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UNEMPLOYMENT INSURANCE RULES, JOBLESSNESS, AND PART-TIME WORK

BY BRIAN P. MCCALL[1]

In most states, unemployment insurance recipients accepting part-time work can earn up to a specific amount (the "disregard") with no reduction in benefits. Benefits are then reduced on a dollar for dollar basis for earnings in excess of the disregard. The disregard varies both across states and within a state over time. This paper analyzes the effects of changes in the disregard on job search behavior. A continuous-time job search model is developed and under general conditions an increase in the disregard is shown to increase both the part-time and overall re-employment hazards. Data from the Current Population Survey's Displaced Worker Supplements are used to test these predictions. Estimates from a competing risks model with correlated risks and time-varying coefficients shows that increasing the disregard significantly increases the conditional probability of part-time re-employment during the first three months of joblessness.

KEYWORDS: Competing risks, disregard, optimal job search, selectivity bias, time-varying coefficients, unmeasured heterogeneity.

1. INTRODUCTION

UNEMPLOYMENT INSURANCE IS ONE of the major social insurance programs in the United States with benefits paid exceeding 25 billion dollars in 1992. Much of the empirical research involving unemployment insurance has focused on the incentive effects of weekly benefit amounts and benefit durations on the job search behavior of recipients (see Atkinson and Micklewright (1991) and Devine and Kiefer (1991) for recent surveys of the literature). Unemployment insurance in the United States is administered at the state level. Since its inception in the 1930's, state laws have had provisions that allowed individuals to continue receiving at least partial benefits while working part-time. Most compensated weeks of unemployment involve payment of full benefits. The fraction of compensated weeks involving partial benefit payments, however, has risen by nearly one-third over the last decade. Originally, the most common provision for partial benefits was to reduce weekly benefits to the point were benefits plus earnings equaled 120 percent of the weekly benefit amount (see Haber and Murray (1966)). Now, most states allow individuals to earn up to a certain amount (termed the "disregard") with no reduction in benefits. Benefits are then reduced on a dollar for dollar basis for earnings exceeding the disregard.

One question that arises is whether the level of the disregard influences an unemployment insurance (UI) recipient's job search behavior. In particular,

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does it affect a UI recipient's willingness to work part-time. This is the primary focus of this paper.

To study the effect of the disregard, this paper first develops a simple job search model which incorporates the institutional features of the unemployment insurance system including those regarding part-time work. When a UI recipient seeks both part-time and full-time work, the model predicts that increasing the disregard will increase the conditional probability of part-time re-employment, with the largest impact coming early in the unemployment spell. In addition, the model predicts that increasing the disregard will lower the expected unemployment durations of UI recipients.

These predictions are tested using micro-data from several years of the Current Population Survey's Displaced Workers Supplement. Previous empirical research on the incentive effects of partial benefit provisions has either focused on a single state in a single year (Munts (1970)), analyzed aggregate state data for a single year (Holen and Horowitz (1974)), or analyzed administrative data from a few states for a single year (Kiefer and Neumann (1979)). Thus one way in which this study differs from its predecessors is by analyzing micro-data from many states over numerous years during which time several states changed their disregard level. This state-time variation in the disregard helps identify the effect of an increase in the disregard on the re-employment behavior of UI recipients.

This study also differs from its predecessors in the econometric methodology that it employs to analyze the incentive effects of the disregard. In this study several competing risks models are estimated.[2] The benefit of this approach is that it allows one to simultaneously analyze the effect of the disregard on the probability of part-time re-employment and expected joblessness durations. The competing risks models used in this paper account for the fact that the data are grouped, risks may be correlated (Han and Hausman (1990) and Sueyoshi (1992)), the coefficients of the regressors may be time-varying, and in some specifications, the potential selectivity-bias (Heckman (1976)) that arises when the unobservable determinants of UI receipt are correlated with the unobservable determinants of joblessness durations and the propensity to accept part-time work. Finally, the estimation methods used in this paper account for the fact that the Displaced Worker Supplements only record the full-time/part-time status of the job worked in the week before the survey.

The empirical findings from the Displaced Worker Supplements provide some support for the theory. An increase in the disregard is found to significantly increase the conditional probability of part-time re-employment for UI recipients during approximately the first three months of joblessness. In model specifications that allow the coefficients of the UI regressors to vary over time, a ten percent increase in the disregard is estimated to increase the probability of part-time re-employment for UI recipients from 3.9 to 5.7 percent. Moreover, a

[2] Some recent analyses of unemployment durations using a competing risks approach include Katz (1986), Han and Hausman (1990), and Fallick (1991).

ten percent increase in the disregard is found to reduce expected joblessness durations from 0.3 to 0.9 percent.

2. THE INCENTIVE EFFECTS OF PART-TIME UNEMPLOYMENT INSURANCE RULES

To analyze the incentive effects of unemployment insurance provisions relating to part-time work on the job search behavior of displaced workers, a continuous-time job search model is developed.[3] To keep the analysis simple it is assumed that there is no wage uncertainty.[4] With no wage uncertainty, the first job offer is accepted. The only source of uncertainty is the elapsed time until a job offer arrives. This is assumed to depend on the intensity of search, which is under an individual's control.

As in Burdett and Mortensen (1978), suppose that when unemployed during the small interval of time $(t, t+dt)$ individuals may devote part of their time towards searching for a job and the remainder of their time towards consuming leisure (or home production). Further, suppose that there are two types of jobs that an individual may look for: full-time and part-time. Let k denote the fraction of time an individual spends consuming leisure, e^p the fraction of time (effort) an individual spends searching for a part-time job, and e^f the fraction of time an individual spends searching for a full-time job. Thus, while unemployed the fraction of time an individual spends consuming leisure equals $1-e^f-e^p$. For simplicity, assume that all part-time jobs pay the wage rate w^p and involve working h^p of the time while all full-time jobs pay the wage rate w^f and involve working h^f of the time. So, individuals are not entirely free to choose their hours of work in this model.

The probability of receiving a type i job offer in the small time interval $(t, t+dt)$ is assumed to be a function, $\lambda^i(e^i(t))\,dt$, of the effort, $e^i(t)$, an individual devotes to searching for a type i job during that interval, $i=f,p$.[5] The probability of receiving more than one job offer during the interval $(t, t+dt)$ is assumed to be negligible, or, more formally, $o_p(dt)$.[6] Furthermore,

[3] For some static equilibrium results concerning the effect of changes in unemployment insurance laws regarding part-time work, see Hotchkiss (1991).

[4] A more elaborate job search model that incorporates wage uncertainty with no recall could be developed. However, the implications for the re-employment hazard would not differ substantially.

[5] This model assumes that there is no "public good" aspect to job search. The empirical predictions of the theory about the effect of the disregard on both the conditional probability of re-employment and the conditional probability of part-time re-employment do not change, however, if seeking a part-time job also increases the probability of receiving a full-time job offer and vice versa, so long as increasing the intensity of part-time search (e^p) does not raise the probability of receiving a full-time job offer by more than it raises the probability of receiving a part-time job offer (and vice-versa).

[6] Recall $o_p(dt)$ implies that

$$\operatorname{plim}_{dt\to 0}\frac{o_p(dt)}{dt}=0.$$

assume that $\lambda^i(0) = 0$, $\lambda^{i\prime}(e^i(t)) > 0$, and $\lambda^{i\prime\prime}(e^i(t)) < 0$, for all $e^i(t)$, $i = f, p$. In this model an individual can seek full-time work while working part-time. Assume that the probability of receiving a full-time job offer in the time interval $(t, t+dt)$ is also a function, $\lambda^{oj}(e^{oj}(t))\,dt$, of the effort, $e^{oj}(t)$, an individual devotes to full-time job search where $\lambda^{oj}(0) = 0$, $\lambda^{oj\prime}(e^{oj}(t)) < 0$, and $\lambda^{oj\prime\prime}(e^{oj}(t)) < 0$, for all $e^{oj}(t)$. Once a full-time job is accepted it is assumed, for simplicity, that an individual remains there indefinitely.

Individuals are assumed to choose their search strategies so as to maximize expected discounted utility:

$$E_0\left(\int_0^\infty u(c(t), k(t))e^{-rt}\,dt\right),$$

where the utility in period $(t, t+dt)$ is approximately $u(c(t), k(t))\,dt$, $c(t)$ denotes the rate of consumption at time t, $k(t)$ denotes the rate of leisure consumed at time t, u is a strictly increasing function that is concave in c and k, r denotes the discount rate and the time horizon is infinite.

The rate of consumption or income flow, $c(t)$, for unemployed individuals who do not receive UI benefits is normalized to zero. Let b denote the benefit rate for unemployed individuals who receive UI benefits. Then, while unemployed, UI recipients have $c(t) = b$ for $0 < t \leq t^e$ and $c(t) = 0$ for $t > t^e$ where 0 represents the start of the unemployment spell and t^e denotes the benefit exhaustion point.[7]

Let $c^p = w^p h^p$ represent an individual's earnings from a part-time job and $c^f = w^f h^f$ represent an individual's earnings from a full-time job. Let y denote the disregard level. The states analyzed in this paper have the following benefit formulas for part-time work:

$$c(t) = \begin{cases} b + c^p & \text{if } c^p < y, \\ b + y & \text{if } y \leq c^p \leq b + ay, \\ c^p & \text{if } c^p > b + ay, \end{cases}$$

where $a = 0$ or 1. Thus, UI recipients still receive full benefits when their earnings from a part-time job are less than the level of disregard y. For earnings above y, UI benefits are reduced on a dollar for dollar basis when $a = 1$. In states where $a = 0$, UI benefits are also reduced on a dollar for dollar basis for earnings exceeding y but are discontinued when earnings exceed the weekly benefit amount.

For a majority of these states, the duration of benefits while working part-time is determined by

$$d^p = \min\left(\max\left(\frac{M - bt'}{b'}, 0\right), 52 - t'\right)$$

where t' denotes the time at which a part-time job is accepted. $M = bt^e$ is the total amount of UI payments an individual is eligible to receive, and b' equals

[7] For most states in the U.S., regular UI benefits are exhausted after 26 weeks.

the benefit rate while working part-time:

$$b' = \begin{cases} b & \text{if } c^p < y, \\ b - (c^p - y) & \text{if } y \le c^p \le b + y, \\ 0 & \text{if } c^p > b + y.^{8,9} \end{cases}$$

For example, if $b' = b$, then $d^p = t^e - t'$. Alternatively, if $b' = .5b$, than $d^p = \min(2(t^e - t), 52 - t')$. The total number of weeks an individual can receive UI benefits (both while unemployed and employed at a part-time job) cannot exceed 52 weeks.[10]

It is assumed that $u(c^p, 1 - h^p) < u(c^f, 1 - h^f)$. Thus individuals prefer to work full-time. This assumption may be unreasonable for some individuals, especially some women (see Blank (1992)). The empirical work below, however, focuses only on workers displaced from full-time jobs.[11]

The decision problem confronting the individual is analyzed using continuous-time stochastic dynamic programming techniques. The optimal solution to such problems can be analyzed by considering the limit of the discrete-time analog (see Plum (1991)). The model developed in this paper is potentially nonstationary (see Van den Berg (1990) for a general analysis of nonstationary search models where decisions are made only at offer arrival times).

Let $V^u(t)$ denote the optimal expected discounted utility at time t of a UI recipient. In addition, let $V^n(t)$ denote the optimal expected discounted utility at time t of a nonrecipient. It is clear that $V^n(t) = V^n$ for all t and $V^u(t) = V^n$ for all $t \ge t^e$. Also, let $V^{pu}(t)$ denote the optimal expected discounted utility of a UI recipient working part-time at time t and $V^{pn}(t) = V^{pn}$ denote the optimal expected discounted utility of a nonrecipient working part-time at time t. Again, it is clear that $V^{pu}(t) = V^{pn}$ when $w^p h^p \ge b + y$ or $t \ge t^{pe}$, where $t^{pe} = t' + d^p$.[12] Finally, let $V^f(t) = V^f$ denote the value function when working full-time.

Taking the limit as the time interval goes to zero of the discrete-time approximation of the Bellman equations yields

$$(1) \qquad V^f = \frac{U(c^f, 1 - h^f)}{r},$$

$$\begin{aligned}(2) \qquad & -V^{pu\prime}(t) + rV^{pu}(t) \\ & = \max_{e^{oj}(t) \ge 0}(U(\max(\min(b + y, c^p + b), c^p), 1 - h^p - e^{oj}(t)) \\ & \qquad + \lambda^{oj}(e^{oj}(t))V^f - \lambda^{oj}(e^{oj}(t))V^{pu}(t)),\end{aligned}$$

[8] For the remainder of this section we will focus only on the case $a = 1$. The results for $a = 0$ are similar.

[9] In the remaining states, the length of time an individual can receive benefits while working a part-time job equals $\min(0, t^e - t')$.

[10] Assuming no supplemental programs are in effect. For the time period and states considered in the empirical work this is a reasonable assumption (see below).

[11] Many of the theoretical findings reported below do not depend on this assumption.

[12] Note that t^{pe} is not deterministic since t' is random.

and

$$\text{(3)} \qquad rV^{pn} = \max_{e^{oj} \geq 0}(U(c^p, 1 - h^p - e^{oj}) + \lambda^{oj}(e^{oj})V^f - \lambda^{oj}(e^{oj})V^{pn}).$$[13,14]

Moreover, V^n satisfies the recursive equation

$$\text{(4)} \qquad rV^n = \max_{e^{fn} \geq 0, e^{pn} \geq 0}(U(0, 1 - e^{fn} - e^{pn}) + \lambda^f(e^{fn})V^f + \lambda^p(e^{pn})V^{pn} - (\lambda^f(e^{fn}) + \lambda^p(e^{pn}))V^n)$$

and $V^u(t)$ satisfies the recursive equation

$$\text{(5)} \qquad -V^{u\prime}(t) + rV^u(t) = \max_{e^f(t) \geq 0, e^p(t) \geq 0}(U(b, 1 - e^f(t) - e^p(t)) + \lambda^f(e^f(t))V^f + \lambda^p(e^p(t))V^{pu}(t) - (\lambda^f(e^f(t)) + \lambda^p(e^p(t)))V^u(t)).$$

Let $e^{f^*}(t)$ and $e^{p^*}(t)$ denote the optimal amounts of effort devoted to seeking full-time and part-time work at time t, respectively. For simplicity, assume throughout the remainder of this section that individuals seek both part-time and full-time jobs while unemployed and seek full-time jobs while working part-time. Then, $de^{f^*}(t)/dV^p(t) < 0$ and $de^{p^*}(t)/dV^p(t) > 0$. Thus, as the value of part-time work increases, the fraction of time spent seeking it increases and the fraction of time spent seeking full-time work decreases. Moreover, it can be shown that the total fraction of time spent seeking work at t increases as $V^p(t)$ increases ($d(e^{f^*}(t) + e^{p^*}(t))/dV^p(t) > 0$), or alternatively, that unemployed individuals reduce their consumption of leisure at time t when $V^p(t)$ increases.

Let $\lambda(t)$ denote the (overall) re-employment hazard function.[15] Thus,

$$\lambda(t) = \lambda^f(t) + \lambda^p(t)$$

where $\lambda^f(t) = \lambda^f(e^{f^*}(t))$ and $\lambda^p(t) = \lambda^p(e^{p^*}(t))$ are the full-time and part-time re-employment hazards, respectively. Then $d\lambda^p(t)/dV^p(t) > 0$, $d\lambda^f(t)/dV^p(t) < 0$, and $d\lambda(t)/dV^p(t) > 0$. So, increasing the value of part-time re-employment at time t increases both the part-time and overall re-employment hazards at time t and decreases the full-time re-employment hazard at time t.

Next, consider the effect of an increase in the disregard on $V^p(t)$. It is clear that for UI recipients with $c^p < y$ or $c^p > b + y$, that $V^p(t)$ is unchanged.

[13] The proofs of the results reported in this section are contained in Appendix A.

[14] It is possible that when individuals receive UI benefits while working part-time they may wish to quit when those benefits start to run out. However, under the assumption that individuals always spend time seeking part-time work while unemployed, quitting is not optimal since $V^u(t) < V^{pu}(t)$.

[15] Let T be a random variable representing the duration of unemployment.Then the re-employment hazard, $\lambda(t)$, is defined as

$$\lambda(t) = \lim_{dt \to 0} \frac{P(t \leq T < t + dt | T \geq t)}{dt}.$$

See Kalbfleisch and Prentice (1980), Kiefer (1988), or Lancaster (1990) for more details.

Suppose then that $\max(\min(b+y, b+c^p), c^p) = b+y$. Also, assume for the moment that $u(c,k) = c + \Phi(k)$. Under these conditions, an increase in y will raise $V^p(t)$ for all $t < t^{pe}(dV^p(t)/dy > 0)$, and, hence, will increase the effort devoted to seeking part-time work $(de^{p^*}(t)/dy > 0)$ and the part-time re-employment hazard $(d\lambda^p(t)/dy > 0)$, will decrease the effort devoted to seeking full-time work $(de^{f^*}(t)/dy < 0)$ and the full-time re-employment hazard $(d\lambda^f(t)/dy < 0)$, and will increase the overall intensity of job search $(d(e^{f^*}(t) + e^{p^*}(t))/dy > 0)$ and the overall re-employment $(d\lambda(t)/dy > 0)$ hazard for $t < t^e$.

For the class of concave utility functions, $dV^p(t)/dy > 0$ when $\max(\min(b+y, b+c^p), c^p) = b+y$ so long as the probability of switching from part-time to full-time work before unemployment insurance benefits run out is sufficiently large.[16] Thus, under fairly general conditions, increasing the disregard increases both the part-time and overall re-employment hazards at least at the start of an unemployment spell. Moreover, if the optimal amount of effort devoted to seeking full-time work while working part-time, $e^{oj^*}(t)$, increases over time and $\partial^2 u/\partial c\,\partial k > 0$, then $\partial^2 V^p(t)/\partial y\,\partial t < 0$ for $t < t^{ep}$. This implies that $d^2\lambda(t)/dy\,dt < 0$, $\partial^2\lambda^p(t)/\partial y\,\partial t < 0$, and $\partial^2\lambda^f(t)/\partial y\,\partial t > 0$ for $t < t^e$ when $\max(\min(b+y, b+c^p), c^p) = b+y$. Under these conditions, increasing the disregard will have its biggest impact on the part-time re-employment hazard at the beginning of the unemployment spell.

To summarize, the theory predicts that, under fairly general circumstances, an increase in the disregard will increase both the part-time and overall re-employment hazards of UI recipients whose part-time earnings fall between y and $b+y$ and will decrease their full-time re-employment hazard. Moreover, the theory predicts that the magnitude of these effects should be greatest at the beginning of the unemployment spell. Unfortunately, the potential part-time earnings of an individual are not observed in the data. Therefore, only the weaker prediction, that increasing the disregard increases both the part-time and overall re-employment hazards of UI recipients and lowers their full-time re-employment hazard, will be tested in this paper.[17] The data used to test these empirical predictions are described next.

3. DATA

The data used to test the model's predictions are derived from the January 1986, 1988, 1990, and 1992 Current Population Survey's Displaced Workers

[16] In this model individuals cannot save. If savings at a zero interest rate is incorporated into the model, then this condition on the probability of switching from part-time to full-time work can be relaxed. If, on the other hand, the duration of UI benefits received on a part-time job is independent of $y(d^p = \min(0, t^e - t'))$, then $dV^p(t)/dy > 0$ for $\max(\min(b+y, b+c^p), c^p) = b+y$ and $t < t^e$ for all concave u.

[17] Recall that the theory abstracts from any potential wage or earnings uncertainty. For UI recipients who are willing to accept (at least some) part-time jobs paying between $(y, y+b)$, an increase in the disregard would make these jobs more attractive. Under suitable conditions, this implies that an increase in the disregard would increase the part-time re-employment hazard of UI recipients.

Supplements (DWS).[18] In the DWS it is not possible to determine whether a respondent was looking for work. Thus, the duration measure more accurately reflects joblessness than unemployment.

To test the predictions of the theoretical model derived above, information is needed about whether the first post-displacement job was part-time or full-time.[19] Further, since individuals may vary their hours worked over their job tenure, what is really required is information about the full-time/part-time (f/p) status at the start of the job.[20] In the DWS, however, measures of f/p status are available only for the job held at the time of the survey. Fortunately, it can be determined whether this job was the first post-displacement job. In the empirical work below, two methods are used to determine whether the first post-displacement job was part-time or full-time. The first method (method 1) designates the first post-displacement job as part-time if an individual was still working at that job at the time of the survey and if he/she reported working less than 35 hours per week in the week before the survey. The DWS also asks respondents who report working less than 35 hours a week whether their job was usually full-time. Thus, the second method (method 2) designates the first post-displacement job as part-time if the individual was still working at that job at the time of the survey, if they reported working less than 35 hours at that job in the week before the survey, and if they reported that this job was not usually full-time. Information from the 1986 Canadian Displaced Worker Survey (CDWS) suggests that both measures accurately reflect the part-time status at the start of the job, although the latter measure contains less measurement error.[21]

To summarize, respondents still working at their first post-displacement job at the time of the survey indicate that their joblessness spell was completed and that the f/p status of this job can be determined. Respondents who had left their first post-displacement job by the time of the survey indicate only that their joblessness spell was completed. The f/p status cannot be determined. Finally, respondents reporting that they did not hold any job since being displaced indicate that their joblessness spell was incomplete.

The econometric methods must then adjust for both right-censoring (see Kalbfleisch and Prentice (1980) and Kiefer (1988)) and an additional form of incomplete data: missing f/p status for some complete joblessness spells. To limit the extent of this latter type of incomplete data, only individuals displaced

[18] The DWS has been used extensively to analyze the costs of job displacement. See Farber (1993) for one recent example of such a study.

[19] In the United States, work of less than 35 hours a week is considered part-time.

[20] In the model developed in Section 2, on-the-job search for a full-time job could occur within a firm.

[21] The CDWS contains data on both the f/p status of the first job worked after being displaced and the number of hours worked in the week prior to the survey. For those still at their first job at the time of the survey, approximately 86 percent had no changed part-time/full-time status. Moreover, the CDWS asked respondents about their usual number of hours worked on the job. When f/p status is determined from this question it coincides with the f/p status at the start of the job 93 percent of the time. It should be noted, however, that in the CDWS work of less than 30 hours a week is considered part-time.

in the year prior to the survey are included in the sample. This also limits the amount of recall bias in the data. The precise manner in which the econometric methods adjust for missing f/p status will be described later.

Only individuals between the ages of 20 and 61 who were displaced from nonagricultural jobs due to plant closure, slack work, or abolished positions are included in the sample. In the DWS, individuals only report whether or not they received UI benefits after being displaced. The DWS does not ask questions pertaining to UI benefit eligibility or the amount of benefits that an individual was eligible to receive. In all states, UI benefit eligibility depends on both the earnings history of an individual and the reason for job loss. No individual in the sample would be disqualified for UI benefits on the basis of why they lost their job. Using the earnings information in the DWS, however, it is difficult to accurately determine whether a displaced worker satisfied the earnings requirements for UI benefit eligibility. To limit the number of ineligibles in the sample, only individuals who lost a full-time job are included in the sample.[22] Moreover, individuals with missing earnings data are excluded.

The weekly benefit amount an individual was eligible to receive is calculated using reported usual weekly earnings in the lost job and information on state benefit formulas contained in U.S. Department of Labor (various issues).[23] Individuals are included in the sample only if they resided in states with unemployment insurance laws that allowed an individual to earn up to some specific amount at a part-time job with no reduction in benefits, with a dollar for dollar reduction in benefits for earnings in excess of this level. Moreover, an individual was included in the sample only if they resided in a state where the disregard was either a fixed dollar amount or based on their weekly benefit amount. Individuals who resided in states where the disregard depended to some extent on their weekly wages in a part-time job are excluded.[24]

The thirty-seven states satisfying these criteria are listed in Table I. This table shows that there is considerable variation in the disregard across states. Approximately 74 percent of the total variation in the disregard is explained by this state variation.[25] The remaining 26 percent is explained by the fact that 7 states

[22] Unfortunately, the DWS only reported years of tenure in the lost job. Although the UI take-up rate of those reporting 0 years tenure in the lost job is significantly lower than those reporting one or more years tenure in the lost job (.36 versus .62), it is still substantial and suggests that a majority of those who report 0 years tenure were eligible for UI benefits. When the earnings requirements reported in Department of Labor (various issues) are converted into a minimum usual weekly earnings equivalent (assuming continuous employment for the year prior to displacement) only 11 people do not satisfy this minimum. The empirical results below do not substantially change when individuals reporting zero years of tenure in the lost job are excluded from the analysis.

[23] A similar imputation method has been used by Portugal and Addison (1990), for example.

[24] In 1988, Minnesota changed their unemployment insurance rule regarding part-time work from one in which the maximum amount of earnings disregarded was $25 to one in which the amount depended on the part-time wage. Therefore this state is excluded.

[25] By comparison, only 23 percent of the variation in weekly benefit amounts is explained by state variation.

TABLE I

STATES INCLUDED IN THE SAMPLE

Census State Code	State	Sample Average (1985 $) Weekly Benefit Amount	Disregard
11	Maine [b]	144.08	17.99
12	New Hampshire[a]	141.45	28.34
13	Vermont	136.33	14.29
14	Massachusetts	192.18	28.78
15	Rhode Island[a, b]	163.78	24.01
22	New Jersey[a]	204.04	40.91
23	Pennsylvania[a]	171.98	68.73
31	Ohio[a]	135.34	26.94
32	Indiana[a]	85.85	17.14
33	Illinois[a]	144.64	72.41
42	Iowa[a]	128.71	32.26
43	Missouri[b]	124.06	13.78
44	North Dakota[a, b]	126.23	65.44
47	Kansas[a, b]	164.48	26.12
51	Delaware[a]	147.81	44.56
52	Maryland	165.05	33.94
54	Virginia	145.08	24.17
55	West Virginia	156.93	23.99
56	North Carolina	136.75	29.80
57	South Carolina[a]	125.35	37.49
58	Georgia	131.32	7.69
59	Florida	140.35	4.80
62	Tennessee	102.30	29.15
63	Alabama	109.54	14.40
64	Mississippi[b]	104.94	4.81
71	Arkansas[a]	132.51	52.98
72	Louisiana[a]	142.72	46.51
73	Oklahoma	151.31	6.74
74	Texas[a]	155.17	38.87
82	Idaho[a]	133.21	66.75
83	Wyoming[a]	156.87	78.55
84	Colorado[a]	174.60	43.77
85	New Mexico[a]	156.79	25.34
86	Arizona[b]	121.35	19.44
87	Utah[a]	152.14	45.63
92	Oregon[a]	179.79	59.82
95	Hawaii	197.17	1.96

[a] Indicates states where the disregard depends on the weekly benefit amount to some extent.
[b] Indicates states that have changed their laws regarding the disregard over the sample period.

changed their disregard level during the sample period and that, in some states, the disregard depends on weekly benefit amounts.

The final sample size is 3343. Table II presents descriptive statistics of the variables used in this study. It is common practice to use the replacement rate (the weekly benefit amount divided by weekly earnings in the lost job) as a measure of benefit generosity (see Devine and Kiefer (1991)). In a similar manner, define the disregard rate as the disregard divided by weekly earnings in

TABLE II

DESCRIPTIVE STATISTICS[a]

Variable	Mean	Standard Deviation
Respondent Received UI benefits (*UI*)	.552	
Replacement Rate (*RR*)	.455	.114
Disregard Rate (*DR*)	.109	.073
Log of Weekly Earnings in Lost Job (1985 dollars)	5.693	.536
Tenure in Lost Job (years)	4.115	5.861
Lost Job Due to Slack Work	.489	
Lost Job Due to Abolished Position	.146	
Expected to Lose Job	.503	
State Unemployment Rate (Percent)	6.552	1.804
Lost Blue-collar Job	.604	
Household Head	.612	
Married	.586	
Children	.450	
Children Five and Under	.196	
Female	.348	
Less than 12 Years Schooling	.281	
More than 12 Years Schooling	.336	
Nonwhite	.139	
Resides in S.M.S.A.	.724	
Region of Residence		
Middle Atlantic	.109	
East North Central	.143	
West North Central	.064	
South Atlantic	.238	
East South Central	.053	
West South Central	.144	
Mountain	.013	
Pacific	.121	
Industry of Lost Job		
Mining	.029	
Construction	.148	
Transportation	.065	
Trade	.184	
Finance, Insurance and Real Estate	.052	
Services	.170	
Public Administration	.010	
Year of Job Loss		
1985	.268	
1987	.218	
1989	.200	
Sample Size	3343	

[a] The excluded categories are New England for region of the country, manufacturing for industry of lost job, plant closing for reason for displacement, 1991 for year of displacement, and twelve years of schooling for years of education.

the lost job. The average replacement rate for the sample is approximately .46 while the average disregard rate for the sample is .11.

Although the original joblessness durations were reported in weeks, the empirical analysis below further groups joblessness durations into two-week intervals. Such grouping serves two purposes. Firstly, it reduces the number of baseline hazard parameters that must be estimated in models with flexible baseline hazards (see Prentice and Gloeckler (1978) and Meyer (1986, 1990)). Secondly, it reduces the potential bias resulting from the heaping of reported joblessness durations at even weeks as is revealed by an examination of the weekly empirical hazard function.

The mean completed duration of joblessness was 9.3 weeks. Figures 1 and 2 present estimates of the overall, full-time, and part-time re-employment hazards for methods 1 and 2 of determining f/p status, respectively, using a discrete-time model which assumes that the differences in the log survivor function follow a fifth order polynomial.[26,27] The overall re-employment hazard is simply the sum of the full-time and part-time re-employment hazards.

In both figures, the full-time and the part-time re-employment hazard functions initially fall as the duration of joblessness lengthens, although in the latter case the decline is more gradual. The full-time hazard then rises moderately, falls, only to rise again quite dramatically towards the end of the sample period. Much of this latter behavior of the full-time re-employment hazard appears to be a result of the spikes in the empirical hazard at time periods 14 (weeks

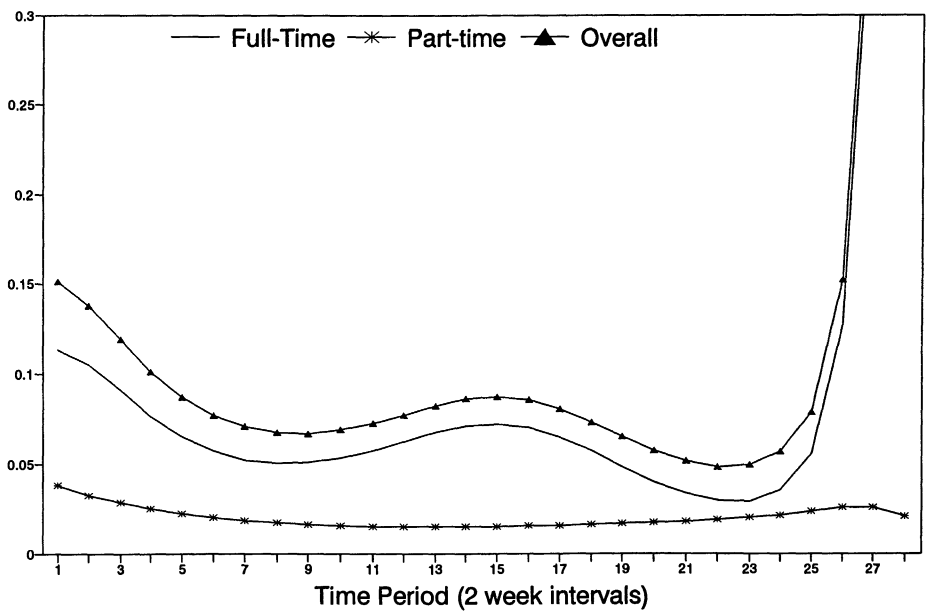

FIGURE 1—Conditional re-employment probabilities: f/p status determined by method 1.

[26] Since the duration data are discrete, hazards refer to conditional re-employment probabilities.

[27] Note that the discrete-time Kaplan-Meier estimates put no functional restrictions on the difference in the log survivor function between consecutive time periods. The estimates in Figures 1 and 2 also correct for right censored (incomplete) spells and for the fact that the f/p statuses of some completed spells are missing. The exact manner in which these missing data are handled is described in the next section.

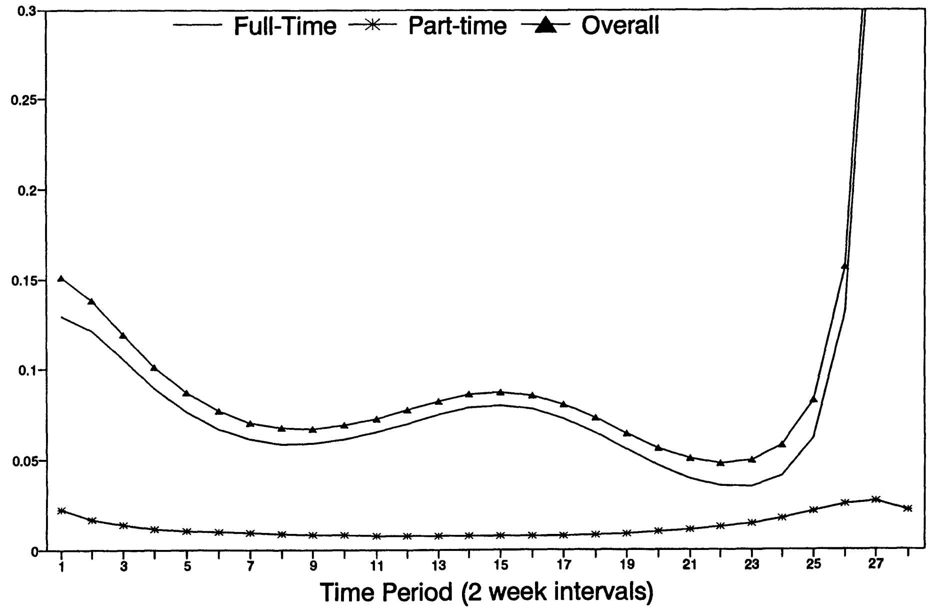

FIGURE 2—Conditional re-employment probabilities: f/p status determined by method 2.

26–27) and 27 (weeks 52–53). After falling, the part-time re-employment hazard then rises in both figures back to approximately its original level in period 1 with a slight decline at the end of the sample period. These smoothed hazards estimates, which correct for right censoring, imply a median joblessness duration of 10.1 weeks.

To get an indication of the effect of the disregard rate on a displaced worker's willingness to accept part-time employment, some simple cross-tabulations are calculated. Table III presents the fraction of individuals working part-time broken down by UI receipt and the level of the disregard rate (DR) among those still employed at their first post-displacement job at the time of the survey. The top panel presents cross-tabulations when f/p status is determined by method 1 while the bottom panel presents cross-tabulations when f/p status is determined by method 2. The fraction of UI recipients with disregard rates greater than .15 who are re-employed in part-time jobs is some 66% (281%) larger than the fraction of UI recipients with disregard rates less than .05 who are re-employed in part-time jobs when f/p status is determined by method 1 (method 2). Although the fraction of nonrecipients with disregard rates greater than .15 who are re-employed in part-time jobs is also larger than the corresponding fraction for nonrecipients with disregard rates less than .05, the increase is not quite as dramatic (38% for method 1 and 39% for method 2). Some caution, however, must be exercised when viewing these results since the distribution of other worker characteristics that affect the probability of part-time re-employment may differ across the different cells in Table III.

TABLE III

FRACTION WORKING PART-TIME AMONG THOSE RE-EMPLOYED BY DISREGARD RATE (DR)[a]

	$DR \leq .05$	$.05 < DR \leq .15$	$DR > .15$
	Part-time in Week Prior to Survey[b]		
Full Sample	.190	.252	.281
UI Recipients	.174	.223	.289
Non-recipients	.200	.277	.276
	Part-time in the Week Prior to Survey and Usually Part-time[b]		
Full Sample	.078	.138	.166
UI Recipients	.052	.114	.198
Non-recipients	.096	.159	.133

[a] Re-employed individuals are defined as all who have found and are still working at their first post-displacement job at the time of survey.

[b] In the first panel, the first post-displacement job is defined as part-time if an individual reports working less than 35 hours in the week prior to the survey. In the second panel, the first post-displacement job is defined as part-time if an individual reports working less than 35 hours in the week prior to the survey and reports that the job is not usually full-time.

4. EMPIRICAL RESULTS

To analyze the effect of the disregard on the re-employment behavior of displaced workers a competing risks model is estimated (see Kalbfleisch and Prentice (1980) or Lancaster (1990) for more complete discussions of competing risks models). Let T_f be the duration of joblessness until full-time re-employment and T_p be the duration of joblessness until part-time re-employment. Define $T = \min(T_f, T_p)$ and let W be an indicator variable which equals 1 if the job is full-time and equals zero if the job is part-time. Only (T, W) is observed. However, given sufficient variation in the vector of regressors x it is possible to uncover the latent joint survivor function, $S(k_f, k_p|x)$, from this "identified minimum" (see Heckman and Honoré (1989) and McCall (1993)).[28] Since the data on joblessness durations are discrete, a grouped data approach is taken (see Prentice and Gloeckler (1978), Meyer (1986, 1990), and Han and Hausman (1990)).

The estimates in Table IV assume that the full-time and part-time re-employment hazards are independent and that the regressors coefficients are time-constant. Moreover, it is assumed that the probability of being re-employed into a type w job in period k conditional on remaining jobless for more than $k-1$ periods, an indicator for UI receipt (UI), the replacement rate (RR), the disregard rate (DR), and a J-dimensional vector of other regressors x, satisfies

$$(6) \qquad P(T_w = k|UI, RR, DR, x, T > k-1) = 1 - \exp(-\exp(\alpha_{wk} + g_w(UI, RR, DR) + \beta_w' x)),$$

[28] The latent joint survivor function $S(k_f, k_p|x) = P(T_f > k_f, T_p > k_p|x)$.

TABLE IV

COMPETING RISKS ESTIMATES
INDEPENDENT RISKS AND TIME-CONSTANT COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Risk Variable	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
RR	0.1315	0.6037	−0.2946	0.9152
	(0.4883)	(1.0392)	(0.5422)	(1.4407)
DR	0.1292	−0.5488	0.3369	−2.3904
	(0.6336)	(1.1156)	(0.5581)	(1.8241)
UI	−1.4325***	−1.0480*	−1.3008***	−2.0688***
	(0.2148)	(0.5472)	(0.2138)	(0.7330)
RRUI	0.8461	−0.7878	0.6554	0.0387
	(0.5314)	(1.2464)	(0.5152)	(1.6726)
DRUI	0.2521	2.6864	−0.2446	6.8964***
	(0.9283)	(1.7091)	(0.8628)	(2.4943)
Log Likelihood	−6512.55	—	−6268.01	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included.

$w = f, p$. Here, α_{wk} are the baseline hazard parameters for risk w and β_w is a J-dimensional vector of parameters measuring the effect of the other regressors on the conditional probability of re-employment into a type w job, $w = f, p$.[29] Analogous to Kennan (1985) and Ham and Rea (1987), the estimates in Table IV further assume that the α_{wk} follow a fifth order polynomial:

$$\alpha_{wk} = \alpha_{w0} + \alpha_{w1}k + \alpha_{w2}k^2 + \alpha_{w3}k^3 + \alpha_{w4}k^4 + \alpha_{w5}k^5, \quad w = f, p. \tag{7}$$

The theory outlined in Section 2 predicted that an increase in the disregard should only increase (decrease) the part-time (full-time) re-employment hazard of UI recipients. Nevertheless, in this study the function g_{wk} is assumed satisfy:

$$g_w(UI, RR, DR,) = \gamma_{1w}UI + \gamma_{2w}RR + \gamma_{3w}DR + \gamma_{4w}RRUI + \gamma_{5w}DRUI \tag{8}$$

where *RRUI* and *DRUI* denote the interaction of the replacement rate with *UI* receipt and the interaction of the disregard rate with UI receipt, respectively,

[29] The other regressors include controls for reason for displacement, state unemployment rate in the year of the job loss, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence, as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household. Since the DWS only reports the year of the job loss, the monthly state unemployment rate could not be included as a (time-varying) regressor.

and γ_{iw} represents the effect of the ith regressor on risk w, $i = 1, \ldots, 5$, $w = f, p$. The effects of the replacement rate and disregard rate on the part-time and full-time re-employment hazards of nonrecipients (γ_{2w} and γ_{3w}) are not restricted to equal zero, a priori, because of the possibility of spillover effects (see Levine (1993)).[30,31] Alternatively, if spillover effects are absent, these coefficients may not equal zero if the model is either misspecified or if the imputation methods described above are seriously flawed.

Individuals with complete joblessness spells who had left their first post-displacement job by the time of the survey (i.e., those with complete spells but missing information about f/p status) are treated in the following manner. Let Q be an indicator variable which equals one for individuals working at their first post-displacement job at the time of the survey, and zero, otherwise. Let $A_w(k|UI, RR, DR, x)$ or simply $A_w(k)$ represent the probability that $T = k$ conditional on UI, RR, DR, and $x, w = f, p$. Finally, let $P(W = 1|T, UI, RR, DR, x, Q)$ or simply $P(W = 1|T, Q)$ be the probability that the first post-displacement job is full-time given T, UI, RR, DR, x, and Q.

It is assumed that Q is noninformative for W. Thus, $P(W = 1|T, Q) = P(W = 1|T)$. Under this assumption, $\log(A_f(k) + A_p(k))$ enters the log-likelihood function for individuals with complete spells but missing f/p status. Appendix B provides more details on the construction of the log-likelihood function for a model which also adjusts for UI selectivity bias (see below).

Turning now to the results in Table IV, an increase in the disregard is found to significantly increase the part-time re-employment hazard of UI recipients when f/p status is determined using method 2. The estimated effect of the disregard on the part-time re-employment hazard of UI recipients, while positive using method 1, is smaller in magnitude and not significantly different from zero. The estimated coefficient of DRUI in the full-time hazard, although negative when method 2 is used to determine f/p status, is not significantly different from zero for either method of determining f/p status. Restricting the coefficients associated with RR and DR (γ_{f2}, γ_{f3}, γ_{p2}, and γ_{p3}) to equal zero is not rejected by the data on the basis of a likelihood ratio (LR) test using either method of determining f/p status.[32] Thus, the replacement rate and disregard rate do not significantly affect the re-employment behavior of nonrecipients.

The competing risk estimates reported in Table V are based on a model with dependent risks. Here, the conditional probability of re-employment into a type w job is modeled by

$$(9) \qquad P(T_w = k|UI, RR, DR, x, \theta_w, T > k-1) \\ = 1 - \exp(-\exp(\alpha_{wk} - g_w(UI, RR, DR) + \beta_2' x)\theta_w),$$

[30] Levine (1993) used state average replacement rates instead of individual replacement rates to measure spillover effects. These two variables, however, have a correlation of .39 in our sample.

[31] Recall that individual replacement rates and (in some cases) disregard rates are based on the (imputed) amount of weekly benefit amounts an individual is eligible to receive.

[32] The LR test statistic (LR(4)) for method 1 equals 1.04 while LR(4) = 2.62 for method 2.

TABLE V

COMPETING RISKS ESTIMATES
INDEPENDENT RISKS AND TIME-CONSTANT COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Risk Variable	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
RR	0.7441	−0.4526	0.3295	0.6474
	(0.8027)	(1.1131)	(0.7673)	(1.4648)
DR	−1.338	−0.9395	−1.1582	−2.2183
	(0.9944)	(1.2471)	(0.9094)	(1.8465)
UI	−3.1991***	−1.8391***	−2.9223***	−2.4863***
	(0.3597)	(0.6265)	(0.3470)	(0.8149)
RRUI	1.8107**	−0.6859	1.4460*	−0.0005
	(0.8127)	(1.3343)	(0.7844)	(1.7553)
DRUI	1.6245	3.2773*	1.0687	6.8625***
	(1.3320)	(1.8414)	(1.2584)	(2.5388)
Log Likelihood	−6429.59	—	−6186.51	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included.

where again α_{wk} satisfies (7) and g_w satisfies (8), $w = f, p$. The univariate random variable θ_w which affects the conditional probability of re-employment into a type w job is assumed to be unobserved, $w = f, p$. The possibility of correlated risks arises in this model when the unobserved random variables θ_f and θ_p are correlated. It is assumed that these unobservables are jointly independent of RR, DR, UI, and the vector of other regressors x. The estimates in Table V model these unobservables by assuming that there are three different pairs $(\theta_{fj}, \theta_{pj})$ in the population which occur with relative frequency p_j, $j = 1, 2, 3$. These six location parameters and two free population proportions are estimated by the data. To ensure identifiability in these models, α_{w0} is fixed at 0, $w = f, p$.

The estimates in Table V are similar to those presented in Table IV. The only differences are that the estimated coefficient of the interaction term *DRUI* is now significantly different from zero at the 10% level when f/p status is determined by method 1 and the estimated coefficient of *RRUI* in the full-time risk increases in size and becomes significantly different from zero. As in the case with independent risks, one cannot reject the null hypothesis that the replacement and disregard rates have no effect on the re-employment behavior of nonrecipients (LR(4) = 3.10 and 3.28 when f/p status is determined by methods 1 and 2, respectively).

TABLE VI

COMPETING RISKS ESTIMATES

DEPENDENT RISKS, *UI* SELECTIVITY CORRECTIONS, AND TIME-CONSTANT COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job					
	Worked Part-time in Week Prior to Survey			Worked Part-time in Week Prior to Survey and Job is Usually Part-time		
Risk Variable	UI Receipt (1)	Full-time (2)	Part-time (3)	UI Receipt (4)	Full-time (5)	Part-time (6)
RR	1.1891	−0.9071	−0.7165	1.2151	−1.3917**	0.5781
	(0.7922)	(0.7416)	(1.0858)	(0.8055)	(0.7032)	(1.5072)
DR	0.8195	−1.0171	−0.4569	0.7802	−0.6355	−1.9752
	(0.8363)	(0.9757)	(1.2056)	(0.8403)	(0.8648)	(1.8524)
UI		−1.8973***	−2.6096***		−1.9654***	−3.3279***
		(0.3381)	(0.5838)		(0.3155)	(0.9092)
RRUI	—	1.2612	0.7632	—	1.3876**	0.7962
		(0.6919)	(1.2140)		(0.6572)	(1.8335)
DRUI	—	1.3229	3.1310*	—	.7633	6.8943***
		(1.1628)	(1.7476)		(1.0567)	(2.5097)
Log Likelihood	−8398.80	—	—	−8154.53	—	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included.

Not all individuals eligible for UI benefits actually apply to receive them.[33] Although the theoretical model outline above did not incorporate the take-up decision, it needs to be accounted for in the estimations since the unobservables determining UI receipt may be correlated with the unobservables θ_f and θ_p.[34] To correct for possible selectivity bias, the estimates in Table VI are derived from maximum likelihood estimation of a joint model of the determinants of UI receipt, the determinants of the joblessness duration following displacement, and the determinants of the f/p status of the first post-displacement job. In this model the full-time and part-time conditional re-employment probabilities are again modeled by equation (9). The conditional probability of UI receipt is modeled by

$$P(UI = 1|z, \theta_u) = 1 - \exp(-\exp(\beta_u' z)\theta_u) \tag{10}$$

where z is an I-dimensional vector of regressors, β_u is an I-dimensional vector of parameters measuring the effects of these regressors on UI receipt, and θ_u is

[33] Fifty-five percent of the sample reported receiving UI benefits. See Blank and Card (1991) for an analysis of the decline in insured unemployment in the 1980's.

[34] If the cost of filing a UI claim is denoted by s, then an eligible individual will file a UI claim if $s < V^u(0) - V^n$. From an econometrician's standpoint then, any observables determining $V^u(0)$ and/or V^n will impact both the probability of UI receipt and joblessness durations.

an unobserved random variable which is possibly correlated with θ_f and θ_p but which is assumed to be distributed independently of z. As was the case in Table V, the estimates in Table VI assume that individuals can be one of three unobserved types each with a distinct triplet of location parameters $(\theta_{uj}, \theta_{fj}, \theta_{pj})$ and occurring in the population with relative frequency p_j, $j = 1,2,3$.

A regressor that is contained in z but is not contained in $(RR, DR, x')'$ would help distinguish the true effect of UI receipt on the conditional probabilities of full-time and part-time re-employment from any spurious effect induced by unobserved heterogeneity. Unfortunately, in this case such an exclusion restriction is not apparent. An exclusion restriction, however, is not necessary to nonparametrically identify the joint distribution of $(\theta_u, \theta_f, \theta_p)$ and the effect of UI receipt on the conditional probabilities of full-time and part-time re-employment in the model described by equations (9) and (10). Under fairly general conditions, sufficient variation in (RR, DR, x') would suffice to identify the model. Intuitively, this follows from the fact that $P(UI=1|z)$ is a nonlinear function of z for any joint distribution of the unobservable (see McCall (1992, 1993)).

The estimates produced by the selectivity-corrected competing risks models are similar to those produced by the competing risks model with dependent risks. An increase in the disregard significantly increases the part-time re-employment hazard of UI recipients, although the significance is only at the 10% level when method 1 is used to determine f/p status in the lost job. Again, one cannot reject the null hypothesis that the replacement rate and the disregard rate have no effect on the re-employment behavior of nonrecipients ($LR(4) =$ 4.74 and 7.22 when f/p status is determined by methods 1 and 2, respectively).

Additional Robustness Checks

To further check the robustness of the empirical results, several additional models were estimated.[35] It is possible that whether or not an individual leaves their first post-displacement job by the time of the survey depends on the job's f/p status. When Q is uninformative for W,

$$(11)\qquad P(W=1|T=k, Q=1) = P(W=1|T=k) = \frac{A_f(k)}{A_f(k)+A_p(k)}.$$

To check whether the empirical findings for the disregard depend on this assumption, it was instead assumed that

$$(12)\qquad P(W=1|T=k, Q=1) = \frac{w_f A_f(k)}{w_f A_f(k) + w_f A_p(k)}$$

where w_f and w_p are weights. Under these circumstances, the term $\log(w_f A_f(k) + w_p A_p(k))$ enters the log-likelihood function for individuals with missing f/p

[35] To reduce the amount of computations all of the estimates reported in this section are produced by the competing risks model with dependent risks described in (9).

TABLE VII

COEFFICIENT ESTIMATES OF *DRUI* UNDER DIFFERENT ASSUMPTIONS ABOUT THOSE WHO HAVE COMPLETE JOBLESSNESS SPELLS BUT HAVE LEFT THE FIRST POST-DISPLACEMENT JOB BY THE TIME OF THE SURVEY[a, b]

Risk Weights Full-time/Part-time	Definition of Part-time Work in First Post-Displacement Job: Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
1. 1.0/0.0	1.8630	2.5919	1.6563	5.3966**
	(1.2153)	(1.9704)	(1.1756)	(2.7606)
2. 1.0/0.5	1.7987	2.9020	1.4632	5.9858**
	(1.2682)	(1.9387)	(1.2101)	(2.7042)
3. 1.0/1.0	1.6245	3.2773*	1.0687	6.8625***
	(1.3320)	(1.8414)	(1.2584)	(2.5388)
4. 0.5/1.0	1.3145	3.4306**	0.5854	6.9114***
	(1.4152)	(1.6284)	(1.3279)	(2.1730)
5. 0.0/1.0	1.0279	2.8106**	0.7492	3.5218***
	(1.4866)	(1.2527)	(1.4176)	(1.2947)
6. 0.93/1.0	1.5956	3.3210*	1.0189	6.9369***
	(1.3407)	(1.8235)	(1.2644)	(2.5135)

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] Estimates derived from a competing risks model with dependent and time-constant coefficients. The discrete baseline hazard parameters for the full-time and part-time risk are assumed to follow fifth order polynomials. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included.

status. Table VII presents competing risk estimates under alternative assumptions about w_f and w_p. To aid the comparison, the third row of Table VII replicates the estimates reported in Table V which assume that Q is uninformative for $W(w_f = w_p = 1)$.

The Canadian Displaced Worker Survey asks questions about the f/p status of both the first post-displacement job and the job held at the time of the survey.[36] Thus, it is possible to get an estimate of w_f (w_p is fixed at 1) using this data. The sixth row of Table VII reports the estimated coefficients for *DRUI* applying this estimated weight.[37]

As Table VII shows, the qualitative conclusions reached from Table V do not depend on the assumption $w_f = w_p = 1$. When method 2 is used to determine f/p status, the disregard has a positive and significant effect on the part-time

[36] Note that in the CDWS, work of less than 30 hours a week is considered part-time.

[37] A consistent estimate of w_f is found by minimizing

$$\sum_{i=1}^{n}\left[I(W_i=1)-\frac{w_f\hat{A}_f(K_i)}{w_f\hat{A}_f(K_i)+\hat{A}_p(K_i)}\right]^2$$

where the sum is taken over all individuals in the CDWS who lost their job in 1985 and left the first post-displacement job by the time of the survey and $\hat{A}_w$ is based on a competing risks model with independent risks and exponential baseline hazards, $w = f, p$.

TABLE VIII

COEFFICIENT ESTIMATES OF *DRUI* FOR COMPETING RISKS MODELS WITH DEPENDENT RISKS AND TIME-CONSTANT COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Specifications	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
1. Full Sample	1.7541	3.0051*	1.0839	6.9238***
Flexible Baseline Hazards	(1.3462)	(1.8059)	(1.2415)	(2.6334)
2. Full Sample	2.0051	3.7197*	1.4325	6.2649**
State Fixed Effects[c]	(1.3627)	(1.9829)	(1.3106)	(2.7416)
3. Full Sample[d]	0.7442	3.0966*	0.4776	5.7755**
Time-Varying UI Variables	(0.8618)	(1.7138)	(0.9058)	(2.3488)
4. Blue-Collar Sample[e]	3.4064**	2.1082	0.9994	6.7136*
	(1.6073)	(2.5672)	(1.6980)	(3.9185)
5. White-Collar Sample[f]	−0.3143	5.4593*	−0.3187	6.6100
	(2.0318)	(3.2405)	(1.9866)	(4.1765)
6. Female Sample[g]	−1.8972	8.5715***	1.4462	4.3018
	(2.1044)	(3.1981)	(1.9459)	(3.9719)
7. Male Sample	1.5614	3.9601	−0.0504	12.8923***
	(1.6371)	(2.8084)	(1.5982)	(4.6570)
8. UI Recipient Sample[h]	−0.2342	1.4443	−0.6762	4.9266**
	(0.7826)	(1.3502)	(0.7139)	(1.9791)

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] Unless stated otherwise, the discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials and controls for *UI*, *RR*, *RRUI*, *DR*, reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household as included.

[c] Individuals residing in Hawaii were excluded from the estimates reported in columns (1) and (2) while Hawaii and Maine were excluded from the estimates reported in columns (3) and (4).

[d] Only time-varying controls for *UI*, *RRUI*, and *DRUI* were included.

[e] The estimates in columns (3) and (4) combine the Service, Financial-Insurance-Real Estate, and Public Administration industries.

[f] The estimates in columns (3) and (4) limit the region effects to: South, West, North-Central, and Northeast.

[g] The estimates combine Mining and Manufacturing industries and Service and Public Administration Industries.

[h] In this row, coefficient estimates are for the *DR* variable. Estimates assume independent risks and limit region effects to: South, West, North-Central, and Northeast.

re-employment hazard of UI recipients for all combinations of w_f and w_p. When method 1 is used to determine f/p status, the disregard has a positive and (weakly) significant effect on the part-time re-employment hazard of UI recipients for all combinations with $w_p = 1$. For combinations with $w_p < 1$, the point estimates for *DRUI* remain large and positive in the part-time re-employment hazard. The standard errors, however, increase so that the estimates are not significantly different from zero at conventional levels.

It is possible that misspecification of the baseline hazard function would bias the estimates. Flexible competing risks models estimate separate baseline hazard parameters for each time period and thereby place no functional restrictions on the baseline hazard parameters α_{wk} (see Han and Hausman (1990)). The first row of Table VIII presents coefficient estimates for *DRUI* from a flexible competing risks model with dependent risks. These estimates are similar to those reported for *DRUI* in Table V.

Only regional fixed-effects are controlled for in the estimates reported in Table IV through VI. As the second row of Table VIII shows, however, controlling for state fixed-effects yields similar conclusions.[38]

The estimates in Tables IV through VI do not make any attempt to incorporate differences in benefit duration. In the theoretical model presented above, however, the UI variables affect the re-employment behavior of individuals only while they are receiving benefits. To account for benefit duration, time-varying UI variables were constructed that "turn on" only while an individual is receiving benefits. Unfortunately, the DWS does not provide detailed information on the duration of benefit receipt. Thus, the maximum benefit duration was set at the maximum duration of regular benefits for the state of residence. This was 26 weeks for all states in the sample except Massachusetts which was 30 weeks. Data obtained from the Advisory Council on Unemployment Compensation, which indicate the dates when federal or state benefit extension programs were in effect, show that the fraction of individuals in the sample that actually received any extended benefits was probably very small.[39,40] Moreover, it is unclear whether any individuals were aware of the possibility of receiving extended benefits when their joblessness spell began. For these reasons, no attempt is made to adjust benefit durations for extended benefits.

The third row of Table VIII presents estimates from competing risks models that instead include three time-varying regressors which equal *UI*, *RRUI*, and *DRUI* when an individual receives benefits and zero, otherwise. Again, the conclusions do not meaningfully change.[41]

Rows four through seven of Table VIII present separate competing risks estimates for blue-collar workers, white-collar workers, men, and women, respectively. As the table shows, the point estimates of the *DRUI* coefficient in the part-time risk are positive for all groups using either method of determining f/p

[38] Individuals residing in Hawaii (Maine and Hawaii) were dropped from the state fixed-effect estimations when f/p status was determined by method 1 (method 2) due to the lack of individuals re-employed into part-time jobs in this (those) state(s).

[39] I am grateful to Daniel McMurrer for providing me this data.

[40] Extended benefits from the Emergency Unemployment Compensation Act became available on November 17, 1991. Although some of the individuals in the sample who were displaced in 1991 may have received extended benefit under this program, it is difficult to determine who actually did since only the year of displacement is reported in the DWS. Prior to 1991, Idaho had supplemental programs operating during March through June of 1985 and 1987, and Louisiana had a supplemental program operating from January until March of 1987. Since individuals in the sample lost their job at some point during 1985 or 1987, however, they would not have exhausted their regular benefits during this period. Maine, Vermont, Massachusetts, Rhode Island, West Virginia, and Oregon had supplemental programs operating at some point during 1991. With the exception of Rhode Island, only a small fraction of individuals in our sample who resided in these states could have possibly received any extended benefits since these supplemental programs operated mainly during the first six months of the year.

[41] The coefficients of *RR* and *DR* are restricted to zero in these estimations.

status. Moreover, the coefficient estimates for *DRUI* in the part-time risk are larger when method 2 is used to determine f/p status except for women.[42]

Focusing on the estimates in columns (3) and (4), the coefficient estimate for *DRUI* in the part-time risk is similar for the white-collar and blue-collar samples while substantially larger for men than for women. The coefficient estimates for *DRUI* in the full-time risk, although negative for white-collar workers and men, are imprecise.

Finally, it is possible that the variables *UI*, *RRUI*, and *DRUI* do not adequately account for the differences between UI recipients and nonrecipients. The last row of Table VIII, presents competing risks estimates based on the sample of UI recipients. The coefficient associated with the disregard rate in the part-time re-employment hazard continues to be positive and significant when method 2 is used to determine f/p status.[43] While still positive, the estimated coefficient of the disregard rate in the part-time re-employment hazard is insignificantly different from zero when method 1 is used to determine f/p status.

Time-Varying Coefficients

The estimates above restrict the coefficients of the *UI* regressors to be time-constant. The theory presented in Section 2, however, implies that the effect of the disregard on the conditional probability of re-employment into a part-time job has its largest effect at the beginning of the unemployment spell. The estimates in Tables IX through XI differ from those in IV and VI in that they allow the coefficients of all the *UI* variables to differ between periods 1, 2 through 6, 7 through 11, and 12 through 27.[44,45]

Comparing the log-likelihoods reported at the bottom of Tables IV through VI with those reported at the bottom of IX through XI, the null hypothesis of time-constant coefficients is rejected at the 1 percent significance level for all model specifications. The null hypothesis that the disregard rate and replacement rate have no effect on the re-employment behavior of nonrecipients is not rejected at the 10 percent significance level for all model specifications except that which assumes risks are independent and uses method 2 to determine f/p

[42] When the coefficients of *DR* and *RR* are restricted to zero the coefficient estimate for *DRUI* in the part-time risk is independent of which method is used to determine f/p status (4.32 for method 1 versus 4.30 for method 2) for the women sample. Large differences in the coefficient estimate for *DRUI* in the part-time risk, however, persist for the other groups. In no case are these coefficient restrictions rejected by the data.

[43] Only competing risks models with independent risks (one unobserved type) were supported by the UI recipient data.

[44] Recall that the duration data are grouped into two week periods; thus the coefficients are allowed to vary between weeks 0–1, 2–11, 12–21, and 22–53.

[45] Competing risks models which instead allow the coefficients of the *UI* variables to vary over time in a quadratic fashion yield similar conclusions.

status. In this case, the null hypothesis is rejected at the 5% significance level ($LR(16) = 26.50$).[46]

As can be seen from Tables IX through XI, the effect of the disregard on the conditional probability of part-time re-employment of UI recipients is largest during weeks 2 through 11 (periods 2–6) of their joblessness spell. In fact, the point estimates for weeks 12 and 23 (periods 7–11) are negative for all model specifications. The coefficient estimates for *DRUI* in the part-time re-employment hazard for weeks 24 through 53 (periods 12–27), while positive, are imprecise. On the other hand, the coefficient estimates for *DRUI* in the part-time re-employment hazard are positive for weeks 0 through 1 (period 1), although they are less than half as large as the coefficient estimates for weeks 2 through 11 (periods 2–6) and are not significantly different from zero. Nevertheless, restricting the coefficients for *DRUI* in both the part-time and full-time re-employment hazard for weeks 0 through 1 to equal the coefficients for weeks 2 through 11 and restricting the coefficients for weeks 12 through 53 to equal zero is not rejected by the data.[47]

The coefficients for *DRUI* in the full-time hazard continue to be imprecisely estimated. Thus the empirical results suggest that the disregard effects only the part-time re-employment hazard and primarily during the first eleven weeks of joblessness.

Elasticity Estimates

Tables XII and XIII present point estimates of the elasticity of expected duration and the elasticity of the probability of part-time re-employment with respect to the disregard. The elasticities reported in columns (1) and (2) are based on estimates which use method 1 to determine f/p status while the estimates reported in columns (3) and (4) are based on estimates which use method 2 to determine f/p status. The elasticities reported in Table XII use competing risks estimates which restrict the coefficients of the *UI* variables to be time-constant while the elasticities reported in Table XIII use competing risks estimates which allow the coefficients of the *UI* variables to vary over time. For all rows except row 4(b), the relevant policy experiment is an increase in the

[46] In estimates with independent risks and using method 1 to determine f/p status, $LR(16) = 22.42$. For models with dependent risks (dependent risks and *UI* selectivity corrections), $LR(16) = 14.80$ and 17.76 (16.92 and 22.32) when method 1 and method 2 are used to determine f/p status, respectively.

[47] In competing risks models (with dependent risks) that already restrict the effect of the disregard rate and replacement rate on both the conditional probability of part-time and full-time re-employment of nonrecipients to zero, further restricting the coefficients of *DRUI* in both the part-time and full-time re-employment hazards to be the same in weeks 0 through 11 and zero thereafter, is not rejected by the data on the basis of likelihood ratio tests ($LR(5) = 1.84$ and 4.00 when methods 1 and 2 are used to determine f/p status, respectively).

TABLE IX

COEFFICIENT ESTIMATES OF *DRUI* FOR COMPETING RISKS MODEL WITH INDEPENDENT RISKS AND TIME-VARYING COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Risk Time Period	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
1	−0.0889	2.1514	0.0592	2.6088
	(3.1716)	(3.9577)	(2.8255)	(4.9672)
2–6	−0.3973	5.1414**	−1.1629	14.1705***
	(1.3028)	(2.3554)	(1.1679)	(3.7659)
7–11	1.0551	−3.0838	1.3063	−5.3685
	(2.7785)	(4.5777)	(2.5327)	(7.6343)
12–27	−1.0597	6.0404	0.4698	0.1492
	(3.2523)	(6.5789)	(3.1317)	(8.9231)
Log Likelihood	−6398.28	—	−6156.09	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. The coefficients of the variables *UI*, *RR*, *DR*, and *RRUI* are allowed to differ between periods 1, 2–6, 7–11, and 12–27. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expecting to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included. The effects of these latter controls are restricted to be time-constant.

disregard of a UI recipient. Any effect on the probability of UI recipiency is ignored.

Estimates are derived by calculating the change in the probability of part-time re-employment (expected duration) for every UI recipient for a 10 percent increase in the disregard. The resultant changes are then averaged across all UI recipients. This is equivalent to the change in the average probability of part-time re-employment (expected duration) for UI recipients for a 10 percent increase in the disregard. Finally, the change in the average is divided by the average probability of part-time re-employment (expected duration) to obtain a percentage change.

The elasticities reported in the first four rows of Tables XII and XIII use competing risks model estimates that are based on the full sample. The elasticities reported in rows 5 through 8 of Tables XII and XIII use competing risks model estimates that are based on particular groups of individuals (for example, women). Moreover, in these rows the average effect of a 10 percent increase in the disregard is calculated using only the UI recipients of the particular group.

In row 4(a) of Tables XII and XIII, the possibility that the distribution of the unobservables (θ_f, θ_p) for UI recipients differs from the distribution of unobservables for the population of displaced workers is accounted for by using the

TABLE X

COEFFICIENT ESTIMATES OF *DRUI* FOR COMPETING RISKS MODEL WITH DEPENDENT RISKS AND TIME-VARYING COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Risk Time Period	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
1	0.3907	2.0867	0.3512	2.7928
	(3.2890)	(4.1239)	(2.9645)	(5.0287)
2–6	1.0271	5.7104**	0.1506	14.4451***
	(1.5875)	(2.7009)	(1.5368)	(4.0263)
7–11	2.1960	−2.7893	2.2973	−5.1842
	(3.2192)	(4.8049)	(3.0368)	(7.7810)
12–27	0.7704	7.1537	1.4441	1.0212
	(3.9148)	(6.6579)	(3.775)	(8.8586)
Log Likelihood	−6383.38	—	−6141.88	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. The coefficients of the variables *UI*, *RR*, *DR*, and *RRUI* are allowed to differ between periods 1, 2–6, 7–11, and 12–27. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expecting to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included. The effects of these latter controls are restricted to be time-constant.

selectivity-corrected estimates reported in Tables VI and XI.[48] The point estimates reported in parentheses of Tables XII and XIII are based on competing risks models which restrict the coefficients associated with the *RR* and *DR* variables to zero.[49]

The point estimates of the elasticity of expected duration with respect to the disregard, reported in rows 1 through 4(a) of columns (1) and (3) of Tables XII and XIII, are uniformly negative and range from −0.03 to −0.11.[50] For the elasticities reported in column (3) of Table XIII, this translates into an estimated decrease in expected joblessness durations of between 2.5 and 6.2 days for a 50 percent increase in the disregard, with the largest point estimate being produced by a competing risks model which corrects for selectivity bias and which restricts the coefficients for the *RR* and *DR* variables equal to zero. A 50 percent increase in the disregard amounts to an average increase of $16 per week (1985 dollars).

[48] The distribution of the three pairs of unobservables (θ_f, θ_p) conditional on $UI = 1$ and x is found by a simple application of Bayes' Theorem.

[49] As noted above these restrictions are not rejected by the data.

[50] Using a discrete-time hazard model that assumes that the baseline hazard parameters follow a fifth order polynominal as in (7) and models unobserved heterogeneity using a three mass-point mixing distribution, *DRUI* is found to have a significantly positive effect (one-tailed test) on the overall re-employment hazard.

TABLE XI

COEFFICIENT ESTIMATES OF *DRUI* FOR COMPETING RISKS MODEL WITH DEPENDENT RISKS, *UI* SELECTIVITY CORRECTIONS AND TIME-VARYING COEFFICIENTS[a, b]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Risk Time Period	Full-time (1)	Part-time (2)	Full-time (3)	Part-time (4)
1	0.6292	1.7647	0.4327	2.2620
	(3.3468)	(3.9703)	(2.9664)	(4.9501)
2–6	0.8008	5.3939**	−0.3757	14.7511***
	(1.5071)	(2.4064)	(1.3443)	(3.8816)
7–11	1.1313	−3.0985	1.4475	−5.0353
	(2.8125)	(4.7423)	(2.5932)	(7.9360)
12–27	−0.5665	5.6334	0.9165	−0.4944
	(3.3309)	(6.8246)	(3.2253)	(9.2437)
Log Likelihood	−8361.49	—	−8118.73	—

[a] Standard errors are in parentheses. One, two, or three asterisks indicate significance at the 10%, 5%, or 1% significance level, respectively.

[b] The discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials. The coefficients of the variables *UI*, *RR*, *DR*, and *RRUI* are allowed to differ between periods 1, 2–6, 7–11, and 12–27. Controls for reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expecting to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are also included. The effects of these latter controls are restricted to be time-constant.

Using competing risks model estimates based on the full sample, the calculated elasticity of the probability of part-time re-employment with respect to the disregard is positive for UI recipients in all but one case (always positive) when the coefficients associated with *DR* and *RR* variables are unrestricted (restricted to equal zero). The elasticities estimates range from −0.04 to 0.23 (0.39 to 0.57) when method 1 (method 2) is used to determine f/p status. For the elasticities reported in rows 1 through 4(a) of column (4) of Table XIII, this translates into an increase in the probability of part-time re-employment of between 0.03 and 0.04 for a 50 percent increase in the disregard. Again, the largest point estimate is produced by a competing risks model which corrects for selectivity bias and restricts the coefficients for the *RR* and *DR* variables to equal zero.

The elasticities reported in rows 5 and 6 of Tables XII and XIII are derived from competing risks model estimates using the blue-collar and white-collar samples, respectively. The point estimates of the elasticity of the probability of part-time re-employment with respect to the disregard are substantially larger for white-collar UI recipients than for blue-collar UI recipients for all specifications. The elasticity estimates reported in rows 7 and 8 of Tables XII and XIII, on the other hand, are based on separate competing risks model estimates for men and women, respectively. The point estimates of the elasticity of the probability of part-time re-employment with respect to the disregard are similar

TABLE XII

EXPECTED DURATION AND PROBABILITY OF PART-TIME RE-EMPLOYMENT ELASTICITY ESTIMATES FOR COMPETING RISKS MODELS WITH TIME-CONSTANT COEFFICIENTS[a]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Specifications	Expected Duration (1)	Probability of Part-time Re-employment (2)	Expected Duration (3)	Probability of Part-time Re-employment (4)
1. Independent Risks	−0.0707	0.1733	−0.0702	0.4898
	(−0.0730)	(0.1913)	(−0.0765)	(0.5297)
2. Dependent Risks	−0.0403	0.1922	−0.0428	0.5229
	(−0.0493)	(0.2020)	(−0.0527)	(0.5546)
3. Flexible Baseline Hazards	−0.0389	0.1713	−0.0378	0.5070
	(−0.0844)	(0.1745)	(−0.0477)	(0.5321)
4. UI Selectivity-Corrected Estimates				
(a) Conditional on UI Receipt	−0.0597	0.2228	−0.8800	0.4765
	(−0.1007)	(0.1932)	(−0.1064)	(0.4945)
(b) Unconditional on UI Receipt	0.0146	0.0972	−0.0124	0.1723
	(−0.0396)	(0.1013)	(−0.0371)	(0.2588)
5. Blue-Collar Sample[b]	−0.0962	−0.0247	−0.0866	0.2538
6. White-Collar Sample[c]	−0.0100	0.4836	−0.0104	0.7883
7. Female Sample[d]	0.0057	0.1553	−0.0340	0.2916
8. Male Sample	−0.0423	0.1468	−0.0406	1.0306
9. UI Recipient Sample[e]	−0.0145	0.1427	−0.0245	0.5654

[a] Unless stated otherwise, the discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials and controls for *UI*, *RR*, *RRUI*, *DR*, reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displaced, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are included. The estimates in parentheses restrict the coefficients of *RR* and *DR* to zero.

[b] The estimates in columns (3) and (4) combine the Service, Financial-Insurance-Real Estate, and Public Administration industries.

[c] The estimates in columns (3) and (4) limit the region effects to: South, West, North-Central, and Northeast.

[d] The estimates combine Mining and Manufacturing industries and Service and Public Administration industries.

[e] Estimates assume independent risks and limit region effects to: South, West, North-Central, and Northeast.

for male and female UI recipients when method 1 is used to determine f/p status. When method 2 is used to determine f/p status, however, the point estimates for male UI recipients are considerably larger than those for female UI recipients.

The point estimates in row 9 of Tables XII and XIII are calculated from competing risks estimates based on the UI recipient sample. These elasticity estimates are similar to those derived from competing risks model estimates based on the full sample.

Row 4(b) of Tables XII and XIII reports elasticity estimates for the policy experiment of increasing the disregard for a displaced worker by 10 percent. These elasticities, which are based on the UI-selectivity-corrected model esti-

TABLE XIII

EXPECTED DURATION AND PROBABILITY OF PART-TIME RE-EMPLOYMENT ELASTICITY ESTIMATES FOR COMPETING RISKS MODELS WITH TIME-VARYING COEFFICIENTS[a]

	Definition of Part-time Work in First Post-Displacement Job			
	Worked Part-time in Week Prior to Survey		Worked Part-time in Week Prior to Survey and Job is Usually Part-time	
Specifications	Expected Duration (1)	Probability of Part-time Re-employment (2)	Expected Duration (3)	Probability of Part-time Re-employment (4)
1. Independent Risks	−0.0519	0.2131	−0.0495	0.5161
	(−0.0586)	(0.2266)	(−0.0581)	(0.5695)
2. Dependent Risks	−0.0330	0.1672	−0.0324	0.3872
	(−0.0.426)	(0.1748)	(−0.0434)	(0.4076)
3. Flexible Baseline Hazards	−0.0455	0.1916	−0.0528	0.4605
	(−0.0616)	(0.1817)	(−0.0611)	(0.5004)
4. UI Selectivity-Corrected Estimates				
(a) Conditional on UI Receipt	−0.0339	−0.0428	−0.0665	0.4034
	(−0.0757)	(0.1301)	(−0.0891)	(0.4051)
(b) Unconditional on UI Receipt	−0.0042	−0.0522	−0.0084	0.0416
	(−0.0285)	(0.0316)	(−0.0324)	(0.1363)
5. Blue-Collar Sample[b]	−0.0734	−0.0098	−0.0783	−0.1446
6. White-Collar Sample[c]	−0.0078	0.4114	−0.0379	0.6499
7. Female Sample[d]	−0.0342	0.1413	−0.0482	0.1849
8. Male Sample	−0.0556	0.1468	−0.0457	1.0678
9. UI Recipient Sample[e]	−0.0024	0.0157	−0.0242	0.3091

[a] Unless stated otherwise, the discrete baseline hazard parameters for the full-time and part-time risks are assumed to follow fifth order polynomials and controls for *UI*, *RR*, *RRUI*, *DR*, reason for displacement, state unemployment rate, tenure in the lost job, industry of the lost job, log of weekly earnings in the lost job, year of displacement, region of residence as well as dichotomous variables indicating expected to be displace, nonwhite, married, female, children, young children, less than 12 years education, more than 12 years education, smsa, and head of household are included. Moreover, the coefficients of *UI*, *RR*, *DR*, *RRUI*, and *DRUI* are allowed to vary between periods 1, 2–6, 7–11 and 12–27. The estimates in parentheses restrict the coefficients of *RR* and *DR* to zero.

[b] The estimates in columns (3) and (4) combine the Service, Financial-Insurance-Real Estate, and Public Administration industries.

[c] The estimates in columns (3) and (4) limit the region effects to: South, West, North-Central, and Northeast.

[d] The estimates combine Mining and Manufacturing industries and Service and Public Administration industries.

[e] Estimates assume independent risks and limit region effects to: South, West, North-Central, and Northeast.

mates reported in Tables VI and XI, allow for the increase in the disregard to endogenously affect the probability of UI receipt. Estimates are derived by calculating the change in the probability of part-time re-employment (expected duration) for a 10 percent increase in the disregard for every individual in the sample and then averaging. Since an increase in the disregard has only a small effect on the probability of UI receipt and a statistically insignificant effect on the re-employment behavior of nonrecipients, it is not surprising that the magnitudes of these elasticities are substantially less than those calculated for UI recipients.[51]

[51] Increasing the disregard by 10 percent increases the probability of UI receipt for the average displaced worker by substantially less than 0.01.

5. CONCLUSIONS

This paper analyzes the incentive effects of unemployment insurance provisions relating to part-time work. A continuous-time job search model in which individuals seek both part-time and full-time work predicts that increasing the disregard will increase the part-time and overall re-employment hazards and decrease the full-time re-employment hazards of UI recipients. In addition, the model predicts that these effects are largest at the beginning of the unemployment spell.

These predictions are tested using data from the Current Population Survey's Displaced Workers Supplements. Estimates from several different competing risks model specifications show that an increase in the disregard significantly increases the conditional probability of part-time re-employment during approximately the first three months of joblessness. For UI recipients, a 50 percent increase in the disregard is found to increase the probability of part-time re-employment from between 0.03 and 0.04 and decrease expected joblessness durations from between 2.5 and 6.2 days.

Since UI receipt is not significantly affected by an increase in the disregard, it may be tempting to conclude from this latter finding that an increase in the disregard will lower a state's benefit payments. Such a conclusion is premature for several reasons. First, benefit payments to part-time workers increase when the disregard is increased. In addition, individuals may remain in these part-time jobs longer when the disregard is increased. Unfortunately, the DWS lacks information on how long individuals remain in their first post-displacement job. Second, the empirical results are based on a sample of full-time displaced workers who were displaced either because of a plant closing, slack work, or abolished position; the incentive effects of changes in the disregard on other types of unemployed workers eligible for UI benefits may differ. Third, any employer's response (e.g. changes in layoff policy) to a change in the disregard has been ignored. Future research needs to address these issues before any definitive conclusions are reached or policy recommendations are made.

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APPENDIX A

This appendix derives some of the theoretical results reported in Section 2. Equation (2) can be derived using delta arguments. The value function at time $t-dt$ associated with working a part-time job satisfies

$$\text{(A1)} \qquad V^{pu}(t-dt)=\frac{U(\max(\min(b+y,c^{p}+b),c^{p}),1-h^{p}-e^{oj*}(t-df))\,dt}{1+r\,dt}$$
$$+\frac{\lambda^{oj}(e^{oj*}(t-dt))\,dt\,V^{f}}{1+r\,dt}+\frac{(1-\lambda^{oj}(e^{oj*}(t-dt))\,dt)V^{pu}(t)}{1+r\,dt}+o_{P}(t).$$

Rearranging (A1) and taking limits as $dt \to 0$ gives (2). The derivation of equations (3), (4), and (5) are similar.

Next it is shown that $de^{f*}(t)/dV^p(t) < 0$, $de^{p*}(t)/dV^p(t) > 0$, and $d(e^{f*}(t) + e^{p*}(t))/dV^p(t) < 0$. The maximization problem on the right-hand side of equation (5) in the text determines the optimal values $e^{f*}(t)$ and $e^{p*}(t)$. These optimal values satisfy the Kuhn-Tucker conditions

$$\text{(A2)} \qquad -U_k(b, 1 - e^{f*}(t) - e^{p*}(t)) + \lambda^{f\prime}(e^{f*}(t))(V^f - V^u(t)) \le 0,$$

$$\text{(A3)} \qquad (-U_k(b, 1 - e^{f*}(t) - e^{p*}(t)) + \lambda^{f\prime}(e^{f*}(t))(V^f - V^u(t)))e^{f*}(t) = 0,$$

$$\text{(A4)} \qquad -U_k(b, 1 - e^{f*}(t) - e^{p*}(t)) + \lambda^{p\prime}(e^{p*}(t))(V^p(t) - V^u(t)) \le 0,$$

and

$$\text{(A5)} \qquad (-U_k(b, 1 - e^{f*}(t) - e^{p*}(t)) + \lambda^{p\prime}(e^{p*}(t))(V^p(t) - V^u(t)))e^{p*}(t) = 0.$$

When $e^{f*}(t) > 0$ and $e^{p*}(t) > 0$, how $e^{f*}(t)$ and $e^{p*}(t)$ change with $V^p(t)$ can be analyzed by implicit differentiation of (A2) and (A4) (which become equalities). This leads to the following system of equations:

$$\text{(A6)} \qquad \begin{bmatrix} u_{kk} + \lambda^{f\prime\prime}(e^{f*}(t))(V^f - V^u(t)) & u_{kk} \\ u_{kk} & u_{kk} + \lambda^{p\prime\prime}(e^{p*}(t))(V^p(t) - V^u(t)) \end{bmatrix} \begin{bmatrix} de^{f*}(t)/dV^p(f) \\ de^{p*}(t)/dV^p(t) \end{bmatrix} = \begin{bmatrix} 0 \\ -\lambda^{p\prime}(e^{p*}(t) \end{bmatrix}$$

where u_{kk} is evaluated at $(b, 1 - e^{f*}(t) - e^{p*}(t))$. Solving this system for $de^{f*}(t)/dV^p(t)$ and $de^{p*}(t)/dV^p(t)$ yields

$$\text{(A7)} \qquad \frac{de^{f*}(t)}{dV^p(t)} = \frac{u_{kk}\lambda^{p\prime}(e^{p*}(t))}{D}$$

and

$$\text{(A8)} \qquad \frac{de^{p*}(t)}{dV^p(t)} = \frac{-(u_{kk} + \lambda^{f\prime\prime}(e^{f*}(t))(V^f - V^u(t)))\lambda^{p\prime}(e^{p*}(t))}{D},$$

where D is the determinant of the 2×2 matrix in (A6):

$$\text{(A9)} \qquad D = u_{kk}(\lambda^{f\prime\prime}(e^{f*}(t))(V^f - V^u(t)) + \lambda^{p\prime\prime}(e^{p*}(t))(V^p(t) - V^u(t))) + \lambda^{f\prime\prime}(e^{f*}(t))\lambda^{p\prime\prime}(e^{p*}(t))(V^p(t) - V^u(t))(V^f - V^u(t)) > 0.$$

When $e^{f*}(t) > 0$ and $e^{p*}(t) > 0$, $V^f - V^u(t) > 0$ and $V^p(t) - V^u(t) > 0$. Thus, since $u_{kk} < 0$, $\lambda^{f\prime} > 0$, $\lambda^{f\prime\prime} < 0$, $\lambda^{p\prime} > 0$, and $\lambda^{p\prime\prime} < 0$, $de^{f*}(t)/dV^p(t) < 0$, $de^{p*}(t)/dV^p(t) > 0$, and $d(e^{f*}(t) + e^{p*}(t))/dV^p(t) > 0$.

To show that $d\lambda(t)/dV^p(t) > 0$ note that

$$\text{(A10)} \qquad \frac{d\lambda(t)}{dV^p(t)} = \lambda^{f\prime}(e^{f*}(t))\frac{de^{f*}(t)}{dV^p(t)} + \lambda^{p\prime}(e^{p*}(t))\frac{de^{p*}(t)}{dV^p(t)} = \frac{\lambda^{p\prime}(e^{p*}(t))(u_{kk}(\lambda^{f\prime}(e^{f*}(t)) - \lambda^{p\prime}(e^{p*}(t))) - \lambda^{f\prime\prime}(e^{f*}(t))(V^f - V^u(t))}{D} > 0$$

when $\lambda^{f\prime}(e^{f*}(t)) < \lambda^{p\prime}(e^{p*}(t))$. Using the first order conditions (A2) and (A4), it is clear that $\lambda^{f\prime}(e^{f*}(t)) < \lambda^{p\prime}(e^{p*}(t))$ if $V^f > V^p(t)$.

To derive conditions under which $dV^p(t)/dy > 0$ when $c(t) = b + y$, consider sample paths where at time t a UI recipient is working part-time and either UI benefits are exhausted before a full-time job is found or UI benefits are not exhausted before a full-time is found. For the former sample paths, total UI payments are invariant to changes in y while for the latter sample paths total UI payments increase with y. Now consider a small increase in the disregard from y' to y''. Fix the on-the-job effort function at its optimal level for y'. For all sample paths where a full-time job is found before benefits are exhausted when the disregard equals y'', a higher discounted utility is received when the disregard equals y'' than when it equals y'. Next, consider all sample paths where benefits are exhausted when the disregard equals y''. For these sample paths, an individual would either exhaust UI payments when the disregard equals y' or not. In either case, however, their total discounted UI payments are strictly larger when the disregard equals y'' than when it equals y', if the discount rate is positive. It is clear then that individuals with utility functions of the form $u(c,k) = c + \Phi(k)$, $\Phi'' < 0$, are strictly better off when the disregard equals y''.

Now for individuals with utility functions of the form $u(c,k) = \gamma(c) + \Phi(k)$ with $\gamma'' < 0$ and $\Phi'' < 0$, it can be shown that along paths which exhaust UI benefits:

$$\text{(A11)} \qquad \frac{d\left(\int_{\infty}^{t} U(c(t), k^*(t)) e^{-rt}\, dt\right)}{dy} = e^{-rt'}\left[\gamma'(b+y)\left(\frac{1-e^{-rd^p}}{r}\right) - \frac{\gamma(b+y)+\gamma(c^p)}{b+y-c^p} e^{-rd^p} d^p\right]$$

where $k^*(t) = 1 - h^p - e^{oj*}(t)$ and recall that t' is the time at which a part-time job is found and d^p equals the duration of benefits in the part-time job. Since γ is strictly concave,

$$\text{(A12)} \qquad \gamma'(b+y) - \frac{\gamma(b+y)-\gamma(c^p)}{b+y-c^p} < 0$$

and

$$\text{(A13)} \qquad \frac{d\left(\int_{\infty}^{t} U(c(t), k^*(t)) e^{-rt}\, dt\right)}{dy} < 0$$

if the difference in (A12) is sufficiently large. However, as long as individuals are either not very risk averse or the probability of exhausting UI benefits at a part-time job before finding a full-time job is sufficiently small, $dV^p(t)/dy > 0$. For general nonseparable concave U, $dV^p(t)/dy > 0$ as long as the probability of exhausting UI benefits before moving from part-time to full-time work is sufficiently small.

To show that $d^2V^p(t)/dy\,dt < 0$ when $c(t) = b+y$, $\partial^2 u/\partial c\,\partial k > 0$, and $e^{oj*}(t)$ is increasing over time, note that by delta arguments it can be shown that

$$\text{(A14)} \qquad \frac{\partial^2 V^p(t)}{\partial y\,\partial t} = (r + \lambda^{oj}(e^{oj*}(t)))\frac{dV^p(t)}{dy} - u_c(b+y, 1-h^p-e^{oj*}(t)).$$

However, $dV^p(t)/dy$ is bounded above by $u_c(b+y, 1-h^p-e^{oj*}(t))/(r+\lambda^{oj}(e^{oj*}(t)))$ when $e^{oj*}(t)$ is increasing with t and $\partial^2 u/\partial c\,\partial k > 0$. This follows by noting that $u_c(b+y, 1-h^p-e^{oj*}(t))/(r+\lambda^{oj}(e^{oj*}(t)))$ would be the derivative of expected discounted utility with respect to y when UI benefits are continued indefinitely and the rate at which a full-time job is found is fixed at $\lambda^{oj}(e^{oj*}(t))$. That $\partial^2\lambda^p(t)/\partial y\,\partial t < 0$ follows from

$$\text{(A15)} \qquad \frac{\partial^2 \lambda^p(t)}{\partial y\,\partial t} = \frac{d\lambda^p(t)}{dV^p(t)}\frac{\partial^2 V^p(t)}{\partial y\,\partial t}$$

and (A14).

APPENDIX B

This appendix derives the likelihood function for the selectivity-corrected competing risks model described in this paper. Since the duration data are grouped, a discrete-time approach is taken. Let T_f and T_p be discrete random variables representing the joblessness duration until a full-time job is found and the joblessness duration until a part-time job is found, respectively. The joint survivor function conditional on θ_f, θ_p, UI, RR, DR, and x, $S(k_f, k_p|UI, RR, DR, x, \theta_f, \theta_p) = Pr(T_f > k_f, T_p > k_p|UI, RR, DR, x, \theta_f, \theta_p)$, is assumed to have the following form:

$$S(k_f, k_p|UI, RR, DR, x, \theta_f, \theta_p) = \exp\left(-\theta_f \sum_{r=1}^{k_f} \exp(\alpha_{fk} + g_{fk}(UI, RR, DR) + \beta'_{fk}x) - \theta_p \sum_{r=1}^{k_p} \exp(\alpha_{pk} + g_{pk}(UI, RR, DR) + \beta'_{pk}x) \right).$$

The parameters α_{wk} are the baseline hazard parameters and may be interpreted as

$$\alpha_{wk} = \log\left(\int_{k-1}^{k} \lambda_w(t)\,dt \right)$$

where $\lambda_w(t)$ is the underlying continuous-time baseline hazard function, $w = f, p$. In the polynomial selectivity-corrected competing risks models,

$$\alpha_{wk} = \alpha_{w0} + \alpha_{w1}k + \alpha_{2w}k^2 + \alpha_{w3}k^3 + \alpha_{24}k^4 + \alpha_{w5}k^4, \quad w = f, p.$$

For identifiability purposes α_{f0} and α_{p0} are both fixed at 0. This is done rather than restricting the two means of the bivariate mixing distribution. The function g_{wk} takes the form specified in equation (8) and the vector of parameters β_{wk} represents the possibly time-varying effects of the other regressors on the re-employment risk to a type w job, $w = f, p$. In the paper, however, $\beta_{wk} = \beta_w$, $w = f, p$. For identifiability purposes, it is assumed that the effect of the regressors within each time period is constant. This assumption is not overly restrictive when the time periods are relatively short.

Recall that only the joblessness duration, $T = \min(T_f, T_p)$ and the type of job accepted are observed. Define

$$\begin{aligned} A_f(k|\theta_f, \theta_p) &= S(k, k|\theta_f, \theta_p) - S(k+1, k|\theta_f, \theta_p) - .5(S(k, k|\theta_f, \theta_p) \\ &\quad + S(k+1, k+1|\theta_f, \theta_p) - S(k, k+1|\theta_f, \theta_p) - S(k+1, k|\theta_f, \theta_p)), \\ A_p(k|\theta_f, \theta_p) &= S(k, k|\theta_f, \theta_p) - S(k, k+1|\theta_f, \theta_p) - .5(S(k, k|\theta_f, \theta_p) \\ &\quad + S(k+1, k+1|\theta_f, \theta_p) - S(k, k+1|\theta_f, \theta_p) - S(k+1, k|\theta_f, \theta_p)), \\ A_o(k|\theta_f, \theta_p) &= S(k, k|\theta_f, \theta_p) - S(k+1, k+1|\theta_f, \theta_p), \end{aligned}$$

and

$$A_c(k|\theta_f, \theta_p) = S(k, k|\theta_f, \theta_p),$$

where the dependence of these functions on UI, RR, DR, and x has been omitted for notational simplicity. The function A_f represents the probability of accepting a full-time job in period k and A_p represents the probability of accepting a part-time job in period k. The term

$$.5(S(k, k) + S(k+1, k+1) - S(k, k+1) - S(k+1, k)),$$

which appears in both A_f and A_p, is an adjustment that is made because durations are measured in discrete-time. Finally, A_o represents the probability of observing a completed joblessness spell K periods long and A_c represents the probability that a joblessness spell lasts more than k periods.

Unemployment insurance receipt is also incorporated into the statistical model as well as the possibility that the unobservables which effect UI receipt are correlated with the unobservables which determine T_f and T_p. It is assumed that

$$P(UI=1|\theta_u)=1-\exp(-\theta_u \exp(\beta_u' z))$$

where z is an I-dimensional vector of regressors (and does not include a constant term), the I-dimensional vector of parameters β_u measures the effects of the regressors on UI receipt, and θ_u is an unmeasured random variable affecting UI receipt which is possibly correlated with θ_f and θ_p but is distributed independently of z.

Now, let

$$B_w(k, UI=1|\theta_u, \theta_f, \theta_p)=A_f(k|\theta_f, \theta_p)P(UI=1|\theta_u), \qquad w=f,\ p,\ o \text{ and } c.$$

The joint distribution of the unobservables $(\theta_f, \theta_p, \theta_u)$ is modeled by assuming that there are J distinct but unobserved types of individuals in the population where individual j is characterized by the triplet of location parameters $(\theta_{fj}, \theta_{pj}, \theta_{uj})$ and occurs in the population with relative frequency p_j, $j=1,\ldots,J$,

$$\sum_{j=1}^{J} p_j = 1.$$

Mixing over all possible types of individuals yields the following unconditional probabilities:

$$B_w(k, UI=1)=\sum_{j=1}^{J} p_f B_w(k, UI=1|\theta_{fj}, \theta_{pj}, \theta_{uj}), \qquad w=f,\ p,\ o \text{ and } c.$$

The functions $B_w(k, UI=0)$ are defined similarly except that $1-P(UI=1|\theta_{uj})$ replaces $P(UI=1|\theta_{uj})$, $w=f$, p, o, and c.

Let K_n represent the joblessness duration of individual n and let UI_n be an indicator variable that equals one if this individual receives UI benefits and equals zero, otherwise, $n=1,\ldots,N$, where N is the sample size. In addition, δ_{fn} is an indicator variable that equals one if an individual n accepts a full-time job and equals zero, otherwise, δ_{pn} is an indicator variable which equals one if an individual n accepts a part-time job and equals zero, otherwise, $n=1,\ldots,N$. Finally, δ_{on} is an indicator that equals one if the spell of individual n is complete but the type of job accepted is unobserved, and equals zero, otherwise, and δ_{cn} is an indicator that equals one if the spell of individual n is incomplete or right censored and equals zero, otherwise, $n=1,\ldots,N$. Then, the log likelihood function of the selectivity-corrected competing risks model is given by

$$\begin{aligned}\log L = \sum_{n=1}^{N} & UI_n \delta_{fn} \log(B_f(K_n, UI=1)) + UI_n \delta_{pn} \log(B_p(K_n, UI=1)) \\ & + UI_n \delta_{on} \log(B_o(K_n, UI=1)) + UI_n \delta_{cn} \log(B_c(K_n, UI=1)) \\ & + (1-UI_n)\delta_{fn} \log(B_f(K_n, UI=0)) + (1-UI_n)\delta_{pn} \log(B_p(K_n, UI=0)) \\ & + (1-UI_n)\delta_{on} \log(B_o(K_n, UI=0)) + (1-UI_n)\delta_{cn} \log(B_c(K_n, UI=0)).\end{aligned}$$

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