WORKING PAPER #212
INDUSTRIAL RELATIONS SECTION
PRINCETON UNIVERSITY
AUGUST 1986

First Draft: November 1984
Second Draft: March 1985
Third Draft: July 1986

UNEMPLOYMENT INSURANCE AND MALE UNEMPLOYMENT DURATION IN CANADA

by

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*The authors are indebted to John Abowd, Joseph Altonji, Joshua Angrist, Christopher Flinn, Morley Gunderson, George Jakubson, John Micklewright, Angelo Melino, Geert Ridder, Daniel Sichel, Gary Solon, Jim Symons, Robert Topel and an anonymous referee for helpful comments and discussions. We are also grateful to seminar participants at Cornell, Princeton, University College (London), and the NBER's Summer Workshop on Economic Fluctuations. We are responsible for any errors. A substantial portion of this research was carried out while Ham was a visitor in the Department of Economics and Industrial Relations Section, Princeton University. He would like to thank both for their hospitality. Research support from Employment and Immigration, Canada and the Industrial Relations Section is gratefully acknowledged.

ABSTRACT

A model of unemployment duration is estimated with weekly micro data on a sample of Canadian men during the 1975 through 1980 period. Entitlement provisions in the unemployment insurance program and demand conditions are found to have a significant impact on the probability of leaving unemployment. The probability of a worker leaving unemployment declines with duration of unemployment, holding unemployment insurance entitlement constant. When entitlement is allowed to vary, the probability of leaving first falls and then generally rises with unemployment duration as the declining entitlement induces a greater willingness to accept offers or search more intensively. These results are robust to alternative specifications of duration dependence and to allowing for person-specific unobserved heterogeneity.

1. Introduction

Over the past decade there has been considerable interest in the impact of unemployment insurance on unemployment duration. In Canada the introduction of a more generous unemployment insurance system in 1971 led to a great deal of research, but this research was largely based on aggregate data.¹ The changes in the unemployment insurance system that took place in 1976, 1977 and 1979 have been analyzed in only one study.² More recent research in the United States and the United Kingdom has been based on microeconomic data, which allows researchers to estimate the hazard rate, the probability of exiting unemployment conditional on the length of the current unemployment spell. This study estimates the hazard rate using weekly micro data on a sample of Canadian men for 1975 through 1980. The results indicate the effect of unemployment insurance on expected unemployment duration and the extent of duration dependence in the hazard rate.

The paper proceeds as follows. In Section 2 we outline the economic model of unemployment duration, and in Section 3 we describe our econometric approach for estimating the model. In Section 4 we describe the data set that we use, the basic specification of the independent variables and some data problems that we encounter. Our empirical results are presented in Section 5.

In the estimation we investigate the effect of: i) changing the form of the duration dependence; ii) allowing for lagged duration dependence and iii) introducing unobserved heterogeneity. We find that entitlement provisions in the unemployment insurance system and demand conditions have a significant impact on the probability of leaving unemployment.³ We also find that holding entitlement constant,

the hazard function falls with unemployment duration over most of the relevant range. However, when one includes the effect of changing entitlement, the hazard first falls and then generally rises with duration. We are unable to measure the impact of unemployment benefits given the lack of independent variation in this variable.

2. Economic Model

Canada's unemployment insurance program is a key part of the social insurance system, and like any insurance program it can potentially influence the behaviour of those who are insured. There have been numerous statutory changes in the program since 1975, but two areas of change were most likely to have influenced the duration of unemployment.

1) Benefit Rate Reductions: The benefit rate for those with dependents was lowered from 75 percent to 67 percent of insurable earnings, effective January 1976, and the benefit rate for all claimants was lowered from 67 percent to 60 percent of insurable earnings, effective January 1979.

2) Weeks of Entitlement: The number of weeks of benefit entitlement at the start of a claim is based on previous weeks worked and the aggregate unemployment rate. Prior to September 1977 the entitlement was based on the national unemployment rate and the difference between the national unemployment rate and the regional unemployment rate, as well as weeks worked. Beginning in September 1977, the entitlement was determined by weeks worked and the absolute level of the regional unemployment rate. These changes reduced the average period of entitlement, particularly in low unemployment rate regions.

Economic theory suggests that these changes reduced the number of people claiming benefits and the average duration of unemployment. The weekly nature of the data allow us to analyze the impact of unemployment insurance in the context of a job search model.⁴ We assume that an individual faces a distribution of potential job offers. If an offer is received, he must accept or reject the offer. Once rejected, an offer disappears. The offer will be accepted if the wage equals or exceeds the individual's reservation wage, chosen such that the expected utility from accepting the offer is equal to the expected utility of remaining unemployed. The latter includes any additional leisure associated with unemployment and any unemployment insurance benefits. Unemployment insurance benefits also raise the value of employment because any subsequent unemployment is compensated. This fact led Mortensen (1977) to suggest that an increase in benefits lowers the probability of leaving unemployment at the early stages of unemployment but raises the probability near the end of entitlement period.

Since we have weekly data, we can allow the worker's reservation wage and arrival rate of job offers to vary from week to week. Consider an individual (denoted i) whose unemployment spell began at calendar time t_{oi} , and who has been unemployed for $t-1$ weeks. If each week he receives exactly one job offer (distributed independently over time), then the probability of his finding employment during the current week, conditional on being unemployed for the previous $t-1$ weeks, equals

$$(1) \quad \lambda_i(t_{oi}, t) = 1 - F(w_i^*(t_{oi} + t))$$

where λ_i is the hazard rate, $w_i^*(t_{oi} + t)$ is the worker's reservation wage in week t of the spell and $F(\cdot)$ is the distribution function

for wage offers. (Assume for now that the worker cannot be recalled to his previous job.) In the more general case, the rate of job offers to the individual is a random variable $\alpha_i(t_{oi} + t)$ and the hazard function will depend on both the rate of job arrivals and the reservation wage

$$(2) \quad \lambda_i(t_{oi}, t) = \lambda(\alpha_i(t_{oi} + t), w_i^*(t_{oi} + t)) .$$

Given our data, we follow much of the literature and estimate a reduced form version of (2). We assume that the rate of job arrivals and the reservation wage are a function of demographic variables, demand conditions, unemployment insurance parameters and the current duration of unemployment. (Unemployment insurance parameters may affect the rate of job arrivals through search intensity.) Given a particular functional form for the hazard in (2), one can then derive expressions for the density function of a completed spell and the expected duration of unemployment. These are given below in equations (7) and (10), respectively.

An important issue in studies of unemployment duration is that of duration dependence: how does the probability of leaving unemployment change with the current length of the spell? We distinguish between two sources of duration dependence in the data. The first of these arises because a worker's entitlement (the number of weeks of benefits remaining to the worker) falls as the duration of the spell increases. A worker's entitlement is expected to have an important influence on the unemployed worker's willingness to accept jobs or to search intensively. During the initial phases of unemployment the recipient has a cushion of

unemployment benefits should the search take longer than expected. As his entitlement decreases, the cost of rejecting a current offer is higher because of the higher probability that there will be a period of uncompensated unemployment. Consequently, the claimant will lower his reservation wage as duration increases and his entitlement decreases, and the probability of leaving unemployment will increase as entitlement falls. This effect is analyzed by Mortenson (1977). Since we can calculate a worker's current entitlement it is possible to measure the duration dependence effect arising from declining entitlement, and we will refer to this as the entitlement effect.

The second source of duration dependence that we measure is the net effect of a number of individually unobservable factors which are likely to lead to the exit probability depending on the duration of the spell, holding entitlement constant. For example, unemployed individuals may adjust downwards their perception of the wage offer distribution when they fail to receive offers above their acceptance wage. Further, assets may decline and the value of leisure may fall as unemployment persists.⁵ On the other hand, extended unemployment may signal to prospective employers that the job applicant is inferior to those with less unemployment, and thus the rate of job arrivals may fall. We will refer to the net impact of these unobserved factors on the hazard function as the direct duration dependence effect. The total duration dependence effect in the data will be the combination of the direct duration effect and the entitlement effect.

Of course, the interpretation of the duration dependence parameters is complicated considerably once one acknowledges the possibility

of unobserved differences across workers. Substantial theoretical results have been obtained on the issue of distinguishing heterogeneity effects from duration dependence effects.⁶ However, distinguishing between these effects empirically can be difficult. We should note that the entitlement effect, because it is based on observable data and is identified from variation in initial entitlement, should be less sensitive to the treatment of unobserved heterogeneity.

3. Econometric Approach

3.1 Specification of the Hazard Function

Over the last few years, economists investigating unemployment duration have increasingly turned to discrete and continuous time hazard models.⁷ In these models, the researcher parametrizes the hazard function. This function is the conditional probability of leaving unemployment at time t in the unemployment spell, given that the current unemployment spell has lasted until time t . The popularity of hazard models in economic research arises from at least three factors. First, these models can easily incorporate economic variables that change over time. Second, hazard models can incorporate incomplete (censored) unemployment spells. Third, the use of these models allows one to examine how the probability of leaving unemployment changes with the duration of the spell.

In the present study, we use a discrete time waiting model previously estimated by Nickell (1979) to analyze unemployment duration.⁸ For a worker whose unemployment spell begins at calendar time t_{0i} , we assume that the conditional probability of the individual leaving unemployment in week t of the spell, given that he has been unemployed for the previous $t-1$ weeks, takes the form

$$(3) \lambda_i(t_{oi}, t) = \{1 + \exp - [\theta + h(t) + z(s(t_{oi} + t)) + \delta t \cdot s(t_{oi} + t) + \beta' X_i(t_{oi} + t)]\}^{-1}.$$

In (3), θ is a constant, $h(t)$ represents the impact of duration on the hazard function, holding entitlement, $s(t_{oi} + t)$, constant, while the function $z(s)$ captures the effect of changing entitlement on the hazard function. Further, β is a vector of parameters, δ is the coefficient on the entitlement-duration interaction, and $X_i(t_{oi} + t)$ is a vector of additional variables (at calendar time $(t_{oi} + t)$) reflecting other provisions of the unemployment system, demographic variables and demand conditions. These variables change over time with different frequencies. Unemployment rates are monthly and annual, age changes every year, and unemployment insurance parameters can change in any week.

3.2 Specification of Duration Dependence and Entitlement Effects

The total measure of duration dependence in the hazard, $d\lambda/dt$, is determined by the sum of the direct duration dependence effect, $dh/dt + \delta s$, and the entitlement effect, $dz/dt = (dz/ds + \delta t) \cdot ds/dt$. We noted above that learning about the wage distribution, a declining value of leisure, and diminishing assets during the spell can lead to a positive direct effect, while a signalling effect that resulted in a declining rate of job offer arrivals could result in a negative direct effect. Thus the overall sign of the direct effect is ambiguous, and may well change over the unemployment spell. To allow for these possibilities we generally assume that $h(t)$ can be described by a sixth-order polynomial

$$(4) \quad h(t) = \sum_{i=1}^6 \gamma_i t^i .$$

(This is the highest order polynomial that the data can support in terms of producing relatively precise coefficients for the duration terms.) To investigate the robustness of our results we also specify $h(t)$ as a step function.

Since we have a stronger a priori expectation on the entitlement effect, we assume that $z(s)$ is quadratic

$$(5) \quad z(s) = \alpha_1 s + \alpha_2 s^2 .$$

Given that s depends on t , it is useful to explicitly discuss the identification of the entitlement parameters α_1 and α_2 .

Consider, in (3), the portion of the term in square brackets relating to t , t^2 , $z(s)$ and $t \cdot s$

$$(6) \quad k = \gamma_1 t + \gamma_2 t^2 + \alpha_1 s + \alpha_2 s^2 + \delta s \cdot t .$$

Noting that $s = IE - t$, where IE is entitlement at the beginning of the spell, rewrite (6) as

$$(6') \quad k = (\gamma_1 - \alpha_1)t + (\gamma_2 + \alpha_2 - \delta)t^2 + \alpha_1 IE + \alpha_2 IE^2 + (\delta - 2\alpha_2)IE \cdot t .$$

If we rewrite the hazard in terms of t and IE (instead of in terms of t and s), we will obtain exactly the same estimates of the entitlement parameters α_1 and α_2 . (The other coefficients in (6') reflect the total duration effect.) Thus, the identification of α_1 and α_2 comes from variation in IE (and IE^2), not from terms in t . In our data, IE varies with national and regional unemployment, an indi-

vidual's employment prior to his current unemployment claim, and previous unemployment spells covered by the current claim.

3.3 Likelihood Function and Expected Duration

To avoid the problem of a spell starting before the sampling period, as well as other initial conditions issues, we follow Heckman and Singer's (1984a) suggestion and consider only spells which begin after the start of the sampling period, January 1975. Given this sampling plan, there are two types of unemployment spells which must be considered in forming the likelihood function. For some spells, we observe that the spell ended in calendar week $t_{oi} + t_i^*$, and thus the duration of the spell is t_i^* weeks. In this case, the probability of the completed spell is

$$(7) \quad g_i(t_{oi}, t_i^*) = \left[\prod_{t=1}^{t_i^*-1} (1 - \lambda_i(t_{oi}, t)) \right] \lambda_i(t_{oi}, t_i^*) .$$

Thus the probability of a completed spell is the conditional probability of leaving in week t times the product of the conditional probabilities of not leaving in each of the previous t_i^*-1 weeks. For incomplete spells, the completed length of the spell is censored at $\bar{t}_i$, and we know only that the spell lasted longer than $\bar{t}_i$. In this case, the contribution to the likelihood function is the survivor function, or probability that the spell lasted more than $\bar{t}_i$ weeks,

$$(8) \quad (1 - G_i(t_{oi}, \bar{t}_i)) = \prod_{t=1}^{\bar{t}_i} (1 - \lambda_i(t_{oi}, t)) .$$

The likelihood function is then

$$(9) \quad L = \prod_{i \in C} g_i(t_{oi}, t_i^*) \prod_{i \in IN} (1 - G_i(\bar{t}_{oi}, t_i)) ,$$

where C denotes completed unemployment spells and IN denotes incomplete or censored spells. Parameter estimates are obtained by maximizing (9) with respect to θ , β , γ and α . Given these parameter estimates we can numerically calculate the expected duration of unemployment using

$$(10) \quad E(D) = \sum_{t=1}^{\infty} t g(t_o, t) ,$$

where the X variables are assumed to take their mean values in (10). The effect of changing the X variables on the expected duration of unemployment can be obtained by numerically differentiating (10).

3.4 Temporary Layoffs

To this point, we have ignored the possibility of the worker being recalled to his previous job. In the presence of temporary layoffs, it is desirable to distinguish the probability of leaving unemployment through recall from the probability of leaving through finding a new job. Katz (1986) provides an econometric framework for doing this given our data. Unfortunately, temporary layoffs comprise less than 20% of our completed spells, and we cannot precisely estimate the recall hazard given our specification. In Appendix 1 we report the results of estimating a parsimonious version of the new job and recall hazards using Katz's approach. Since the estimates of the parameters of the recall hazard are based on relatively few transitions, these results should be viewed with some caution. For the moment we simply note that the data

strongly reject the hypothesis that the parameters of the new job and recall hazards are equal.⁹ In the remainder of the paper, we focus solely on estimating the probability of leaving unemployment through finding a new job. In this estimation, spells that end in temporary recall are treated as censored with respect to the new job hazard and (8) is the contribution to the likelihood for such spells.

3.5 Heterogeneity

In forming (8), (9) and (10) we have assumed that there are no unobserved differences across workers. We now introduce unobserved heterogeneity. In the logit specification of the hazard (3), a natural means of doing this is through the constant term θ . We consider two separate specifications for the distribution of θ . First, we assume that θ is distributed independently across unemployment spells. Second, we make the more plausible assumption that θ is distributed independently across workers, but is fixed across spells for a given worker. In each case, we follow Heckman and Singer (1984a, 1984c), (see also Nickell (1979)), and assume that θ is drawn from a discrete distribution with points of support $\theta_1, \theta_2, \dots, \theta_K$ and associated probabilities $P_1, P_2, \dots, P_K$, respectively,¹⁰ where $P_K = 1 - \sum_{k=1}^{K-1} P_k$.

In the first case (where θ is independent across spells), the contribution to the likelihood for a completed spell is

$$(11a) \quad g_i^*(t_{oi}, t_i^*) = \sum_{k=1}^K P_k g_i(t_{oi}, t_i^*; \theta_k)$$

where

$$(11b) \quad g_i(t_{oi}, t_i^*; \theta_k) = \left[\prod_{t=1}^{t_i^*-1} (1 - \lambda_i(t_{oi}, t; \theta_k)) \right] \lambda_i(t_{oi}, t_i^*; \theta_k) .$$

The contribution of censored spells to the likelihood function is defined in a similar fashion.

To form the likelihood in the second case, (where θ is constant across an individual's spells), assume that we observe n_{1i} complete spells and n_{2i} incomplete spells for individual i . Define $g_{is}(\theta)$ as the density of the s^{th} completed spell conditional on θ , and $(1 - G_{is'}(\theta))$ as the survivor function for the s'^{th} censored spell conditional on θ . Define

$$(12a) \quad L_i(\theta_k) = \prod_{s=1}^{n_{1i}} g_{is}(\theta_k) \prod_{s'=1}^{n_{2i}} (1 - G_{is'}(\theta_k)) .$$

Then the contribution to the likelihood function for individual i is

$$(12b) \quad L_i^* = \sum_{k=1}^K P_k L_i(\theta_k) .$$

IV. Data and Empirical Specification

Our data set consists of a random sample of males drawn from the Canadian Employment and Immigration Longitudinal Labour Force File for the period January 1975 through December 1980. The data are weekly observations on the labor market status of workers based on administrative records. While this data set provides labor market histories at a fine level of time aggregation, it also has some of the disadvantages of data sets compiled from administrative records. In particular, in many cases it is not possible to determine the labour force status in

the weeks after unemployment insurance benefits are exhausted. For such workers the contribution to the likelihood function is given by (8). Similar missing data problems arise in many U.S. studies (e.g., Solon (1985)). We focus on spells that begin with insured unemployment.¹¹ In comparison to much of the U.S. data, the data set we use has the advantage that we can distinguish between different unemployment spells covered by the same claim (on the unemployment insurance system) but has the disadvantage that it contains less demographic information on individuals.

An important independent variable in any study of unemployment duration is the real level of benefits. There are several reasons to include also the worker's previous real weekly wage as an explanatory variable. First, the monetary cost of unemployment can be proxied by this wage minus benefits. Second, the previous wage can act as a proxy for the mean of the offer distribution. Finally, this variable can be used as a control for heterogeneity. Unfortunately, in our data, previous wages and benefits are closely linked. In Canada benefits are a constant fraction (the benefit ratio) of insurable earnings, which equal the previous weekly wage up to a maximum.¹² This maximum is known as the maximum insurable earnings. Wages above this maximum are not insured. Benefits for workers with wages above the maximum equal the benefit ratio times maximum insurable weekly earnings. Except for legislative changes in the benefit ratio, variation between benefits and wages comes from individuals with weekly wages above the maximum insurable earnings.

Weekly wages are available for all workers in 1975 and 1976.

Unfortunately, after 1976 only data on insurable earnings are available, and this has the effect of censoring wages for those with earnings above the maximum insurable level. Given the role that these workers play in identifying the respective coefficients on wages and benefits, this is a serious data problem. To deal with this problem, we impute missing wages for these workers in 1977-1980 using a procedure due to Hsiao and Mountain (1985). Specifically, we estimate Tobit equations for weekly wages for each of the years 1977-1980 using observations on all workers.¹³ We then replace missing values of the weekly wage for those above the maximum in each year with the expected value of the wage conditional on the independent variables and the information that weekly wages were greater than maximum (weekly) insurable earnings. However, it should be noted that this imputation procedure tends to reduce the variation in the wage, and thus it is likely to be more difficult to identify the wage and benefit coefficients than if we had complete wage data.

Legislative changes in the benefit ratio over the sample period also produce independent variation in benefits. First, the benefit ratio was lowered in January 1976 from .75 to .67 for those with dependents. (Those without dependents had a benefit ratio of .67.) Second, the benefit ratio was reduced for all workers from .67 to .60 in January 1979. These changes improve our chances of isolating the respective impact of wages and benefits on the hazard function. However, our ability to exploit these legislative changes is limited by the fact that the first modification affects only a relatively small fraction of our sample in one year, while the second change could be completely captured

by a simple time dummy for 1979 and 1980.¹⁴

As noted above, a worker's entitlement should also have an important effect on his probability of leaving employment. For a given individual on the first spell of an unemployment claim, entitlement equals initial entitlement minus duration.¹⁵ Initial entitlement indicates the maximum number of weeks that an individual's spell will be covered by unemployment insurance. For example, if a worker enters the spell with an initial entitlement of thirty-six weeks, after fifteen weeks of unemployment he has twenty-one weeks of remaining entitlement. A worker's initial entitlement is a function of his past employment duration, the amount of entitlement used in previous spells of unemployment in the current claim, and the unemployment rate in the individual's region. This function does not differ across the country, but was changed (legislatively) over the sample period. As a result, there is substantial variation in initial entitlement both across individuals at a point in time and over time.

In the basic specification, two unemployment rates are used to proxy the demand conditions facing an unemployed worker. (Intuitively, one can think of these as directly affecting the arrival rate of job offers.) First, a monthly provincial unemployment rate is entered as an independent variable. It is important to control for differing demand conditions across the country, since initial entitlement is a function (through the legislation) of regional unemployment. (Regions are more narrowly defined than provinces in the unemployment insurance legislation.) If these demand conditions are omitted, one could overestimate the impact of entitlement. On the other hand, provincial

unemployment rates will, in part, reflect the benefit and entitlement provisions, and thus may capture some of the effect of these parameters of the insurance system. An annual industrial unemployment rate at the two-digit level forms the second proxy for demand conditions. We use the United States unemployment rate instead of the Canadian rate for the individual's two-digit industry in order to have a demand variable that is less dependent on the Canadian unemployment insurance system.¹⁶

As noted above, little demographic information is available in the data set. Data on age are available, and we enter age dummies in our specification of the hazard.

V. Empirical Results

To summarize, our basic specification contains entitlement, entitlement squared, benefits, previous wages, provincial unemployment rate, the U.S. industrial unemployment rate and age dummies. Seasonal dummy variables are also included. In what follows, we refer to these as the independent variables. Direct duration dependence is measured by a sixth-order polynomial in duration. The functional form of the hazard is given by (3). We begin by ignoring heterogeneity. The goal is to estimate the parameters of the probability of leaving unemployment through finding a new job. The sample consists of observations on males aged 18 to 64 years.

To summarize the data, in Table 1 we present the Kaplan-Meier empirical transition or hazard rates. (See Kalbfleish and Prentice (1980, p.16).) These are simply the ratio of the number of spells ending in week t to the number of spells lasting at least t weeks.

Table 1

Empirical Hazard Function^{a/}

Duration	Hazard	Standard Error ^{b/}	Duration	Hazard	Standard Error ^{b/}
1	.171	.012	27	.030	.017
2	.170	.013	28	.011	.011
3	.098	.012	29	.071	.028
4	.066	.010	30	.014	.014
5	.062	.012	31	0. <u>c/</u>	---
6	.076	.011	32	.030	.021
7	.051	.013	33	.066	.032
8	.068	.010	34	.018	.018
9	.041	.013	35	.037	.026
10	.064	.011	36	.020	.020
11	.038	.013	37	.093	.044
12	.048	.013	38	.025	.025
13	.059	.014	39	.061	.042
14	.048	.013	40	.071	.049
15	.026	.010	41	.080	.054
16	.054	.015	42	0.	---
17	.054	.016	43	0.	---
18	.043	.015	44	0.	---
19	.017	.010	45	.055	.054
20	.065	.019	46	.133	.088
21	.039	.016	47	0.	---
22	.056	.019	48	.100	.095
23	.040	.017	49	.111	.105
24	.026	.015	50	0.	---
25	.009	.009	51	0.	---
26	.038	.018	52	.5	.25
			53	1. <u>d/</u>	---

^{a/} Except for the Appendix Table, all hazard functions refer to the new job hazard.

^{b/} If λ_t is the hazard, then $V(\lambda_t) = \lambda_t (1-\lambda_t)/n_t$, where n_t is the number of spells that last at least t weeks.

^{c/} Indicates zero transitions in this week.

^{d/} Reflects 1 spell ending.

The means and standard deviations of the variables are given in Table 2.¹⁷

The result of estimating the basic specification of the hazard function have been placed in column 1 of Table 3. Note first that both entitlement and entitlement squared are individually significantly different from zero, while this is not true for duration-eligibility interaction coefficient. Ignoring the later coefficient, the results suggest that effect of entitlement on the hazard is negative for those with less than 42 weeks of entitlement. Thus entitlement has the expected sign except for those with high levels of initial entitlement in the early portion of their unemployment claim. (Maximum initial entitlement for an unemployment claim is approximately 52 weeks, and a claim cannot last for more than a calendar year.) The regional and industrial unemployment rates have the expected (negative) sign, although only the industrial unemployment rate is significantly different from zero. The coefficients on the age dummies indicate that those aged forty to forty-nine have, everything else held equal, the lowest probability of leaving unemployment for a new job. This result is not surprising. What is, perhaps, surprising is the result that those fifty and over, and particularly those sixty and over, appear to have the highest probability of leaving unemployment. This may reflect individuals with a relatively short horizon setting a relatively low reservation wage.¹⁸

The previous wage enters with a negative coefficient that is statistically significant. The estimated coefficient on benefits has the wrong sign but is very imprecise. We interpret these results as indicating that there is insufficient variation in benefits (independent of

Table 2
Summary Statistics

Variable	Mean	Standard Deviation
Weekly Wages/CPI ^{a/}	3.077	2.642
Weekly Benefits/CPI	1.656	0.378
Provincial Unemployment Rate	8.386	2.566
Industrial Unemployment Rate	8.272	3.373
Entitlement in First Week of Spell (weeks)	36.340	13.893
<u>Age Dummies:</u>		
Under 25 years	0.362	0.481
25-29 years	0.157	0.364
30-34 years	0.124	0.329
35-39 years	0.099	0.299
40-44 years	0.037	0.189
45-49 years	0.035	0.189
50-54 years	0.047	0.212
55-59 years	0.051	0.220
<u>Seasonal Dummies:</u>		
Winter	0.267	0.442
Spring	0.198	0.399
Summer	0.230	0.421
<u>Duration (Weeks):</u>		
All Spells ^{b/}	9.224	10.846
Spells Ending in a New Job	7.370	9.177
Spells ending in a Recall	9.208	10.150

Number of Spells: 1058
Number of Individuals: 282
Number of Spells Ending in a New Job: 754
Number of Spells Ending in a Recall: 130

^{a/} CPI equals 100 in 1981 and changes monthly.

^{b/} Includes censored spells.

Table 3

Alternative Specifications of the Hazard Function

	(1)	(2)	(3)	(4) ^{a/}	(5)	(6) ^{b/}	(7)
Constant	.7217 (.4016)	-.1939 (.3801)	.1114 (.3365)	.8793 (.3911)	.6507 (.3999)	1.8094 (.4677)	1.6878 (.4712)
Weekly Benefits/CPI	.0425 (.1242)	.0364 (.1239)	.0417 (.1237)	.0416 (.1244)	-.0641 (.1135)	-.0795 (.1594)	-.1082 (.1588)
Weekly Wages/CPI	-.0494 (.0251)	-.0492 (.0250)	-.0491 (.0250)	-.0507 (.0251)	---	-.0794 (.0514)	-.0834 (.0509)
Provincial Unemployment Rate	-.0099 (.0171)	-.0087 (.0170)	-.0087 (.0169)	-.0172 (.0123)	-.0036 (.0169)	.0068 (.0200)	.0002 (.0204)
Industrial Unemployment Rate	-.0324 (.0123)	-.0330 (.0122)	-.0324 (.0123)	-.0309 (.0123)	-.0327 (.0122)	-.0292 (.0156)	-.0510 (.0206)
<u>Age Dummies:</u>							
Under 25 years	-.5867 (.1497)	-.6028 (.1487)	-.5860 (.1498)	-.6379 (.1521)	-.5145 (.1478)	-.8333 (.1709)	-.8345 (.1721)
25-29 years	-.5783 (.1637)	-.5927 (.1627)	-.5806 (.1637)	-.6193 (.1646)	-.5095 (.1619)	-.6721 (.1849)	-.6933 (.1870)
30-34 years	-.4451 (.1769)	-.4426 (.1757)	-.4416 (.1770)	-.4862 (.1774)	-.3885 (.1754)	-.5431 (.1997)	-.5362 (.2011)
35-39 years	-.5696 (.1808)	-.5772 (.1796)	-.5550 (.1806)	-.6035 (.1811)	-.4969 (.1791)	-.5170 (.1995)	-.5279 (.1999)
40-44 years	-.8586 (.2543)	-.8665 (.2536)	-.8582 (.2544)	-.8792 (.2545)	-.7960 (.2532)	-.8155 (.3042)	-.7442 (.3066)
45-49 years	-.8063 (.2709)	-.8011 (.2698)	-.8059 (.2709)	-.8728 (.2749)	-.7375 (.2700)	-.9969 (.3429)	-.9908 (.3460)
50-54 years	-.3665 (.2246)	-.3738 (.2236)	-.3632 (.2246)	-.4221 (.2256)	-.3036 (.2236)	-.5862 (.2634)	-.5963 (.2640)
55-59 years	-.3665 (.2246)	-.2906 (.2303)	-.3017 (.2317)	-.3475 (.2325)	-.2747 (.2310)	-.3167 (.2538)	-.3000 (.2544)

-continued-

Table 3 (Continued)

<u>Seasonal Dummies:</u>							
Winter	.2637 (.1159)	.2533 (.1152)	.2641 (.1160)	.2445 (.1135)	.2540 (.1159)	.1410 (.1350)	.1468 (.1353)
Spring	.6091 (.1170)	.6254 (.1165)	.6128 (.1171)	.5976 (.1161)	.5947 (.1167)	.5382 (.1351)	.5407 (.1355)
Summer	.4748 (.1189)	.4824 (.1187)	.4866 (.1190)	.4684 (.1190)	.4668 (.1188)	.3532 (.1370)	.3463 (.1372)
Entitlement ^{c/} (10 week units)	-.6060 (.1478)	-.6175 (.1473)	-.6059 (.1181)	-.6444 (.1501)	-.5998 (.1477)	-.7394 (.1193)	-.7907 (.1215)
Entitlement **2	.0707 (.0232)	.0729 (.0231)	.0708 (.0205)	.0775 (.0236)	.0690 (.0231)	.0776 (.0125)	.0790 (.0125)
Duration (10 week units)	-5.4774 (1.1539)	-1.5696 (.2388)	Step Function ^{d/}	-5.5229 (1.1543)	-5.4841 (1.1535)	-6.4772 (1.3839)	-6.4717 (1.3854)
Duration x Entitlement	.0087 (.0592)	-.0017 (.0611)	---	.0274 (.0606)	.0062 (.0591)	.0486 (.0526)	.0555 (.0530)
Duration **2	6.7565 (2.3570)	.2554 (.0453)	---	6.7564 (2.3575)	6.7387 (2.3561)	8.3781 (2.8556)	8.2224 (2.8581)
Duration **3	-4.5416 (2.0002)	---	---	-4.5380 (2.0008)	-4.5214 (1.9994)	-5.9552 (2.4539)	-5.8219 (2.4559)
Duration **4	1.5770 (.7923)	---	---	1.5779 (.7926)	1.5689 (.7919)	2.1473 (.9830)	2.0981 (.9839)
Duration **5	-.2691 (.1459)	---	---	-.2696 (.1459)	-.2675 (.1458)	-.3737 (.1829)	-.3654 (.1831)
Duration **6	.0179 (.0101)	---	---	.0179 (.0101)	.0178 (.0101)	.0250 (.0127)	.0245 (.0128)
Duration Previous Spell	---	---	---	---	---	-.2548 (.0481)	-.2726 (.0490)
Entitlement Previous Spell	---	---	---	---	---	---	.0784 (.0402)
Industry Unemployment Previous Spell	---	---	---	---	---	---	.0319 (.0186)
Log Likelihood	-2453.5	-2465.1	-2452.6	-2452.6	-2457.3	-1762.6	-1759.4

Notes to Table 3

- a/ Regional unemployment rate replaces provincial unemployment rate.
- b/ Sample size consists of 786 spells based on observations from 199 individuals. Of these spells, 547 were completed.
- c/ In this and following tables (except Table 4), entitlement and duration are scaled by .1. Thus, entitlement squared equals weeks of entitlement squared divided by 100.
- d/ The coefficients (standard errors) are: 2 weeks, .0421 (.1254); 3 weeks, -.5957 (.1564); 4 weeks, -1.0202 (.1890); 5 weeks, -1.0724 (.1999); 6 weeks, -.8566 (.1926); 7 weeks, -1.3127 (.2364); 8 weeks, -1.0074 (.2190); 9 weeks, -1.5415 (.2793); 10 weeks, -1.0874 (.2393); 11-12 weeks, -1.5311 (.2213); 13-14 weeks, -1.3610 (.2164); 15-17 weeks, -1.6588 (.2133); 18-20 weeks, -1.7663 (.2397); 21-25 weeks, -2.0171 (.2354); 26-30 weeks, -2.1964 (.2832); 31-39 weeks, -2.3302 (.2719); 40+ weeks, -2.0691 (.3279).

wage rates) to precisely estimate the effect of this variable. (We investigate dropping the wage below.)

The polynomial in duration is difficult to interpret from simply viewing the coefficients. To describe the direct duration dependence, in Figure 1 we have plotted the hazard function holding entitlement constant. (This is given by the solid line. Duration is measured in weeks.) We see that the hazard falls with duration until 46 weeks. It rises after this point, but we should note that we have very little data on durations of this length. In Figure 2, we have plotted the hazard function allowing entitlement to change with duration. Thus this graph demonstrates the total duration dependence effect. The hazard falls with duration until 24 weeks, at which point it begins to rise with duration. The change at 37 weeks reflects the exhaustion of benefits in this example.

For comparison to earlier work, in column 2 of Table 3 we report the results of assuming that the duration dependence is described by linear and quadratic terms in duration only. This change has very little effect on the coefficients of the independent variables. However, Figures 1 and 2 suggest that this modification does affect somewhat the pattern of direct and total duration dependence. (If we restrict $h(t)$ and $z(s)$ to be linear, both the direct and total duration dependence effects are significantly negative.)

To check the robustness of our results, we replaced the sixth-order polynomial with a step function using time dummies. Define a dummy variable D_j such that $D_j = 1$ if t lies in the interval, (a_j, b_j) and zero otherwise. Then

FIGURE 1
HAZARD FUNCTION-ALTERNATIVE SPECIFICATIONS OF DURATION DEPENDENCE
(ENTITLEMENT HELD CONSTANT)

FIGURE 2

HAZARD FUNCTION-ALTERNATIVE SPECIFICATIONS OF DURATION DEPENDENCE
(ENTITLEMENT VARIES WITH DURATION)

$$(13) \quad h(t) = \sum_{j=1}^J \mu_j D_j .$$

We chose eighteen time dummies. The first seventeen of these each represent at least three percent of the weekly observations, while the final (residual) dummy covers only two percent of these observations.¹⁹ These results have been placed in column 3 of Table 2. (Footnote d of the table contains the parameter estimates and standard errors for the coefficients on the time dummies). The results for the independent variables are remarkably close to those reported in columns 1 and 2 of this table. In Figure 1, we have plotted the hazard function holding entitlement constant, while in Figure 2 we illustrate the total duration dependence effect by allowing entitlement to change with duration. The plots of the hazard function for this specification are quite close to those for the sixth-order polynomial.²⁰

In Table 4, we present the effect on expected duration (calculated from (10)) of a change in an independent variable holding age at 40-44 years and the other variables at their mean values. We report the results for each of the three parameterizations of $h(t)$. Since the results are similar, we focus on the sixth-order polynomial. Mean expected duration for this age group is 20.67 weeks, while expected duration equals 15.17 weeks when the age dummies are set equal to the proportion of the sample in the respective category. An increase in initial entitlement of one week raises expected duration by .33 weeks. Alternatively, a one percentage point increase in the industrial unemployment rate leads to an increase in expected duration of .53 weeks, while a one percentage point increase in the provincial

Table 4

The Expected Duration of Unemployment and the
Effect of Changes in the Independent Variables on the Expected Duration

	Form of Duration Dependence		
	Sixth-Order	Quadratic	Step-Function
A. Change in Independent Variable	Change in Expected Duration^{a/}		
Initial Entitlement (+1 week)	+ .33	+ .35	+ .26
U.S. Industrial Unemployment (+1 percentage point)	+ .53	+ .58	+ .31
Provincial Unemployment (+1 percentage point)	+ .22	+ .15	+ .09
Age (40-44 to 45-49 years)	- .86	- 1.11	- .53
B. Expected Duration			
Age 40-44 years	20.77	20.78	18.55
Mean of Age Dummies ^{b/}	15.17	15.11	14.57

^{a/}Duration measured in weeks.

^{b/}Age dummies set equal to proportion of sample in respective category.

unemployment rate is estimated to raise expected duration by .22 weeks. A change in age group from 40-44 years to 45-49 years is predicted to reduce expected duration by .86 weeks.²¹

We also considered a number of further modifications to the specification of the hazard while maintaining the assumption that $h(t)$ is a sixth-order polynomial. In column 4 of Table 3, the (monthly) provincial unemployment is replaced by the regional unemployment rate actually used to calculate initial entitlement. This has very little effect on the parameter estimates and strongly suggests that the estimated impact of entitlement on the hazard does not simply reflect regional differences in demand conditions.

Given the lack of independent variation between wages and benefits, in column 5 of Table 3 we investigate the effect of dropping the wage. In this specification, benefits reflect movements in the two variables. The benefits coefficient now has the expected sign, but is still quite imprecisely estimated.

In column 6 of Table 3, duration in the previous unemployment spell is added as an independent variable to investigate lagged duration dependence (Heckman and Borjas (1980)). As in the Narendranathan, Nickell and Stern (1983) study, we find that the coefficient on lagged duration is quite significant but that inclusion of this variable does not have a substantial effect on the other parameter estimates. The lagged duration term can proxy either the true impact of lagged duration or simply that portion of unobserved differences which is constant across spells. Since workers do not face a stationary environment, conditions facing the worker during his previous unemployment spell may

also be relevant. As a first approach to this issue, we added the industrial unemployment rate and entitlement at the beginning of the previous spell (column 7).²² Both of these additional variables are statistically significant (at the 10 percent level) and enter with a sign opposite to that of lagged duration. Since an increase in these variables would be expected to raise lagged duration, the signs on the coefficients of the lagged variables act analogously to what one would find in a regression model with positive autocorrelation.

We also investigate several further modifications to the basic specification. (These are not reported in the tables.) First, wages and benefits are interacted with entitlement in an attempt to capture Mortenson's (1977) suggestion that the impact of higher benefits will change as a worker's remaining entitlement falls. Second, we interact the wages and benefits with a dummy for those under 25 years of age, since Narendranathan, et al. (1985) found that younger workers were much more responsive to changes in benefits. None of these interactions are individually or jointly significantly different from zero. Third, we follow the suggestion in Nickell (1979), Topel (1984) and Narendranathan, et al. (1985), and impute a wage for all workers.²³ The coefficient and normal statistic on the imputed wage essentially becomes zero, and the other parameter estimates are very close to those obtained when the wage is dropped.

Finally, we investigate the sensitivity of our results to allowing for unobserved heterogeneity. Consider first the case where the heterogeneity term θ is distributed independently across spells and is constant only within a spell. We first assume that θ is drawn from a

two-point distribution and that $h(t)$ is a sixth-order polynomial. These results are placed in column 1 of Table 5. The log-likelihood rises by 27.2, leading to a likelihood ratio statistic of 54.4, which is considerably larger than any reasonable critical value of $\chi^2(2)$.²⁴ However, the addition of the heterogeneity terms leads to imprecise estimates of the duration coefficients. A Wald test of the joint hypothesis that the coefficients on the fifth- and sixth-order terms are zero yielded a test statistic of only .51, while a test of the hypothesis that the last three terms of the polynomial jointly have zero coefficients produced a statistic of 20.78.²⁵ We re-estimated the model using a fourth-order polynomial for direct duration dependence and the estimates are in column 2 of Table 5. For comparison, column 3 contains the corresponding estimates when one ignores heterogeneity. Note first that dropping the fifth- and sixth-order terms (conditional on two-point heterogeneity) in duration results in a likelihood ratio test statistic of 0.4, which is quite close to the Wald statistic reported above. Second, a comparison of columns 2 and 3 indicates that allowing for two-point heterogeneity raises the log likelihood by 29.0. This modification appears to affect the duration coefficients (this is discussed below), and, to a lesser degree, the coefficients on the age dummies. The coefficient on the duration-eligibility interaction rises considerably in absolute value and statistical significance. The latter change tends to strengthen the entitlement effect, since the linear entitlement estimate is unchanged and the coefficient on entitlement squared only rises slightly. We added a third mass point, but this led to an increase in the log-likelihood of only .8.

Table 5

Alternative Specifications of Heterogeneity

	Heterogeneity Constant Within a Spell		No Heterogeneity	Heterogeneity Constant for an Individual (Fourth-Order Polynomial)	
	Sixth-Order Polynomial	Fourth-Order Polynomial	Fourth-Order Polynomial	2 Points of Support	3 Points of Support
Weekly Benefits/ CPI	-.0804 (.1609)	-.0818 (.1603)	.0384 (.1240)	-.0312 (.1494)	.2246 (.1536)
Weekly Wages/ CPI	-.0794 (.0317)	-.0798 (.0317)	-.0491 (.0244)	-.0466 (.0273)	-.0448 (.0278)
Provincial Unem- ployment Rate	-.0187 (.0211)	-.0182 (.0210)	-.0098 (.0171)	-.0246 (.0209)	-.0230 (.0215)
Industry Unem- ployment Rate	-.0403 (.0142)	-.0403 (.0142)	-.0328 (.0127)	-.0405 (.0141)	-.0312 (.0141)
<u>Age Dummies:</u>					
Under 25 years	-1.0862 (.2056)	-1.0866 (.1867)	-.5919 (.1495)	-.6144 (.2339)	-.2192 (.2245)
25-29 years	-.8792 (.2056)	-.8816 (.2046)	-.5813 (.1636)	-.7541 (.2247)	-.3413 (.2351)
30-34 years	-.6284 (.2181)	-.6327 (.2170)	-.4489 (.1767)	-.7498 (.2014)	-.3022 (.2347)
35-39 years	-.7667 (.2105)	-.7695 (.2094)	-.5773 (.1805)	-.8446 (.2023)	-.3352 (.2383)
40-44 years	-1.2101 (.3101)	-1.2148 (.3106)	-.8621 (.2541)	-1.0714 (.2555)	-.5907 (.2815)
45-49 years	-1.2977 (.3219)	-1.2985 (.3203)	-.8094 (.2707)	-1.1994 (.3237)	-.7449 (.3390)
50-54 years	-.9982 (.3167)	-.9975 (.3166)	-.3664 (.2244)	-.9570 (.3067)	-.5225 (.3135)
55-59 years	-.2663 (.3388)	-.2656 (.3331)	-.3192 (.2316)	-.3021 (.2604)	-.2434 (.2787)

-continued-

Table 5 (Continued)

<u>Seasonal Dummies:</u>					
Winter	.2185 (.1334)	.2205 (.1332)	.2641 (.1158)	.2706 (.1207)	.2841 (.1214)
Spring	.6046 (.1385)	.6007 (.1384)	.6080 (.1170)	.6642 (.1221)	.6834 (.1229)
Summer	.3032 (.1372)	.3035 (.1369)	.4734 (.1188)	.4656 (.1233)	.4981 (.1247)
Entitlement (10 week units)	-.6048 (.1942)	-.6022 (.1940)	-.6056 (.1476)	-.6291 (.1582)	-.5836 (.1618)
Entitlement **2	.0817 (.0292)	.0816 (.0291)	.0705 (.0231)	.0807 (.0248)	.0740 (.0253)
Duration (10 week units)	-1.4463 (1.2832)	-.7328 (.5787)	-3.4491 (.4889)	-2.9628 (.5081)	-2.5760 (.5193)
Duration x Entitlement	-.6317 (.1075)	-.6365 (.1064)	.0106 (.0590)	-.0022 (.0637)	-.0218 (.0642)
Duration **2	4.0928 (2.5591)	2.3528 (.5127)	2.220 (.4551)	2.0655 (.4663)	1.7528 (.4729)
Duration **3	-2.4136 (2.1826)	-.8689 (.1853)	-.6578 (.1588)	-.6350 (.1624)	-.5422 (.1637)
Duration **4	.7096 (.8664)	.0908 (.0203)	.0677 (.0174)	.0669 (.0178)	.0578 (.0179)
Duration **5	-.1132 (.1597)	---	---	---	---
Duration **6	.0077 (.0110)	---	---	---	---
θ_1	1.4921 (.4962)	1.4061 (.4798)	.5162 (.3872)	1.1614 (.4387)	1.6646 (.4581)
θ_2	-2.8537 (.6702)	-2.9798 (.6447)	---	.0344 (.4284)	-.0560 (.4883)
θ_3	---	---	---	---	-.9694 (.4734)
P_1	.7784 (.0176)	.7789 (.0173)	1.0	.5035 (.0102)	.0200 (.0006)
P_2	---	---	---	---	.5611 (.0171)
Log Likelihood	-2426.3	-2426.5	-2455.5	-2431.5	-2418.8

We next assume that θ is constant across an individual's unemployment spells. We could not reject the hypothesis that $h(t)$ is a fourth-order polynomial when compared to a sixth-order polynomial, assuming two points of support. (The likelihood ratio statistic is only 0.4.) These results are contained in column 4. With the possible exception of the age dummies, assuming this form of heterogeneity has little effect on the coefficients of the independent variables. The results for the case where θ has a three-point distribution are given in column 5, since the addition of the two parameters (θ_3 and P_2) increased the log-likelihood by 12.7.²⁶ In terms of the independent variables, the addition of the third mass point affects only the coefficients on the age dummies.

Thus, among the independent variables only the coefficients on the age dummies are affected by allowing for heterogeneity. In particular, our results for demand conditions and entitlement are robust to allowing for unobserved heterogeneity.

In Figure 3 we plot the expected value of the hazard (holding entitlement constant) for the different heterogeneity specifications. In each case, a fourth-order polynomial is used to describe duration dependence. Figure 4 contains the corresponding plots when entitlement varies with duration. The hazard functions for the no heterogeneity case and the cases where the heterogeneity component is fixed for an individual are quite similar, and generally replicate the results in Figures 1 and 2. (The hazard for the fourth-order polynomial without heterogeneity in Figure 4 is somewhat flatter than the sixth-order polynomial in Figure 2.) However, the hazard functions for the case where

FIGURE 3

HAZARD FUNCTION - ALTERNATIVE HETEROGENEITY SPECIFICATIONS
(ENTITLEMENT CONSTANT)

FIGURE 4

HAZARD FUNCTION - ALTERNATIVE HETEROGENEITY SPECIFICATIONS
(ENTITLEMENT VARIES WITH DURATION)

the heterogeneity component is assumed to be independently distributed across spells are implausible. While we have no clear cut explanation for the time pattern in this latter case, it may reflect the inadequacy of the strong independence assumption. Alternatively, this result also could indicate the greater ease of identifying the model which assumes that the heterogeneity component is fixed for an individual since we have multiple spell data.

6. Conclusion

By applying a discrete time duration model to microeconomic data on unemployment spells in Canada, we are able to isolate the impact of entitlement, demand conditions, and age on the probability of leaving unemployment. The results indicate that unemployment insurance entitlement has a substantial effect on the employment duration, even for those who do not ultimately exhaust their benefits. In the absence of entitlement effects, the probability of a worker leaving unemployment declines with duration of unemployment. When entitlement is allowed to vary, the probability of leaving unemployment first falls and then generally rises with unemployment duration as the declining entitlement induces a greater willingness to accept offers or search more intensively. These results are robust to alternative specification of duration dependence and to allowing for person specific heterogeneity.

Footnotes

¹An exception is Glenday and Alam (1982).

²Beach and Kaliski (1983) analyze the 1979 changes using aggregate gross flow data.

³Initial entitlement refers to the number of weeks of unemployment insurance benefits that an individual is entitled to at the beginning of an unemployment spell, while entitlement refers to the number of weeks of coverage remaining at any point in the spell. In some U.S. studies, initial eligibility and eligibility are used in place of initial entitlement and entitlement respectively. We have adopted the former terminology since it corresponds to that used by Canadian policy makers. For the same reason, we use benefit rate instead of replacement rate.

⁴Our data require that we ignore the possibility of leaving the labour force. This is not a severe restriction for men.

⁵This assumes that there is some leisure in periods of unemployment.

⁶See, for example, Elbert and Ridder (1982), Heckman and Singer (1984b, 1984c), Lancaster and Nickell (1980), and Ridder and Verbakel (1984).

⁷Studies estimating discrete time models include Atkinson et al., (1984), Kiefer and Neumann (1979, 1981) and Nickell (1979). Continuous time models are investigated by Burdett et al., (1984), Flinn and Heckman (1982, 1983), Heckman and Singer (1984a), Lancaster (1979), Lancaster and Nickell (1980), Narendranathan et al., (1985), Solon (1985), and Topel (1984), among others. For important earlier studies

using a different methodology, see Classen (1979) and Ehrenberg and Oaxaca (1976).

⁸Flinn and Heckman (1982) note that distrete-time models cannot deal with time intervals shorter than that upon which the estimation is based. Since we use weekly data, as a practical matter we do not believe that this is a serious problem in our study.

⁹Topel (1984) and Katz (1986) also obtain this result.

¹⁰In forming standard errors and test statistics, we ignore problems arising from the fact that K is determined by the data.

¹¹Available evidence suggests that insured spells represent the great majority of all spells of unemployment in Canada. See Ashenfelter and Card (1986, Table 9a).

¹²In that there is no regional variation in the benefit ratio, the Canadian system is closer to the British system than that in the U.S.

¹³The explanatory variables are 1976 weekly earnings, age and age-squared. (For cases where 1976 weekly earnings were missing, we estimate equations based on age, age-squared, and regional dummies.) The equations are estimated in logarthimic form.

¹⁴This suggests an alternative sampling strategy could be used to investigate the role of the benefits in determining the value of the hazard function. One could focus on 1975 and 1976 data, and oversample unemployment spells for those with dependents. This would allow one to simulate an experiment, where workers with dependents are the treatment group while those without dependents act as the control group. This approach would be quite similar to Solon's (1985) use of the tax-back provision introduced in the U.S. in 1979 for high income recipients of

unemployment insurance payments. Benefits are considered taxable income in Canada and in 1979 the government began placing a higher tax on U.I. payments to high income individuals. However, we cannot isolate which individuals were affected by this change because of the censoring of the wage data.

¹⁵Of course, actual entitlement equals zero if duration is greater than initial entitlement.

¹⁶The U.S. unemployment rates are taken from unpublished Department of Labor annual (December) averages.

¹⁷Given a benefit or replacement ratio of .6, the standard deviation of benefits may appear surprisingly small when compared to that for wages. However, one would expect the standard deviation of benefits to be less than .6 of that of wages because earnings are only insured to maximum insurable earnings. For example, let $w \sim N(\mu, \sigma^2)$. Define a variable B such that $B = .6w$ if $w < \mu$ and $B = .6\mu$ if $w \geq \mu$. (In Canada, maximum insurable earnings equal (approximately) the mean of weekly earnings.) Then the standard deviation of B equals $.35\sigma$, which is considerably closer to the ratio in Table 2.

¹⁸For an individual to have a transition, he must report earnings in the next week, and thus this result does not represent transitions into retirement being erroneously coded as a transition into a new job.

¹⁹Each week from 2 to 10 has its own dummy. We use dummies for the following intervals: 11-12 weeks; 13-14 weeks; 15-17 weeks; 18-20 weeks; 21-25 weeks; 26-30 weeks; 31-39 weeks; and 40 weeks and over.

²⁰We also investigated several specifications for $h(t)$ using cubic splines with known knot points. In no case did this modification affect

the coefficients or standard errors of the independent variables. For the case that we plotted, the pattern of direct and total duration dependence was quite close to that for the sixth-order polynomial.

²¹An increase in the benefits of \$10 (1981\$) is calculated to decrease duration by .1 weeks. However, this calculation is essentially meaningless, given the imprecise estimate of the coefficient for this variable.

²²This approach is quite similar to that suggested in Heckman and Borjas (1980).

²³Regression equations are estimated for 1975 and 1976, while Tobit equations are estimated for 1977 through 1980. The explanatory variables are age, age-squared and regional dummies. We investigated adding industrial dummies to the wage equation, but they were not (jointly) significantly different from zero in the wage equation. As noted above, data on education or occupation are unavailable, and this clearly restricts our ability to identify the wage coefficient using an imputed wage. It is interesting to note that if we enter the replacement ratio instead of wages and benefits, we obtain a coefficient of .670 with a standard error of (.362) when we do not impute the wage, and a coefficient (and standard error) of -.294 (.326) when we impute the wage.

²⁴In using $\chi^2(2)$ as a guide, we are ignoring the facts that: if $\theta_1 = \theta_2$, then P_1 is unidentified; if $P_1 = 1$, θ_2 is unidentified; and that testing $P_1=1$ involves problems of testing whether a parameter is on the boundary of the parameter space.

²⁵The critical value of $\chi^2(2)$ at the 5% level is 5.99, while the

critical value of $\chi^2(3)$ at the 5% level is 7.81.

²⁶We found that the data would not support a fourth mass point. When we tried to estimate this model, the third and fourth mass points converged to the same value. This result occurred with different optimization algorithms and different starting values.

Appendix

Variable	Recall Hazard ^{a/} (1)	New Job Hazard ^{b/} (2)	Employment Hazard ^{c/} (3)
Constant	-8.4473 (1.0409)	-.2000 (.4044)	-1.7681 (.3658)
Weekly Benefits/CPI	1.0385 (.3201)	.0732 (.1232)	.1926 (.1114)
Weekly Wages/CPI	-.0240 (.0412)	-.0464 (.0248)	-.0409 (.0211)
Provincial Unemployment Rate	.0410 (.0349)	-.0002 (.0161)	.0064 (.0144)
Industrial Unemployment Rate	.0063 (.0264)	-.0371 (.0121)	-.0286 (.0181)
Age ÷ 10	1.4281 (.4732)	-.4244 (.1953)	-.1349 (.1756)
(Age ÷ 10)**2	-.1585 (.0580)	.0645 (.0241)	.0289 (.0216)
Winter	.2943 (.2758)	.2272 (.1147)	.2295 (.1040)
Spring	.7541 (.2741)	.6117 (.1160)	.6097 (.1046)
Summer	.6320 (.2842)	.4738 (.1179)	.4756 (.1067)
Entitlement (10 week units)	-.1704 (.0775)	-.2158 (.0324)	-.1975 (.0290)
Duration (10 week units)	-.9950 (.2618)	-1.7033 (.1200)	-1.5333 (.1068)
Duration**2	.1677 (.0663)	.2941 (.0302)	.2643 (.0270)
Log Likelihood ^{d/}	-662.3	-2474.2	-3419.0

^{a/}Spells ending in a transition to a new job treated as censored.

^{b/}Spells ending in a recall treated as censored.

^{c/}Hazard functions constrained to be the same.

^{d/}Likelihood ratio test statistic for hypothesis that the two hazard functions are the same equals 545. The critical value of $\chi^2(13)$ equals 22.4 at the 5% level.

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