Inflow into Unemployment: Employment Spells and Unemployment Insurance

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Abstract

This paper uses U.S. micro level data on employment durations to quantify the effect of potential Unemployment Insurance (UI) entitlement on job separations. Economic theory motivates estimation of a competing risk hazard model for quits and layoffs. The estimation procedure simultaneously allows for unobserved heterogeneity, defective risks and sample selection into future spells. It also uses alternative assumptions about agents' ability to determine eligibility for future UI claims. Empirical results suggest that being entitled to UI compensation raises the layoff hazard, but workers with higher levels of potential UI entitlement do not appear to be more likely to get laid off.

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1 Introduction

While there have been numerous studies estimating the effect of the unemployment insurance (UI) system on duration of unemployment, there has been no empirical work analyzing the effect of UI on employment *durations* in the United States.[1] This gap in the literature is somewhat surprising since there are at least two theoretical arguments for why we would expect UI to affect employment durations. First, the implicit contract literature suggests that unemployment insurance makes layoffs more likely (e.g. Feldstein 1976, Baily 1977). Second, job search models show that workers with generous UI coverage will search less intensively while unemployed. Below we show that the optimal firm response to this behavior, in the presence of demand fluctuations and firm specific human capital, is for the firm to lay off workers with high levels of UI entitlement and recall workers as they approach exhaustion of their benefits.[2] This paper therefore analyzes the effect of unemployment insurance on unemployment inflow using a micro data set on employment durations. The empirical motivation for analyzing employment durations as opposed to cross-sectional data comes from the fact that the amount of potential UI compensation, as well as the demand conditions, varies over the duration of individual employment spells

The lack of research on this topic is likely caused by the fact that large micro data sets on employment durations and UI compensation are scarce. We use a data set which consists of a dislocated workers survey, augmented with information on the amount of UI compensation individuals can expect to receive if they are laid off or quit. Unemployment compensation provisions, including the trigger dates of various extended benefit programs, are coded for over five years for seven states. The resulting multiple spell, event history data

[1] The only studies looking at employment durations we are aware of are Baker and Rea (1993) and Christofides and McKenna (1996). Both analyze the effect of Canadian UI eligibility requirements. The closely related U.S. cross-sectional work of Anderson and Meyer (1994) is discussed in section 2.

[2] There is also extensive research focusing on the layoff effect of UI taxes (see section 2). Analyzing this issue is beyond the scope of the present paper, but it is addressed in Jurajda (1997).

set is unusually rich in terms of the variation of entitlement and benefit levels.

A worker who quits will generally not be entitled to UI compensation. In the presence of a positive layoff probability, delaying a quit to non-employment will provide the worker with a chance of getting laid off and obtaining UI coverage. Thus, one may expect the opposite entitlement effects when comparing layoff and quit decisions. We therefore analyze quits and layoffs separately using a competing risk duration model. The use of hazard models in analyzing duration data has become widespread, and accounting for unobserved heterogeneity is now a standard part of hazard estimation sensitivity analysis. The estimation procedure used here allows for the effects of unobserved heterogeneity in a number of ways and controls for sample selection into multiple spells, a potentially important issue in the estimation of duration models. Using multiple spell data on employment durations provides greater variation and improves identification of the unobserved heterogeneity distribution. The use of this type of data, however, raises the possibility of selection bias; i.e., the workers who have multiple employment spells may be a non-random sample. To control for this problem, we estimate employment and unemployment durations jointly while allowing the unobserved heterogeneity to be correlated across these spells. Thus, the estimated model is a multiple-spell, multiple-state competing risk duration model with unobserved heterogeneity. Finally, the estimated unobserved heterogeneity models naturally extend to account for the possibility of defective risks (absorption states).[3]

Due to the nature of the UI system, any attempt to evaluate the effects of UI on economic outcomes has to rely on arbitrary assumptions about how agents form expectations of the available UI compensation. In this paper, we examine the robustness of the empirical results with respect to different assumptions about how firms and workers account for UI rules when determining *eligibility for future UI claims*. This issue has not been addressed previously. The type of assumption one makes in the estimation significantly affects the

[3]The estimation procedure allows for zero probability of quitting for a fraction of the sample.

levels of the explanatory variable of interest–UI entitlement. In the empirical analysis we therefore compare results based on the assumption that future UI eligibility is ignored to results based on the assumption that future UI eligibility is taken into account.

The empirical results suggest that being entitled to UI compensation significantly increases the layoff hazard. In contrast to theoretical prediction, however, neither the length of potential UI entitlement nor the dollar amount of UI benefits, conditional on being positive, affect the layoff probability. The probability of a quit is not affected by any of the UI system parameters. Some of the layoff hazard parameters are sensitive to introducing a typical unobserved heterogeneity distribution.[4] Further, the layoff hazard point coefficient of the eligibility dummy approximately doubles in size compared to the no-heterogeneity estimate when we control for selection bias *and* allow for the possibility of defective risks. The results differ across the two alternative assumptions on how agents account for future UI eligibility in some of the estimated specifications. They are very similar, however, in two important cases: when we do not control for unobserved heterogeneity and when we use the most general form of heterogeneity distribution.

The paper proceeds as follows. Section 2 discusses previous work and Section 3 models firm employment decisions. The data set is described in Section 4. Section 5 presents the econometric approach together with the empirical results. Section 6 concludes.

2 Previous Work

The strand of economic literature focusing on temporary layoffs and UI starts with the analyses of implicit contract models by Feldstein (1976) and Baily (1977). In these models, firms facing competitive labor markets have to offer employment contracts which provide workers with a market-determined level of expected utility. In Feldstein's 1976 model, the

[4]Estimated sample likelihoods strongly support both specifications with unobserved heterogeneity and their defective risk extensions.

imperfect experience rating of firms creates a subsidy to layoffs,[5] which makes layoffs more likely in the presence of product demand fluctuations. Baily (1977) shows that increases in the level of UI compensation cause firms to increase layoffs. Since workers with UI coverage are better protected against prolonged spells of unemployment, the layoff probability becomes an increasing function of UI compensation.[6]

Implicit contract models assume workers' utility level is exogenous and endogenize the level of wages. On the other hand, models focusing on the adjustment cost aspect of UI taxes[7] (e.g. Card and Levine 1992) take wages as exogenous. Firms are at least partially responsible for UI benefits paid to their former employees. A typical adjustment cost model would therefore imply that more generous UI coverage leads to lower risks of layoff, contrary to predictions of implicit contract models. Anderson and Meyer (1994) extend the adjustment cost model to include the compensation package concept of implicit contract models. While all models predict a negative relation between the degree of experience rating and layoffs, their analysis allows for different effects of UI compensation.

The theoretical work discussed above has motivated a number of empirical studies. Typically, these studies use CPS cross-sectional data sets (e.g. Topel 1983, Card and Levine 1992) and suggest that the unemployment inflow effect of UI is potentially quite large because of imperfect experience rating. Anderson and Meyer (1994) analyze cross-sectional data sets based on the Continuous Wage and Benefit History (CWBH) survey to quantify the effects of experience rating *and* the level of potential UI coverage on the incidence of layoffs. While they confirm previous findings of large experience rating effects, they obtain conflicting estimates of the effect of potential UI coverage.

[5]The U.S. system of levying unemployment insurance tax based on an employer's unemployment experience is called experience rating.

[6]The implicit contract analysis is generalized by Burdett and Hool (1983), who incorporate the optimal contract determination into a bargaining problem of firms and workers, and by Haltiwanger (1984), who analyzes a *multiperiod* contract model, allowing for the interaction of stock adjustment and factor utilization decisions.

[7]UI taxes make employment adjustment costly because of experience rating.

Baker and Rea (1993) and Christofides and McKenna (1996) analyze the effect of Canadian UI eligibility rules to identify spikes in the employment hazard (i.e. the hazard of leaving employment) in the first week of eligibility. Canadian eligibility rules depend on local economic conditions, but Baker and Rea (1993) are able to untangle this dependency by using a unique change in the eligibility formula orthogonal to changes in the economic environment. Their results indicate a significant increase in the employment hazard in the week in which individuals qualify for UI compensation. Other than including three dummy variables capturing UI eligibility,[8] they do not control for the level of available UI compensation. In particular, they do not control for the dollar amount of UI benefits available and for changes in the maximum amount of UI entitlement.

This paper extends the existing literature by analyzing the effect of UI on employment *durations* in the U.S., using different, rich sources of variation in UI compensation,[9] and considering quits and layoffs separately.

3 A Dynamic Model of Layoffs and Recalls

This section investigates a dynamic decision problem of a price-taking, profit-maximizing firm deciding on the employment status of a fixed roster of workers.[10] The firm is assumed to know the workers' optimal job search strategies. In the presence of training costs, the firm responds to demand fluctuations by laying off workers with greater entitlement and recalling workers as they approach exhaustion of their benefits. Even though the model assumes

[8]The first one equals one in the week when a given worker becomes eligible. The second indicates that the worker's entitlement is between the minimum and maximum value. The third dummy variable equals one when the worker has attained the maximum potential entitlement.

[9]Anderson and Meyer (1994) use state variation in entitlement and benefits coming from the high quarter wage and base period earnings which, together with the state level of the maximum benefit amount, are used to determine regular benefit amount and duration.

[10]A similar assumption of a fixed roster of workers was used in most previous studies, e.g. Feldstein (1976), Card and Levine (1992), to narrow the model's focus to temporary layoffs. Long term worker-firm attachments are motivated by the existence of firm specific human capital (Becker 1962) and by implicit contract models (Azariadis 1975), where firms provide workers with insurance against labor market fluctuations.

exogenous wages, similar to the adjustment cost analyses discussed in Section 2, it results in predictions similar to those of the implicit contract models. It borrows from job search theory, both in terms of motivation and modeling technique, to extend the adjustment cost argument to a non-stationary dynamic framework.

A standard result from the job search literature is that the probability of a worker on layoff finding a job with another firm is a decreasing function of the length of the remaining entitlement period. Hence, assuming that firms take workers' search strategies into account, laying off a worker with a high value of potential UI entitlement is less costly for the firm since such a worker will be less likely to find a new acceptable job with an alternative employer.[11] If recovery occurs, the firm simply recalls the worker on layoff instead of incurring the training costs of hiring a new worker. The maintained assumption here is that workers do not take the optimal recall strategy of firms into account when optimizing their search behavior.

The modeling strategy follows Mortensen's (1977) analysis. Demand fluctuations are modeled as independent draws of firm specific marginal revenue product M from a time constant density $f(M) = F'(M)$.[12] An *unemployed* worker finds a job with an alternative employer with per period probability $q(\tau)$, where $\tau \in [0, T]$ is the remaining UI entitlement and

$$\frac{dq(\tau)}{d\tau} < 0$$

(Mortensen 1977). Since the profit function from having the worker employed is monotonically increasing in the value of the marginal revenue product, there is an optimal layoff stopping rule m_E, such that, for a given τ, the firm decides to keep the worker at all $M \geq m_E$ and to layoff otherwise. Similarly, corresponding to the profit function from having a worker

[11]This point was originally made by Pissarides (1982) in his stationary model of recall behavior.

[12]See also Tannery (1993). The firm observes new values of marginal revenue product even for workers on layoff. All workers in a given firm have the same value of M at each point in time. Adding a person specific component to M would make the model more realistic but would not affect qualitative results of the analysis.

on layoff, there is an optimal stopping value for recalls m_L. Assume that workers are eligible for the maximum potential UI entitlement of T immediately after being hired or recalled.[13] Finally, assume that the firm pays a fixed cost C_L for laying off each worker. Let $\beta(h) = e^{-rh}$ be the discount factor, where r is the rate of time preference and h is the time increment. Appealing to Bellman's principle of optimality, we can write the firm's profit value function π_E from employing a given worker as

$$\pi_E(M,T) = \max_{m_E}\{h(M-w) + \beta(h)[(1-h\alpha)\pi_E(M,T) \tag{1}$$
$$+h\alpha\int_{m_E}^{\infty}\pi_E(\hat{M},T)dF(\hat{M}) + h\alpha F(m_E)(\pi_L(T) - C_L)]\},$$

where w denotes wages, α stands for the probability of a new value of M arriving, and $\pi_L(T)$ is the profit value function from having a worker with UI entitlement of T on layoff. The firm's objective consists of the per period profit rate $M-w$ and the discounted future profits in three possible states. First, with per period probability $1-h\alpha$, there is no change in M, and the firm faces a similar optimization problem next period. Second, the firm evaluates the expected profits resulting from the arrival of a new value of M above the layoff threshold. Third, with probability $h\alpha F(m_E)$, a below-the-threshold value of M arrives and the worker is laid off.

Next, let us write the firm profit function from having the worker on layoff with residual entitlement $\tau \in [h,T]$

$$\pi_L(\tau) = \max_{m_L}\beta(h)\{hq(\tau)\pi_N + (1-hq(\tau))[(1-h\alpha)\pi_L(\tau-h) \tag{2}$$
$$+h\alpha\int_{m_L}^{\infty}\pi_E(\hat{M},T)dF(\hat{M}) + h\alpha F(m_L)\pi_L(\tau-h)]\},$$

where π_N stands for the profit from not having the worker available (i.e. the optimal profit when the unemployed worker quits to another firm), which is strictly lower than $\pi_L(\tau)$ $\forall\tau$,

[13]This assumption is made for analytical convenience and reflects an extreme limiting case of UI eligibility rules, where UI entitlement depends on earnings and job duration.

assuming sufficiently high training costs of hiring new workers.[14] With probability $1-hq(\tau)$, the worker does not quit and the firm evaluates profits in three possible future states. If the current value of marginal revenue product remains unchanged, the worker stays unemployed and draws UI so that his entitlement decreases. If a high enough new value of M arrives, the worker is recalled and the firm collects the appropriate profit π_E. Finally, if the new value is below the recall threshold, the worker remains unemployed and UI entitlement decreases by the amount of time spent in unemployment.[15]

The optimal layoff stopping value of marginal revenue product $m_E(T)$ is implicitly defined by $\pi_E(m_E(T), T) = \pi_L(T) - C_L$. It follows that

$$\frac{\partial \pi_E(m_E(T), T)}{\partial M} \frac{\partial m_E(T)}{\partial T} + \frac{\partial \pi_E(m_E(T), T)}{\partial T} = \frac{d\pi_L(T)}{d\tau}.$$

Applying the envelope theorem to equation 1 gives

$$\frac{\partial \pi_E(m_E(T), T)}{\partial T} = \frac{\alpha F(m_E(T))}{r + \alpha F(m_E(T))} \frac{\partial \pi_L(T)}{\partial \tau} < \frac{d\pi_L(T)}{d\tau},$$

and since π_E is increasing in M, the per period layoff rate $\alpha F(m_E(T))$ is an increasing function of available UI compensation.

The model also provides a prediction for the effect of UI entitlement on recall decisions, i.e. for the properties of the optimal recall threshold. The optimum recall stopping value of marginal revenue product $m_L(\tau)$ is implicitly defined by $\pi_E(m_L(\tau), T) = \pi_L(\tau)$.[16] Using a

[14] The firm's layoff costs in terms of UI taxes are ignored as we do not focus on the effect of experience rating in this analysis. See Jurajda (1997) for a similar model allowing for non-zero experience rating. Also, note that $\pi_L \geq \pi_N$ even in the absence of training costs since the firm is always free to hire an outside worker.

[15] The profit value function from a laid off worker who has exhausted UI benefits can be defined in a similar fashion.

[16] Since π_E is monotonically increasing in its first argument, the layoff threshold and the recall threshold at T would coincide if $C_L = 0$. Also, note that if $C_L = 0$, firms would like to recall *and* layoff *all* unemployed workers in the same instant to increase their entitlement to T, which would lower the probability of losing them to another firm. This scenario reflects the limiting assumption of no eligibility requirements. In order to keep the model's solution well defined, we have to assume that $C_L \geq \pi_L(T) - \pi_L(0)$ so that the net gain from such an action would be negative.

similar argument as in the layoff case, it follows that

$$\frac{\partial \pi_E(m_L(\tau),T)}{\partial M}\frac{\partial m_L(\tau)}{\partial \tau}=\frac{d\pi_L(\tau)}{d\tau},$$

and since π_E is increasing in M we conclude that $m_L(\tau)$ increases in τ if π_L is increasing in τ. This last condition follows from differentiating 2 with respect to τ:

$$[r+q(\tau)+\alpha(1-F(m_L(\tau)))]\frac{d\pi_L(\tau)}{d\tau}=\frac{dq(\tau)}{d\tau}[\pi_N-\pi_L(\tau)].$$

The recall probability $\alpha[1-F(m_L(\tau))]$ is therefore a decreasing function of the remaining UI entitlement. The model motivates layoff and recall hazard estimation much the way job search models motivate new job unemployment hazards.

A worker who quits will generally not be entitled to UI compensation, and so one would expect opposite entitlement effects when comparing layoff and quit decisions. Clearly, there will be no effect of UI compensation on job-to-job quits. Quits to non-employment are present in the job matching models (e.g. Jovanovic 1979). If there is a positive probability of getting laid off, it could pay off for a worker contemplating a quit to non-employment to stay employed one more period, since by doing so he could get laid off and be qualified for UI coverage. The higher the available UI compensation, the stronger the incentive to wait for (or induce) layoff. Workers with high entitlement can therefore be expected to be less likely to quit.[17]

4 Data Description

The data employed in this paper comes from the Trade Adjustment Assistance (TAA) Survey. Implemented in 1974, the TAA program was intended to compensate workers harmed

[17] Using a similar argument as in the layoff analysis, one would expect that of the workers who temporarily prefer non-employment to working and ask the firm for a layoff in order to be qualified for UI compensation, those with a higher level of UI entitlement would be more likely to succeed, since the firm would be less worried about losing them to an alternative employer. Hence, one could expect those who actually quit for nonemployment to have low values of potential UI compensation.

by market fluctuations resulting from a rise in imports.[18] The data was collected from retrospective interviews with individuals who became unemployed in the mid 1970s. This information was merged with UI claims records. The data comes from seven states[19] and covers the period up to 1979. The TAA recipients were entitled to extensions of the regular UI entitlement of up to 52 weeks. Also, their replacement ratio (i.e. the ratio of UI benefits to wages on the last job) was set at 70% as opposed to the 50% typical of regular UI. Both regular UI recipients and TAA recipients are included in the sample.[20] The combination of TAA and UI recipients leads to a rich variation in UI entitlement and benefits. The other attractive feature of this sample is that it covers a period with many dramatic changes in UI entitlement, caused by various extended coverage programs being triggered on and off. Further, to the best of our knowledge, it is the only U.S. data set on employment durations.

During the sample period there were two types of extended coverage programs in effect: the Extended Benefits program and the Federal Supplemental Benefits program. These programs trigger on and off based on state and national insured unemployment rates. The State-federal Extended Benefits program triggers both at state and national levels and adds up to 13 weeks of UI benefits (50% beyond the state potential duration). The Federal Supplemental Benefits program extended the previous entitlement by up to 26 additional weeks of UI compensation. It was enacted at the national level and the number of extra weeks of UI differed both across states and over time.[21] The two programs could therefore change the typical 26 weeks of regular UI entitlement by as much as 39 weeks. Most of the empirical leverage necessary for the identification of the entitlement effect comes from these programs, as well as from the combination of UI and TAA recipients.

[18] The program was amended several times and is expected to be amended again at its current expiration date in 1998.

[19] California, Indiana, Massachusetts, New York, Ohio, Pennsylvania and Virginia.

[20] For a thorough description of the data and for information about the TAA program, see Corson and Nicholson (1981).

[21] A brief description of these programs can be found in Jurajda (1997).

Note that potential entitlement can also be quite low in some cases. Consider a worker who is recalled or finds a new job only a few weeks prior to exhausting UI benefits. Before he accumulates enough earnings to be eligible for the full UI entitlement, the worker faces the possibility of layoff with a low value of entitlement left from the previous spell of unemployment. The existence of the UI benefit year is another source of variation in potential entitlement. The UI benefit year starts when a UI claim is filed, at which moment the initial entitlement is determined based on the eligibility requirements. If a worker becomes employed after a few weeks of unemployment, a large amount of entitlement remains available for the duration of the UI benefit year. However, potential entitlement for those workers with only a few weeks left in their UI benefit year can be less than the remaining (non-collected) part of their initial entitlement. Hence, potential entitlement can also vary with time left in a UI benefit year.

From the initial sample of 1501 individuals, we drop those who do not start an employment spell during the sample frame *and* report being out of the labor force.[22] We also omit cases in which the initial unemployment spell was in fact a period of reduced hours. Finally, inconsistent and missing data records were deleted, yielding a sample of 1245 men and women. The empirical analysis is conducted on a subsample of 808 men. The data is recorded for a period of about 3.5 years for each individual. The initial spell of unemployment is followed by an employment spell for all 808 workers. Approximately 50% of the first employment spells are censored and about half of the subsequent unemployment spells end in another employment. Moreover, about 10% of workers experience three employment spells within the sample frame.[23] The existence of this group of individuals with short employment and unemployment durations suggests the possibility of substantial unobserved

[22]The information on dropout from the labor force is unusual in data sets used in duration analysis, where distinguishing unemployment from out-of-labor-force is usually a problem.

[23]This group of individuals has lower than average durations of both unemployment and employment. Only about 10% of those who enter second and third jobs are construction workers.

heterogeneity and non-random selection into subsequent employment spells. This issue will be explored in the empirical analysis.

Table 1 shows the data means at the first week of spells for all 808 men. The averages for unemployment spell benefits and entitlement in the current UI claim are taken over UI recipients only. The non-recipients consist primarily of people who have quit their previous jobs. The low average values of UI benefits and entitlement in the first week of *employment* spells come from the fact that, at the beginning of a spell, individuals are often not eligible for UI compensation either because they do not have enough earnings to qualify or because they have exhausted their UI entitlement for a given UI benefit year. While the standard deviations of the UI variables are already quite high, they reflect only the cross-sectional variation in the first week of each spell. Additional time variation comes mostly from the extended coverage programs, which change the amount of available compensation even for spells in progress.

The simplest approximation to the underlying hazard functions which ignores both observed and unobserved differences in the population is provided by the Kaplan-Meier empirical hazards. A basic set of empirical hazards is presented in Appendix A, which contains the overall unemployment empirical hazard with one standard deviation bounds. It also presents empirical hazards for the employment spells (overall and competing risks), and reveals differences between layoffs and quits (the layoff hazard is larger than quit hazard in the first 40 weeks of duration) as well as spikes at approximately one year of duration, reflecting perhaps the end of a probation period or recall bias.[24]

The data set contains information on the level of initial entitlement and benefits only for the first unemployment spell. We impute both (i) the potential entitlement for the employment spells and (ii) the actual entitlement levels for the second and third unemployment

[24] Recall bias occurs when individuals who do not recall the exact duration of their employment spell report approximate duration rounded to the closest six-month period, for example.

spell from the state specific UI laws and the individual data. To impute the UI compensation, we use the level of initial entitlement in the first unemployment spell and follow each individual over time, determining the level of entitlement in each week based on the individual's employment history, information on the reason for job separation (i.e. quit as opposed to layoff [25]), UI eligibility requirements and the effective trigger dates of extended benefits programs. In the imputation procedure we assume that workers file UI claims whenever they are entitled to do so. When determining eligibility we assume that wages do not change on the job (only accepted wages are reported). Using predicted values of UI entitlement instead of actual ones is a potential drawback of the data. Note, however, that workers or firms contemplating a transition out of employment will have to use their own prediction of potential entitlement based on a similar information set. Thus, we would argue that our prediction of the potential UI compensation should not significantly affect the results, at least in the employment spells. The information sources used in imputing UI compensation are listed in Jurajda (1997).

One important question arising when imputing potential entitlement values is whether workers and firms are able to determine the UI eligibility for *future* UI claims. For example, is a recently recalled worker with only 10 weeks of entitlement left from his spell of unemployment able to predict that if he were laid off at that time, he would (after exhausting the remaining 10 weeks of entitlement) become eligible for another UI claim? If so, then the value of potential entitlement should equal the sum of the remaining UI compensation from the existing UI claim, plus the initial UI entitlement a newly eligible worker would obtain at the beginning of a new UI claim. This assumption on potential entitlement seems reasonable since all of the workers in the sample went through the process of filing the initial UI claim at the beginning of the sample frame and, therefore, should have at least some understanding

[25]There were only a few cases of an individual being fired for cause, and they are omitted in the empirical analysis.

of what the UI eligibility requirements are. Similarly, firms can be assumed to know the UI rules as they face layoff decisions on a regular basis. Assuming that UI eligibility rules are well known, an employed worker who becomes eligible for a new UI claim during his current UI claim will have higher potential UI entitlement than a worker who has been on a job for over one year. Taking future repeated UI claims into account therefore breaks the usual positive relationship between the level of potential UI entitlement and job duration. On the other hand, it may be that firms and especially workers are somewhat myopic in measuring potential UI entitlement. In the estimation we therefore allow for alternative assumptions on whether individuals account for UI eligibility rules when determining future entitlement.

The advantage of analyzing displaced workers (especially those with multiple spells) is that the focus is on individuals most likely affected by the amount of potential UI compensation. There are large groups of individuals in whose employment history UI entitlement plays no role. These individuals, who have close to zero lifetime weeks of unemployment, will most likely be entitled to the maximum unemployment benefits throughout their careers. Yet, they may never become unemployed. In future work it would be desirable to work also with a large representative sample of the population. In the present analysis, which is the first to look at the effects of UI on employment durations in the U.S., we start by examining the more likely UI sensitive fraction of the population.

5 Estimation and Results

A typical job search model derives the per period escape rate out of unemployment as a function of the remaining UI compensation. Job search models therefore naturally motivate the estimation of unemployment hazard functions, which parametrize the probability of leaving unemployment at each time period.[26] Similarly, estimation of the employment quit process has been motivated by on-the-job search models (e.g. Burdett 1978). Finally, the

[26] For a survey of search approach empirical literature, see Devine and Kiefer (1991).

model of optimal layoff decisions (discussed in Section 3) results in per period layoff rates and motivates estimation of a layoff hazard function. The reduced-form hazard model used here therefore estimates the conditional probability of (i) finding a job while unemployed or (ii) losing a job while employed. The resulting estimates for employment or unemployment durations can be interpreted as approximations to the comparative statics implied by a corresponding model of job separations or job search. The theoretical considerations presented in Section 3 also point to a differential effect of UI on quits and layoffs and lead to a competing risks estimation of employment hazard functions.[27]

5.1 Econometric Model

The duration model builds upon the concept of a *hazard* function, which is defined as the probability of leaving a given state at duration t conditional upon staying there up to that point. Using this definition one can build a likelihood function for the observed durations and estimate it using standard methods. However, it is well known that in the presence of unobserved person specific characteristics affecting the probability of exit, all of the estimated coefficients will be biased. To control for unobserved factors, we follow the flexible approach of Heckman and Singer (1984). The strategy is to approximate any underlying distribution function of unobservables by estimating a discrete mixing distribution $p(\theta)$ of an unobserved heterogeneity term θ as a part of the optimization problem. This approach was applied for example by Ham and LaLonde (1996) and McCall (1996).

More specifically, let $\lambda_j(t, x_t|\theta_k^j)$ be the conditional probability (hazard) of leaving a given state at time (duration) t for someone with person specific characteristics x_t, conditional upon this person having the unobserved factor θ_k^j, $k=1,2,...,N_\theta^j$. The j subscript stands for the different ways of leaving a given state and serves, therefore, as a state subscript as well.

[27] In the unemployment hazard we do not differentiate between recalls and new job findings since this issue has been analyzed extensively in the existing literature (e.g. Katz and Meyer 1990).

For example one can leave employment through a quit or through a layoff, in which case $j \in \{q, l\}$. This is often referred to as a competing risk model. In what follows, we work in discrete time with weekly hazards in logit specification:

$$\lambda_j(t, x_t|\theta_k^j) = \frac{1}{1 + e^{-h_j(t, x_t|\theta_k^j)}}, \tag{3}$$

where

$$h_j(t, x_t|\theta_k^j) = r_j(e_t, \alpha_j) + \beta_j' z_t + g_j(t, \gamma_j) + \theta_k^j. \tag{4}$$

Here, $x_t' = (e_t, z_t')$, $r_j(e_t, \alpha_j)$ denotes a function of remaining entitlement e_t, the vector z_t includes levels of benefits, wages, demographics and time changing demand measures.[28] Finally, $g_j(t, \gamma_j)$ is a function capturing the duration dependence.

To give an example of how the sample likelihood is evaluated using the concept of a hazard function, assume away any complications arising from the competing risks for now. Let λ denote the overall hazard out of a given state. In the absence of any unobserved heterogeneity, the likelihood function contribution of a single employment spell which ended at duration t would be

$$L_e(t) = \lambda(t, x_t) \prod_{v=1}^{t-1} [1 - \lambda(v, x_v)].$$

In a competing risks specification with layoff and quit hazards (not allowing for unobserved factors), the unconditional probability of someone leaving employment through a quit at duration t would become

$$L_e^q(t) = \lambda_q(t, x_t) \prod_{v=1}^{t-1} [1 - \lambda_q(v, x_v)][1 - \lambda_l(v, x_v)],$$

where λ_q and λ_l denote the quit and layoff hazards respectively. Similarly, for someone who gets laid off in week t of an employment spell, the likelihood contribution becomes

$$L_e^l(t) = \lambda_l(t, x_t) \prod_{v=1}^{t-1} [1 - \lambda_q(v, x_v)][1 - \lambda_l(v, x_v)].$$

[28] In order to streamline notation, we do not use individual i subscript in any of the formulas.

Hazard models are natural candidates for dealing with the problem of right-censoring. For an employment spell which is still in progress at the end of our sampling frame (i.e. no transition out of employment has been observed), one enters the survival probability

$$S_e(T) = \prod_{v=1}^{T} [1 - \lambda_q(v, x_v)][1 - \lambda_l(v, x_v)].$$

Here, T denotes the highest duration at which we observe the spell in progress and $S_e(T)$ gives the probability of a given spell lasting at least T periods. The sample likelihood then equals the product of individual likelihood contributions. Now, if we introduce the unobserved heterogeneity, the likelihood function contribution for someone leaving $e$mployment at duration t by way of a layoff would be

$$L_e^l(t) = \sum_{k=1}^{N_\theta^l} \sum_{m=1}^{N_\theta^q} p(\theta_k^l, \theta_m^q) L_e^l(t|\theta_k^l, \theta_m^q), \tag{5}$$

where $p(\theta_k^l, \theta_m^q)$ is the probability of having the unobserved components θ_k^l and θ_m^q in the layoff and quit hazards respectively, and where

$$L_e^l(t|\theta_k^l, \theta_m^q) = \lambda_l(t, x_t|\theta_k^l) \prod_{v=1}^{t-1} [1 - \lambda_l(v, x_v|\theta_k^l)][1 - \lambda_q(v, x_v|\theta_m^q)]. \tag{6}$$

The previous discussion focuses on examples with a single spell of each type. Equation 7 gives the likelihood contribution of a person with two completed spells of employment. The first spell starts in week $t+1$ and ends with a layoff in week s (at duration $s-t$); the second spell starts in week $r+1$ and ends with a quit in week w (at duration $w-r$).

$$L(s, w) = \sum_{k=1}^{N_\theta^q} \sum_{m=1}^{N_\theta^l} p(\theta_k^q, \theta_m^l) L_e^l(s|\theta_k^q, \theta_m^l) L_e^q(w|\theta_k^q, \theta_m^l) \tag{7}$$

Here θ^q and θ^l denote the unobserved terms entering quit and layoff hazards respectively and

$$L_e^l(s|\theta_k^q, \theta_m^l) = \lambda_l(s, x_s|\theta_m^l) \prod_{v=t+1}^{s-1} [1 - \lambda_q(v, x_v|\theta_k^q)][1 - \lambda_l(v, x_v|\theta_m^l)], \tag{8}$$

$$L_e^q(w|\theta_k^q,\theta_m^l) = \lambda_q(w, x_w|\theta_m^l) \prod_{v=r+1}^{w-1} [1-\lambda_q(v, x_v|\theta_k^q)][1-\lambda_l(v, x_v|\theta_m^l)]. \quad (9)$$

In order to control for selection bias, the unemployment and employment hazards have to be estimated *jointly*. One has to take into account the joint density of the unobservables across the two hazards, denoted by $p(\theta^u, \theta^e)$. Suppose we want to estimate a competing risks specification for quits and layoffs jointly with an overall hazard for unemployment. The likelihood contribution of someone leaving the first unemployment spell after t weeks, then getting laid off after $s-t$ weeks on a job and staying in the second unemployment spell till the date of the interview, say at $T-s$ weeks into the last spell, then becomes

$$L^{u,l,u}(t,s,T) = \sum_{k=1}^{N_\theta^u}\sum_{m=1}^{N_\theta^q}\sum_{n=1}^{N_\theta^l} p(\theta_k^u,\theta_m^q,\theta_n^l) L_u(t|\theta_k^u) L_e^l(s|\theta_m^q,\theta_n^l) S_u(T|\theta_k^u), \quad (10)$$

where

$$L_u(t|\theta_k^u) = \lambda_u(t, x_t|\theta_k^u) \prod_{v=1}^{t-1} [1-\lambda_u(v, x_v|\theta_k^u)]. \quad (11)$$

The employment contribution, L_e^l is defined in equation 8 and finally

$$S_u(T|\theta_k^u) = \prod_{v=s+1}^{T} [1-\lambda_u(v, x_v|\theta_k^u)] \quad (12)$$

is the survivor function expressing the probability of a given spell lasting at least T periods.

One can compute individual contributions to the sample likelihood for other labor market histories in a similar way. The number of points of support of the distribution of unobservables (N_θ^u, N_θ^q and N_θ^l) is assumed to be finite and is determined from the sample likelihood. Note the assumption of θ^u, θ^q and θ^l staying the same across multiple unemployment and employment spells respectively. Detailed estimation strategy issues are discussed below.

5.2 Employment Hazard Estimates

The employment hazard empirical specifications capture the effect of explanatory variables on the length of an employment spell, which does not correspond to the cummulated job duration (seniority) for recalled workers. Our focus is on the effect of UI on job separations, and not on the issue of seniority. The amount of potential UI compensation -which is computed for each individual at each point in time- is based on the length of employment spells and earnings in the base period and does not depend on the duration of a specific worker-firm employment relationship.

We start by estimating the employment hazard functions with no unobserved heterogeneity. In terms of the notation introduced in Section 5.1, $\theta_k = \theta \ \forall k$. Table 2 contains the estimates for the competing risks employment hazard functions based on assuming firms and workers do *not* take eligibility for future UI claims into account. Let us first discuss the layoff hazard estimates. In column (1) we control for the potential UI compensation by including a dummy variable equal to one in each week when a given worker would be entitled for UI in the case of a layoff. We also control for the potential dollar amount of weekly UI benefits. Being entitled to UI compensation significantly raises the layoff hazard. The negative estimate of the potential benefits coefficient contradicts the economic intuition of our model but is not precisely estimated. Higher benefits lead to lower risks of layoff in the adjustment cost models (e.g. Card and Levine 1992).

Next, we allow for effects of the *length* of available entitlement, conditional on the worker being eligible. Specifically, we add a step function in the value of entitlement. The base case are those with more than 52 weeks of available UI compensation.[29] The table also reports the fraction of weekly observations covered by each of the entitlement steps. Column (2) lists the estimated coefficients which indicate that, conditional on eligibility, the amount of

[29] We have also estimated specifications including a dummy indicating the first week when a worker becomes eligible, but the estimated coefficient never reached conventional levels of statistical significance. This might suggest that in the U.S., unlike in Canada (see Baker and Rea 1993), the agents' ability to precisely impute the timing of eligibility is low. Alternatively, the optimal job duration in the U.S. could be longer than that required for UI eligibility even in firms which are engaged in temporary layoff strategies, perhaps because of lower volatility of demand and consequently lower layoff pressures during periods of low demand. Finally, U.S. firms might be less willing to keep workers they intend to lay off permanently just to ensure their UI coverage.

entitlement plays no role in the firm's layoff decisions as the steps in entitlement are neither individually nor jointly significant.[30] The estimated quit hazard function is presented in Column (3). Being entitled to UI compensation has no effect on the quit probability. UI compensation played no significant role in any of the quit hazards we have estimated.

Part of the entitlement variation comes from various extended benefits programs which trigger on and off at different points in time across states. The actual trigger dates of these programs depend on the level of the state or national insured unemployment rate. Properly controlling for the demand side effects is therefore important for disentangling demand effects from the effect of longer entitlement. To measure demand effects we use the monthly state unemployment rate average and deviation from this state specific mean. We also use the industry specific national monthly unemployment rate.[31] Controlling for demand conditions was successful in that all of the significantly estimated demand effect coefficients have the expected sign. Higher levels of the state unemployment rate (in deviation from a mean) significantly raise the layoff probability. Averages of the state unemployment rates contain state specific long-term levels of unemployment and could be confounded by other time-invariant state specific effects (there are no state dummies in any of the specifications). This coefficient is not precisely estimated in the layoff hazard, while the variable significantly reduces the quit hazard. Workers appear to be more cautious about quitting their jobs in regions with persistently high unemployment rates.

We also control for a standard set of demographic regressors including the TAA dummy, which equals to one when the worker can receive TAA compensation in the case of a layoff. The probability of exit from a given state is also allowed to vary with seasonal effects by adding a set of quarterly dummies to each specification. In all hazards we control for the industry class and a set of year dummies. TAA workers are less likely to quit, while the effect

[30]We have experimented with different choices of the base case and the finding of no significant impact of any of the entitlement steps was robust to the base case choice.

[31]We experimented with other demand measures with no impact on the estimates of interest.

on the layoff hazard is not precisely measured conditional on the industry unemployment rate and a set of industry dummies. Workers with higher wages are significantly less likely to exit their jobs in both employment hazards. Highly educated workers are significantly more likely to quit their jobs but are less likely to be laid off. Age plays an important role in both hazards, reducing the likelihood of a quit and affecting the layoff decisions in a nonlinear way where both younger and older workers are at a higher risk of layoff. Being a union member has a large and significant effect on reducing both of the hazards. If the current employment spell is in fact a recall spell, the probability of being laid off is higher, while quits become less likely.

The effect of spell duration on the transition probabilities is specified as a step function in duration, with each step chosen to cover at least 5% of transitions.[32] Such flexible parametrization should avoid any influence of the duration dependence specification on estimation of other coefficients. A full set of the baseline hazard estimates for columns (1) and (3) is reported in Table 8 in Appendix B.

Next, unobserved heterogeneity is allowed for in the estimation procedure. Controlling for unobserved person specific characteristics has been important in a number of empirical applications (e.g. Ham and LaLonde 1996), and we carry out a sensitivity analysis of using different distributional assumptions for the heterogeneity terms. The primary tool for dealing with unobserved factors is a heterogeneity distribution which uses N-tuples of unobserved factors (McCall 1996), where N is the number of hazard functions to be estimated. First, we estimate the employment competing risks with a 2-tuple distribution, allowing the unobserved factors in the layoff and quit hazards to be correlated. Second, reemployment and job exit processes create correlation between unobserved characteristics in different types of spells. Thus, the employment hazard, the unemployment hazard functions and the unob-

[32]For a similar approach see Meyer (1990). In the specifications with no unobserved heterogeneity, we also experimented with richer specifications using 2.5% steps in duration, with no effect on the parameters of interest.

served heterogeneity distribution is estimated jointly, allowing for a full correlation structure of the unobservables. This general type of heterogeneity is parametrized using the 3-tuple distribution described in Table 3, where u, l and q denote overall unemployment hazard, layoff and quit employment hazards, respectively. K denotes the number of estimated points of support of the mixing distribution. Our example of a likelihood contribution from equation 10 for someone leaving the first unemployment spell after t weeks , then getting laid off after $s-t$ weeks on a job and staying in the second unemployment spell till the date of the interview, say at $T-s$ weeks into the last spell, now becomes

$$L^{u,l,u}(t,s,T)=\sum_{k=1}^{K}p(\Theta_k)L_u(t|\theta_k^u)L_e^l(s|\theta_k^q,\theta_k^l)S_u(T|\theta_k^u). \tag{13}$$

Table 4 reports the layoff UI coefficients from the heterogeneity estimation. We have estimated both i) specifications allowing for the amount of entitlement and ii) specifications conditional on only the eligibility dummy. The no-heterogeneity results suggest using the more parsimonious specification. Further, in most specifications the entitlement steps were not jointly significant. Hence, we present the parsimonious estimation here and report the results including the step function in entitlement in Table 9 in Appendix B. The quit hazard UI coefficients were not significant in any of the specifications and are not reported. We also do not report the demographic and demand coefficients, which were not affected by introducing heterogeneity except as noted below. Column (1) is taken from Table 2 for comparison. The estimates from the specifications with 2-tuple heterogeneity distribution (quit and layoff) are presented in column (2). Introducing unobserved heterogeneity was strongly supported by the estimated sample likelihood.[33] Although the UI parameters are not affected by introducing the 2-tuple heterogeneity, both the recall and union dummy estimates in the layoff hazard increase by more than four times the size of their standard

[33]Log-likelihood improved by 47.2 when going from no heterogeneity to 2 points of support for 2-tuples when there were 3 more coefficients to be estimated. To make this comparison to the joint log-likelihood of quits and layoffs from colum (2), one has to sum up the quit and layoff no-heterogeneity log-likelihoods, which were estimated separately.

errors. None of the quit hazard coefficients was sensitive to unobserved factors. Column (3) contains the estimates from a specification where sample selection is controlled. The employment durations are estimated jointly with the overall unemployment hazard using the 3-tuples heterogeneity distribution from Table 3 with two points of support (i.e. $K = 2$). The positive layoff effect of being eligible increases slightly, but correcting for selection bias was not very important as none of the coefficients moved by more than the size of their standard errors.

When searching for additional (more than 2) points of support for 3-tuple heterogeneity, the likelihood was unbounded in large negative values of one of the heterogeneity terms in the quit hazard. This suggested estimation of a defective risk model, with a heterogeneity distribution parametrizing the probability of never leaving employment through a quit. Further motivation for this type of estimation comes from the empirical hazard literature, which argues that for processes in which the probability of exit is very low, one should reflect this fact in the estimation by parametrizing the probability of never leaving a given state.[34] Heckman and Walker (1990) use the general framework developed in Heckman and Singer (1984) to allow for defective risks in the context of unobserved heterogeneity in a continuous time, single exit model. Here, a similar approach is applied to discrete time estimation with multiple exits.

There is a natural way of incorporating defective risk probabilities into the N-tuple heterogeneity distribution. In doing so one retains the richness of the estimated heterogeneity distribution while adding a new dimension to it. The empirical strategy used here is to estimate as many points of support for the usual N-tuple heterogeneity as possible and then substitute a fixed large negative value for those unobserved factors which pointed in the direction of the defective risk in the previous estimation. This large negative $\theta = -M$ is

[34] For example, Schmidt and Witte (1989) look at the probability of returning to prison for a sample of formerly arrested individuals. They parametrize both the probability of eventual return and the timing of return.

not to be estimated, and only the probability of having this unobserved factor, i.e. of never leaving a given state, enters the maximization problem.[35] All other explanatory variables are excluded from the hazard with the absorption θ. This strategy incorporates the traditional defective risk (absorption state, stayer) model into a competing risk setting with unobserved heterogeneity.

For example, suppose that we only estimate the absorption probability for the quit hazard with a corresponding θ^l and that there are two points of support for the N-tuple heterogeneity with a full set of thetas (θ^l, θ^q) to be estimated. In terms of equation 7, $N_\theta^q = 2, N_\theta^l = 3$. There are three probabilities to be estimated (which requires only two parameters). Equation 7 is used with $p(\theta_1^q, \theta_1^l)$ and $p(\theta_2^q, \theta_2^l)$. For $p_s = p(\theta^q = -M, \theta_3^l)$, the likelihood contribution in case of a layoff would be

$$L_e^l(t|\theta^q = -M, \theta_3^l) = \lambda_l(t, x_t|\theta_3^l) \prod_{v=1}^{t-1} [1 - \lambda_l(v, x_v|\theta_3^l)], \tag{14}$$

and it would equal zero if the transition were a quit. Finally, to give an example of joint estimation using the hypothetical observed employment history from equation 10, the joint likelihood contribution would now be

$$\begin{aligned} L(t,s,T) = & \sum_{k=1}^{N_m} p(\Theta_k) L_u(t|\theta_k^u) L_e^l(s|\theta_k^q, \theta_k^l) S_u(T|\theta_k^u) + \\ & \sum_{j=1}^{N_s} p(\Theta_j) L_u(t|\theta_j^u) L_e^l(s|\theta^q = M, \theta_j^l) S_u(T|\theta_j^u), \end{aligned} \tag{15}$$

with N_m denoting the number of points of support for the full heterogeneity with movers (i.e. the number of 3-tuples of heterogeneity factors), N_s denoting the number of points of support for the stayers with degenerate heterogeneity (i.e. the number of 2-tuples of heterogeneity factors with $\theta^q = -M$), and $L_e^l(s|\theta^q = -M, \theta_j^l)$ being defined in Equation 14. The reported specifications actually include a heterogeneity distribution with two points of

[35] We use $M = 100$ in the estimation, which sets the (quit) hazard at 10^{-43}.

support for the full N-tuples and with two quit absorption probabilities in order to allow for more than one layoff heterogeneity type (θ^l) with zero probability of a quit.[36]

These defective risks heterogeneity results are presented in column (4) of Table 4. The estimated likelihood function improves upon the maximized value of the model with no absorption state heterogeneity. The estimate of the eligibility dummy increases by approximately the size of its standard error when compared to the joint specification in column (3). The positive eligibility coefficient is now almost twice as large as the no-heterogeneity estimate in column (1), although still within two standard deviations. Except for the increase in the potential benefits coefficient, the other estimates were almost identical to those from the more conventional models. The estimated probability of never leaving employment through a quit is 0.08 (with a corresponding standard deviation of 0.021).

The specifications in columns (3) and (4) involve also estimating the overall unemployment hazard. A set of coefficients with no unobserved heterogeneity is reported in Table 7 in Appendix B. Large values of entitlement and higher UI benefits make unemployed workers less likely to leave unemployment. Such findings are in accord with both job search models and empirical unemployment duration literature. None of the unemployment hazard coefficients was sensitive to introducing the heterogeneity factors.

Finally, the specifications based on allowing workers and firms to account for the possibility of future multiple UI claims are presented in Table 5. The reported means of the entitlement and eligibility dummies show that the imputation procedure now makes more workers eligible and increases the average amount of available entitlement. When we control for the effect of UI eligibility and benefits on employment durations, the estimates in columns (1) and (3) are not affected by the different assumptions regarding future claims. Being entitled to UI compensation makes quits less likely, but the effect is not precisely esti-

[36]Searching for additional points of support for this most general heterogeneity distribution resulted in trivial increases of the log-likelihood.

mated. Column (2) lists estimates which control for the length of available UI entitlement. Compared to Table 2, accounting for future claims affects the parameters of interest as the eligibility dummy coefficient is now relatively small and insignificant.[37] When controlling for unobserved heterogeneity we again estimate both specifications with and without the step function in entitlement. The entitlement steps are not jointly significant at the 5% level in six out of the total of eight estimated layoff hazards (four for each assumption about agents' ability to impute eligibility for future UI spells). Moreover, most of the estimated UI coefficients in these richer specifications are imprecisely estimated, and we conclude that the data does not allow separate identification of both the eligibility and entitlement effects. We proceed with the more parsimonious parametrization and report heterogeneity estimation which controls for the length of entitlement in Appendix B.

Introducing unobserved heterogeneity in columns (2) to (4) of Table 6 quantitatively affects the eligibility coefficient. When we estimate the two employment hazards jointly, allowing for a 2-tuple correlated heterogeneity distribution, the coefficient becomes smaller and insignificant. Controlling for the effects of sample selection in column (3), however, raises the estimated eligibility effect by more than one standard error size and the defective risk (stayer) heterogeneity estimation in column (4) confirms the large significant positive effect of UI eligibility on layoffs.[38] The behavior of all remaining coefficients was similar to the estimation based on not accounting for future UI claims. The unemployment hazard was not materially affected by the type of assumption regarding future UI claims and is reported in Appendix B.[39]

[37]Further, given that we control for eligibility, having 14 to 26 weeks of available entitlement significantly raises the layoff hazard. The joint likelihood ratio test for the four entitlement steps, however, does not reach conventional levels of statistical significance.

[38]The defective risk heterogeneity estimation with entitlement steps in Table 10 also confirms the large positive effect of eligibility from the more parsimonious specification of column (4).

[39]Note that most of the unemployment data comes from the initial unemployment spells. Since the data set does not include information on the employment histories preceding the initial spell of unemployment, we can only control for the possibility of multiple UI claims in the second and third spell of unemployment. The extent to which the value of entitlement is affected by the future claim assumption in the unemployment

6 Conclusion

Empirical evidence on the effect of UI coverage on employment durations is scarce. The present study employs methods similar to those used in the unemployment duration literature to examine how the UI system affects duration of employment. It uses variation in entitlement levels and benefits (both over time and cross-sectional) to quantify the effect of UI coverage on terminating job spells. Unemployment and employment spells are analyzed jointly in order to control for selection into multiple spells. The empirical results suggest that eligibility for UI compensation significantly raises the probability of a layoff. Conditional on eligibility, however, neither the length nor the dollar amount of the UI compensation to which workers are entitled appear to affect the risks of layoff. No aspect of UI affects the probability of a quit in any of the estimated specifications.

Further, we find a relatively small effect of sample selection. This finding is reassuring for empirical applications which use multiple spell unemployment data to estimate the effect of the UI system on outflow from unemployment (e.g. Ham and Rea 1987). Our most general heterogeneity specification also allows for the possibility of defective risks, an important consideration when the probability of a particular type of exit is very low for a fraction of the sample, as is the case with quits in the current study.

This paper also focuses on how different assumptions about the ability of firms and workers to impute available UI compensation affect the estimation. Alternative assumptions about determining eligibility for future UI claims affected the results in some of the specifications, but the estimated layoff effect of UI eligibility is almost identical in two important cases: when we do not control for unobserved worker specific characteristics and when we use the most general heterogeneity distribution.

Our theoretical model predicts that the layoff probability should increase with the length

hazards is, therefore, much smaller compared to the employment hazards where we have enough information to impute future UI claims even in the first employment spells.

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Table 1: Individual and Spell Characteristics

Unemployment Spells

Number of Individuals = 808 Number of Spells = 1375

Variable	Mean	Std.Dev.	Dummy Variable	Mean
Education	11.4	(2.45)	Union	0.52
Age	34.6	(12.0)	Previous Recall	0.46
Previous Wage	242.	(105.)	UI Non-recipient	0.15
UI Entitlement($if > 0$)	52.5	(13.8)		
UI Benefits ($if > 0$)	113.	(34.3)		
Duration	34.9	(46.3)		

Employment Spells

Number of Individuals = 808 Number of Spells = 1074

Wage	251.	(116.)	Recall	0.61
Potential Entitlement	22.2	(23.9)	Quit	0.25
Potential UI Benefits	62.3	(61.4)		
Duration	85.1	(72.3)		

(Duration and entitlement are in weeks; all wages and benefits are weekly measures in 1975 dollars.)

Table 2: Employment Hazard Estimates Not Accounting for Future UI Claims

Type of Hazard			Layoff	Layoff	Quit
Variable		[Mean]	(1)	(2)	(3)
Potential UI Eligibility, Benefits and Weeks of Entitlement					
Eligibility Dummy		[0.93]	0.448**	0.449*	0.00156
			(0.207)	(0.248)	(0.296)
Entitlement	39 to 52	[0.30]	—	0.0776	—
				(0.154)	
	27 to 39	[0.11]	—	-0.197	—
				(0.192)	
	14 to 26	[0.11]	—	0.0974	—
				(0.186)	
	01 to 13	[0.07]	—	-0.0407	—
				(0.199)	
Weekly Benefits $*10^{-2}$			-0.200	-0.196	-0.0708
			(0.159)	(0.159)	(0.279)
Demographics					
TAA Dummy			0.111	0.0964	-1.11***
			(0.164)	(0.166)	(0.434)
Weekly Wage $*10^{-2}$			-0.141***	-0.140***	-0.443***
			(0.559)	(0.0558)	(0.114)
Years of Education			-0.0307	-0.0311	0.124***
			(0.0206)	(0.0206)	(0.0365)
Age			-0.0704**	-0.0724***	-0.0195**
			(0.0271)	(0.0272)	(0.00813)
Age SQ $*10^{-2}$			0.0842***	0.0869***	—
			(0.0332)	(0.0333)	
Union Dummy			-1.06***	-1.06***	-0.758*
			(0.137)	(0.137)	(0.451)
Recall Spell Dummy			0.390***	0.392***	-1.01***
			(0.151)	(0.151)	(0.379)
Demand Conditions					
Average of State Unemployment			-0.416	-0.297	-2.05**
Rate $*10^{-1}$			(0.532)	(0.547)	(0.889)
Deviations from Average State			1.34**	1.33**	0.437
Unemployment Rate $*10^{-1}$			(0.593)	(0.594)	(0.987)
Industry National			0.805***	0.810***	0.481
Unemployment Rate $*10^{-1}$			(0.217)	(0.218)	(0.400)
Log-Likelihood			-2735.4	-2733.3	-993.13

Standard errors in parentheses. All specifications include annual, quarterly and industry dummies as well as a step function in duration, which is reported for Columns (1) and (3) in Table 8 in Appendix B. $*$ denotes significance at 10% level; $**$ denotes significance at 5% level; $***$ denotes significance at 1% level

Table 3: Heterogeneity Distribution with 3-tuples

$p(\Theta_1)$	$\Theta_1 = \{\theta_1^u, \theta_1^l, \theta_1^q\}$
$p(\Theta_2)$	$\Theta_2 = \{\theta_2^u, \theta_2^l, \theta_2^q\}$
...	...
$p(\Theta_K)$	$\Theta_K = \{\theta_K^u, \theta_K^l, \theta_K^q\}$

Table 4: Employment Hazard Estimates with Unobserved Heterogeneity not Accounting for Future Claims

Heterogeneity	No	2-tuple	Joint	Joint/Stayer
Variable	(1)	(2)	(3)	(4)
Eligibility Dummy	0.448**	0.487*	0.539*	0.841***
	(0.207)	(0.277)	(0.290)	(0.278)
Weekly Benefits $*10^{-2}$	-0.200	-0.255	-0.283	-0.414*
	(0.159)	(0.210)	(0.218)	(0.222)
Log-Likelihood	-2735.4	-3681.3	-8171.8	-8129.3

Standard errors in parentheses. For a standard set of regressors see Table 2 .

Table 5: Employment Hazard Estimates Accounting for Future UI Claims

Type of Hazard			Layoff	Layoff	Quit
Variable		[Mean]	(1)	(2)	(3)
Potential UI Eligibility, Benefits and Weeks of Entitlement					
Eligibility Dummy		[0.95]	0.452**	0.299	-0.275
			(0.225)	(0.255)	(0.307)
Entitlement	39 to 52	[0.34]	—	0.151	—
				(0.133)	
	27 to 39	[0.12]	—	0.115	—
				(0.160)	
	14 to 26	[0.06]	—	0.460**	—
				(0.191)	
	01 to 13	[0.01]	—	-0.242	—
				(0.340)	
Weekly Benefits $*10^{-2}$			-0.196	-0.193	0.184
			(0.160)	(0.159)	(0.276)
Demographics					
TAA Dummy			0.109	0.180	-1.17***
			(0.164)	(0.171)	(0.434)
Weekly Wage $*10^{-2}$			-0.142***	-0.143***	-0.484***
			(0.563)	(0.0562)	(0.114)
Years of Education			-0.0310	-0.0316	0.122***
			(0.0206)	(0.0206)	(0.366)
Age			-0.0713***	-0.0715***	-0.0208***
			(0.0271)	(0.0272)	(0.0815)
Age SQ $*10^{-2}$			0.0851***	0.0861***	—
			(0.0332)	(0.0333)	
Union Dummy			-1.06***	-1.06**	-0.762*
			(0.137)	(0.137)	(0.451)
Recall Spell Dummy			0.386***	0.368**	-1.00***
			(0.151)	(0.152)	(0.380)
Demand Conditions					
Average of State Unemployment			-0.378	-0.144	-1.92*
Rate $*10^{-1}$			(0.531)	(0.549)	(0.885)
Deviations from Average State			1.33**	1.29**	0.365
Unemployment Rate $*10^{-1}$			(0.593)	(0.593)	(0.987)
Industry National			0.802***	0.794***	0.481
Unemployment Rate $*10^{-1}$			(0.218)	(0.217)	(0.402)
Log-Likelihood			-2735.8	-2731.8	-992.77

Standard errors in parentheses. All specifications include annual, quarterly and industry dummies as well as a step function in duration. $*$ denotes significance at 10% level; $**$ denotes significance at 5% level; $***$ denotes significance at 1% level

Table 6: Employment Hazard Estimates with Unobserved Heterogeneity Accounting for Future Claims

Heterogeneity	No	2-tuple	Joint	Joint/Stayer
Variable	(1)	(2)	(3)	(4)
Eligibility Dummy	0.452**	0.384	0.777***	0.810***
	(0.225)	(0.273)	(0.274)	(0.291)
Weekly Benefits $*10^{-2}$	-0.196	-0.118	-0.167	-0.280
	(0.160)	(0.196)	(0.198)	(0.211)
Log-Likelihood	-2735.8	-3681.7	-8172.8	-8127.0

Standard errors in parentheses. For a standard set of regressors see Table 5 .

Appendix

A Kaplan-Meier Empirical Hazards

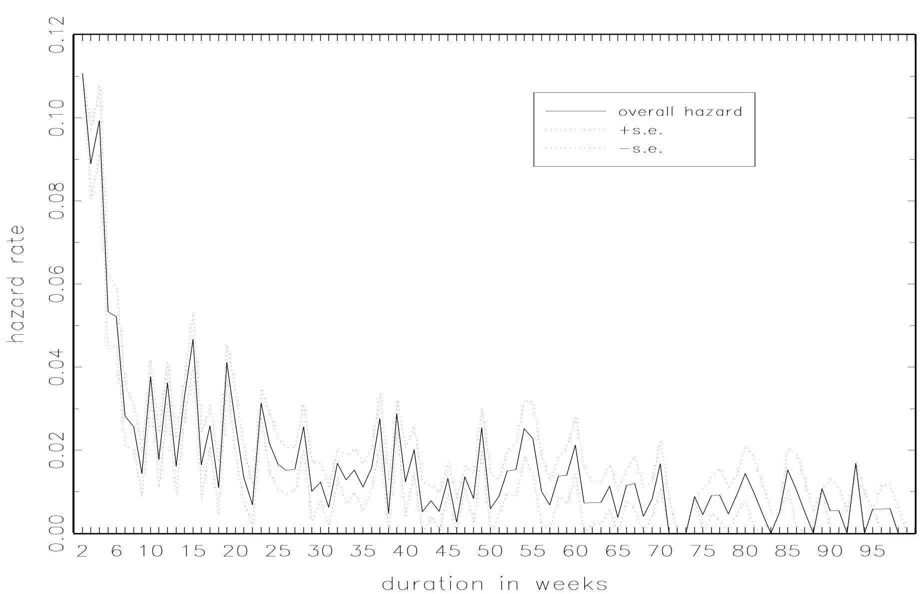

Figure 1: Overall Empirical Hazard for Unemployment Spells

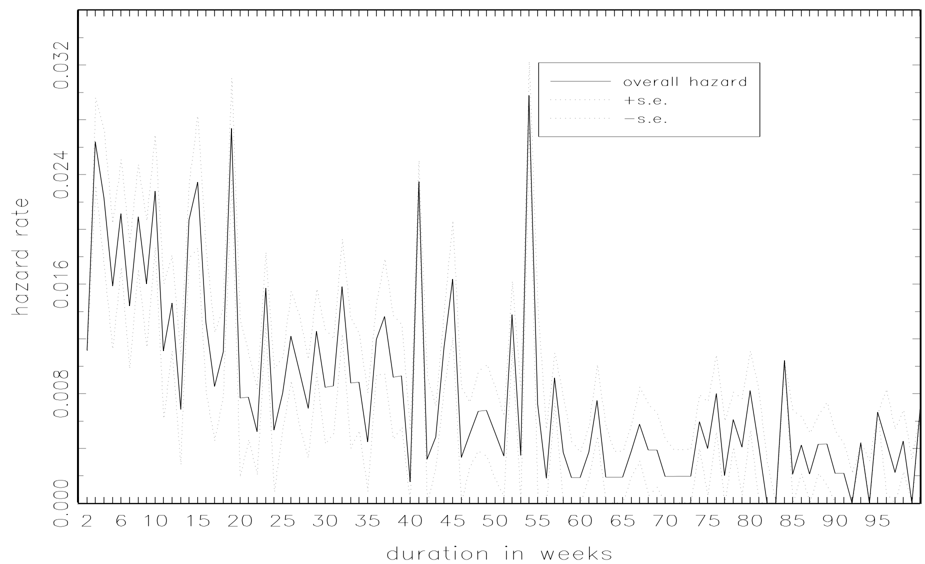

Figure 2: Empirical Hazards for Employment Spells: Overall Hazard

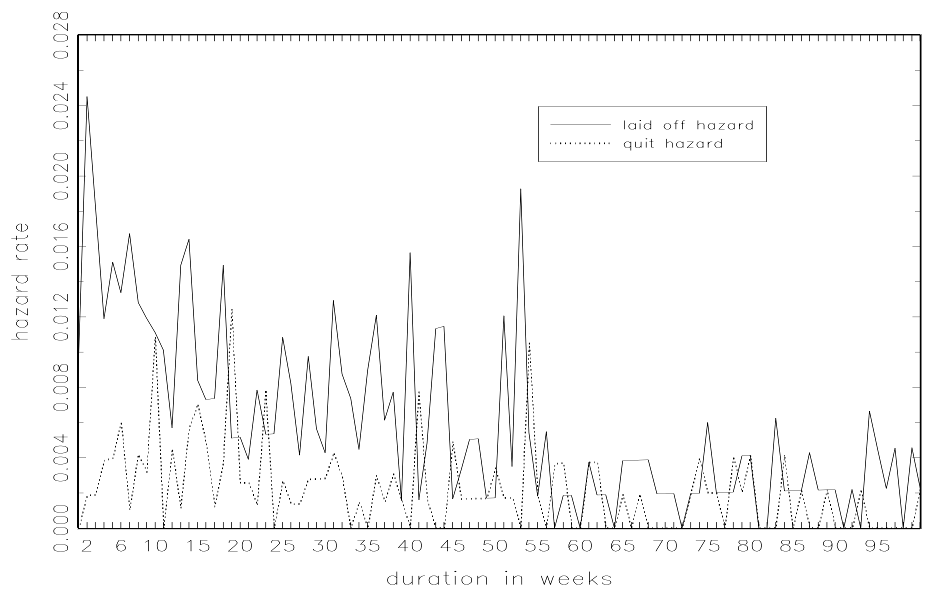

Figure 3: Empirical Hazards for Employment Spells: Competing Risks

B Additional Hazard Function Estimates

Table 7: Unemployment Hazard Estimates with No Unobserved Heterogeneity

Accounting for Future Claims		No		Yes
Variable	[Mean]	(1)	[Mean]	(2)
Weeks of Remaining UI Entitlement and UI Benefits				
Entitlement over 52	[0.08]	-0.585**	[0.09]	-0.671***
		(0.222)		(0.224)
39 to 52	[0.11]	-0.346*	[0.12]	-0.404*
		(0.210)		(0.213)
27 to 39	[0.12]	-0.666***	[0.12]	-0.751***
		(0.219)		(0.221)
14 to 26	[0.11]	-0.676***	[0.11]	-0.706***
		(0.220)		(0.223)
01 to 13	[0.09]	-0.215	[0.08]	-0.0915
		(0.216)		(0.218)
No Eligibility Dummy	[0.25]	-0.816***	[0.23]	-0.847***
		(0.162)		(0.165)
Weekly Benefits $*10^{-2}$		-0.0576		-0.0510
		(0.131)		(0.131)
Demographics				
TAA Dummy		0.107		0.109
		(0.110)		(0.110)
Weekly Wage $*10^{-2}$		0.0519		0.0513
		(0.0354)		(0.0355)
Years of Education		-0.0208		-0.0205
		(0.0139)		(0.0139)
Age		0.0176		0.0171
		(0.0182)		(0.0182)
Age SQ $*10^{-2}$		-0.0199		-0.0191
		(0.0228)		(0.0228)
Previously Recalled Dummy		-0.369***		-0.365***
		(0.0721)		(0.0721)
Demand Conditions				
Average of State Unemployment		0.375		0.434
$*10^{-1}$		(0.407)		(0.408)
Deviations from Average State		-1.32***		-1.35***
$*10^{-1}$		(0.395)		(0.394)
Log-Likelihood		-4493.0		-4492.5

Standard errors in parentheses. All specifications include annual, quarterly and industry dummies as well as a step function in duration.

Table 8: Baseline Hazard Estimates for Table 2

Variable	*Layoff Hazard* (1)		*Variable*	*Quit Hazard* (3)	
Weekly Baseline Hazards					
Duration 3 to 4	-0.669	(0.248)	Duration 5 to 7	0.272	(0.420)
5 to 6	-0.136	(0.255)	8 to 13	0.437	(0.358)
7 to 8	-0.126	(0.256)	14 to 15	0.808*	(0.432)
9 to 11	-0.353	(0.251)	16 to 18	0.813**	(0.398)
12 to 14	-0.236	(0.247)	19 to 22	0.430	(0.424)
15 to 18	-0.478**	(0.232)	23 to 29	-0.100	(0.441)
19 to 24	-1.06***	(0.263)	30 to 36	-0.0258	(0.456)
25 to 28	-0.689***	(0.271)	37 to 41	0.399	(0.460)
29 to 33	-0.772***	(0.262)	42 to 49	-0.714***	(0.271)
34 to 38	-0.753***	(0.265)	50 to 54	0.651	(0.632)
39 to 44	-0.779***	(0.258)	55 to 64	0.102	(0.471)
45 to 53	-0.999***	(0.253)	65 to 78	-0.229	(0.470)
54 to 71	-1.77***	(0.288)	79 to 105	-0.521	(0.456)
72 to 92	-1.95***	(0.282)	106 to 154	-0.895*	(0.493)
93 to 122	-2.22***	(0.283)	155 and over	-2.32 ***	(1.07)
123 to 170	-2.62***	(0.287)		—	
171 and over	-3.01***	(0.424)		—	

Standard errors in parentheses.

Table 9: Employment Hazard Estimates with Unobserved Heterogeneity Not Accounting for Future Claims

Heterogeneity / *Variable*	No (1)	2-tuple (2)	Joint (3)	Joint/Stayer (4)
Eligibility Dummy	0.449*	0.503	0.556*	0.930***
	(0.248)	(0.320)	(0.337)	(0.335)
Entitlement 39 to 52	0.0776	0.186	0.194	0.126
	(0.154)	(0.183)	(0.193)	(0.200)
27 to 39	-0.197	-0.189	-0.170	-0.206
	(0.192)	(0.222)	(0.233)	(0.256)
14 to 26	0.0974	0.0213	0.0399	0.0440
	(0.186)	(0.228)	(0.240)	(0.243)
01 to 13	-0.0407	-0.210	-0.200	-0.175
	(0.199)	(0.240)	(0.258)	(0.268)
Weekly Benefits $*10^{-2}$	-0.196	-0.292	-0.327	-0.492**
	(0.159)	(0.210)	(0.219)	(0.229)
Log-Likelihood	-2733.3	-3678.2	-8168.6	-8126.7

Standard errors in parentheses. For a standard set of regressors see Table 2 .

Table 10: Employment Hazard Estimates with Unobserved Heterogeneity Accounting for Future Claims

Heterogeneity / *Variable*	No (1)	2-tuple (2)	Joint (3)	Joint/Stayer (4)
Eligibility Dummy	0.299	0.0900	0.196	0.638*
	(0.255)	(0.321)	(0.342)	(0.338)
Entitlement 39 to 52	0.151	0.312**	0.338**	0.201
	(0.133)	(0.147)	(0.159)	(0.165)
27 to 39	0.115	0.285	0.318	0.115
	(0.160)	(0.187)	(0.201)	(0.212)
14 to 26	0.460**	0.636***	0.667***	0.522**
	(0.191)	(0.229)	(0.246)	(0.254)
01 to 13	-0.242	-0.193	-0.205	-0.226
	(0.340)	(0.402)	(0.445)	(0.537)
Weekly Benefits $*10^{-2}$	-0.193	-0.0922	-0.165	-0.298
	(0.159)	(0.197)	(0.210)	(0.205)
Log-Likelihood	-2731.8	-3675.9	-8163.0	-8123.0

Standard errors in parentheses. For a standard set of regressors see Table 5 .