032 (.054)
Log. Likelihood	-5781.3	-7461.5	-6498.6	-4798.5
Prob > Chi2	.074	0	0	0
Same Coefficient, Same Int.				
Log. Likelihood	-5908.1	-7500.1	-6525.4	-4798.6
Prob > Chi2	0	0	0	0

In this table, we use likelihood ratio tests which look at the decrease in the ability of the model to explain the data when certain algebraic restrictions are added. This measured by the magnitude of the decline in the "Log Likelihood" value (show in the table) when restrictions are imposed. Statistical methods ar eused to determine whether the losses in explanatory power associated with these restrictions are large enough to be "significant". If losses are not significant, we cannot reject the restrictions. Following accepted practice in econometrics, we reject restrictions when the value for "Prob > Chi2" is smaller than 0.05.

D. The effect of UI on the timing of claims

Few variables have a significant effect upon the length of time between job loss and the filing of a UI claim.

In this section, a proportional hazards analysis of the duration of the time elapsed between lay-off and filing a UI claim is conducted. Proportional hazards analysis allows us to look at how the probability of claiming benefits changes as the length of period of joblessness increases. We assume that there is an underlying "baseline" pattern to the way that the claim probability varies with time and then allow for observable characteristics to move this pattern around in a proportional way. In this analysis, positive coefficients increase the probability of claiming at each point in time[6]. This shortens the expected wait before claiming since it is more likely that a person claims early.

This analysis uses the 1995 COEP data for which more specific information regarding delays is available than was the case for 1993. The two columns of Table 6 conduct this analysis both without and with persons who do not claim. Non-claimants are treated as censored observations since failure to claim does not preclude an eventual claim after a lengthy delay. It may seem natural to only use claimants since we do not know how long non-claimants may wait or even whether they ever will claim. It is also possible to use information from claimants and non-claimants by using estimation techniques developed for situations with "censored" observations. The two sets of results are both provided to see if conclusions are sensitive to the presence or absence of non-claimants.

Few variables have a significant effect upon the length of the time between job loss and the filing of a UI claim. Sex has a weak effect with women having a negative coefficient. A positive effect for seasonal workers is found only when non-claimants are included in the proportional hazards analysis. It could be that seasonal workers move more quickly than other workers to claim UI benefits. This suggests that among those who do claim benefits, seasonal workers do not differ in the timing of their UI claims. It is only once eligible non-claimants are added that seasonal workers are seen to behave differently.

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[6] More correctly, it is the *conditional* probability of claiming after a given time given that the person has not claimed earlier, known as the "hazard rate", that increases.

Table 6
The Timing of UI Claims
Cox analysis of the time between layoff and UI claim by Eligible Unemployed
(standard errors in parentheses)

	Excluding Non Claimants		Including Non Claimants	
Age	-0.001	(0.002)	0.003	(0.002)
Married	0.026	(0.036)	-0.028	(0.036)
Minority	-0.013	(0.042)	-0.011	(0.042)
Disabled	-0.004	(0.069)	0.019	(0.069)
Sex	-0.082	(0.035)	-0.131	(0.034)
Interview in English	-0.037	(0.073)	-0.097	(0.072)
Schooling: *				
Some Secondary	-0.011	(0.085)	0.077	(0.085)
High School Degree	0.024	(0.060)	0.160	(0.059)
Some College	-0.020	(0.056)	0.060	(0.056)
College Degree	-0.058	(0.081)	-0.024	(0.081)
Some University	-0.062	(0.066)	0.069	(0.066)
Other Training	-0.062	(0.087)	-0.053	(0.087)
Province: **				
PEI	0.182	(0.243)	-0.165	(0.242)
Nova Scotia	0.010	(0.105)	0.044	(0.103)
New Brunswick	0.062	(0.086)	0.097	(0.086)
Ontario	-0.007	(0.078)	-0.067	(0.076)
Manitoba	0.000	(0.142)	-0.135	(0.141)
Saskatchewan	-0.029	(0.143)	-0.248	(0.142)
Alberta	-0.111	(0.094)	-0.097	(0.093)
British Columbia	-0.007	(0.091)	-0.048	(0.090)
N.W.T. and Yukon	-0.261	(0.512)	-0.703	(0.510)
Wage Lost	0.000	(0.001)	-0.002	(0.001)
Job Lost Unionized	-0.006	(0.037)	0.081	(0.036)
Received Notice	0.058	(0.040)	0.005	(0.040)
Had recall Date	0.075	(0.041)	0.010	(0.041)
Job Lost Seasonal	-0.009	(0.038)	0.091	(0.038)
Weeks of Benefit Entitlement	0.004	(0.002)	0.006	(0.001)
UI Rate (1)	0.019	(0.038)	0.034	(0.038)
Regional Unemployment Rate	-0.013	(0.006)	-0.007	(0.006)

* Missing: University degree

** Missing: Quebec

7. Quantifying the Effects of C-17 on Eligibility and Benefit Take-Up

In this section, we use the behavioral relations of the previous section to analyze the following "what if" questions:

- how would eligibility for benefits have changed if the post-C-17 (1995) link between personal/job characteristics had applied in 1993.

- how would take-up of benefits have been different in 1993 if the 1995 link between observable characteristics and take-up had applied in 1993.

These issues are examined in Table 7 which presents results for the four possible combinations of 1993 and 1995 UI rules and 1993 and 1995 samples of unemployed persons. The first column uses 1993 COEP characteristics and looks at the probability of being eligible for UI using 1993 and 1995 rules. Interestingly, eligibility probabilities change very little with either set of rules and this is true for the weighted and unweighted samples. Weeks of entitlement for benefits fall when 1995 rules replace those in place in 1993 but the effect is relatively small. As a consequence of these changes, take-up rates fall but not by a great deal.

For the 1995 COEP sample, the 1995 reforms again have little effect upon the probability of being eligible for UI. The largest impact of policy in this table is observed when 1995 benefit entitlement rules replace the 1993 rules and the characteristics of the 1995 sample are used. Despite this change in weeks of benefits, however, the take-up probability effect is small for this sample.

Taken together, these results have some interesting implications. The take-up results show that while there may have been a small decrease in the perceived adequacy of UI benefits, take-up was not greatly affected by UI reform. This is despite reductions in the length of maximum benefit entitlement periods that were in some cases quite large. One interpretation of this could be that large changes in benefit entitlements did not lower take-up probabilities to a greater extent because of off-setting changes in other factors. What these other factors are is not immediately clear. A somewhat opposed interpretation is that claimants were receiving more than was needed to make benefits attractive. Of course, this is to be expected since we would like benefits to more than compensate for the costs of claiming.

This means that C-17 may have achieved its goals ... relatively few unemployed persons were rendered ineligible for benefits but those qualifying could no longer finance an extended period of joblessness following a short period of employment.

Another curious results is that C-17 reforms apparently had a greater impact on benefit entitlement for eligible persons than on the probability that a given person is eligible for benefits. This may be surprising in light of the fact that C-17 reforms were in large part designed to exclude low-attachment workers from benefit eligibility. This result could mean that most workers were able to increase employment lengths prior to job loss so as to preserve benefit eligibility, a conjecture supported by work of authors such as Christofides and McKenna (1996) who have documented previous examples of employment durations being tailored to meet the eligibility requirements of the UI system. This view may be further supported by the observation that weeks of benefit entitlement fell for many eligible persons, suggesting that a significant number of the eligible were minimally qualified. This means that C-17 may have achieved its goals in what is arguably the best possible way - relatively few unemployed persons were rendered ineligible for benefits but those qualifying could no longer finance an extended period of joblessness following a short period of employment.

Table 7
The Impact of C-17 Changes in Rules on Eligibility and Take-up
(Standard errors in parentheses)

	Coep 93		Coep 95	
	Not Weighted	Weighted	Not Weighted	Weighted
Probability of Being Eligible				
1993 rules	.927	.879	.829	.823
1995 rules	.926	.878	.828	.822
Weeks of Benefit Entitlement				
1993 rules	40.27	35.65	34.59	34.12
1995 rules	34.67	32.14	25.69	24.44
Mean Predicted Take-up				
1993 rules	.872		.789	
	(.100)		(.069)	
1995 rules	.869		.771	
	(.093)		(.078)	

8. The Effect of C-17 on the Financial Status of the UI Account

An important goal of this research is to provide information regarding the effect of changes in UI due to C-17 and the state of the UI account. This goal is addressed in Table 8 which looks at amounts of benefit dollars potentially payable under two hypothetical scenarios as well as the amounts actually paid. We look at maximum potential pay-out (benefits payable if all eligible persons use their maximum benefit entitlement), maximum claimed pay-out (benefits payable to those who actually claim benefits if they use their maximum benefit entitlement) as well as actual dollars paid[7].

... a large fraction of this potential dollar reduction is due to shorter benefit entitlement periods ...

The first column of Table 8 indicates the dollar amounts implied by the 1995 C-17 rules. When we apply C-17 rules to the characteristics of the 1995 COEP sample we find that the job losers of the COEP 1995 sample generated a maximum potential pay-out for the UI account of $46,366,613[8] . This potential pay-out exceeds the benefits actually paid to this group for two reasons: some eligible workers do not claim benefits at all and some of those who do claim will find a job before the end of their benefit entitlement period.

The second column of this table shows that the maximum potential pay-out would have been $64,535,398 had the pre-C-17 rules been applied to a group with the same demographic characteristics as the 1995 COEP sample. The amount obtained using the C-17 rules thus represents a 28.2 percent fall in maximum potential pay-outs due to the following factors: reduced eligibility, lower replacement rates, a drop in "cycling" behaviour, shorter benefit entitlement periods as well as interactions between these effects. A breakdown of the dollar reduction into these categories shows that, as our econometric analyses of eligibility probabilities suggested, a large fraction (61 percent) of this potential dollar reduction is due to shorter benefit entitlement periods, the drop in replacement rates accounts for seven percent of the fall in the maximum potential pay-out and reduced eligibility explains less than one percent of the dollars potentially saved.

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7 These measures are discussed in greater detail in section 4 of the report.

8 This liability is for the 1995 COEP sample rather than the entire Canadian economy. To extrapolate to the unemployed as a whole it would be necessary to multiply by sample weights. A more straightforward approach is to simply discuss changes in terms of percentage as well as dollar effects and this method is used here.

The second row of Table 8 examines the effect of benefit take-up on the potential liability of the COEP 1995 sample for the UI account. In the first column, actual reported take-up of benefits for the 1995 sample is used to calculate a maximum claimed pay-out for those unemployed persons who actually claimed benefits. The failure of some individuals to claim benefits produces reduction in the amount potentially payable of $10,586,221 (a 22.8 percent drop). Of course, the actual savings due to non-claimants is less than this, particularly since eligible non-claimants tend to have shorter unemployment durations than those who do claim benefits.

To calculate an analogous maximum claimed pay-out for the 1995 COEP sample using the hypothetical scenario that the 1993 rules were still in place, the maximum claimed benefits using 1995 rules was adjusted by multiplying by 1.023. This adjustment reflects the fact that our empirical analysis of take-up probabilities showed that the more generous pre-C-17 system produced a take-up probability that was, on average, 2.3 percent higher than with C-17[9]. Accordingly, with 1995 rules rather than 1993 rules, the maximum amount of benefits potentially payable to claimants falls even more than did the total potential pay-out to the eligible (by 29.4 percent for pay-outs to claimants versus 28.2 percent for the eligible) because under C-17 a greater proportion of eligible persons did not claim their benefits.

The final row of Table 8 examines dollars that would actually be paid given the length of unemployment spells in the 1995 COEP. The dollar amount paid is $20,864,655, slightly over half of the maximum amount that claimants were entitled to receive. This reflects the fact that some people find new jobs quickly while others use up their entire benefit entitlement. It is interesting that while dollars paid are 58 percent of dollars claimed under the C-17 rules, when the pre-C-17 rules are applied this falls to 52 percent. In other words, for a large number of persons neither the C-17 nor the pre-C-17 limit on benefit weeks payable was a binding constraint because they found a job quickly. On the other hand, enough people did exhaust benefits that there was still a 20.7 percent fall in benefits paid due to C-17[10]. The fact that many people find jobs before exhausting benefits means that while C-17 had the potential to lower pay-outs to claimants by 29.4 percent if all claimants exhausted their benefits, the actual drop in dollars paid was slightly lower at 20.7 percent.

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[9] This adjustment factor is obtained from Table Six. The mean predicted take-up rates for the 1995 COEP sample are 0.789 with the pre-C-17 rules and 0.771 with C-17. An increase from 0.771 to 0.789 reflects a 2.3 percent change.

[10] This COEP-based analysis of changes in benefits paid is consistent with administrative figures for benefits paid. Between June 1994 and June 1995, total UI benefits paid went down by 19.4 percent.

Table 8
The Impact of C-17 Reforms on the Financial State of the UI Account
Estimations Based on Survey Data - COEP 95

	1995 Rules	1993 Rules	% Difference
Maximum Potential Pay-out (1)	46,366,613	64,535,398	-28.2
Difference due to:			
Reduction in Replacement Rate		**-1,330,763**	
Reduction in Eligibility		-44,622	
Reduction in Number of Cyclers (2)		-5,712,726	
Remainder: Reduction in weeks of benefits plus cross effects		-11,080,674	
Maximum Claimed Pay-out (3)	35,780,392	50,696,231	-29.4
Dollars Actually Paid (4)	20,864,655	26,303,737	-20.7
Difference due to:			
Reduction in Number of Cyclers		-605,509	-18.8
Reduction in Eligibility		-6,532	
Reduction in Replacement Rate		-581,966	
Increase in Number of Exhaustions		-1,783,459	
Remainder (decrease in weeks of exhaustees and cross effects)		-2,461,616	

(1) Estimated maximum number of weeks payable x estimated average weekly payment (Maximum: $448). This figures represents the potential benefit pay-out if all eligible persons claim benefits through to the end of their benefit entitlement period.

(2) The authors define cyclers as those individuals who file a new claim as soon as the one they either filed or was ongoing at the time of their job loss is exhausted.

(3) The amount of the potential maximum pay-out once take-up rates are applied. To account for behavioral changes following the changes in rules, the amount actually claimed under 93 rules is multiplied by 1.023 to account for the increase in take up which would resulted from the increased generosity.

(4) Here, weeks actually paid are the minimum of weeks payable or actual unemployment duration for persons who did not exhaust their benefits.

9. Conclusions

This study has found some evidence that the reforms of Bill C-17 provoked some small decline in the probability that workers eligible for UI benefits actually go on to claim these benefits. The raw data, adjusted only for sampling weights, show a fall from 80 percent to 75 percent in the self-reported claim rate for eligible persons. Statistical analysis shows that other factors such as a fall in unemployment rates also explain part of the change so the raw figure above is an upper bound on the effect from C-17. Larger changes were observed for weeks of benefit entitlement due to C-17, however. These changes mean that the UI account experienced a significant reduction in its financial liability as a result of the changes to the structure of benefit entitlement periods and eligibility criteria. Interestingly, there was not a large reduction in the probability of eligibility. Rather, C-17 saved money primarily through decreases in maximum durations of benefits.

These changes mean that the UI account experienced a significant reduction in its financial liability as a result of the changes to the structure of benefit entitlement periods and eligibility criteria.

An unanswered question of this research is what were the consequences of unchanged eligibility rates combined with shorter benefit entitlement periods. One possibility is that some unemployed workers were eventually re-employed in jobs other than their traditional seasonal job. Also, it is possible that some workers suffered drops in income after their first post-C-17 period of unemployment. It is only after this period that we might see movement of workers to other forms of income support such as welfare. These questions merit further analysis.

The study found that C-17 had a large effect upon dollars paid from the UI account. We were able to hold constant the economic environment and look at how many dollars the 1995 COEP sample would have received under the 1993 rules. We find that C-17 accounted for a 20.7 percent drop in benefits paid, mostly because of shorter benefit entitlement periods. It is worth stressing that none of this fall is due to improved economic conditions since it holds constant the characteristics of the 1995 COEP.

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