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Unemployment Insurance and Joblessness in Portugal

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Abstract

Using data from the Portuguese quarterly employment surveys, 1992-96, the present chapter examines the effect of unemployment insurance (UI) on escape rates from unemployment. The first concern is to identify the manner in which UI influences the time profile of these broad transitions. To this end, two different measures of UI are used, namely, receipt of benefits and time to benefit exhaustion. The second concern is to incorporate destination state into the analysis. Four such destination states are identified: full-time employment, part-time employment, "discouragement," and inactivity. The goal of the exercise is not merely to obtain to refined estimates of the effect of UI (and the other covariates) on unemployment but also to assist in the design of policy.

I. Introduction

The effect of unemployment insurance (UI) in prolonging jobless duration is well established. Thus, job search theory has addressed the role of UI in elevating the reservation wage (e.g. Mortensen, 1970), and a burgeoning empirical literature has amply confirmed its prediction that unemployment spells will lengthen in consequence (see the survey by Devine and Kiefer, 1991). The possibility that UI will in addition evince time-varying effects has also been recognized, but in a much sparser literature (e.g. Fallick 1991; Belzil, 1992; Meyer, 1990; Narandranathan and Stewart, 1993). But to our knowledge there is almost no evidence as to the influence of UI on the choice of destination status out of unemployment. [1] The identification of time-varying effects and the role of destination status provide the twin focuses of the present inquiry.

The importance of a time-varying effect is that it provides a check on the implication of the UI-augmented job search model that escape rates from unemployment should be lower for those workers who are eligible for benefits but should increase as the exhaustion point approaches (since the value of being unemployed drops). Empirical estimation of these time-varying effects yields more information on worker behavior and hence assists in the design of policy (i.e. limiting the duration of benefits).

The importance of destination status for its part is that unemployed workers attach different utilities to the alternatives to unemployment. In each period, the unemployed worker has to define a search strategy and adopt a decision rule appropriate to the maximization of his or her utility function. If the destination states that we consider here - full-time employment, part-time work, "discouragement" (see

[1]The sole exceptions would appear to be McCall's (1996, 1997) studies of the role of state unemployment insurance rules in the U.S. and Canada permitting part-time work on the transition rates out of unemployment into part-time and full-time employment. Narandranathan and Stewart (1993) also look at time-varying effects of UI on transitions into employment and out of the labor force.

below), and inactivity - are indeed the outcome of distinct behavioral choices, de facto aggregation over these states will provide an inadequate portrait of the unemployment experience of individuals. Specifically, both the aggregate hazard function and the estimated regression will compound different, perhaps even opposing, influences. By accounting for such differences our empirical model aims to obtain a better understanding of transitions out of unemployment. More concretely, different cause-specific hazard functions may be expected to characterize, say, transitions into full-time work on the one hand and part-time work on the other. Transitions into part-time work may only be observed after all hope of getting a full-time job is extinguished. In consequence, rising hazard rates might be found in the data after a possibly protracted interval of low escape rates.

The effect of the regressors may also differ materially once destination status is incorporated into the analysis. UI itself is the most subtle but yet most direct case in point. UI may be expected to influence the destination state because it enters as a negative, possibly time-varying cost in the individual utility function. In the most practical terms, if the individual is collecting benefits in a system that does not permit them to paid in conjunction with part-time work, then it is unlikely that the data will indicate many transitions out of unemployment into part-time work until benefit exhaustion. Again, the decision to exit the labor force could be regarded as a timing phenomenon by those who are loosely attached to the labor force. Receipt of UI by workers who optimally time their exit - and (re)entry - into the labor force is likely to be associated with a continued claiming of benefits beyond the date that corresponds with effective withdrawal.

The use of Portuguese data is also of wider interest. First of all, the Portuguese labor market has been singled out as having distinct barriers to reemployment that might be expected to amplify the impact of UI on joblessness. Second of all, the Portuguese case is distinguished by period heterogeneity. That is, there occurred a sea change in the UI system in 1989, that is, shortly before our sample period begins. The administrative changes in question liberalized eligibility requirements for both regular unemployment benefits and unemployment social assistance and, most importantly, increased the maximum duration of benefits. Third, the design of the household survey has a quasi-longitudinal capacity that allows us to properly identify the out-of-unemployment transitions of both UI recipients and non-recipients, without the need to rely solely on stock sample information which would require outside information on inflow rates into unemployment and the assumption of a particular unemployment duration distribution. Arguably, it should therefore be even easier to track the effects of UI than in other economies where there have occurred fewer overt changes in UI systems.

II. Data

Prior to discussing the dataset used here, we discuss some background information relevant to our inquiry. Descriptive information on the course of Portuguese unemployment and unemployment insurance is provided in Table 1. It can be seen that the unemployment rate has risen by more than two thirds since 1992. The mean (elapsed) duration of unemployment has increased continuously, and the distribution of unemployment has changed fairly profoundly. Thus, short-term unemployment (3 months or less) has nearly halved to approximately 20 percent of the total, while long-term unemployment (over 12 months) has risen by somewhat more than 60 percent. Today, two in five workers are now out of work for more than one year.

[Table 1 near here]

The proportion of workers covered by the UI system - currently about one-third of the unemployed - has not risen since 1993.[2] Equally, both the maximum duration of benefits and the replacement rate have remained the same since 1989, in the range of 10-30 months and 65 percent, respectively.[3] It is therefore the pronounced rise in the number of unemployed workers and in unemployment duration that explain why outlays on UI have almost tripled in nominal terms between 1992 and 1996. Those outlays include both regular unemployment insurance benefits and unemployment assistance benefits. Unemployment insurance benefits were first introduced in 1985 with quite stringent eligibility requirements. Unemployment

[2] In computing the coverage rate we exclude those individuals who had exhausted their UI benefits at the time of survey. In other words, we are considering as "covered" only those individuals who responded affirmatively to the question: "Are you currently receiving unemployment benefits?"

[3] The lower and upper limits are determined by age: 10 months for those aged less than 25 years and 30 months for those over 55 years of age.

assistance benefits can be drawn on a means-tested basis (if the household per capita income is less than 80 percent of the minimum wage) for workers ineligible of regular UI and for exhaustees.[4] With this survey we cannot distinguish between the two categories, although we can predict whether an individual is eligible for regular UI or UI assistance based on the tenure of the previous job (see below).

With these preliminaries behind us, we now turn to the disaggregated data that form the basis of the present inquiry, namely the Portuguese quarterly employment surveys (Inquérito ao Emprego) for 1992-96, administered by the Instituto Nacional de Estatística. Note that the choice of period is dictated by changes in survey methodology after the first quarter of 1992, including new sampling procedures and revised definitions of employment, unemployment, and inactivity.

Each quarter, the INE inquires of a random sample of individuals their current labor market status and past labor history. In this sense, just like the U.S. Current Population Survey (CPS), the Portuguese employment survey samples the population of members of a state at a given time and observes their elapsed durations. This sampling plan is called *stock sampling,* and the elapsed (necessarily incomplete) durations are referred to as *backward recurrence times.* It is well known that the distribution of elapsed durations of a stock (of, say, unemployed) gives a distorted image of the distribution of complete durations of a flow of entrants (into the unemployment state). This is because the former sampling plan oversamples long durations (*length biased sampling*) and underestimates their mean duration due to the fact that the spells are ongoing. Such problems can, however, be partially overcome via a joint modeling of the elapsed duration distribution, the probability of being sampled, and the history of flows into the state. This procedure may still impose too much structure on the data and require information that is unavailable (e.g. with respect to flows of entrants). A feasible, and much simpler, alternative is available if the members of a state at a given time are observed over a fixed time interval. Observation over a fixed interval allows us to obtain information about the remaining duration (or *forward recurrence time*) that, conditional on elapsed duration, is distributed as the entrant conditional (on elapsed duration) density function (see Lancaster, 1990).

[4] See Bover, Garcia-Perea, and Portugal (1998) for a detailed description of the Portuguese UI system.

As noted earlier, this nationally representative survey has a quasi-longitudinal capacity - one-sixth of the sample rotating out each quarter - allowing us to track transitions out of unemployment for up to six quarters, and thus enabling us to pursue the conditional approach. In other words, transition rates can be obtained simply by identifying those unemployed individuals, and their elapsed duration in a given quarter, who move out of unemployment over the subsequent quarter.

The destination state of once-unemployed workers can also be identified. For the present purposes, we shall distinguish between four states, namely, full-time employment, part-time work, "discouragement," and inactivity. The first two categories are self explanatory, but some clarification is required for the states of discouragement and inactivity. Discouraged workers are defined as those individuals who, although they did not search for work in the prior 30-day interval, nevertheless responded that they would like a job. In every other respect, however, they are the same as the economically inactive category.

Focusing for the moment upon unemployment, each survey contains information on the length of the current unemployment spell in months, and the unemployment insurance benefit status of the individual. Those identified as UI recipients in the survey will in fact include both those enjoying full benefits and also others in receipt of a lower order level of unemployment benefits which we will term "social assistance" (Subsídio Social de Desemprego). From the UI rules and information contained in the survey on the individual's tenure on his or her previous job, we can further subdivide the recipient category into these two groups. Similarly, given that the maximum duration of UI benefits depends solely on the age of the individual, we can also obtain a reliable measure of the remaining life of benefits (i.e. the time to benefit exhaustion).

In other words, drawing on the administrative rules and the data contained in the survey, we deploy a number of separate UI measures: a dummy variable indicating whether or not the individual is a recipient (UIR); the remaining weeks of benefit entitlement (maximum duration less elapsed duration) (TIMEEX); and, using information on tenure on the previous job, two variables indicating eligibility for either full benefits (ELEG) or lower order social assistance (SOCIAL).[5] For two of these variables - the binary UI measure and time to exhaustion of benefits - we allow for time-varying effects.

The survey contains in addition to the unemployment, destination status, and UI information, data on the individual's age, marital status, disability, level of schooling, tenure on the lost job, number of jobs held (and whether the individual is a new entrant to the labor market), occupational status, reason for job loss, and region of residence, inter al. Descriptive statistics are provided in Appendix 1.

The sole restrictions placed on the data were that the individual be unemployed at the time of the survey, aged between 16 and 64 years, and resident in continental Portugal. Given the possibility of attrition, we also ensured that individuals appearing in two contiguous surveys with the same ID number were in fact the same individual. The sample size is 15734, comprising 7544 males and 8190 females.

III. Methodology

The basic empirical model used here exploits the particular nature of our data set. It will be recalled that we can follow the individual over a period of up to 6 quarters. Given this sample plan, unemployment transitions are observed over a fixed interval of 3 months.

We begin by considering the definition of the overall hazard function for a given set of covariates X, supposing failure time is continuous. Specifically, the instantaneous probability of exiting unemployment at t, given that the individual stayed unemployed until t, is expressed as

$$\theta(t;X) = \lim_{\Delta t \to 0} \frac{P(t \leq T < t + \Delta t | T \geq t, X)}{\Delta t} \tag{1}$$

[5] Because jobs other than that previously held may also contribute to the fulfillment of the eligibility

The instantaneous escape rate function to destination j , j=1, 2, …,w (in our case, j denotes the four destination states of full-time employment, part-time work, discouragement, and inactivity), has been called the "cause-specific hazard function". It is similarly defined as:

$$\theta_j(t;X) = \lim_{\Delta t \to 0} \frac{P(t \leq T < t + \Delta t, J = j | T \geq t, X)}{\Delta t}, \tag{2}$$

which yields

$$\theta(t;X) = \sum_{j=1}^{w} \theta_j(t;X). \tag{3}$$

In this framework, the relevant information is given by the observed unemployment duration, the mode of exit out of unemployment and, for continued unemployment, the indication of an incomplete duration (i.e. the censoring indicator). Our estimation is thus based on data provided by the triplet (T, J, C; X), where C is the indicator of a censored duration.

Assuming proportionality of (possibly time-varying) covariate effects, the cause-specific hazard function in equation (2) can be rewritten for the arbitrary baseline hazard $\theta_{0j}(t)$

$$\theta_j(t;X) = \theta_{0j}(t)\exp\left[X(t)\beta_j\right]. \tag{4}$$

Recall, however, that our information on elapsed duration of unemployment is grouped into monthly intervals, while transitions can solely be identified over a fixed interval of 3 months. A convenient discrete model with multiple modes of failure can be achieved via a straightforward generalization of the grouped duration model offered by Prentice and Gloeckler (1978).

Consider a time axis that is divided into K intervals by points c_1, c_2, … ,and c_{K-1}, and let M=m denote the occurrence of an exit in the interval $\left[c_{t-1}, c_{t+3}\right)$, where t is the realization of a discrete random unemployment duration variable $M \in (1, \ldots, K)$. For a worker who is still unemployed in the m interval, the conditional probability of exiting unemployment into destination r over the next 3-month interval can be written

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requirements, our procedures imply a conservative bound for eligibility.

$$h_r(m|X) = \frac{\exp\left[\lambda_{m_r} + X(m)\beta_r\right]}{\sum_{j=1}^{w} \exp\left[\lambda_{m_j} + X(m)\beta_j\right]} \left\{1 - \exp\left[-\sum_{j=1}^{w} \exp\left[\lambda_{m_j} + X(m)\beta_j\right]\right]\right\}, \tag{5}$$

where the baseline hazard function is absorbed into the parameters $\lambda_{m_r} = \log\left[\int_{c_{m-1}}^{c_{m+3}} \theta_{0r}(t)\right]$.

With random censoring, the likelihood contribution of an individual i who remained unemployed at the beginning of interval m, is given by

$$L_i = h_r(m_i|X_i)^{\delta_i} \prod_{m=1}^{m_i-1} (1 - \sum_{j=1}^{w} h_j(m_i|X_i)), \tag{6}$$

where δ_i denotes the occurrence of a terminating event (a complete duration).

The model being estimated here is analogous to a discrete choice model where the multinomial choice is applied to a sequence of categorical variables that identify the current state (Jenkins, 1995). The advantage of this approach is that it provides a simple interpretation of the regression results in terms of the hazard model. In other words, the estimated model can be viewed as both a discrete choice model and as a duration model.[6]

Our model specifies the conditional (on elapsed unemployment duration) probability of exiting unemployment. Despite some efficiency loss in estimation, this approach has the advantage over the unconditional densities specification of not having to deal with the evolution of inflows into unemployment.

A recurring issue in unemployment duration analysis is the role of unobserved individual heterogeneity. If one seeks to incorporate unobserved heterogeneity into this model, one has also to consider correlated "risks" in the sense that that the error

[6]Strictly speaking, our hazard model is to be distinguished from a competing risks model specification. The latter approach involves defining a number of distinct latent durations according to each risk, where we only observe completed duration for one of the risks, namely, that with minimum duration. In such models, developed in the biostatistics literature, it is sensible to assume that all risks are always present. In our case, however, it is not clear that the assumption holds for labor market decisions and in particular for the case of flows out of unemployment. Our approach is therefore best viewed in the context of the literature on (independent) cause-specific hazard functions (Kalbleisch and Prentice, 1980).

terms capturing heterogeneity may themselves be correlated. Establishing whether or not this is the case would be a difficult and cumbersome procedure, and it is not clear that the identification requirements could be met. It would, of couse, be feasible and straightforward to model unobserved individual heterogeneity assuming uncorrelated error terms, but there is little justification for this essentially artificial procedure.

Finally, turning to the issue of implementation, we note that the baseline hazard comprises eleven intervals. The choice of intervals is dictated by the relative frequency of observations (at risk) within each monthly cell.

IV. Findings

We first consider the probability of escaping unemployment at the most general level, that is, not distinguishing at this point between destination states. A preliminary indication of the effect of UI on hazard rates is provided in Figure 1. Despite the narrowing in the difference between the escape rates of recipients and nonrecipients through time, there is prima facie evidence of marked and persistent disincentive effects of UI.

[Figure 1 near here]

The fitted discrete duration model is reported in Table 2. As discussed earlier, the baseline hazard function is specified as a piecewise-constant function with eleven intervals The first six intervals correspond to calendar months, the next two intervals are three months each, while the ninth and tenth intervals are are thirteen to eighteen and nineteen to twenty-four months, respectively. The final interval thus pertains to twenty-five months and above.

[Table 2 near here]

The coefficient estimates show the effect of the regressors in (proportionally) shifting the baseline hazard up or down. The coefficient estimate for the UI dummy indicates that receipt of UI benefits decreases by 25 percent the chance of exiting from unemployment. For now, it is supposed that this disincentive effect is constant through time, although we will shortly relax this assumption.

The effects of the other covariates on escape rates might briefly be described. First, it is apparent that males have modestly higher escape rates than females (around 8.2 percent higher). Second, the impacts of AGE, TEN, and DISABLE are thoroughly conventional. Each is associated with a reduction in escape rates. AGE and TEN may be expected to lower escape rates by reason of their effect on reservation wages, whereas the effect of DISABILTY most probably operates through a lower arrival rate of job offers. Third, the directional effects of SCHOOL and MARRIED on escape rates are both positive, but only the former is statistically significant at conventional levels. Interestingly, and consistent with analysis of U.S. displaced worker data (e.g. Portugal and Addison, 1995), these results are reversed if the regression is run for males alone. Fifth, the opposite signs on the (number of) JOBS and FIRST JOB variables provide some indication that greater labor market knowledge raises escape rates. That said, note that the coefficient estimate for the argument proxying broad occupational status, WCOLL, is both negative and statistically significant. This latter result, however, stems from the inclusion of females within the sample. Sixth, whether a worker lost his job as a result of a mass layoff or by reason of the termination of a fixed term contract (respectively, LAYOFF and ENDFT) is apparently not material to escape rates. Finally, the coefficient estimates associated with the four year dummies broadly indicate that flows out of unemployment are procyclical (1992, a boom year, being the omitted category), while those for the four regional dummies reflect the persistence of unemployment rate differences across different areas of the country (the high escape rates of the Alentejo region [the omitted category] are noteworthy in this regard).

[Table 3 near here]

We next consider alternative specifications for the effect of UI on escape rates. For this purpose, we first add to our previous specification the alternative unemployment insurance measure TIMEEX, namely, the time to benefit expiration. This substitution is charted in the second column of Table 3, the first column entry merely reiterating the coefficient estimate for the UI recipiency variable, UIR, from Table 2. It is clear that escape rates decline substantially, the further away one is from expiration. Specifically, the escape rate declines by 2.6 percent for each month of remaining entitlement.

The next innovation, given in the third column of the table, is to substitute two measures of UI for one, namely, the imputed receipt of full UI benefits (ELEG) and the imputed receipt of entitlement to the second-order benefits of social assistance (SOCIAL). It is apparent that imputed UI receipt has the stronger effect. The relative magnitude of the coefficient estimates reflects the fact that regular benefits are both larger and paid for a longer period than social assistance.

The balance of the material in Table 3 is reserved for the time-varying effects of unemployment insurance. We deploy two UI measures, namely, UIR (actual UI receipt) and TIMEEX. We allow each to have different effects over the spell of unemployment. The fourth and fifth columns of Table 3 give separate results for the conventional UI variable; the first using exactly the same intervals as those defining the baseline hazard, and the second aggregating over those intervals to yield a somewhat more parsimonious specification. The results in the fourth column of the table indicate UI effects that persist up to two years. The effect is not monotonic, however, and as expected the pattern of coefficient estimates mirrors the gap between the empirical hazards of Figure 1. The findings in the fifth column, using a smaller number of intervals, confirm these results.

The final column of Table 3 provides further information on the effects of unemployment insurance. This redefinition of the UI variable indicates rather strong disincentive effects the longer is the period to expiration - at least over the first four intervals. For example, if there is just under 18 months of remaining entitlement, the recipient is 38 percent less likely to escape from unemployment than his uninsured counterpart. At one year the difference is 34 percent, falling to 26 percent at six months and to only 3 percent at one month.

Next consider the issue of destination status. Sample means of jobless duration, UI recipiency, and time to exhaustion of benefits by destination status are provided in Table 4. Comparing the still unemployed (in the next quarter) with individuals in our four destinations, it can be seen that their elapsed unemployment duration is much larger, with the notable exception of part timers. The proportion of UI recipients is also much greater among the unemployed.

[Table 4 near here]

From Table 4 it can also be seen that the most common form of transition is from unemployment to full-time employment. In terms of elapsed duration, full-time employment is associated with the lowest joblessness, and part-time employment with the longest (though note the comparatively small number of transitions into part-time work). In terms of UI recipiency, transitions to part-time work record the lowest coverage and transitions to full-time employment the highest. As far as time to benefit exhaustion is concerned, no clear pattern emerges although there is the suggestion that part-time workers use up most of their benefit entitlement.

[Table 5 near here]

The object of the regression analysis is to establish how UI affects the probability of entering a particular destination state. Summary results for two UI measures are given in panel (a) of Table 5. A general opening observation is that the disincentive effects of UI are evident across all destination states, even if there is considerable variation in this effect. Beginning with the UIR variable, the most striking result is the magnitude of the disincentive effect for part time work and, to a lesser extent, inactivity. Specifically, recipients are respectively 4.6 times and 1.7 times less likely than are nonrecipients to enter these states. Neither result is surprising since insured workers have reservation wages that typically exceed the part-time wage, while for those who transition out of the labor force) the explanation would seem to be linked to optimal timing considerations (see below).

Before turning to consider the summary results for our alternative UI measure - time to exhaustion of benefits - we note that full results for the UIR model are provided in Appendix 2. The most notable results are as follows. First, as expected, males are much more likely than females to enter full-time employment and correspondingly less likely to enter the other destinations; in particular, discouragement is a relatively unlikely destination state. Second, older workers are less likely to transition into full-time employment than are their younger counterparts but, unlike longer-tenured workers, they are not also more prone to be discouraged. Third, better educated workers (if not white-collar employees per se) are more likely to transition into full-time employment and disabled workers to enter part-time employment. Fourth, those looking for their first job are much less likely to find full-time employment and much more likely to end up discouraged or inactive that other job seekers. Fifth, those who previously had a full-time job are more likely to be employed than discouraged/inactive, and this distinction is palpably more important than the broad type of job found (i.e. full-time employment vs. part-time work). Finally, although a poor macro environment seems to hurts job prospects it is not clear that destination rates are affected in a consistent manner.

The time to expiration measure provides us with a similar description of the role of UI. Thus, disincentive effects are again observed for all transitions, and the pattern of coefficient estimates for TIMEEX very closely track those obtained for UIR. Once TIMEEX substitutes for UIR, the model has a somewhat better fit.

In panel (b) of Table 4 we present results for our redefined UI variable that allows for time-varying effects of the time to exhaustion variable. This is our preferred specification. Again, the general conclusion as to the disincentive effects of UI on entry into all destination states is borne out. The specific results are as follows. Perhaps the most obvious result is that the effect of UI on exiting into full-time employment is fairly strong during the first four intervals reported in the table. Over that period, the disincentive effect of UI declines sharply and becomes positive, if not statistically significant, at one month to exhaustion. For part-time employment, on the other hand, disincentive effects are reported throughout with no clear suggestion that escape rates rise at exhaustion. If these results are as expected, the findings for discouragement and inactivity are rather disappointing: only for the latter are there any indications of a rise in hazard rates at exhaustion. On this evidence at least it seems hard to conclude that there is any consistent effect of UI.

[Figure 2 near here]

The baseline hazard functions corresponding to the specification contained in the lower panel of Table 5 are charted in Figure 2. In each case, the functions describe the experience of an individual possessing the sample characteristics with respect to all the covariates but who is assigned a zero value for the UI dichotomous variables. The results for transitions into full-time employment, shown in panel (a) of the figure, point to a near continuous decline in escape rates with rising jobless duration. This negative duration dependence phenomenon will be familiar to students of U.S. jobless duration data. The U-shape of the baseline hazard function for transitions into part-time work, shown in panel (b), is consistent with the following story: those who desire part-time employment ab initio manage to locate such jobs rather rapidly, while others less enamored of part-time work only reluctantly take part-time employment after unsuccessful search for a preferred, full-time job. As for the baseline hazards applicable to the two remaining destination states, the picture is more clouded. Evidently, those whom we have identified as "discouraged" appear not to fit the stereotype; that is, there is little evidence of rising escape rates over the course of the jobless spell. Such individuals may indeed be prepared to work but at a wage the market is unprepared to pay, at least currently. The not-dissimilar pattern reported in panel (d) would seem to make the same point more generally: labor market withdrawal is not an end state realized after all else has been tried.

Summarizing, we have found that UI is a disincentive to all but continued unemployment. In other words, disincentive effects are observed across all destination states. Second, UI influences the choice of destination state, since if it failed to do so we would observe the same UI effect across all such states. Third, the disincentive effect is strongest for part-time work, followed by labor force withdrawal, and then discouragement. Accordingly, it is weakest for full-time employment.

V. Conclusions

This Chapter has used a newly-released datatset for Portugal to investigate the effect of UI on unemployment duration. The strengths of the new dataset, aside from it being a nationally representative survey with the usual array of demographic and human capital variables, are that it contains information on UI recipiency, has a quasi-longitudinal capacity, and allows us to identify destination states out of unemployment. Also notable is the fact that the sample period marks a major liberalization of the Portuguese UI system.

The analysis has exploited a conditional hazard model that avoids the pitfalls of other approaches (and datasets), stemming from the fact that UI recipient status may be affected by elapsed unemployment duration. We refer here to the problem of endogenous sample stratification bias that dogs more conventional exercises based on observed differences between the jobless durations of UI recipients and nonrecipients. Our approach gets around this problem because we are able to observe elapsed duration *and* transitions out of unemployment for *both* recipients and non-recipients which enable us to estimate fairly conventional discrete choice models.

Apart from its interest in time-varying effects, the major innovation of the Chapter is its incorporation of destination status into the analysis. This innovation accommodates the potentially different search strategies of unemployed workers. Failure to differentiate between types of transition out of unemployment will compound likely heterogeneous effects, and impart bias to estimates of the impact of UI on unemployment duration. Our methodology does not incorporate a formal model of sequential search behavior and thus has to be interpreted as a reduced form approach. However, without additional information on the search behavior of the individuals, it is not clear how to implement such a model without imposing an undesirable amount of structure.

Strong evidence of the disincentive effects of UI was detected in the data. This result obtained both in general and for each of the four transition states identified here. That is, access to UI benefits and time to benefit exhaustion were found to depress escape rates. Time-varying effects were also evident in the data for each UI measure. In general, the effect of UI tends to decrease with elapsed duration of unemployment. This pattern is clearly influenced by the maximum duration of benefits as evinced by the coefficient estimates associated with the time to benefit exhaustion.

Moreover, different effects were observed by destination state. The UI effect was huge for transitions into part-time employment, and it was both differentiated and substantive (albeit smaller) for the other three destinations. It was smallest for those who transitioned into full-time employment. Turning to the shape of the baseline hazards, the declining hazard rates of those eventually exiting into full-time employment point to the role of stigmatization and/or human capital depreciation in diminishing reemployment prospects. There is also evidence of duality for the part-time jobs destination state, likely reflecting fairly rapid job finding among those who want such jobs and more protracted unemployment among those who may not.

Interestingly, there is only limited evidence to favor real discouragement among those we classified as discouraged. That is, there was only a modest tendency toward rising escape rates into the "discouragement" destination state over the course of the unemployment spell and as far as the more general classification of inactivity is concerned, the broad picture is one of declining hazard rates.

Finally, from a policy perspective, perhaps the most obvious conclusion is that the duration of benefits should be shortened, even if such a move would have to be accompanied by other policy initiatives (see below). Another alternative would be to reduce the average replacement rate. Here our evidence might suggest that changes in duration rather than replacement rate would have the bigger bang per escudo, based on the rather modest observed difference in disincentive effect as between full UI benefits on the one hand and social assistance on the other. Another fairly obvious conclusion stemming directly from our analysis would be that serious attention be given to the possibility of allowing part timers to draw benefits when working part time.

It would be idle to suggest that a study such as this could offer solutions to the problems of long-term unemployment per se. Employment and retraining subsidies for the truly disadvantaged may have an important role to play here, even if the U.S. experience does not exactly encourage a sanguine view of past programs (e.g. LaLonde, 1996).

Our analysis offers only the broadest of hints as to the problem - as manifested in the declining hazard rates associated with exiting unemployment into full-time employment - not the policy response. What our analysis does address is the disincentive effects of UI for the generality of workers. Policies that ameliorate this disincentive are within its purview. In addition to the measures identified earlier, there is at least the suggestion that Portugal might wish to experiment with reemployment bonuses, whereby those obtaining jobs more rapidly than heretofore share some of the UI savings. There are risks with this strategy, the most obvious one being moral hazard: some workers who are eligible for benefits but do not currently draw them might be induced to do so in order to qualify for the bonus were the scheme introduced.

A further option would be to significantly expand job search assistance programs. The U.S. evidence suggests that rather modest job search assistance programs can both yield savings to the public purse and benefit workers, although there is some debate as to whether it is the measures themselves or the more stringent application of the job search test that is responsible for the beneficial outcomes (Meyer, 1995).

A final option, suggested by this latter remark, is that the employment agency actively consider tightening the job search requirement. All such measures are secondary to root and branch reform of the UI system, but the protracted nature of Portuguese unemployment suggests that they are worth investigating.

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Figure 1: Empirical Hazard Function by Unemployment Insurance Recipiency Status

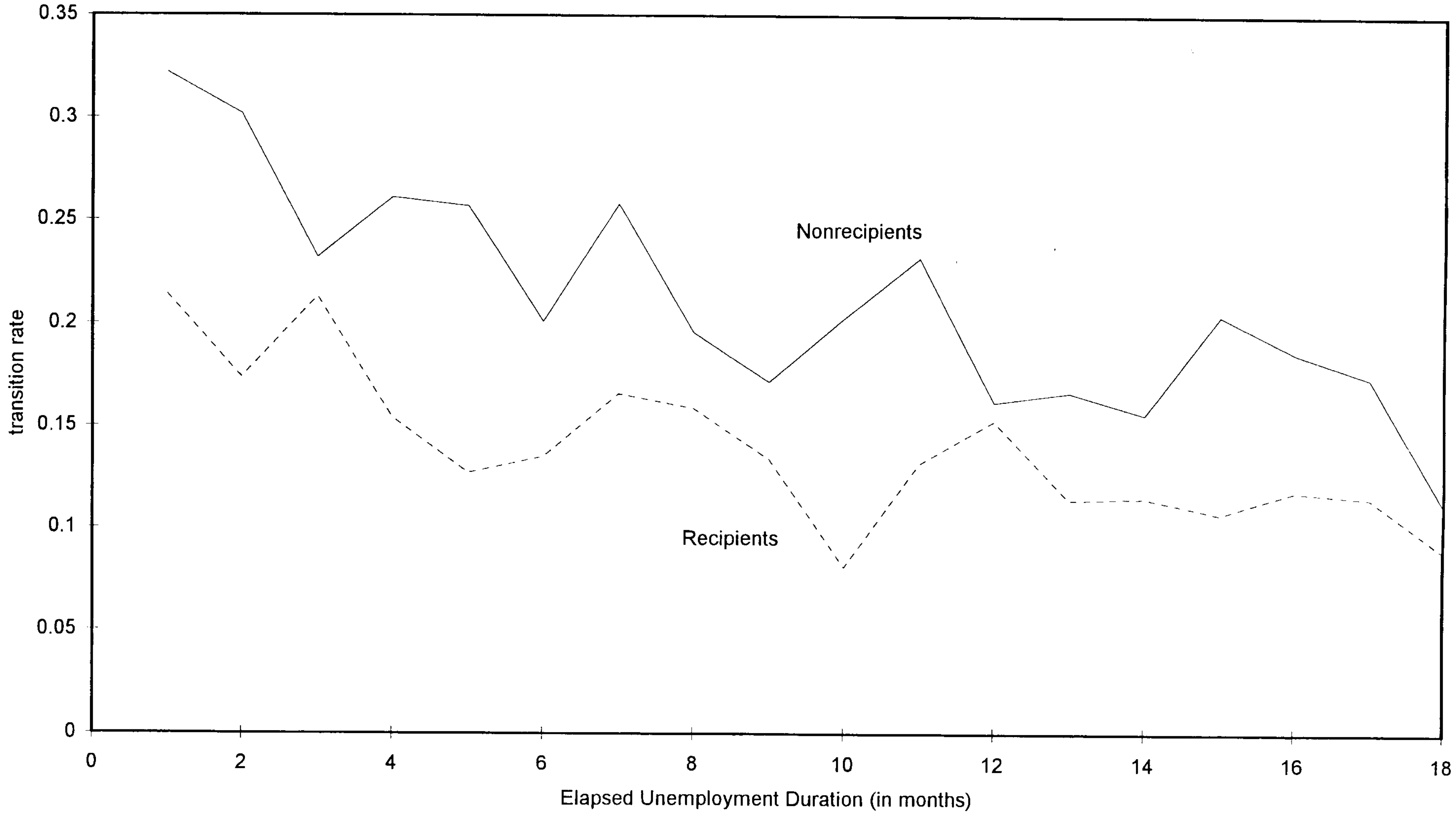
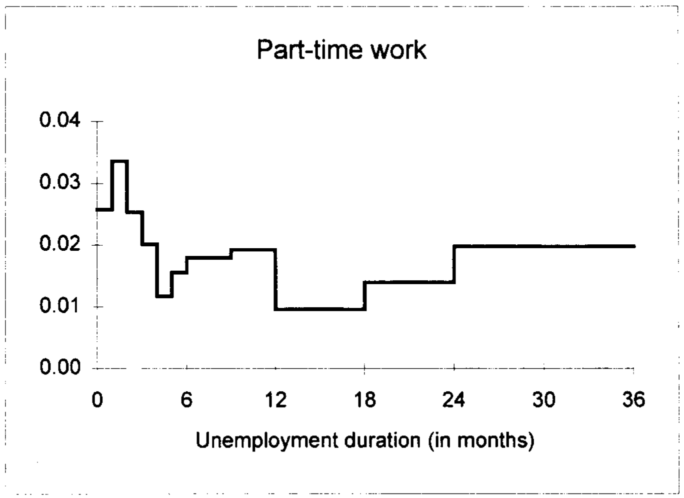


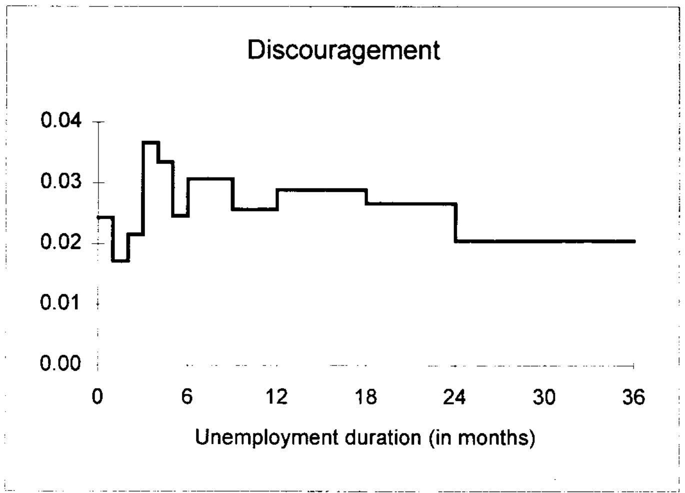
Figure 2: Baseline Hazard Functions by Destination State

Panel (a)

Panel (b)



Panel (c)



Panel (d)

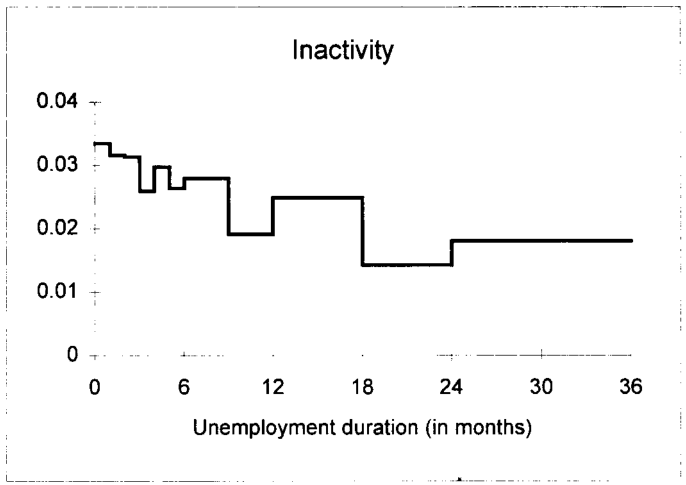


Table 1: Unemployment and Unemployment Insurance in Portugal, 1992-96

	1992	1993	1994	1995	1996
Unemployment rate (percent)	4.1	5.5	6.8	7.2	7.3
Average unemployment duration (months)	12.2	11.6	12.3	14	15
Proportion looking for work for 3 months or less (percent)	36.2	29.6	27.2	21.8	20.8
Proportion looking for work for more than 12 months (percent)	24.9	28.9	32.6	38.3	40.4
Unemployment insurance (UI) coverage (percent)	24	34.4	33.8	34.1	31.2
UI replacement rate (percent)	65	65	65	65	65
UI benefit duration (months)	10 to 30	10 to 30	10 to 30	10 to 30	10 to 30
Average monthly unemployment benefits (thousand escudos)					
Full UI benefits	55.6	66.2	71.2	73.1	74.9
Social assistance	35.4	39.0	42.2	41.8	41.6
Numbers entitled to					
Full UI benefits (December in each year)	82,049	110,534	109,763	101,439	98,505
Social assistance (December in each year)	47,115	60,640	65,393	74,964	83,569
UI outlays (billion escudos)	46.6	106.3	127.5	132.5	132.4
UI expenditures as a proportion of GDP (percent)	0.41	0.80	0.90	0.86	0.80

Table 2: Estimated Piecewise-Constant Hazards Regression, Aggregate Model

Variable	Coefficient Estimate
UIR	-0.291
	(0.048)
MALE	0.079
	(0.038)
AGE	-0.010
	(0.002)
SCHOOL	0.019
	(0.006)
TEN	-0.011
	(0.004)
JOBS	0.012
	(0.003)
WCOLL	-0.115
	(0.057)
MARRIED	0.032
	(0.047)
DISABILITY	-0.487
	(0.220)
FIRSTJOB	-0.178
	(0.062)
LAYOFF	-0.014
	(0.066)
ENDFT	0.082
	(0.047)
YEAR 93	-0.156
	(0.063)
YEAR 94	-0.147
	(0.063)
YEAR 95	-0.294
	(0.066)
YEAR 96	-0.142
	(0.068)
NORTH	-0.311
	(0.058)
CENTER	-0.009
	(0.073)
LISBOA	-0.346
	(0.057)
ALGARVE	-0.317
	(0.084)
Log-likelihood	-7465.312

Asymptotic standard errors in parenthesis

Table 3: Summary Results of the Effect of Unemployment Insurance on Transitions Out of Unemployment

Variable	Specification (1)	(2)	(3)	(4)	(5)	(6)
UIR	-0.291					
	(0.048)					
TIMEEX		-0.026				
		(0.004)				
ELEG			-0.423			
			(0.065)			
SOCIAL			-0.311			
			(0.091)			
Recipient Elapsed Duration						
1 month				-0.371		
				(0.114)		
2 months				-0.547		
				(0.137)		
3 months				0.036		
				(0.139)		
4 months				-0.527		
				(0.168)		
5 months				-0.672		
				(0.178)		
6 months				-0.317		
				(0.176)		
7-9 months				-0.242		
				(0.118)		
10-12 months				-0.266		
				(0.128)		
13-18 months				-0.272		
				(0.140)		
19-24 months				-0.244		
				(0.180)		
25 months or more				0.304		
				(0.149)		
Recipient Elapsed Duration						
1-6 months					-0.388	
					(0.062)	
7-12 months					-0.253	
					(0.088)	
13-18 months					-0.272	
					(0.140)	
19 months or more					-0.060	
					(0.118)	
Recipient Time to Exhaustion						
1-2 months						-0.034
						(0.169)
3-5 months						-0.296
						(0.118)
6-11 months						-0.414
						(0.073)
12-17 months						-0.479
						(0.094)
18-23 months						-0.392
						(0.112)
24 months or more						-0.336
						(0.160)
Log-likelihood	-7465.312	-7460.552	-7458.113	-7449.713	-7459.540	-7452.336

Note: The full array of covariates are given in Table 1

Table 4: Mean Values of Elapsed Duration and Unemployment Insurance Status by Destination State

	Destination state[a]				
Variable	Full-time employment	Part-time work	Discouragement	Inactivity	Unemployed
DURATION	8.717	13.995	14.475	10.778	13.533
UIR	0.278	0.093	0.251	0.187	0.342
TIMEEX	10.060	8.950	9.012	9.453	10.028
Number of events	2087	216	339	347	12749

Note: [a]Individuals that exit unemployment into any of the four categories in the subsequent quarter

Table 5: Summary Results of the Effect of Unemployment Insurance on Transitions Out of Unemployment by Destination State

	Transition to:							
Variable	Full-time employment		Part-time work		Discouragement		Inactivity	
Panel (a)								
UIR	-0.130		-1.533		-0.324		-0.511	
	(0.055)		(0.247)		(0.143)		(0.156)	
TIMEEX		-0.013		-0.118		-0.035		-0.044
		(0.004)		(0.024)		(0.012)		(0.013)
Log-likelihood	-5905.8	-5904.0	-1096.3	-1102.7	-1576.8	-1574.5	-1628.4	-1627.6
Panel (b)								
Recipient Time to Exhaustion								
1-2 months		0.071		-1.112		-0.375		0.131
		(0.195)		(1.009)		(0.586)		(0.460)
3-5 months		-0.189		-0.991		-0.312		-0.488
		(0.136)		(0.593)		(0.347)		(0.393)
6-11 months		-0.245		-2.625		-0.324		-1.006
		(0.080)		(0.714)		(0.226)		(0.301)
12-17 months		-0.329		-1.862		-0.233		-0.854
		(0.110)		(0.591)		(0.236)		(0.336)
18-23 months		-0.090		-2.567		-1.312		-0.495
		(0.126)		(1.013)		(0.468)		(0.347)
24 months or more		-0.125		-0.741		-0.780		-0.512
		(0.187)		(0.616)		(0.532)		(0.484)
Log-likelihood		-5899.5		-1091.8		-1572.4		-1623.0

Note: The full array of covariates are given in Table 2

Appendix 1: Definition of Variables and Descriptive Statistics by Unemployment Insurance Recipiency

Variable	Recipient mean	Nonrecipient mean
UNOUT transition out of unemployment =1, 0 otherwise	0.146	0.210
DURATION length of unemployment in months	10.540	13.995
TIMEEX time to benefit exhaustion (in months)	10.103	
ELEG eligible for full UI =1, 0 otherwise	0.629	
SOCIAL eligible for lower UI assistance =1, 0 otherwise	0.157	
AGE age in years	39.073	29.816
SCHOOL years of schooling completed	5.278	7.278
TEN years of tenure on previous job	8.088	3.125
JOBS number of previous jobs	3.547	2.314
MALE =1 if male, 0 otherwise	0.514	0.463
WCOLL =1 if white collar employee, 0 otherwise	0.239	0.180
MARRIED =1 if married, 0 otherwise	0.713	0.421
DISABILITY =1 if disabled, 0 otherwise	0.009	0.011
FIRSTJOB =1 if looking for first job, 0 otherwise		0.253
LAYOFF =1 if job lost by reason of mass layoff, 0 otherwise	0.291	0.080
ENDFT =1 if job lost through termination of a fixed-term contract, 0 otherwise	0.322	0.261
[Omitted category: OTHER REASONS]		
YEAR 93 =1 if 1993, 0 otherwise	0.254	0.233
YEAR 94 =1 if 1994, 0 otherwise	0.248	0.234
YEAR 95 =1 if 1995, 0 otherwise	0.242	0.224
YEAR 96 =1 if 1996, 0 otherwise	0.173	0.173
[Omitted category: YEAR 92]		
NORTH =1 for the North region, 0 otherwise	0.381	0.326
CENTER =1 for the Center region, 0 otherwise	0.081	0.092
LISBOA =1 for the Lisboa and Vale do Tejo region, 0 otherwise	0.351	0.370
ALGARVE =1 for the Algarve region, 0 otherwise	0.071	0.070
[Omitted category: ALENTEJO]		
n	5105	10629

Appendix 2: Estimated Piecewise-Constant Hazards Regression by Destination State

	Transition to			
Variable	Full-time employment	Part-time work	Discouragement	Inactivity
UIR	-0.130	-1.533	-0.324	-0.511
	(0.055)	(0.247)	(0.143)	(0.156)
MALE	0.289	-0.387	-0.671	-0.161
	(0.045)	(0.148)	(0.121)	(0.112)
AGE	-0.014	0.004	0.003	-0.003
	(0.003)	(0.008)	(0.007)	(0.007)
SCHOOL	0.028	0.018	-0.024	0.006
	(0.007)	(0.022)	(0.018)	(0.017)
TEN	-0.026	-0.014	0.024	0.014
	(0.005)	(0.014)	(0.008)	(0.009)
JOBS	0.019	-0.020	-0.030	-0.006
	(0.004)	(0.020)	(0.017)	(0.016)
WCOLL	-0.153	-0.123	0.091	-0.065
	(0.068)	(0.210)	(0.168)	(0.169)
MARRIED	0.025	0.263	0.122	-0.174
	(0.056)	(0.171)	(0.136)	(0.143)
DISABILITY	-0.556	0.701	-1.105	-0.676
	(0.270)	(0.514)	(1.004)	(0.712)
FIRSTJOB	-0.442	-0.213	0.349	0.482
	(0.077)	(0.226)	(0.175)	(0.168)
LAYOFF	0.111	0.048	-0.385	-0.320
	(0.078)	(0.253)	(0.193)	(0.208)
ENDFT	0.138	0.273	-0.263	-0.172
	(0.054)	(0.170)	(0.148)	(0.153)
YEAR 93	-0.217	-0.123	-0.097	0.305
	(0.077)	(0.225)	(0.166)	(0.191)
YEAR 94	-0.048	-0.085	-0.836	0.039
	(0.076)	(0.226)	(0.198)	(0.200)
YEAR 95	-0.260	-0.168	-0.463	-0.084
	(0.080)	(0.234)	(0.184)	(0.206)
YEAR 96	-0.069	-0.446	0.349	0.069
	(0.082)	(0.264)	(0.190)	(0.209)
NORTH	-0.249	0.027	-1.004	0.030
	(0.070)	(0.225)	(0.167)	(0.185)
CENTER	-0.046	0.469	0.063	-0.370
	(0.089)	(0.256)	(0.174)	(0.269)
LISBOA	-0.338	-0.029	-0.892	0.177
	(0.069)	(0.220)	(0.159)	(0.179)
ALGARVE	-0.180	-0.768	-0.607	-0.471
	(0.097)	(0.421)	(0.234)	(0.303)
Log-likelihood	-5905.771	-1096.275	-1576.820	-1628.372