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ABSTRACT

Unemployment Insurance and Subsequent Job Duration: Job Matching vs Unobserved Heterogeneity*

The relationship between unemployment benefit duration, unemployment duration and subsequent job duration is investigated using a multi-state duration model with state specific unobserved heterogeneity. I allow maximum benefit duration to be correlated with unemployment duration as well as accepted job duration. I examine two potential explanations for the relationship between unemployment and job spell durations; UI benefits increase job matching quality vs unobserved heterogeneity. I find that the escape rate out of unemployment seems to raise significantly within 5 weeks of benefit termination and new jobs accepted within this 5 week period seem to have a higher dissolution rate. At the same time, unobserved heterogeneity is also found to explain the correlation between unemployment and job duration. Various simulations indicate that increasing the maximum benefit duration by one week will raise expected unemployment duration by 1 to 1.5 days and expected job duration by 0.5 to 0.9 day.

JEL Classification: J64, J65

Keywords: Unemployment insurance, unemployment duration, job duration, job matching

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* This is a revised version of the papers “Contiguous Duration Dependence and Nonstationarity in Job Search” and ”Contiguous Duration Dependence in Unemployment and Accepted Job Spells.”

1. Introduction

The effect of Unemployment Insurance (UI) on the labor market is one of the most active areas of research in modern labor economics. The payment of UI benefits increases the welfare of risk averse workers affected by adverse employment shocks. At the same time, UI benefits raise the reservation wages of the unemployed and are therefore expected to increase the quality of subsequent job matches. However, UI systems are known to face severe moral hazard problems. Although the payment of UI benefits is typically contingent on active job search, the managers of UI programs are usually unable to observe the actual search effort invested by the unemployed[1]. As a consequence, the effects of Unemployment Insurance benefits on the duration of unemployment has become one of the most widely studied issues in applied labor economics (see Atkinson and Micklewright, 1991, for a survey).[2] In particular, several economists have investigated the effects of an increase in benefit level and in maximum benefit duration on the exit rate out of unemployment. It is generally admitted that an increase in benefit generosity reduces the hazard rate out of unemployment. In particular, empirical work by Meyer (1990) and Han and Hausman (1990) have found spikes in the escape rate out of unemployment when unemployment benefits lapse[3]. In both cases, empirical work was based on the specification of a hazard function which incorporates time varying regressors while allowing a non-parametric estimation of the baseline hazard function.

The effects of UI on labor markets are not limited to the duration of unemployment. UI can also affect job duration as well as employment duration. For instance, it is well known that the payment of UI benefit based on accumulated

[1]For an analysis of moral hazard and Unemployment Insurance in the dynamic macroeconomic literature, see Hansen and Imrohoroglu (1992).

[2]Although most efforts devoted to the effects of UI on labor markets have focused on the duration of unemployment, early attempts to capture the impact of UI on the labor markets seem to have focused on the effects of UI benefits on re-employment earnings (see Ehrenberg and Oaxaca, 1976 or Classen, 1977).

[3]¿From the job search literature, it is well known that when individuals claim UI benefit for a limited period only, their reservation wages decline until benefit termination and individuals escape unemployment at an increasing rate. Theoretical work by Mortensen (1990) shows that, in a short period before exhaustion, the reservation wage may fall sharply.

weeks of employment can affect the rate at which workers quit their job and the frequency of temporary layoffs (Baker and Rea, 1998 and Christofides and McKenna, 1996). Equally, UI can affect the quality of subsequent job matches by preventing the unemployed to accept job offers that are not commensurate with their qualifications. As a consequence the links between UI and the incidence of unemployment need to be examined. Although the effects of UI on unemployment duration is relatively well documented, very little is known about the effects of UI on the quality of labor market adjustments. This paper proposes an analysis of the effects of UI benefit duration from an angle which has, so far, been neglected in the literature[4]. The empirical analysis is based on the idea that entitlement to UI benefits for a limited period should not only affect the escape rate out of unemployment but can also create a relationship between completed unemployment duration and subsequent job duration. If spikes in the escape rate out of unemployment are explained by a significant decrease in reservation wages, individuals, who are close to benefit termination, might accept jobs which they are likely to quit in the future. Given that better job matches are less likely to dissolve, this suggests that poorer matches are made at the time of UI benefit termination. As appealing at it is, this matching hypothesis has to be confronted to the "unobserved heterogeneity" hypothesis. The correlation between unemployment duration and subsequent job duration would be spurious if individual unobserved heterogeneity affecting unemployment duration is correlated with unobserved heterogeneity affecting job duration. For this reason, the econometric model must be constructed so that both hypotheses can be confronted.

In what follows, I consider subsequent job duration as a measure of job match quality. I use a multi-state hazard model to investigate how the subsequent job hazard function is related to unemployment duration as well as unemployment benefit duration. Formally, the model has 2 states; unemployment and, subsequent job. However, in order to pay attention to the potential endogeneity of the maximum benefit duration period, I also model maximum benefit duration accumulated by the time of the job separation as a separate duration observed before the onset of the unemployment spell.[5] Although maximum benefit duration is

[4] To my knowledge, Belzil (1990) is the first example of empirical analysis of the effects of UI on job/employment stability.

[5] In this paper, I use the term "maximum benefit duration" to designate the initial benefit entitlement period and the term "potential benefit duration" to designate the number of weeks of benefit remaining at a given point in time.

determined by government rules and individuals cannot influence its length once unemployed, the maximum benefit duration accumulated by an individual is function of endogenous variables such as weeks of employment and therefore affects the decision to quit. While it is true that empirical analysis of the effect of UI benefits on re-employment outcomes is typically based on sample of workers who have been laid off, many individuals willing to quit their job might have implicit agreements (with their employer) to be laid off.[6] In such a case, it is possible that maximum benefit duration accumulated is affected by individual behaviour and therefore correlated with unobservables affecting unemployment duration and job duration.

In order to correct for the potential endogeneity of the maximum benefit duration, I also use a hazard specification for benefit duration and I allow unobserved heterogeneity affecting benefit duration to be correlated with unobserved heterogeneity in both unemployment duration and accepted job duration. I compare the results obtained when maximum benefit duration is assumed to be exogenous (the conventional approach). The analysis builds on Belzil (1990 and 1995) in which the effects of UI benefits and completed unemployment duration on the incidence of unemployment is investigated using Canadian administrative data.[7] The data comes from administrative file of the Canadian UI program and is quite similar to the data used in Belzil (1996, 1995). More details are found in Section 3.

Overall, the results indicate that both hypotheses contribute to explain the observed correlation between unemployment duration and accepted job duration. The escape rate out of unemployment seems to raise significantly within 5 weeks of benefit termination and new jobs accepted within this 5 week period seem to have a higher dissolution rate. At the same time, there is a strong negative correlation between unobserved heterogeneity affecting unemployment duration and unobserved heterogeneity affecting subsequent job duration. These results are very robust; they are true whether or not previous earnings are included in the unemployment hazard function and whether or not maximum benefit duration is allowed to be correlated with unemployment duration and job duration.

[6]During the period covered by the data, the unemployed quitting their job without just cause could have been penalized for a period as long as 6 weeks.

[7]In Belzil (1990, 1995), I have assumed that the maximum benefit duration period was exogenous such at it is usually done in the literature and restricted the analysis to parametric forms of unobserved heterogeneity.

The results of various simulations indicated that increasing the maximum benefit duration by one week will raise expected unemployment duration by 1 to 1.5 days and expected job duration by 0.5 to 0.9 day. In other words, the increase in unemployment duration is 50% to 100% higher than the increase in accepted job duration.

The paper is organized as follows. In the next section (Section 2), the empirical specification is discussed. Section 3 is devoted to the presentation of the data while the results are in Section 4. In Section 5, I illustrate the implications of the structural parameters with unemployment and accepted job hazards and use various model specifications to simulate the effects of an increase in benefit duration on average unemployment duration and average accepted job duration. The conclusion is in Section 6.

2. Econometric Specification

The analysis presented in this paper is based on a multi-state duration model which is used to estimate the effects of unemployment insurance benefit duration on subsequent job match quality (subsequent job duration). The definition of job duration which I use is the waiting time until the subsequent job (following a spell of unemployment) is terminated voluntarily.[8] The model has the following features;

- The distribution of subsequent job duration depends on completed unemployment duration (contiguous duration dependence) and the dependence is specified such that I can distinguish between a potential benefit duration effect and a general unemployment duration effect.
- The distribution of unemployment spell durations depends on the maximum benefit duration accumulated at the time of the job separation as well as the potential benefit period (as the unemployment spell progresses).
- In the case where I also model maximum benefit duration (the initial condition), its distribution must depend on regressors that are exogenous.[9]

[8]However, I will relax this assumption later and treat jobs terminated by layoffs as completed spells.

[9]Although it would be possible to use previous job duration as a regressor to determine benefit duration, this would mean using a control variable (potentially endogenous).

- All episodes; benefit duration (when endogenous), unemployment duration and accepted job duration are stochastically related through unobserved heterogeneity.

Although it would be possible to make use of previous job duration (actually used to calculate the maximum benefit periods) when modelling benefit duration, the use of job duration in the benefit duration equation would imply that I am conditioning on a variable which is clearly endogenous. For this reason, I prefer to use variables such as age, experience and industrial sector in the benefit duration equation. Another approach could be to model only unemployment and accepted job duration but allow the distribution of unobserved heterogeneity (population proportions for various type) to be function of the maximum duration of benefit.

2.1. Modeling the Hazard Functions

In order to estimate the effect of benefit duration on subsequent job duration I have to model three separate durations; maximum benefit duration, unemployment duration and subsequent job duration (for those who have found reemployment). Because semi-parametric estimations of each hazard function would require the estimation of a very large number of parameters on top of the already large number of regression parameters which I have to estimate, I restrict myself to parametric representation of the baseline hazard functions. I however estimate the distribution of the unobserved heterogeneity terms using a flexible method.

As a starting point, I assume that maximum benefit duration at the start of a spell of unemployment is potentially endogenous. I assume that observed benefit duration (t_{υ}) takes discrete values but is generated from a continuous random variable, denoted τ_{υ}, which follows a Weibull Proportional Hazards model;

$$h_{\upsilon}(\tau_{\upsilon} \mid \varepsilon^{\upsilon}) = \exp(Z_{\upsilon}^{\prime}\eta_{\upsilon})\alpha\tau_{\upsilon}^{\alpha-1}\varepsilon^{\upsilon} \tag{2.1}$$

where Z_{υ} is a vector of time (duration) invariant exogenous regressors to be discussed below, η_{υ} is a vector of parameters. The term ε^{υ} plays the role of unobserved heterogeneity. Clearly,

$$\Pr(\tau_{\upsilon} \geq t_{\upsilon} + 1 \mid \tau_{\upsilon} \geq t_{\upsilon}) = \exp\{-\exp(\mathrm{Z}_{\upsilon}'\eta_{\upsilon} + \log \varepsilon^{\upsilon} + h_0^*(t_{\upsilon}, \alpha))\} \tag{2.2}$$

where $h_0^*(t_{\upsilon}, \alpha) = \log((t_{\upsilon} + 1)^{\alpha} - t_{\upsilon}^{\alpha})$ and Z_i^{ν} contains Age, Experience and industrial classification dummies. The industrial groups are Primary, Construction, Transportation, Trade, Finance, Services and Administration (Manufacturing is the reference group).

As a second step, I model the hazard function for the duration of unemployment. This is the hazard function given maximum benefit duration t_v. As this stage, it is useful to discuss duration dependence. Duration dependence can arise because the data generating process implies duration dependence or because a co-variate is itself changing with elapsed duration. Simultaneously, spurious duration dependence in unemployment can also be caused by unobserved (neglected) heterogeneity. The specification of a hazard function where potential benefit duration changes every week will capture duration dependence in the search behavior of the unemployed which is explained by UI benefit exhaustion. Other duration effects will typically be captured in the baseline hazard. Normally, the baseline hazard should be estimated non-parametrically such as in Meyer (1990) or Han and Hausman (1990). However, in the present case, the consideration of three durations (benefit duration, unemployment duration and accepted job duration) renders a semi-parametric approach quite difficult.

For these reasons, I specify the unemployment hazard function as proportional hazards model with a time varying covariate (potential benefit duration) model and unobserved heterogeneity and a Weibull (rather than non-parametric) baseline hazard.[10] Denoting unemployment duration (in continuous time) by τ_u, letting Z^u denote the set of time-invariant covariates and X(.) the potential benefit duration at a given point in time, it is easy to see that

$$\Pr(\tau_u \geq t_u + 1 \mid \tau_u \geq t_u) = \exp\{-\exp(\mathrm{Z}_u\text{'}\eta_u + \delta X(t_u) + \log \varepsilon^u + h_0^*(t_u;\beta))\} \quad (2.3)$$

where $\mathrm{h}_0^*(t_u;\beta) = \log((t_u+1)^\beta - t_u^\beta)$ and where β is a parameter to be estimated. To do so, I assume that UI benefit entitlement changes between intervals (from one week to another) but remains constant within each interval (between t_u and $\mathrm{t}_u + 1$). The vector of time invariant regressors (Z_u) includes Age, Experience and industrial classification. The time varying regressor, X_t (potential benefit duration) is simply the difference between maximum benefit duration (t_v) minus

[10]Note that most theoretical models based on job search or job matching arguments predict that the job hazard rates are declining with tenure. Empirical evidence also suggests that job exit rates are declining with tenure (see Devine and Kiefer, 1991, for a survey).

elapsed unemployment duration in discrete time (t_u), that is

$$X(t_u) = Max(t_v - t_u, 0)$$

For observations censored at t_u,

$$\Pr(\tau_u \geq t_u) = \prod_{s=1}^{t_u-1} \exp(-\exp(\mathrm{Z}_u\text{'}\eta_u + \delta X(s) + \log \varepsilon^u + h_0^*(s;\beta))) \tag{2.4}$$

In the sample which I analyzed, the average potential benefit period is around 30 weeks while the maximum is 50 (for 7 individuals). Because UI officials report that recorded unemployment duration is likely to be inaccurate for those who have very long durations, I censor every durations at 50 weeks.

In order to take into account that the effect of a decrease in one week of benefit entitlement may change as benefit termination approaches, I estimate a more general specification where the effects of potential benefit period is allowed to vary over the duration of unemployment. This can be accomplished if I use the following set of variables; $\mathrm{X}(\mathrm{t}_u)$, X_{20-29}, X_{10-19}, X_{6-9} and X_{1-5}. These variables are defined as follows;

$$X_{20-29} = X(t_u) \text{ if } X(t_u) \leq 29 \text{ and 0 if not}$$

$$X_{10-19} = X(t_u) \text{ if } X(t_u) \leq 19 \text{ and 0 if not}$$

$$X_{6-9} = X(t_u) \text{ if } X(t_u) \leq 9 \text{ and 0 if not}$$

$$X_{1-5} = X(t_u) \text{ if } X(t_u) \leq 5 \text{ and 0 if not}$$

Altogether, these variables measure the remaining weeks of UI benefit at each point in the unemployment spell. With more than 29 weeks remaining, only potential benefit duration ($\mathrm{X}(\mathrm{t}_u)$) take non-zero values. When the number of weeks of UI benefit remaining lies between 20 and 29, both $\mathrm{X}(\mathrm{t}_u)$ and X_{20-29} take non-zero values. As a similar argument is applied to the remaining segments (1 to 5 and 6-9), it follows that the effect of an additional week of UI benefit is captured by the coefficient on X(t_u) in the 30-50 segment, by the sum of the coefficients for X(t_u) and X_{20-29} in the 20-29 segment weeks, by the sum of X(t_u), X_{20-29} and X_{10-19} in the 10-19 segment, by the sum of X(t_u), X_{20-29}, X_{10-19} and X_{6-9} in the 6-9 segment and by the sum of X(t_u), X_{20-29}, X_{10-19}, X_{6-9} and X_{1-5} in the 1-5 segment. As an example, if the sum of all the coefficients from X_{1-5} to $X(t_u)$ is negative, this indicates that, within 5 weeks from benefit termination, the hazard decreases with each additional week of potential (remaining) benefit duration or, in other words, that the hazard increases as the individual approaches benefit termination. Note that, contrary to what I do here, many authors (such as Meyer, 1990) use a specification where the coefficients measure the effect of moving closer to benefit termination. In other words, a positive coefficient reported in Meyer (1990) would be equivalent to a negative coefficient in this study.

Finally, the last component of the model is subsequent job duration. As stated earlier, the accepted job duration is understood as the waiting time until the individual quits the accepted job and it is meant to measure the quality of the job match. I do not model the hazard functions for competing risks since I already have three different durations to model. Accepted job spells terminated by reasons other than a quit are considered as censored accepted job durations. This simply means that a job terminated by layoff was still acceptable on the part of the worker. Given my objective to estimate the effects of UI benefit duration on accepted job duration, I must allow the accepted job hazard to depend on completed unemployment duration as well as a measure of benefit termination. I simply model accepted job duration with a proportional hazard with a Weibull baseline distribution. As I do with benefit duration and unemployment duration, I assume that recorded job duration takes discrete values which are generated from a continuous random variable, τ_j, which hazard function is given by

$$h_j(\tau_j \mid t_u, \varepsilon^j) = \exp(Z_j^{\prime}\eta_j + \zeta_0 \cdot t_u + \zeta_1 \cdot t_{u,50} + \lambda \cdot X(t_u) + \log \varepsilon^j) \cdot \gamma\tau_j^{\gamma-1} \quad (2.5)$$

where Z_j is a vector of regressor including experience, age, industrial classification dummies. The effect of unemployment duration is measured by t_u (when duration is below or equal to 50) and $t_{u,50}$ (a binary variable equal to 1 for those unemployed more than 50 weeks) while X(t_u) is the number of weeks of benefit left when the individual left unemployment to accept a new job. For instance, X(t_u) is 0 for all those who found job following benefit termination.

In order to be more flexible, I can also allow the effect of benefit duration left to vary over the total benefit duration period and make use of the variables X_{20-29}, X_{10-19}, X_{6-9} and X_{1-5} in order to obtain the ex-post effect of potential benefit duration on accepted job duration. As for unemployment hazards, the effects of potential benefit duration on accepted job hazards is measured by the sum of the appropriate coefficients.

2.2. Constructing the Likelihood Function

I construct the likelihood function as if exact spell lengths are unknown but I assume that the interval during which failure time takes place are known. Given a specification for three endogenous variables, benefit duration, unemployment duration and accepted job duration, the likelihood function can easily be constructed[11]. If we assume that conditional on unobserved heterogeneity, benefit duration, unemployment duration and accepted job duration are independent, the likelihood function is simply the product of three individual densities. Noting that benefit duration t_v is completed by definition and that the number of accepted job spells (M) is smaller than the number of unemployment spells (N), the 3 components are as follows

- **Benefit duration**

$$L_v(t_v \quad | \quad \varepsilon^v) = \prod_{i=1}^{N} \left[\prod_{s=1}^{t_{vi}-1} \exp(-\exp(\mathrm{Z}_v\text{'}\eta_v + \log \varepsilon^v + h_0^*(s;\alpha))) \right] .$$
$$\{1 - \exp(-\exp(\mathrm{Z}_v\text{'}\eta_v + \log \varepsilon^v + h_0^*(t_{vi};\alpha)))\}$$

- **Unemployment Duration**

$$L_u(t_u \quad | \quad \varepsilon^u) = \prod_{i=1}^{N} [1 - \exp(-exp(Z_u'\eta_u + \delta X(t_{ui}) + \log \varepsilon^u + h_0^*(t_{ui};\beta)))]^{c_i^u}$$
$$\left[\prod_{s=1}^{t_{ui}-1} \exp(-exp(Z_u'\eta_u + \delta X(s) + \log \varepsilon^u + h_0^*(s;\beta))) \right]$$

[11]See Meyer (1990) for an example devoted to the semi-parametric estimation of the distribution of unemployment spells and the effect of UI benefit duration.

The censoring indicator, c_i^u, equals 1 if a spell is completed (between t_{ui} and $t_{ui}+1$) and 0 if right censored.

- **Accepted Job Duration**

$$L_j(t_{ji} \quad | \quad \varepsilon^j) = \prod_{i=1}^{M} \left[1 - \exp(-exp(Z_j'\eta_j + \zeta_0.t_u + \zeta_1.t_{u,50} + \lambda.X(t_u) + \log\varepsilon^j + h_0^*(t_{ji};\gamma)))\right]^{c_i^j}.$$
$$\left[\prod_{s=1}^{t_{ji}-1} \exp(-exp(Z_j'\eta_j + \zeta_0.t_u + \zeta_1.t_{u,50} + \lambda.X(t_u) + \log\varepsilon^j + h_0^*(s;\gamma)))\right]$$

where c_i^j is the censoring indicator for accepted job duration and is defined similarly as c_i^u. Given these definitions, the conditional likelihood function, $L(t_{vi}, t_{ui}, t_{ji} \mid \varepsilon^v, \varepsilon^u, \varepsilon^j)$, is simply

$$L(t_v, t_u, t_j \mid \varepsilon^v, \varepsilon^u, \varepsilon^j) = L_v(\varepsilon^v).L_u(\varepsilon^u).L_j(\varepsilon^j) \tag{2.6}$$

2.3. Unobserved Heterogeneity

Estimation of the model by likelihood techniques requires that the individual unobserved heterogeneity terms be integrated out. In the paper, I consider the case where ε^u and ε^j follow a bi-variate discrete distribution and where both ε^u and ε^j have two points of support. The distribution is summarized as follows

$$\Pr(\varepsilon^u = \varepsilon_1^u, \varepsilon^j = \varepsilon_1^j) = p_1$$

$$\Pr(\varepsilon^u = \varepsilon_2^u, \varepsilon^j = \varepsilon_1^j) = p_2$$

$$\Pr(\varepsilon^u = \varepsilon_1^u, \varepsilon^j = \varepsilon_2^j) = p_3$$

$$\Pr(\varepsilon^u = \varepsilon_2^u, \varepsilon^j = \varepsilon_2^j) = p_4$$

for $\varepsilon_1^u > \varepsilon_2^u$ and $\varepsilon_1^j > \varepsilon_2^j$. In this case, 4 points of support and three free probabilities need to be estimated (for more details, see van den Berg et Al., 1994). It can be shown that the correlation between ε^u and ε^j can be evaluated by the following expression;

$$Corr(\varepsilon^u, \varepsilon^j) = \frac{p_1 p_{4-} p_2 p_3}{\sqrt{(p_1+p_3)(p_2+p_4)(p_1+p_2)(p_3+p_4)}} \tag{2.7}$$

Finally, when I consider the case where maximum benefit duration is endogenous, I assume that the individual effect in the benefit duration is such that[12]

$$\varepsilon^v = \beta_v . \varepsilon^u$$

The likelihood function to be maximized is the average of (2.6) over the four cases possible.[13] In order to implement the model, I define p_1, p_2, p_3 and p_4 as

$$P_i = \frac{\exp(q_i)}{\sum_{j=1}^{4} \exp(q_j)} for\ i = 1, 2, 3, 4$$

and I fix q_4 to 0. Standard errors for all p's can be obtained using the delta method. It follows that the restriction needed to impose independence (a correlation of 0) between ε^u and ε^j is simply $q_3 = q_1 - q_2$. Testing for independence (given that $\varepsilon_1^u \neq \varepsilon_2^u$ and $\varepsilon_1^j \neq \varepsilon_2^j$) can be achieved with a likelihood ratio statistic which has a χ_1^2 distribution under the null hypothesis.

The model is estimated with the Maximum Likelihood application in Gaussi 3.2.26 on a Pentium 200. The log likelihood function is maximized using the BFGS and the BHHH algorithms.

[12] Given that potential weeks of benefit duration vary with weeks worked during the previous year, it would also be possible to specify the benefit duration unobserved heterogeneity term as a linear function of job duration unobserved heterogeneity.

[13] I have also considered a case where there is a univariate heterogeneity term (for unemployment duration) and where heterogeneity in accepted job and benefit duration are both defined as the product of this heterogeneity term times a loading parameter.

3. The Data

The econometric models, presented in the previous section, are estimated from a panel of Canadian labor force participants which is extracted from the Longitudinal Labor Force File of Employment and Immigration Canada. In what follows, I provide a brief discussion of the Canadian UI system (Section 3.1), a description of the original data set (Section 3.2) and then explain the sampling method (Section 3.3).

3.1. The Canadian Unemployment Insurance System

The Canadian Unemployment Insurance system (now called Employment Insurance) was established in 1940. As most UI systems in western countries, its fund is financed by premia collected from employers and employees. After having remained intact between 1940 and 1971, Canada's UI system was changed substantially in 1971. The increase in coverage and in the benefit rate, which took place in 1971-72, were substantial. At the same time, the maximum benefit period was extended and the minimum period of employment required to qualify for benefits was reduced. Although the 1971-72 changes were partially reversed by changes made between 1977 and 1979 (small reduction in the replacement ratio and in the maximum benefit duration, the Canadian UI system remained quite generous over the period of the current analysis.

The analysis presented in this paper is based on a sample of young males who have experienced a layoff between January 1976 and February 1978. Over this period, the average benefit period was slightly decreased after changes to the UI regulations. This change took place in September 1977. In the sample used in this study, the difference in maximum benefit duration between those who have experienced job separation before September 77 and those after is around 3 weeks.[14] For all those individuals who have experienced a layoff in between January 1976 and February 1978, the benefit rate has however remained constant at 66% of insurable earnings. The maximum insurable earnings are typically adjusted yearly to reflect changes in the average industrial salary. Over the sample period, individuals had to work between 10 to 14 weeks in order to qualify for benefit. The potential benefit duration is calculated from the number of weeks of employment over the 52 week period and the local rate of unemployment at the

[14] For more details, see Belzil (1995).

time. Except for those working in fishing industry (excluded in this study), there are no variations in UI rules according to industry.

3.2. Description of the Data

The data are constructed as an event history data set and covers a period going from January 1972 until December 1984. It contain several pieces of information about employment and unemployment spells of a random sample of Canadian labor force participants. The data are actually based on a merge of several administrative files such as Records of Employment (ROE) and the Unemployment Insurance administrative files and they enable the researcher to recreate the sequence of labor market states occupied by a given individual. As it is usually the case with administrative data, information on insured spells of unemployment (such as benefit durations and the weekly benefit level) are relatively accurate. However, the data are much less reliable when it comes to evaluating the labor market status of those unemployed for a relatively long period, especially those unemployed beyond benefit termination.

The Records of Employment (ROE) identify the reason for separation and provide information about job tenure, age, experience and industry. In Canada, firms are legally required to issue a ROE for every job separation that takes place. The measure of experience available is the total number of weeks of employment from 1972 until 1984. It is therefore reliable for younger workers. The Unemployment Insurance file, along with some partial income tax records file, gives information about potential benefit duration for the unemployed, weekly insurable earnings, unemployment duration, UI benefit level and the number of weeks of benefit entitlement left when a new job is accepted. The employer code available is used to identify individuals who have been laid off and returned with the same employer subsequently.

3.3. Sampling Method

The data set used in this paper contains 2610 individual records of labor market histories for young males who have suffered a job separation between January 1976 and February 1978. Each individual was between 18 and 25 at the time of the job separation and was followed until 1984 through administrative records. Consequently, the number of spells attached to each records varies considerably across individuals. Out of these 2610 records, 1910 are coded as layoffs while 700 individuals quit to become unemployed.

As a first step, I retain the 1910 individuals who have experienced a layoff. As a second step, I exclude the 1001 cases where unemployment was followed by a job with the previous employer as well as the 700 individuals quo quit their job. This is because those individuals who returned to their previous employer are most likely to be on temporary layoffs and are likely to have a distinct search behavior than those who were displaced permanently while those who quit their job are more likely to have a weaker attachment to the labor market.[15] The resulting sample contains 909 individuals who have experienced an involuntary job separation. Most of these individuals (889 cases) have found a new job with a different employer while 20 of them have been lost by the UI authorities which means that I observe no subsequent job duration for these individuals. Given that reported unemployment duration is likely to be unreliable for these individuals, I censor unemployment duration at 50 weeks. In terms of the subsequent jobs, these 909 individual records are classified as followed;

- 464 subsequent job durations which were later terminated by a layoff
- 289 subsequent job durations terminated by a quit.
- 136 subsequent job spells still in progress as of 1984
- 20 cases where the unemployment spell is the last recorded state.

With administrative data, it is quite difficult to collect information on individuals who become non-participants. Among all individuals experiencing a layoff during this period (not only young workers), only 3.7% have actually left the accepted job to leave the labor force. However, as this information is actually estimated from the existence of subsequent employment records, this number can only be viewed as an estimate. As I look only at young males whom are well known to have a high turnover rate, only 15% (136/909) are still employed with same employer at the end of 1984. To summarize, each individual contributes, at most, one unemployment duration-accepted job duration sequence. Some summary statistics are found in Table1.

[15] Belzil (1995) has performed a separate analysis of those who are on temporary layoffs and found that the effect of UI benefit on unemployment duration and re-employment duration (unemployment incidence) differed substantially from the effects obtained for those individuals who accepted a new job. However, benefit duration was treated as an exogenous regressor.

Table 1

Some Sample Statistics

Variable	**Mean**	**Stand. Dev.**
Experience (weeks)	127	61
Previous Earnings (1977 dollars)	240	120
Duration of unemployment (weeks)	14	18
Maximum Benefit Period (weeks)	33	14
Potential Benefit period (at re-employment)	6	3
Unemployment Benefits (1977 dollars)	122	30
Previous Job Duration (weeks)	22	29
Accepted Earnings (current dollars)	223	102
Duration of Accepted job (weeks)	35	65
% in Primary Sector	7.1	-
% in Construction	7.2	-
% in Manufacturing	18.7	-
% in Transportation	11.0	-
% in Trade	13.6	-
% in Finance	11.2	-
% in Services	15.4	-
% in Administration	15.3	-

Note:

Earnings and unemployment benefits are measured in 1977 Canadian dollars. For the period over which job separation took place, the maximum benefit level was around 145$ per week.

4. Empirical results

In this section, the main results are presented. First, in Section 4.1, I consider a model specification where the effect of potential benefit duration on subsequent job duration is assumed to be constant over the spell of unemployment. The flexible model specification is in Section 4.2.

4.1. The Relationship between Potential Benefit Duration and Subsequent Job Duration

I present estimates obtained where maximum benefit duration is endogenous (Table 2A) as well as estimates obtained from a more conventional specification where maximum benefit duration is exogenous (Table 2B). In both cases, I consider a specification where past earnings are introduced (column 1) and one when it is omitted (column 2). I do so because previous earnings are possibly correlated with either $\varepsilon^u, \varepsilon^v$ or ε^j. Finally, the parameter estimates for age, experience and industry dummy regressors are not reported in the main tables to save space but are found in Appendix 1.

The results for the model with endogenous maximum benefit duration are in Table 2A. The estimates in column 1 were obtained when previous earnings were included in the unemployment hazard function. I also use a binary variable for the case where previous earnings are top coded (equal to the maximum insurable earnings). The estimate for β^v (-.70) indicates a negative correlation between unobserved heterogeneity affecting maximum benefit duration and unobserved heterogeneity affecting unemployment duration[16]. The effect of previous earnings

[16] Several experiments have shown that most of the fundamental results regarding the effects of UI benefit duration on unemployment duration and job duration are quite robust to the exclusion of job duration from the benefit duration equation. The estimates that vary most significantly are experience and the correlation between unemployment and benefit duration heterogeneity (β^v). I have also experimented with the possibility that the benefit duration heterogeneity term is specified as a linear function of accepted job duration heterogeneity and previous job duration is used in the benefit duration equation, the correlation between benefit duration heterogeneity and accepted job duration heterogeneity is found to be positive insignificant. When previous job duration is ignored, the correlation between benefit duration heterogeneity and accepted job duration heterogeneity is found to be positive and significant.

on the exit rate out of unemployment is positive (0.31) and indicates that, other things equal, those with higher pre-unemployment earnings leave unemployment faster. The effect of benefit level is negative (-0.25) and indicates that, all else equal, those receiving higher level of benefits tend to leave unemployment more slowly. Both results are standard in the literature. The estimate for β (0.93) indicates the presence of mild negative duration dependence (the hazard function decreases with elapsed unemployment). Finally, the results in Appendix 1 indicate that unemployment hazards increase with age and experience.

As it has been found in the literature, the effect of potential benefit duration (captured in $\delta_{X(t_u)}$ and $\delta_{X(1-5)}$ to $\delta_{X(20-29)}$) varies with the level of potential benefit duration. The value of the coefficient for maximum benefit period (-.0002) indicates that, until 29 weeks to benefit termination, potential benefit duration has practically no impact on unemployment hazards. However, after adding up the appropriate parameters, it can be seen that the escape rate out of unemployment increases as individuals approach benefit termination. In particular, the effect of an additional week of potential benefit duration is particularly strong up to 9 weeks preceding benefit exhaustion, around -0.37 ($-0.028 - .0.062 - .276$) between 1 to 5 weeks, -0.09 between 6 to 9 weeks, and weaker during the period going from week 19 to week 10 prior to benefit termination. A likelihood ratio test for the joint significance of $\delta_{X(1-5)}$ to $\delta_{X(20-29)}$ rejects the null hypothesis that the parameters are equal to 0 at a 2% level (the p value is 0.011).

Turning to the estimates of the effects of benefit duration and unemployment duration on subsequent job hazards, potential benefit duration left when a new job was accepted has a negative effect on subsequent job hazards (-0.038 with a t-ratio of 1.74) but the estimate is not really significant. After controlling for potential benefit duration left, an increase in the duration of unemployment increases subsequent job hazards (the estimate is 0.01 with a t-ratio of 1.26). The scale parameter for subsequent job duration ($\gamma = 0.85$) indicates the presence of negative duration dependence; that is the conditional probability of subsequent job termination is declining sharply. The results, in Appendix 1, indicate that accepted job hazards decrease with age and experience. Finally, the distribution parameters (ε^u and ε^j and their respective probabilities) indicate that there is a negative correlation between unobserved heterogeneity in unemployment duration and subsequent job duration. The correlation is found to be relatively high; -0.52). The illustration of the effects of benefit duration on unemployment and accepted job hazards is delayed to Section 5.

The estimates in column 2 are those obtained when previous earnings have been excluded from the unemployment duration equation. By far, the effect of benefit level on unemployment hazards is the parameter most sensitive to the exclusion of previous earnings. The effect of benefit level goes from -0.25 (when earnings are included) to -0.38 (when earnings are excluded). However, the estimates for the effect of potential benefit duration on unemployment hazards are more or less comparable ($\delta_{X(1-5)}$ is however slightly smaller at -0.21) and the likelihood ratio test for the joint significance of $\delta_{X(1-5)}$ to $\delta_{X(20-29)}$ rejects the null hypothesis that the parameters are equal to 0 also at a 3% level (the p value is 0.016). An increase in potential benefit duration is found to decrease the subsequent job hazard by -0.039 but the estimate is still insignificant (with a t-ratio of 1.54). Finally, the distribution parameters still indicate a negative correlation (-0.47) between unemployment duration and subsequent job duration.

At this stage, I find very little evidence in favor of the matching hypothesis. Potential benefit duration appears unrelated to subsequent job duration as the null hypothesis that benefit duration has no effect on accepted job hazards fails to be rejected. Indeed, the correlation between subsequent job duration and unemployment duration seems to be explained by unobserved heterogeneity. Given that maximum benefit duration is probably exogenous for a certain fraction of the unemployed (and therefore uncorrelated with unobserved heterogeneity), it seems important to evaluate the sensitivity of the results to the endogeneity assumption.

In Table 2B, I have re-estimated the model presented in Table 2A under the more conventional assumption that maximum benefit duration is exogenous. When maximum benefit duration is assumed to be exogenous, the effect of benefit level seems to be weaker (-0.17 in column 1 and -0.25 in column 2). The effects of potential benefit duration are however quite similar to those obtained in Table 2A. There is again a strong decrease in hazard rates when potential benefit duration is higher (-.72 between week 1 and week 5 in column 1 and -0.60 in column 2). Interestingly, the effect of potential benefit duration on subsequent job hazards is even weaker when maximum benefit duration is exogenous. The estimates are still negative (-0.0252 in column 1 and -0.0222 in column 2) and clearly insignificant. After controlling for potential benefit duration, the effect of completed unemployment duration is also quite weak (0.0175 and 0.0203) and relatively insignificant.

To summarize, I find very little evidence in favor of a matching effect of unemployment insurance benefit duration. This seems to be true regardless of the inclusion of previous earnings and the effect of potential benefit duration appears particularly weak when maximum benefit duration is assumed to be exogenous. After investigating these issues with a model specification which restricted the effects of potential benefit duration to be constant over a spell of unemployment, I now turn to a more flexible model specification.

4.2. A Flexible Model for the Effects of Potential Benefit Duration

In what follows, I present the estimates obtained when the effect of potential benefit duration on subsequent job duration is estimated more flexibly. I use a procedure similar to the one used for the effect of potential benefit duration on the unemployment hazard and allow the effect of an additional week of potential benefit duration on accepted on hazards to vary accordingly. Again, I examine a model where maximum benefit duration is endogenous (Table 3A) and a version where maximum benefit duration is exogenous (Table 3B).

Turning to Table 3A (column 1), I find again a negative correlation between benefit duration heterogeneity and unemployment duration heterogeneity ($\beta^v = -0.72$). The effect of previous earnings is still positive (0.31) while an increase in UI benefits decreases the hazard rate out of unemployment (-0.26). The estimates for the effect of potential benefit duration have not changed much; an increase (decrease) in potential benefit duration decreases (increases) unemployment hazard by a factor -0.51 between 1 to 5 weeks from benefit termination (after adding up all relevant parameters) but does not seem to matter really when individuals are more than 10 weeks away from benefit termination. The estimate for β (0.94) indicates that the escape rate out of unemployment is decreasing slightly with elapsed unemployment duration.

The specification used in Table 3A and 3B (with $\lambda_{X(1-5)}..\lambda_{X(20-29)}$) allows me to compare the exit rate out of the subsequent job at various levels of potential benefit duration left when the new job was accepted. I find a pattern relatively similar to the one observed in unemployment hazards; the effect of potential benefit is stronger (negative) around benefit termination. However, the parameter estimates are generally much lower (in absolute values) than the effect of potential benefit duration on unemployment hazards (the $\delta's$). Accepting a new job with one more week of potential benefit duration decreases the dissolution rate of the subsequent job by around -0.025 $(0.010 - 0.035)$ within 5 weeks of benefit termination. However, potential benefit duration appears insignificant (and quite erratic) as we move away from benefit termination; the marginal effect of an additional week of benefit is positive (0.007) between week 6 and week 9 and practically 0 beyond 10 weeks of benefit termination. After controlling for benefit duration, the effect of an additional week of unemployment on subsequent job hazards is positive (0.0154) but insignificant. A likelihood ratio test for the joint significance of the parameters capturing the effects of benefit duration on accepted job hazards ($\lambda_{X(1-5)}$...=...$\lambda_{X(20-29)} = 0$) fails to be rejected at the 5% level but is rejected at the 7% level (p value is 0.068). Like in table 2A and Table 2B, I find a negative correlation (-0.34) between unemployment duration and subsequent job duration unobserved heterogeneity. Finally, the scale parameter of the subsequent job hazard function ($\gamma = 0.8891$) indicates negative duration dependence.

Table 2A
Endogenous Benefit Duration

		(1)	(2)
Benefit duration	Duration dependence (α)	1.0421 (10.34)	1.0418 (12.08)
	β^v	-.7012 (3.12)	-.7453 (3.98)
Unemp. duration	Log Earnings	0.3114 (3.12)	-
	Benefit level (log)	-0.2506 (2.45)	-0.3815 (3.06)
	Duration Dependence	0.9316 (6.89)	0.9224 (8.76)
Unemp. duration	Potential Benefit Duration		
	$\delta_{X(1-5)}$	-0.2762 (4.13)	-0.2134 (3.67)
	$\delta_{X(6-9)}$	-0.0623 (2.21)	-0.0534 (2.65)
	$\delta_{X(10-19)}$	-0.0283 (1.78)	-0.0304 (1.69)
	$\delta_{X(20-29)}$	0.0004 (0.56)	0.0005 (0.52)
	$\delta_{X(t_u)}$	-0.0002 (0.23)	-0.0004 (0.87)
Unobs. hetero			
	Het. Support Points		
	ε_1^u	0.6424 (4.98)	0.6231 (5.01)
	ε_2^u	0.3583 (3.99)	0.3325 (4.18)
	ε_1^j	0.3236 (5.21)	0.2795 (3.22)
	ε_2^j	0.1901 (4.22)	0.1634 (2.87)
	Probabilities		
	P_1	0.1200 (2.04)	0.1500 (2.78)
	P_2	0.3500 (3.04)	0.2900 (2.67)
	P_3	0.4100 (1.65)	0.4500 (3.04)
Correlation	$Corr(\varepsilon^u, \varepsilon^j)$	-0.5236	-0.4703
Accepted Job Dur	Benefit duration left	-0.0386 (1.74)	-0.0390 (1.54)
	Unemployment Duration	0.0104 (1.26)	0.0145 (1.35)
	Unemp dur$_{50-\infty}$	0.5604 (1.76)	0.5463 (1.69)
	Dur. dependence (γ)	0.8518 (5.12)	0.8490 (6.23)
Log Likelihood		-1346.7	-1350.2

Note: Asymptotic t-ratios in parentheses. The p-value for the likelihood ratio test for the null that $\delta_{X(1-5)}..=..\delta_{X(20-29)} = 0$ **was 0.011 in column 1 and 0.006 in column 2.**

Table 2B
Exogenous Benefit Duration

		(1)	**(2)**
Benefit duration	Duration Dependence (α)		
	β^v		
Unemp. duration	Log Earnings	0.4122 (2.37)	-
	Benefit Level (log)	-0.1723 (1.98)	-0.2456 (3.04)
	Duration Dependence (β)	0.9003 (5.75)	0.8934 (7.46)
Unemp. duration			
	Remaining Benefit Duration		
	$\delta_{X(1-5)}$	-0.3216 (3.79)	-0.2865 (3.25)
	$\delta_{X(6-9)}$	-0.2913 (2.32)	-0.2543 (2.12)
	$\delta_{X(10-19)}$	-0.1002 (1.98)	-0.0578 (2.00)
	$\delta_{X(20-29)}$	0.0056 (0.91)	0.0068 (1.08)
	$\delta_{X(t_u)}$	-0.0176 (0.67)	-0.0074 (0.62)
Unobs. hetero			
	Het. Support Points		
	ε_1^u	0.4623 (4.03)	0.5745 (4.76)
	ε_2^u	0.3743 (3.45)	0.3425 (2.97)
	ε_1^j	0.2957 (3.87)	0.2956 (2.56)
	ε_2^j	0.2005 (3.12)	0.2004 (2.29)
	Probabilities		
	P_1	0.1005 (1.56)	0.1734 (2.05)
	P_2	0.4734 (4.03)	0.4328 (3.97)
	P_3	0.3645 (4.28)	0.3845 (2.95)
Correlation	$Corr(\varepsilon^u, \varepsilon^j)$	-0.5061	-0.7012
Accepted job dur	Benefit duration left	-0.0252 (1.27)	-0.0222 (1.46)
	Unemployment Duration	0.0175 (1.69)	0.0203 (1.84)
	Unemp dur$_{50-\infty}$	0.3645 (1.89)	0.3326 (1.92)
	Dur. dependence (γ)	0.8435 (5.28)	0.8406 (5.25)
Log likelihood		-1373.5	-1376.7

Note: Asymptotic t-ratios in parentheses. The p-value for the likelihood ratio test for the null that $\delta_{X(1-5)}..=..\delta_{X(20-29)} = 0$ **was 0.021 in column 1 and 0.003 in column 2**.

The results obtained in column 2 of Table 3A (when previous earnings are excluded) are quite consistent with those of column 1. Again, an increase (decrease) in potential benefit duration decreases (increases) unemployment hazard by a factor 0.50 around benefit termination (between 1 to 5 weeks from benefit termination) but does not seem to matter really when individuals are more than 10 weeks away from benefit termination. The estimates of the effect of potential benefit duration on subsequent job hazards indicate that most of the matching effects of UI benefits are located between 1 and 5 weeks and 10 to 19 weeks from benefit exhaustion. The test for the joint significance of $\lambda_{X(1-5)}..\lambda_{X(20-29)}$ rejects the null hypothesis at 5% and indicates that there are significant matching effects (the p-value is 0.047).

The results obtained when maximum benefit duration is exogenous are in Table 3B. Again, the effect of potential benefit duration on subsequent job hazards is negative and significant as individuals approach benefit termination but, except for $\delta_{X(1-5)}$ and $\delta_{X(6-9)}$, the other parameter estimates attain a very low level of significance. The matching effects of UI benefit duration appear to be the largest in Table 3B. In column 1, an additional week of potential benefit duration reduces the subsequent job hazard by -0.07 within 5 weeks of benefit termination and -0.04 between week 6 and week 9. The other parameter estimates do not appear to have changed much and there is still a strong negative correlation between unemployment unobserved heterogeneity and subsequent job heterogeneity. The exclusion of previous earnings (column 2) did not change the results very much. The effects of benefit duration on subsequent job duration is again around -0.07 within 5 weeks from benefit exhaustion and -0.04 between 6 weeks and 10 weeks of benefit termination. In both cases, the likelihood ratio test for the joint significance of $\lambda_{X(1-5)}$ to $\lambda_{X(20-29)}$ is rejected at the 5% level (the respective p values are 0.053 and 0.042).

As the results indicate that the correlation between completed unemployment duration and subsequent job duration is typically negative, it appears important to examine the level of significance of the correlation between unemployment and subsequent job heterogeneity. Independence between ε^u and ε^j can be tested by imposing the following restriction $p_1 \cdot p_4 - p_3 \cdot p_2 = 0$. To do so, I have re-estimated all model specifications of Table 3A and Table 3B under the maintained hypothesis that ε^u and ε^j are independent. In all cases, the null hypothesis are strongly rejected[17]. The likelihood ratio statistics obtained for Table 3A are 8.1 (column

[17]These parameter estimates are however not reported here.

1) and 7.2 (column 2). In Table 3B, the likelihood ratio statistics are 9.1 (column 1) and 7.8 (column 2). In all cases, the test statistics exceed a χ^2_{1df} critical value at any reasonable level of confidence.

Overall, when the effects of potential benefit duration is estimated flexibly, I find support for a positive effect of unemployment benefit duration on subsequent job match quality. The results indicate that both the effects of benefit duration on unemployment hazards and accepted job hazards are located mostly within 10 weeks of benefit termination. More precisely, while potential benefit duration raises unemployment duration, it also raises subsequent job duration. At the same time, I find strong evidence of a negative (and significant) correlation between unemployment duration and job duration. Interestingly, the results are more or less invariant to the allowance for potential endogeneity of the maximum benefit duration.

As the model used to obtain the estimates is highly non-linear, I now turn to a descriptive presentation of the results which can illustrate how potential benefit duration affects unemployment and accepted job hazards. In particular, I shall investigate the relative importance of the matching hypothesis vs the unobserved heterogeneity hypothesis and simulate how an increase in benefit duration can affect both unemployment and accepted job durations.

Table 3A
Flexible Model-Endogenous Benefit Duration

		(1)	**(2)**
Benefit duration	Duration dependence (α)	1.3961 (5.32)	1.0392 (7.34)
	β^v	-0.7204 (2.31)	-0.7193 (2.87)
Unemp. duration	Log Earnings	0.3136 (2.56)	-
	Benefit level (log)	-0.2615 (2.86)	-0.2980 (2.55)
	Duration Dependence (β)	0.9389 (7.34)	0.9375 (6.88)
Unemp. duration			
	Potential Benefit Duration		
	$\delta_{X(1-5)}$	-0.2512 (2.56)	-0.2519 (2.76)
	$\delta_{X(6-9)}$	-0.2043 (1.98)	-0.2045 (2.03)
	$\delta_{X(10-19)}$	-0.0578 (1.22)	-0.0534 (1.28)
	$\delta_{X(20-29)}$	0.0005 (0.49)	0.0005 (0.97)
	$\delta_{X(t_u)}$	-0.0010 (1.06)	-0.0014 (1.00)
Unobs. heterogeneity			
	Hetero. Support Points		
	ε_1^u	0.6718 (2.64)	0.6238 (2.88)
	ε_2^u	0.3214 (3.12)	0.3674 (3.01)
	ε_1^j	0.4916 (1.98)	0.5512 (2.24)
	ε_2^j	0.2006 (1.56)	0.2411 (1.82)
	Probabilities		
	P_1	0.1300 (2.60)	0.1600 (2.75)
	P_2	0.3700 (2.11)	0.3800 (2.02)
	P_3	0.3000 (1.59)	0.29001.76)
Correlation	$Corr(\varepsilon^u, \varepsilon^j)$	-0.3400	-0.3300

Table 3A-Continued

Accepted job duration			
	Duration Dependence (γ)	0.8891 (5.34)	0.8795 (6.02)
	Unemp. Duration	0.0154 (1.69)	0.0162 (1.72)
	Unemp. Dur_{50}	0.2936 (2.42)	0.2893 (1.89)
	Benefit Duration Left		
	$\lambda_{X(1-5)}$	-0.0345 (1.88)	-0.0239 (1.90)
	$\lambda_{X(6-9)}$	0.0103 (1.21)	0.0253 (1.18)
	$\lambda_{X(10-19)}$	-0.0032 (1.12)	-0.0313 (1.08)
	$\lambda_{X(20-29)}$	-0.0001 (0.78)	-0.0001 (0.59)
	$\lambda_{X(t_u)}$	0.0002 (0.34)	0.0002 (0.39)
Log likelihood		-1315.6	-1319.7
LR test for corr($\varepsilon^u, \varepsilon^j = 0$)		8.1	7.2

Notes:
Asymptotic t-ratios in parentheses.
The p-value for the likelihood ratio test for the null that $\lambda_{X(1-5)}..=..\lambda_{X(20-29)} = 0$ was 0.068 in column 1 and 0.047 in column 2.
The likelihood Ratio statistic is for the null that ε^u and ε^j are independent.

Table 3B
Flexible Model-Exogenous Benefit Duration

		(1)	(2)
Benefit duration	Duration dependence (α)	-	-
	β^v	-	-
Unemp. duration	Log Earnings	0.3822 (2.88)	-
	Benefit level (log)	-0.2834 (3.01)	-0.2935 (2.92)
	Duration Dependence (β)	0.9256 (6.68)	0.9221 (6.58)
Unemp. duration			
	Remaining Benefit Duration		
	$\delta_{X(1-5)}$	-0.3076 (3.01)	-0.3325 (2.64)
	$\delta_{X(6-9)}$	-0.1978 (2.27)	-0.2032 (2.36)
	$\delta_{X(10-19)}$	-0.0567 (2.02)	-0.0612 (1.93)
	$\delta_{X(20-29)}$	0.0126 (1.18)	0.0139 (0.85)
	$\delta_{X(t_u)}$	-0.0307 (0.83)	-0.0298 (0.56)
Unobs. heterogeneity			
	Hetero. Support Points		
	ε_1^u	0.5387 (2.89)	0.5832 (3.12)
	ε_2^u	0.3628 (2.28)	0.3394 (2.42)
	ε_1^j	0.3923 (3.18)	0.3623 (2.90)
	ε_2^j	0.2638 (2.02)	0.2341 (1.79)
	Probabilities		
	P_1	0.0645 (2.00)	0.0589 (1.83)
	P_2	0.3745 (2.57)	0.4005 (3.24)
	P_3	0.3834 (3.04)	0.3905 (2.36)
Correlation	$Corr(\varepsilon^u, \varepsilon^j)$	-0.6020	-0.6201

Table 3B**-Continued**

Accepted job duration			
	Duration Dependence (γ)	0.8471 (4.24)	0.8395 (5.15)
	Unemp. Duration	0.0256 (1.68)	0.0200 (1.68)
	Unemp. Dur_{50}	0.2015 (1.67)	0.1956 (1.74)
	Benefit Duration Left		
	$\lambda_{X(1-5)}$	-0.0247 (1.88)	-0.0327 (1.90)
	$\lambda_{X(6-9)}$	-0.0393 (1.68)	-0.0356 (1.65)
	$\lambda_{X(10-19)}$	-0.0067 (1.02)	-0.0074 (1.39)
	$\lambda_{X(20-29)}$	0.0012 (0.57)	0.0059 (0.95)
	$\lambda_{X(t_u)}$	0.0056 (0.89)	0.0059 (0.78)
Log likelihood		-1343.9	-1348.0
LR test for $corr(\varepsilon^u, \varepsilon^j = 0)$		9.1	7.8

Notes:
Asymptotic t-ratios in parentheses.
The p-value for the likelihood ratio test for the null that $\lambda_{X(1-5)}..=..\lambda_{X(20-29)} = 0$ was 0.0531 in column 1 and 0.042 in column 2.
The likelihood Ratio statistic is for the null that ε^u and ε^j are independent.

5. Illustrating the Effects of Benefit Durations on Unemployment and Accepted Job Hazards

In this section, I present a summary of predicted unemployment and accepted job hazards (Section 5.1). The predicted hazards are used to illustrate how benefit duration affects unemployment hazards and accepted job hazards. After illustrating the behaviour of empirical hazards, I compare the variance in predicted job durations explained by variations in benefit durations as opposed to the variance explained by unobserved heterogeneity. Finally, in Section 5.2, I simulate the effects of increasing maximum benefit duration on mean unemployment and mean accepted job durations

5.1. Predicted Unemployment and Accepted Job Hazards

In order to illustrate the implications of the parameter estimates of Table 3A and Table 3B for unemployment and accepted job hazards, I report predicted hazards in Table 4A (for the parameters of Table 3A where benefit duration is endogenous) and Table 4B (for the parameters of Table 3B where benefit duration is exogenous). The predicted hazards are reported for various levels of potential benefit duration ranging from 40 weeks to 1 week. The observed characteristics affecting the hazard rates are set at their sample average (see Table 1). The variation in unemployment hazards (in the first and third column of Table 4A and Table 4B respectively) is therefore explained simultaneously by a decrease in remaining benefit duration and negative duration dependence in unemployment. The reported accepted job hazards (in column 2 and column 4 of Table 4A and Table 4B) are computed in week 1 of the accepted job. As for unemployment hazards, accepted job hazards are computed by setting observed characteristics to their sample average. The observed variations is therefore due only to variations in potential benefit duration at the time when a new job was accepted.

The results for unemployment hazards indicate that, at high levels of benefit duration (beyond 20 weeks), benefit duration has virtually no effect on unemployment hazards. As a consequence, predicted unemployment hazards are initially decreasing (because of negative duration dependence) and tend to increase within 20 weeks of benefit termination.[18] Between 10 weeks to 1 week from benefit

[18]The effects of potential benefit duration on unemployment hazards would clearly be higher if, unlike what was done in table 4A and 4B, I had fixed the level of the Weibull baseline hazard

termination, unemployment hazards raise from 0.06 to 0.08. Although the parameters of the accepted job hazard functions (presented in the previous section) indicated a smaller marginal effect of potential benefit duration than for unemployment hazards, the effect of potential benefit duration on accepted job hazards are comparable. This is a reflection of the fact accepted job hazards are computed in the first week of the accepted job and are therefore not affected by duration dependence.[19] The predicted job hazards (Table 4A and Table 4B) indicate that the job matching effects are present within 10 weeks of benefit termination; the accepted job hazards remain around 0.04 between 40 and 20 weeks prior to benefit termination and can eventually raise up to 0.09.

Given that the correlation between unemployment duration and accepted job duration is explained partly by differences in benefit duration and partly by unobserved heterogeneity, it is natural to investigate the relative importance of these two competing explanations for the observed correlation between unemployment and accepted job duration. First, I have computed an expected job duration for every individual (given maximum benefit duration) after having fixed individual characteristics and unobserved heterogeneity to their respective average value.[20] Then, in order to capture the importance of benefit duration in explaining variations in mean accepted job duration, I computed the standard deviation of the expected accepted job durations. The resulting value is an indication of the job duration variability explained by individual variations in maximum benefit duration. In order to investigate the importance of unobserved heterogeneity, I computed an expected accepted job duration for a maximum benefit entitlement of 33 weeks (the sample average) for a representative individual (after having fixed individual characteristics to their sample average) and computed the variability due to unobserved heterogeneity by letting the values of job duration heterogeneity vary according to the estimated probabilities.

The results, found in Table 4C, tend to indicate the relative importance of unobserved heterogeneity and individual variations in maximum benefit period in explaining variations in expected job durations. The estimates of column 1

function to a fixed level.

[19]Whether accepted job hazards are computed at the first week of the accepted job or later, is irrelevant. I have computed predicted job hazards at week 10 and found a similar tendency in the results.

[20]As there exist no closed-form solution for the mean of a Weibull random variable, I used the computed densities for each discrete interval of one week and used the mid-point of each interval as a support point.

indicate that, when unobserved heterogeneity is fixed to a common value (as well as observed characteristics), the standard deviations in mean unemployment durations is around 3 weeks. When benefit duration is fixed to 33 weeks and observed characteristics to sample averages, standard deviations are around 5 weeks. Overall, unobserved heterogeneity appears to account for a larger fraction of mean job durations than UI benefit potential durations.

Table 4A
Flexible Model-Endogenous Benefit Duration
Predicted Unemployment and Accepted Job Hazards

	(1) **Table 3A (1)**	(2) **Table 3A (1)**	(3) **Table 3A (2)**	(4) **Table 3A (2)**
Earnings	**Unemp. Haz.** included	**Acc. Job Haz.** included	**Unemp. Haz.** excluded	**Acc. Job Haz.** excluded
Weeks left				
1 week	0.0841	0.0314	0.0851	0.0787
5 weeks	0.0713	0.0291	0.0724	0.0707
10 weeks	0.0603	0.0290	0.0601	0.0594
15 weeks	0.0593	0.0289	0.0590	0.0420
20 weeks	0.0516	0.0287	0.0520	0.0390
25 weeks	0.0498	0.0265	0.0501	0.0389
30 weeks	0.0501	0.0261	0.0497	0.0389
35 weeks	0.0531	0.0260	0.0492	0.0388
40 weeks	0.0562	0.0258	0.0501	0.0387

Note: Mean unemployment hazards are computed for a maximum benefit duration

of 40 weeks (at sample average of observed characteristics) and averaged over unobserved heterogeneity types. Mean accepted job hazards are computed in the first week of the accepted job (at sample average of observed characteristics) and averaged over unobserved heterogeneity types.

Table 4B
Flexible Model-Exogenous Benefit Duration
Predicted Unemployment and Accepted Job Hazards

	(1) **Table 3B (1)**	(2) **Table 3B (1)**	(3) **Table 3B (2)**	(4) **Table 3B (2)**
Earnings	**Unemp. Hazards** included	**Job Hazards** included	**Unemp. Hazards** excluded	**Job Hazards** excluded
Weeks left				
1 week	0.0859	0.0916	0.0867	0.0902
5 weeks	0.0719	0.0812	0.0780	0.0846
10 weeks	0.0610	0.0740	0.0701	0.0750
15 weeks	0.0584	0.0700	0.0654	0.0691
20 weeks	0.0506	0.0651	0.0589	0.0550
25 weeks	0.0456	0.0642	0.0430	0.0540
30 weeks	0.0430	0.0613	0.0429	0.0513
35 weeks	0.0542	0.0579	0.0512	0.0512
40 weeks	0.0604	0.0576	0.0612	0.0512

Note: Mean unemployment hazards are computed for a maximum benefit duration of 40 weeks and averaged over all individuals for the most common heterogeneity types. Mean accepted job hazards are computed in the first week of the accepted job and averaged over unobserved heterogeneity types.

Table 4C
Sources of variations in Mean Accepted Job Durations

		Source of Variation	
		Benef. Dur.	**Unobs. hetero.**
		(1)	(2)
	Earnings	St. Dev.	St. Dev.
Table 2A	Included	3.1 weeks	5.6 weeks
Table 2A	excluded	2.8 weeks	5.3 weeks
Table 2B	Included	3.2 weeks	5.0 weeks
Table 2B	excluded	2.9 weeks	4.9 weeks
Table 3A	Included	3.3 weeks	5.0 weeks
Table 3A	excluded	3.0 weeks	4.8 weeks
Table 3B	Included	3.0 weeks	5.2 weeks
Table 3B	excluded	2.8 weeks	5.0 weeks

Note: In column 1, mean job duration is computed for every individual (given individual maximum benefit duration) after having fixed individual characteristics and unobserved heterogeneity to their respective average value. In column 2, mean job duration is computed given a maximum benefit entitlement of 33 weeks (at the same level of observed characteristics). The only source of variation is therefore unobserved heterogeneity.

5.2. The Effects of Benefit Duration on Unemployment and Accepted Job Durations: Some Simulations

After having investigated the importance of benefit duration versus unobserved heterogeneity, it seems important to investigate the effects of increasing maximum benefit duration by an additional week. I have performed simulations for two different individuals; one entitled to 25 weeks and one entitled for a total a 45 weeks. In both cases, I have computed the increase in mean unemployment duration and mean accepted job duration following an increase of 1 week in maximum benefit duration. I have performed those simulations for every model specification reported in the paper.

Overall, the results indicate that the unemployment duration elasticity ranges between 0.20 and 0.15. In terms of weeks, these estimates indicate that increasing potential benefit duration by an additional week will mean unemployment duration by 1.5 to 1.05 days[21]. The effects of benefit duration on accepted job durations are typically smaller. They range between 0.07 to 0.13. Increasing potential benefit duration by 1 week will typically increase mean job duration by less than 1 day; the increase in job duration ranges between 0.9 to 0.49 day. To summarize, although both the unemployment duration and accepted job duration effects of an increase in benefit duration are small, the increase in unemployment duration is 50% to 100% higher than the increase in accepted job duration. The negative effects of benefit duration therefore tend to dominate the positive effects.

Table 5
Simulating the Effects of an Increase in Benefit Duration

	$\frac{\Delta\%Unemp.Duration}{\Delta\%Benefit.Duration}$		$\frac{\Delta\%Job.Duration}{\Delta\%Benefit.Duration}$	
	(1)	(2)	(3)	(4)
Ben. dur	**25 weeks**	**45 weeks**	**25 weeks**	**45 weeks**
Table 2A, col 1	0.19	0.15	0.10	0.07
Table 2A, col 2	0.19	0.16	0.11	0.07
Table 2B, col 1	0.18	0.15	0.10	0.08
Table 2B, col 2	0.17	0.15	0.10	0.07
Table 3A, col 1	0.21	0.19	0.13	0.10
Table 3A, col 2	0.22	0.19	0.12	0.10
Table 3B, col 1	0.19	0.16	0.13	0.11
Table 3B, col 2	0.20	0.15	0.13	0.10

Note: Simulations have been computed at the maximum benefit duration of 25 weeks and 45 weeks.

[21]The effects of benefit duration on unemployment duration appear consistent with simulation results reported by various authors. For instance, Ham and rea (1987) report that an increase of 1 week in benefit duration increases unemployment duration by 0.16 to 0.20 week. For a review, see Devine and Kiefer (1991). As far as I know, the effect of benefit duration on job duration cannot be compared to any study.

Finally, all the results reported previously are based on the hypothesis that the matching effects of UI benefits are measured by the waiting time until an individual will quit the post unemployment job. In order to check the robustness of the results to the definition of a job termination , I have extended the analysis found in Table 3A (the model with endogenous benefit duration where the effect of benefit duration on job duration is estimated flexibly) to the case where accepted job duration is defined as the waiting time until the accepted job is terminated either by a layoff or a quit. This can be justified by the fact that the distinction between quits and layoffs can sometimes be insignificant.

The results are in Appendix 2 (Table A4) and the simulated effects of potential benefit duration on unemployment and job duration are in Table A3. Despite the fact that accepted job duration is now re-defined, the changes in the estimates of benefit duration on the accepted job duration hazard function are quite marginal. Indeed, the effects of potential benefit duration are slightly smaller in Table A4 than in Table 3A. Consequently, the simulations, reported in Table A3, indicate that the matching effects of UI are relatively small. The elasticities (ranging between 0.08 and 0.11) indicate that an increase of 1 week in benefit duration will increase accepted job duration by 0.4 to 0.8 day.

6. Conclusion

It is generally admitted that UI benefit generosity induces those who have lost their job to remain unemployed for longer periods. The effects of UI benefit generosity on the quality of labor market adjustments are however not as clear. In this paper, I have tried to measure the effects of potential UI benefit duration on the quality of subsequent job matches. More precisely, I have investigated whether the correlation between completed unemployment duration and subsequent job duration is explained by a job matching effect or simply by unobserved heterogeneity.

Overall, the results indicate that both hypothesis contribute to explain the observed correlation between unemployment duration and accepted job duration. The escape rate out of unemployment seems to raise significantly within 5 weeks of benefit termination and new jobs accepted within this 5 week period seem to have a higher dissolution rate. However, at the same time, there is a strong negative correlation between unobserved heterogeneity affecting unemployment duration and unobserved heterogeneity affecting subsequent job duration. These results are very robust. They are true whether or not previous earnings are included and whether or not maximum benefit duration is allowed to be correlated with unemployment duration and job duration. The results of various simulations indicated that increasing the maximum benefit duration by one week will raise expected unemployment duration by 1 to 1.5 days and expected job duration by 0.5 to 0.9 day. The increase in unemployment duration is therefore higher than the increase in accepted job duration and the disincentive effects of benefit duration seem to dominate the matching effects.

It is well known that an increase in unemployment hazards can be explained by increase search efforts, decrease in reservation wages or implicit recall arrangements between workers and firms. As workers having returned to the same employer are eliminated from the sample, the fact that jobs accepted with lower benefit periods tend to be terminated faster is perhaps explained by a decline in reservation wages, which is more pronounced in a short period prior to benefit termination. At the same time, the strong negative correlation between unemployment and accepted job durations explained by unobserved heterogeneity is harder to explain. Perhaps, the most credible explanation lies in the existence of moral hazard. If individuals, who have a strong taste for leisure or household production, tend to exhaust their benefits and accept jobs that last long enough to re-qualify for UI benefits, the data would disclose negative correlation between unemployment job duration and accepted job duration. This suggests an avenue for future research. Structural job search models, which are estimated from micro-data and based on the maintained hypothesis that individuals start searching once they have lost their job, should perhaps be modified to take into account that household production (or leisure) is a substitute to job search activities.

7. References

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Appendix 1

Table A1-Remaining Parameter Estimates

		Table 2A (1)	Table 2A (2)	Table 2B (1)	Table 2B (2)
Ben. dur.					-
	Primary	-0.0823 (1.64)	-.0836 (1.79)	-	-
	Construc.	-0.0314 (1.43)	-0.0359 (1.27)	-	-
	Transport.	0.0453 (1.70)	0.0512 (1.78)	-	-
	Trade	0.0812 (2.05)	0.0734 (1.73)	-	-
	Finance	0.1621 (1.97)	0.1679 (1.96)	-	-
	Services	0.0423 (1.06)	0.0423 (1.32)	-	-
	Adminin	0.0523 (1.67)	0.0623 (1.70)	-	-
	Age	0.1945 (3.03)	0.2134 (2.89)	-	-
	Exper.	0.0312 (1.89)	0.0404 (2.02)	-	-
Unem. dur.					
	Primary	-0.0467 (1.69)	-0.0467 (1.80)	-0.0502 (1.69)	-0.0889 (1.78)
	Constru.	-0.0230 (1.56)	-0.0245 (1.39)	-0.0289 (1.03)	-0.0319 (1.50)
	Transport.	0.0435 (1.67)	0.0439 (1.79)	0.0459 (2.00)	0.0479 (2.32)
	Trade	0.0803 (2.05)	0.0834 (1.94)	0.0821 (1.89)	0.0803 (2.34)
	Finance	0.1934 (1.96)	0.1845 (2.00)	0.1823 (2.05)	0.1835 (1.97)
	Service	0.0465 (1.55)	0.0478 (1.90)	0.0477 (1.37)	0.0459 (1.76)
	Admin.	0.0497 (2.07)	0.0506 (1.86)	0.0537 (1.62)	0.0578 (1.87)
	Age	0.1789 (2.05)	0.1856 (2.67)	0.1965 (1.99)	0.1756 (3.24)
	Exper.	0.0423 (2.67)	0.0432 (2.11)	0.0356 (2.56)	0.0409 (3.00)
Job dur.					
	Primary	0.1224 (1.65)	0.1278 (1.78)	0.1323 (1.48)	0.1267 (1.68)
	Construc.	0.2345 (1.79)	0.2236 92.45)	0.2045 (1.76)	0.2312 (1.90)
	Transport	-0.0647 (2.04)	-0.0589 (1.99)	-0.0579 (1.94)	-0.0612 (2.11)
	Trade	0.0523 (0.89)	0.0538 (1.05)	0.0578 (1.22)	0.0534 (0.68)
	Finance	-0.0478 (1.54)	-0.0467 (1.66)	-0.0478 (1.69)	-0.0458 (1.55)
	Service	-0.0823 (1.97)	-0.0867 (2.33)	-0.0876 (2.67)	-0.0839 (1.94)
	Admin.	-0.1023 (2.33)	-0.1056 (2.56)	-0.1068 (2.41)	-0.1055 (2.16)
	Age	-0.0237 (2.89)	-0.0267 (2.69)	-0.0268 (3.02)	-0.0267 (2.78)
	Exper.	-0.0578 (3.10)	-0.0634 (2.95)	-0.0638 (2.69)	-0.0628 (3.11)

Note: Asymptotic standard errors in parentheses.

Table A2-Remaining Parameter Estimates

		Table 3A (1)	Table 3A (2)	Table 3B (1)	Table 3B (2)
Ben. dur.					-
	Primary	-0.0743 (1.84)	-.0776 (1.78)	-	-
	Construc.	-0.0378 (1.83)	-0.0376 (1.56)	-	-
	Transport.	0.0462 (1.82)	0.0544 (1.88)	-	-
	Trade	0.0722 (2.16)	0.0734 (1.70)	-	-
	Finance	0.1693 (1.79)	0.1628 (1.55)	-	-
	Services	0.0391 (1.48)	0.0445 (1.66)	-	-
	Adminin	0.0393 (1.60)	0.0333 (1.80)	-	-
	Age	0.1744 (2.54)	0.1854 (2.49)	-	-
	Exper.	0.0377 (2.05)	0.0455 (2.15)	-	-
Unem. dur.					
	Primary	0.0123 (1.60)	0.0246 (1.57)	-0.0102 (1.28)	-0.0188 (1.30)
	Constru.	-0.0230 (1.56)	-0.0378 (1.79)	-0.0334 (1.53)	-0.0345 (1.84)
	Transport.	0.0457 (1.68)	0.0404 (1.69)	0.0478 (2.08)	0.0421 (2.12)
	Trade	0.0356 (2.15)	0.0432 (2.08)	0.0734 (1.84)	0.0793 (2.34)
	Finance	0.1134 (1.96)	0.1247 (2.20)	0.1004 (2.19)	0.1136 (1.86)
	Service	0.0823 (1.65)	0.0734 (1.30)	0.0774 (1.67)	0.0592 (1.78)
	Admin.	0.0576 (2.34)	0.0693 (1.86)	0.0706 (1.92)	0.0646 (1.89)
	Age	0.1180 (1.65)	0.1256 (2.05)	0.1264 (1.79)	0.1356 (2.20)
	Exper.	0.0634 (2.28)	0.0657 (2.03)	0.0792 (2.06)	0.0659 (2.10)
Job dur.					
	Primary	0.1034 (1.83)	0.0784 (1.56)	0.0945 (1.83)	0.10037 (1.64)
	Construc.	0.2646 (1.93)	0.2556 (2.28)	0.2742 (1.66)	0.2634 (1.88)
	Transport	-0.0247 (1.09)	0.0159 (1.29)	-0.0579 (1.47)	-0.01612 (1.11)
	Trade	0.0836 (1.29)	0.0836 (0.59)	0.0854 (1.57)	0.0903 (1.60)
	Finance	-0.0448 (1.67)	-0.0484 (1.80)	-0.0415 (1.49)	-0.0499 (1.49)
	Service	-0.0382 (1.87)	-0.0367 (2.37)	-0.0387 (2.87)	-0.0289 (1.88)
	Admin.	-0.1626 (2.03)	-0.1156 (1.50)	-0.1039 (2.01)	-0.1444 (1.38)
	Age	-0.0199 (2.01)	-0.0346 (2.22)	-0.0204 (2.84)	-0.0359 (2.78)
	Exper.	-0.0628 (2.74)	-0.0823 (2.35)	-0.0733 (2.49)	-0.0823 (2.74)

Note: Asymptotic standard errors in parentheses.

Appendix 2

All the results reported previously were based on the assumption that the quality of a job match was measured by the waiting time until an individual quit the post unemployment job. This implies that jobs terminated by a layoff (464 cases) were treated as censored spell. As many economists have claimed that the distinction between quits and layoffs can sometimes be insignificant, one might argue that treating all jobs terminated by a layoff as a censored spell might seriously over estimate job duration and affect the results reported previously. In order to check the robustness of the results, I have extended the analysis found in Table 3A (the model with endogenous benefit duration where the effect of benefit duration on job duration is estimated flexibly) to the case where accepted job duration is defined as the waiting time until the accepted job is terminated either by a layoff or a quit. The results are in Table A4.

Despite a re-definition of accepted job duration, the parameter estimates for the effects of benefit duration on accepted job duration are not really different. In table A3, I report the simulations for the effect of increasing benefit duration obtained with the parameter estimates of Table A4. The job duration elasticities range between 0.08 and 0.10 and imply, as in Table 5, that increasing potential benefit duration by one week will increase expected job duration by less than 1 day.

Table A3
Simulating the Effects of an Increase in Benefit Duration/Layoffs and Quits

	$\frac{\Delta\% Unemp.Duration}{\Delta\% Benefit.Duration}$		$\frac{\Delta\% Job.Duration}{\Delta\% Benefit.Duration}$	
	(1)	(2)	(3)	(4)
Ben. dur.	**25 weeks**	**45 weeks**	**25 weeks**	**45 weeks**
Table A3, col 1	0.19	0.17	0.10	0.08
Table A3, col 2	0.20	0.17	0.11	0.08

Table A4
Jobs Terminated by Layoffs and Quits

		(1)	**(2)**
Benefit duration	Duration dependence (α)	1.3451 (4.92)	1.0272 (7.04)
	β^v	-0.7088 (2.08)	-0.7207 (2.90)
Unemp. duration	Log Earnings	0.3754 (2.26)	-
	Benefit level (log)	-0.2812 (2.33)	-0.2921 (2.37)
	Duration Dependence (β)	0.9403 (7.34)	0.9407 (6.28)
Unemp. duration			
	Potential Benefit Duration		
	$\delta_{X(1-5)}$	-0.2316 (2.26)	-0.2444 (2.16)
	$\delta_{X(6-9)}$	-0.2136 (2.79)	-0.2073 (2.29)
	$\delta_{X(10-19)}$	-0.0608 (1.32)	-0.0593 (1.40)
	$\delta_{X(20-29)}$	0.0008 (0.44)	0.0006 (0.77)
	$\delta_{X(t_u)}$	-0.0019 (1.34)	-0.0034 (1.10)
Unobs. heterogeneity			
	Hetero. Support Points		
	ε_1^u	0.5523 (2.38)	0.6063 (2.53)
	ε_2^u	0.3187 (2.92)	0.3823 (3.33)
	ε_1^j	0.5012 (2.85)	0.5381 (2.84)
	ε_2^j	0.2040 (1.76)	0.2387 (1.92)
	Probabilities		
	P_1	0.1500 (2.42)	0.1300 (2.66)
	P_2	0.3800 (2.32)	0.4000 (2.41)
	P_3	0.3200 (1.64)	0.2700 (1.80)
Correlation	$Corr(\varepsilon^u, \varepsilon^j)$	-0.3600	-0.3900

Table A4-Continued

Accepted job duration			
	Duration Dependence (γ)	0.8823 (5.03)	0.8721 (5.72)
	Unemp. Duration	0.0182 (1.73)	0.0183 (1.62)
	Unemp. Dur_{50}	0.3004 (2.05)	0.2773 (1.86)
	Benefit Duration Left		
	$\lambda_{X(1-5)}$	-0.0324 (1.83)	-0.0296 (1.79)
	$\lambda_{X(6-9)}$	-0.0010 (1.30)	-0.0025 (1.38)
	$\lambda_{X(10-19)}$	-0.0134 (1.62)	-0.0154 (1.48)
	$\lambda_{X(20-29)}$	-0.0011 (0.48)	-0.0017 (0.89)
	$\lambda_{X(t_u)}$	0.0006 (0.79)	0.0007 (0.83)
Log likelihood		-1316.8	-1320.6
LR test for corr($\varepsilon^u, \varepsilon^j = 0$)		7.3	7.2

Notes:
Asymptotic t-ratios in parentheses.
The p-value for the likelihood ratio test for the null that $\lambda_{X(1-5)}..=..\lambda_{X(20-29)} = 0$ **was 0.059 in column 1 and 0.044 in column 2**.
The likelihood Ratio statistic is for the null that ε^u **and** ε^j **are independent.**

Both columns of Table 3 have been re-estimated under the restrictions that $\beta^v = 0$ and $q_3 = q_1 - q_2$. Altogether, these restrictions imply that unobservables affecting benefit duration, unemployment duration and accepted job duration are independent. When independence is assumed, I find stronger effects of benefit duration left on unemployment hazards (especially within 9 weeks of benefit termination) and stronger effects of benefit duration on accepted job hazard. The values of the likelihood ratio statistics, 9.6 in column 1 (compared to column 1 of Table 3) and 7.6 in column 2 (compared to column 2 of Table 3) indicate a strong rejection of the null hypothesis as the critical value at $\alpha = 0.05$ is 5.99.

Table A1 – Independent Unobserved Heterogeneity

Benefit duration	Duration dependence (α)	1.0234 (4.12)	1.0226 (3.98)
	β^v	-	-
Unemp. Duration	Log Earnings	0.4563 (3.44)	-
	Benefit level (log)	-0.3398 (2.21)	-0.3664 (4.01)
	Duration Dependence (β)	0.8646 (4.99)	0.8587 (5.12)
Unemp. Duration			
	Potential Benefit Duration		
	δ_1	-0.4673 (3.12)	-0.4624 (3.77)
	δ_{2-5}	-0.3345 (2.20)	-0.3298 (2.56)
	δ_{6-9}	-0.2867 (1.89)	-0.2845 (2.02)
	δ_{10-19}	-0.0987 (1.45)	-0.0937 (1.24)
	δ_{20-29}	0.0024 (0.34)	0.0028 (0.76)
	δ_{30-50}	-0.0004 (0.78)	-0.0004 (0.92)

Table $A1$−Continued

Unobs. Heterogeneity			
	Hetero. Support Points		
	ε_1^u	0.8719 (2.27)	0.6238 (2.88)
	ε_2^u	0.3826 (2.10)	0.3674 (3.01)
	ε_1^j	0.7926 (2.87)	0.5512 (2.24)
	ε_2^j	0.5006 (3.05)	0.2411 (1.82)
	Probabilities		
	P_1	0.3500 (2.63)	0.3800 (1.84)
	P_2	0.0900 (1.97)	0.1200 (1.67)
	P_3	-	-
	Correlation	0.00	0.00
Accepted Job Duration			
	Duration Dependence (γ)	0.8721 (4.94)	0.8705 (5.09)
	Unemp. Duration	0.0358 (2.78)	0.0345 (1.80)
	Unemp. Dur_{50}	0.3067 (2.52)	0.3077 (1.83)
	Benefit Duration Left		
	δ_1	-0.0454 (2.28)	-0.0489 (2.40)
	δ_{2-5}	-0.0746 (2.27)	-0.0764 (1.99)
	δ_{6-9}	0.0356 (1.68	0.0349 (1.82)
	δ_{10-19}	-0.0435 (1.56)	-0.0439 (1.35)
	δ_{20-29}	-0.0305 (1.56)	-0.0310 (0.98)
	δ_{30-50}	0.0123 (1.24)	0.0119 (1.33)
Log Likelihood		-1320.4	-1324.2

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