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To Search or Not to Search? The Effects of UI Benefit Extension for the Elderly Unemployed

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ABSTRACT: In Finland the older unemployed can collect unemployment insurance (UI) benefits until retirement, while the entitlement period for younger groups is two years. In 1997 the eligibility age of persons benefiting from extended benefits was raised from 53 to 55. This paper takes advantage of this quasi-experimental setting to identify the effect of extended UI benefits on transitions out of unemployment among the elderly unemployed. We apply a competing risks version of a split population duration model to account for multiple exit routes and the possibility that some of the elderly unemployed may not be active in the labour market due to pension rules. We estimate that roughly half of the workers with extended UI benefits have effectively withdrawn from the labour market. Those who remain active have a similar hazard rate to employment as individuals with the two-year entitlement period, but much lower hazard rates to non-participation and labour market programmes.

Keywords: Unemployment insurance, unemployment duration, early retirement, competing risks models.

TIIVISTELMÄ: Suomessa ikääntyneet työttömät voivat nostaa työttömyyspäivärahaa 60 ikävuoteen asti, jolloin heillä on oikeus siirtyä työttömyyseläkkeelle. Vuonna 1997 laajennetun päivärahaoikeuden alaikäraja nostettiin 53:sta 55:een. Tutkimuksessa hyödynnetään reformia analysoitaessa, miten laajennettu päivärahaoikeus vaikuttaa poistumiin työttömyydestä. Työttömyyden kestoa mallinnetaan kilpailevien riskien split population -mallilla, jossa huomioidaan vaihtoehtoiset poistumisreitit sekä mahdollisuus, että osa ikääntyneistä saattaa passiivisesti odottaa työttömyyseläkkeelle pääsyä. Tulosten mukaan joka toinen laajennetun päivärahaoikeuden piirissä oleva 53–54-vuotias työtön on passiivinen ja käytännössä jo vetäytynyt työmarkkinoilta. Ehdollinen siirtymätodennäköisyys työllistyä niille laajennetun päivärahaoikeuden piirissä oleville, jotka vielä aktiivisesti pyrkivät takaisin työelämään, on samalla tasolla kuin verrokkiryhmällä, jolla ei ole laajennettua päivärahaoikeutta. Toisaalta heidän ehdolliset todennäköisyytensä siirtyä työvoimapoliittisille toimenpiteille ja työvoiman ulkopuolelle ovat verrokkiryhmää selvästi alhaisemmat.

Asiasanat: Ansiosidonnainen työttömyysturva, työttömyyden kesto, varhaiseläke, kilpailevien riskien malli.

Contents

1 Introduction

Unemployment differences between the European countries and the United States have been the focus of much political and academic debate during the past couple of decades. In addition to higher levels of unemployment, the duration of unemployment spells is typically much longer in Europe than in the US (Machin and Manning, 1999). A high incidence of long-term unemployment among older workers and a tendency of workers to leave the labour force at ages several years below the official retirement age are common problems in Europe. The unemployment compensation system, with generous benefit levels and long entitlement periods, is often blamed for being responsible for much of the European unemployment problem (e.g. Ljungqvist and Sargent, 1998). In many European countries the entitlement periods are further extended for older workers and/or particular early retirement schemes are tailored for the elderly unemployed so that unemployment-related benefits effectively provide a particular pathway to early withdrawal from the labour market (Duval, 2003). In Finland unemployment insurance (UI) benefits can be collected for a maximum of two years but an exception is made for the elderly unemployed: workers old enough can receive UI benefits until early retirement via a particular unemployment pension at age 60. Unemployment pension benefit can be received until old-age pension at age 65. The combination of extended UI period and unemployment pension is known as the unemployment tunnel (UT) scheme. Since the level of unemployment pension benefits is comparable to that of UI benefits, this scheme effectively provides an indefinite period of UI benefits for older workers.

This study contributes to the growing literature on the effects of maximum UI duration on transitions out of unemployment. The analysis of unemployment insurance has been the topic of several theoretical and empirical studies in recent years. In the empirical work the focus has been on estimating reduced-form duration models where the restrictions implied by structural job search models are not imposed. A typical problem in most empirical work has been the lack of variation in the maximum duration of UI benefits that can be regarded as independent of other factors determining unemployment duration (Atkinson and Micklewright, 1991; and Holmlund, 1998). A number of recent studies have exploited various policy changes in an attempt to overcome the endogeneity issue (e.g. Hunt, 1995; Winter-Ebmer, 1998; Bratberg and Vaage, 2000; Card and Levine, 2000; Carling et al., 2001; Røed and Zhang, 2003; Lalive and Zweimüller, 2004; Uusitalo and Moisala, 2003; and Kyyrä and Wilke, 2007). In these studies escape rates from unemployment before and after the reform are compared within the group affected. A similar approach is taken in this study.

In Finland the age threshold for the extended UI period was raised from 53 to 55 in 1997. As a result, the entitlement period of UI benefits for workers aged 53-54 at the time of job loss was effectively reduced to two years, while the other groups remained unaffected by the reform. Kyyrä and Wilke (2007) exploit this reform as a quasi-experiment to evaluate the effects of the UT scheme on transitions between unemployment and employment among older workers. They found that disproportionate numbers of dismissals fall on employees who are old enough to be entitled to the extended UI period, and that only a relatively small fraction of the elderly unemployed with extended UI benefits eventually return to employment. Especially large employers tend to exploit this feature of the UI system to get rid of their elderly employees, as a reasonable income level until retirement is fully secured for them. The 1997 reform led to a decrease in the flow to unemployment and an increase in the flow out of unemployment to employment among workers aged 53-54, and thereby reduced unemployment among the elderly.

This study elaborates the effects of extended UI benefits on the flow out of unemployment among the elderly unemployed in a competing risks setting. We use a sample of workers aged 50-54 who became unemployed between 1995 and 1998, drawn from the Finnish Employment Statistics database. Unemployment experiences of the group aged 53-54 are compared under two schemes: the extended UI entitlement period (pre-reform scheme) and the conventional UI period of two years (post-reform scheme). The younger group (aged 50-52) serves as a control group used to eliminate the business cycle effect. Since the unemployment spells can end through the taking of a job, withdrawal from the labour force, or participation in an active labour market programme (ALMP), competing risks are inherent in our data. As pointed out by Carling et al. (1996), the availability of various labour market programmes, in particular, may play an important role in mitigating the incentive effects of UI on the job-finding rate. A novel feature of our analysis is that we explicitly allow for some older workers, registered as unemployed job seekers, to effectively withdraw from job search and simply wait for access to early retirement. If this is the case, the standard duration models that assume all individuals are at risk of experiencing the event of interest are not applicable. Therefore, we apply a competing risks version of a split population duration model that is capable of accounting for this sort of heterogeneity, i.e. the possibility that some individuals do not consider all possible exit routes.[1] The idea is to model simultaneously both the likelihood that the worker is still active in the labour market and the timing of exit to various end-states conditional on being active. This approach allows us to distinguish the participation decision from labour market behaviour in the case of continued search.

The empirical analysis of UI has focused on detecting effects on the hazard rates. The cause-specific hazard function can be used to identify the effect of the UI scheme on the instantaneous transition rate to a particular destination at a given phase of the spell conditional on having not exited from unemployment to *any* destination. As a result, the employment hazard functions do not provide direct information about policy-relevant issues like changes in the overall or cumulative probability of employment, nor in the expected duration of unemployment. These are functions of all cause-specific hazard func-

[1] Kyyrä and Wilke (2007) estimate a Cox proprotional hazard model. They focus on exits to employment and treat observations on transitions to other destinations as censored.

tions, each of which is subject to change in response to the change in the entitlement period. Therefore, we summarize our results from the competing risks analysis in terms of the cumulative incidence functions, which describe the probability of leaving unemployment to a particular destination by a given time. In this way we can assess to what extent the effect of the extended UI benefits on the probability of entering employment stems directly from a change in the employment hazard and indirectly from changes in other cause-specific hazard functions. This is the question of obvious interest because policy makers can also affect the employment probability indirectly, for example, by regulating the availability of labour market programmes over the course of the unemployment spell.

The plan of the paper is as follows. In the next section we give an overview of the topic. Section 3 describes the Finnish unemployment compensation system. In Section 4 we give details of the data and report some descriptive statistics. Section 5 discusses the econometric methods. Section 6 reports our empirical findings. The final section concludes.

2 UI duration in theory and practice

The analysis of unemployment insurance has been an active subject of both theoretical and empirical work for the past two or three decades. The key predictions of theory have found support from the microeconometric analysis of unemployment duration data. Here we briefly discuss the likely effects of maximum UI duration on exits out of unemployment among UI recipients, and give a selective survey of relevant empirical evidence. Needless to say, our focus is highly restrictive, as the existence of UI can affect labour market outcomes in a variety of ways that depend on the institutional features of the UI system under consideration. Atkinson and Micklewright (1991) and Holmlund (1998) discuss the topic in a much wider context.

2.1 Theoretical insights and evaluation issues

Dynamic models of job search have shed light on the ways in which UI can affect unemployment duration through the reservation wage and search effort. Mortensen's (1977) model is the seminal contribution of this branch of the literature. In his model, eligibility for UI requires some previous work experience and UI benefits can be received for a fixed period. Workers are either employed or unemployed. When unemployed, the worker chooses optimal search effort and samples job offers from a known distribution of wage offers using the reservation wage strategy. When employed, the worker faces the risk of becoming unemployed. The transition rate from unemployment to employment increases with search effort (as the arrival rate of job offers increases) and decreases with the reservation wage (as the probability that a received offer is acceptable declines). As the unemployed worker approaches the time when UI benefits will expire, his search effort increases and the reservation wage decreases. After the exhaustion of UI benefits, the worker faces a stationary environment, and hence his search effort and reservation wage do not change anymore. The employment hazard therefore increases up to the point of benefit exhaustion and remains constant thereafter.[2]

The longer entitlement period increases not only the value of being unemployed but also the value of becoming employed with a given wage rate, as unemployment spells in the future will be better compensated for. The relative importance of this second effect, known as the "entitlement effect", increases as the day of benefit exhaustion comes closer. As a result, an increase in the maximum duration of UI benefits reduces the employment hazard over the forepart of the unemployment period but increases it close to and beyond the exhaustion point. In other words, the effect on the employment hazard is predicted to

[2] Mortensen derived a number of other predictions as well. First, the existence of UI benefits increases the transition rates of the unemployed who are not currently eligible for UI benefits. This is so because accepting a job offer qualifies for UI benefits in the future. Second, the time profile of the employment hazard among UI recipients depends on the replacement rate of UI benefits. An increase in the replacement rate increases both the value of search unemployment and the value of accepting a job offer, where the former becomes less important as the exhaustion time gets closer. As a consequence, the higher replacement rate is associated with a steeper hazard rate over the compensated part of the unemployment spell.

change over the course of the unemployment spell, potentially reversing its sign at some point. In the empirical analysis this possibility is sometimes ruled out a priori by imposing the restriction that changes in the length of the entitlement period may lead to level shifts in the underlying hazard function but cannot affect its shape (e.g. Hunt, 1995; and Lalive and Zweimüller, 2004).

Although Mortensen (1977) sophistically incorporates the key institutional features of UI into his model, some important aspects of the real-world labour market are abstracted away. The existence of large worker flows between unemployment, employment, inactivity, and various labour market programmes brings into question the extent to which the predictions of the two-state model are carried over to labour markets where workers can escape unemployment via various routes. Obviously, the generosity of unemployment compensation affects the attractiveness of participation in the labour market. Exhaustion of UI benefits in particular may encourage an unemployed worker to withdraw from the labour force rather than to accept lower wage offers. Moreover, the labour market policy in many countries, especially in the Nordic countries, involves a heavy stress on various labour market programmes. Such programmes are often targeted at the long-term unemployed who are at risk of benefit exhaustion. Participation may be associated with a rather high compensation level (e.g. relief work), may postpone the exhaustion day of UI benefits, or may even provide a way of regaining eligibility for UI. Using a search model, Carling et al. (1996) illustrate how the existence of labour market programmes can mitigate the incentive effects of fixed UI duration on the job-finding rate.

Much of the empirical analysis of the potential UI duration has focused on detecting spikes in the transition rates around exhaustion time. The finding of an increasing exit rate in the vicinity of UI benefit exhaustion is consistent with theory but has only a limited use from the policy point of view. It does not give much guidance as to how transition rates would change if the entitlement periods were subject to change. To address these questions, one needs some variation in the maximum length of UI benefits.[3] In an ideal setting one would have an inflow sample of unemployed workers who were randomly allocated into groups with varying lengths of UI entitlement periods. This sort of randomized data is not available anywhere, and therefore other sources of variation must be exploited in practice. The extent to which the available data provide variability in the benefit duration that is independent of unobserved determinants of unemployment duration is a crucial identification issue.

In many countries, the maximum entitlement length of UI benefits is partly determined by past work history (e.g. the US, Canada, and the Netherlands).[4] As a consequence,

[3] Even if the existence of a spike around benefit exhaustion were the only matter of interest, variation in the benefit duration would be useful to disentangle the exhaustion effects from the duration dependence.

[4] Where the entitlement period is the same for all UI recipients, the hazard rates have sometimes been compared between UI recipients and non-recipients (e.g. Carling et al., 1996). This is problematic, since workers who are not qualified for UI do not generally serve as a valid comparison group. Eligibility for UI benefit typically requires, among other things, previous work experience and hence non-recipients are often a highly selected group of labour market entrants and individuals with an unstable labour market

individuals with better track records in the labour market will be entitled to longer UI benefits periods. This suggests an endogeneity problem at the individual level if the past work history that determines the length of the entitlement period is related to unobserved characteristics that also affect the job-finding rate. Moreover, policy changes may cause changes in the entitlement periods across all UI recipients (e.g. US benefit extension programmes and reforms in Norway) or within some groups (e.g. reforms in Germany and Austria) over time. These policy changes are typically triggered by economic downturns, however. If the benefits are extended during periods of high unemployment, longer UI periods will be available at times when finding a new job is more difficult due to the demand constraint. This leads to an aggregate level endogeneity problem, which may be difficult to overcome unless a suitable control group unaffected by the reform is available to control for business cycle effects. If such a control group is available or the policy change is "exogenous" in the sense that it takes place over a period of stable economic conditions, the policy reforms are potentially very useful for identification purposes.

In the case of multiple exit routes, the expected duration of the unemployment spell and the shape of the employment hazard are not the only questions of interest. From the policy point of view, the likelihood of escaping unemployment through a given route is an equally important question. The goal of policy makers is typically to induce the unemployed to find an acceptable job within a reasonable amount of time, transitions out of the labour force and into labour market programmes being less desired outcomes. The maximum duration of UI benefits is likely to affect all hazard rates, not only the employment hazard. This suggests that a proper way of analysing the role of the UI entitlement period requires simultaneous accounting of all cause-specific hazards. We illustrate this point with a numerical example.

Suppose that when a transition out of unemployment occurs, it can be the result of an exit to employment (e), active labour market programme (p), or out of the labour force (o). The hazard function for destination $k \in \{e, p, o\}$ is denoted by $\theta_k(t)$. The cumulative distribution function (CDF) of unemployment duration is given by

$$F(t) \equiv \Pr(T \leq t) = 1 - \exp\left\{-\int_0^t \theta(u)du\right\},$$

where $\theta(u) \equiv \theta_e(u) + \theta_p(u) + \theta_o(u)$ denotes the overall hazard rate out of unemployment. The likelihood of entering employment within a short time interval $(t, t + dt]$ equals $\theta_e(t)[1 - F(t)]\, dt$, which depends on all cause-specific hazard functions through $F(t)$. It is important to recognize that a policy reform that causes a uniform increase in the employment hazard does not necessarily imply a higher cumulative or overall probability of employment. The cumulative incidence function (CIF) for destination $k \in \{e, p, o\}$ is defined as

$$F_k(t) \equiv \Pr(T \leq t, K = k) = \int_0^t \theta_k(u)[1 - F(u)]\, du,$$

history.

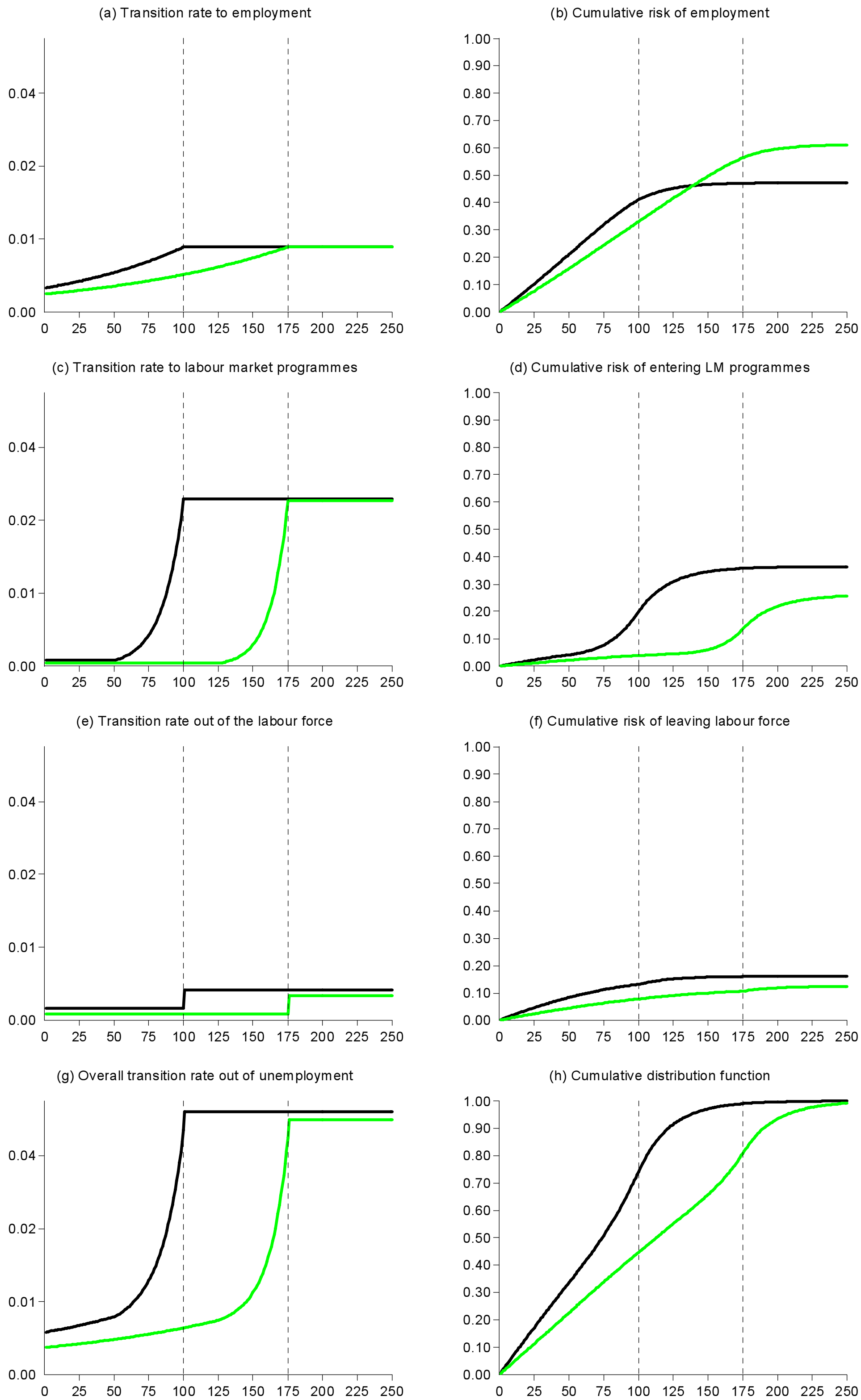
which gives the probability of entering state k by time t (see e.g. Gaynor et al., 1993). It is easy to see that $F(t) = F_e(t) + F_p(t) + F_o(t)$. Moreover, the probability of ever escaping unemployment to state k is given by $\Pr(K = k) = \lim_{t\to\infty} F_k(t)$.

Consider two groups of the unemployed who are otherwise identical but differ in the length of the UI entitlement period. In the first group, UI benefits expire after 100 weeks of unemployment, and in the other group after 175 weeks. Hypothetical hazard functions under the two UI schemes are shown in Figure 1. Although these hazard functions are not derived from any particular theoretical model, their shapes are intuitively reasonable and do not contradict the implications of the search models discussed above. The employment hazard increases up to the point of benefit exhaustion. The hazard rate for transitions to ALMPs is close to zero over the early stages of unemployment but begins to raise rapidly 50 weeks prior to the time when benefits will expire. There is a discrete upward shift in the hazard rate out of the labour force at the week followed by benefit exhaustion. The longer entitlement period leads to lower hazard rates for transitions to ALMPs and out of the labour force, with the level shifts occurring at a later point in the unemployment spell. The employment hazards under both UI regimes are assumed to converge at the same level. This is not strictly consistent with the prediction of Mortensen's model that the hazard rate associated with the longer UI entitlement period should end up at a higher level. The entitlement effect is neglected here to emphasize our main point.

The longer entitlement period is associated with a uniformly lower overall hazard rate (see Figure 1g), which implies a longer average spell duration. Without information on the states occupied after unemployment, there is not much more to be said about the effects of the entitlement period length. In particular, we are unable to assess how the overall incidence of employment changes with the length of the entitlement period. One might guess that it must be lower in the case of the longer entitlement period. From Figure 1b we see that this may not be the case, however. The fraction of the unemployed who will eventually exit to employment is higher among those who are entitled to the longer benefit period. Since the employment hazard of this group never exceeds that of workers with the shorter entitlement period, this outcome is driven by discrepancies in the hazard rates out of the labour force and into ALMPs.

To summarize, competing risks analysis is potentially very useful in evaluating the effects of the length of the entitlement period. This is especially true in the cases where ALMPs play an important role, as in the Nordic countries. Second, the cumulative incidence functions can be used to summarize information from all cause-specific hazard functions in a policy-relevant way. Third, reliable analysis requires independent variation in the entitlement periods across individuals. Fourth, a flexible model specification should be adopted, as no a priori parametric restrictions for the effects of the entitlement period on the hazard functions can be derived from theoretical models.

Figure 1: Entitlement period length and transitions out of unemployment (Note: Black lines correspond to the entitlement period of 100 weeks, and green lines to that of 175 weeks)



2.2 Empirical evidence

In the US the maximum duration of UI benefits varies for various reasons: the length of regular benefits varies across states, the entitlement period may depend on the individual's work history, and benefit periods are occasionally extended (at federal or state level) in respond to slackness in the labour market. Meyer (1990) and Katz and Meyer (1990) exploit these sources of variation for identification. They find that the hazard rate out of compensated unemployment increases sharply in the last few weeks of UI benefit eligibility. Moreover, when the benefit entitlement period is extended, the hazard rate also exhibits a peak around the time when benefits were previously expected to lapse. This finding may indicate that some firms recall their workers who are temporary laid-off at the time of the original benefit exhaustion according to a preplanned recall policy. In other words, during business downturns employers attempt to retain their skilled workers but forward part of the bill to the taxpayers via the UI system. Alternatively, some individuals eligible for extended benefits may fail to claim them because they are not aware of such a possibility or because they managed to arrange the start of a new job at the time when UI benefits were originally expected to run out. Furthermore, the simulations of Katz and Meyer (1990) suggest that a 13-week benefit extension (from 26 to 39 weeks) increases the mean spell duration of compensated unemployment by slightly over 2 weeks.

Card and Levine (2000) take advantage of a politically motivated programme that provided up to 13 weeks of extended benefits for those who exhausted their regular UI benefits in the state of New Jersey. The New Jersey Extended Benefit (NJEB) programme was in effect for a limited period of 6 months. Since the NJEB programme emerged as a result of a unique legislative episode that was unrelated to changes in economic conditions, it caused an exogenous change in the entitlement period. Card and Levine (2000) find only a small increase in the fraction of UI recipients who remained unemployed until the exhaustion of their regular benefits. The authors argue that such a moderate effect is attributable to the short-term nature of the NJEB programme, since many of those affected had been unemployed for a while before the programme was introduced. Their simulations suggest stronger long-term effects: Had the programme affected UI recipients from the beginning of their spell, the average duration of regular benefits would have increased by one week and the fraction exhausting the regular benefits would have risen by 7 percentage points.

The two major data limitations of these US studies are worth emphasizing. First, only unemployment spells covered by the UI system are observed. All spells ongoing at the time of benefit exhaustion are censored, and hence nothing can be said about the transition rate beyond the exhaustion point. Second, the data do not permit a competing risks analysis because it is not possible to distinguish whether a completed spell ended through recall, the taking of a new job, withdrawal from the labour force, or exit to uncompensated unemployment. That is, what has been analyzed is the overall rate out of compensated unemployment up to the point of benefit exhaustion.[5]

Consistent with the US evidence, Ham and Rea (1987) find the conditional probability of leaving unemployment through the finding of a new job to increase just prior to the exhaustion of UI benefits in Canada.[6] Lindeboom and Theeuwes (1993) provide evidence that the overall hazard rate out of unemployment increases sharply as the entitlement period comes to an end in the Netherlands.[7] By analysing the number of search contacts over the course of the unemployment spell, they conclude that the benefit exhaustion effect is mainly due to the declining reservation wage. Micklewright and Nagy (1999) recover a clear spike in the job-finding rate around the time of UI exhaustion in Hungary. They do not find evidence that the transition rate is strongly affected by the probability of being eligible for means-tested social assistance after UI exhaustion. Jenkins and García-Serrano (2004) detect only a moderate increase in the hazard to employment prior to the exhaustion of UI benefits in Spain. Somewhat surprisingly, longer entitlement periods appear to be associated with higher re-employment hazards (Figure 1, p. 255). While this observation was not discussed by the authors, it may be an indication of the endogeneity problem, since there was no clear exogenous variation in entitlement periods in the Spanish data. The length of the entitlement period was completely determined by the number of months for which contributions had been made over the 48-month period prior to unemployment.

Evidence on exhaustion effects from the Nordic countries, where benefit levels are higher, entitlement periods longer, and labour market programmes play a more important role, is less convincing. Carling et al. (1996) find only very weak evidence of an increase in the job-finding rate of UI recipients compared with that of non-recipients around the exhaustion time of UI benefits at 60 weeks in Sweden. There were no significant differences in the hazard rates out of the labour force between the two groups. By contrast, the transition rate to ALMPs among UI recipients increases sharply near the time of UI benefit exhaustion, being almost five times higher than the transition rate of non-recipients immediately after UI benefits have run out. These findings are intuitively reasonable in view of the fact that Swedish labour market policy involves a right to a temporary public job or training course for the unemployed whose benefits lapse.

In Norway, until 1991, UI benefits expired temporarily for 26 weeks after 80 weeks of continuous unemployment. This was then followed by a second covered 80-week unemploy-

[5] Katz and Meyer (1990) provide some indirect evidence that much of the observed effect of the potential duration of UI is expected to arise from differences in recall and job-finding hazards. Using complementary data from the Panel Study of Income Dynamics, they compare the distributions of unemployment spells between UI recipients and non-recipients. This source of data provides information on exit states and the eligibility status of UI benefits but lacks information on the benefit level and length of entitlement periods for UI recipients. Katz and Meyer (1990) find substantial peaks in both the empirical recall rate and the new job-finding rate around the durations when UI benefits are "likely to lapse" for UI recipients but not for non-recipients. These descriptive findings are consistent with a more formal analysis by Han and Hausman (1990) who apply a competing risks duration model to the same data.

[6] The length of the UI entitlement period is determined by the past work history and the local unemployment rate. The mean entitlement period in the data is 36 weeks.

[7] The length of the entitlement period depends on the past work history, the maximum duration being 26 weeks.

ment period with somewhat reduced UI benefits.[8] As a response to increasing long-term unemployment, the intervening period without benefits was first shortened to 13 weeks in 1991 and one year later was abolished entirely. Bratberg and Vaage (2000) and Røed and Zhang (2003) document rises in the transition rate out of unemployment just prior to temporary and permanent benefit exhaustion. These effects are stronger for women than for men (Røed and Zhang, 2003). The reforms were followed by a drop in the employment hazard over the early stages of the unemployment spell but there is no evidence of increases in the employment hazard around benefit exhaustion (80 weeks) prior to the reforms, nor after the reforms (Bratberg and Vaage, 2000). Thus, the observed peaks in the overall hazard rate around (at least temporary) benefit exhaustion are due to the increasing outflow to other end-states than employment.

In the UK, the level of UI benefits does not depend on previous earnings but is a flat rate. If UI benefits expire, the individual can claim means-tested social assistance that is comparable to UI payments and potentially available for an indefinite period. Therefore, there is practically no fall in the level of unemployment compensation when the individual passes from UI benefits to social assistance. Stancanelli (1999) exploits this feature of the UK system in detecting benefit exhaustion effects. She compares hazard rates to full-time employment between those unemployed that should expect their benefits to exhaust and those whose benefit level is unlikely to change due to eligibility for social assistance.[9] Although a larger spike near the time of benefit exhaustion at 52 weeks exists for those whose benefits are going to expire, Stancanelli (1999) does not find statistically significant differences in the hazard functions for the two groups.

Hunt (1995) evaluates the impact of a series of UI entitlement extensions for older workers with much work experience in the former West Germany. These policy measures were motivated by concern about the increasing unemployment rate and long average spell duration of older workers. The number of additional months for UI benefits was tied to the amount of past work experience. As a result of an increase in the maximum UI duration from 12 to 22 months, the hazard rates for workers aged 44 to 48 were found to fall by 46% for transitions to employment and by 63% for transitions out of the labour force. Among older workers between the ages of 49 and 57, only the hazard rate out of the labour force fell by 56% in response to an increase in the entitlement period from 12 to 32 months.

A reform similar to the German case took place in Austria around the turn of the 1990s. The reform was a response to a privatization process of nationalized firms that led to mass layoffs due to plant closures and downsizing especially in the steel industry (Winter-Ebmer, 1998, and Lalive and Zweimüller, 2004). The maximum entitlement period of UI benefits

[8] UI benefits depend on past earnings and the system is universal covering all workers with earnings above a minimum level.

[9] Eligibility for social assistance depends on the level of household savings and spouse's earnings. Since household savings tend to be correlated with the work history of the unemployed individual, the same kind of endogeneity issues arise as in the UI systems where the length of the entitlement period is directly tied to past work history.

was raised from 52 to 209 weeks for workers aged 50 or more with experience above a certain threshold living in certain regions of the country. In terms of pre-reform labour market conditions, the counties selected for the extended benefit periods were similar to other counties but characterized by a larger employment share in the steel industry and hence expected to be hit by a negative demand shock. Winter-Ebmer (1998) finds a decrease in the exit rate to a new job for men, but no notable effects on transitions to recall employment, retirement, or non-participation for either gender.[10] The resulting increase of 5 weeks in men's unemployment duration is very low, given the substantial increase in the maximum benefit duration. Lalive and Zweimüller (2004) exploit variation in the programme eligibility across regions and over time in an attempt to account for policy endogeneity. They conclude that the increase in the UI entitlement period from 30 to 209 weeks declined men's transition rate to employment by 17%.

The focus of Finnish empirical literature has been on the effects of benefit levels (e.g. Kettunen, 1993, Lilja, 1993, Kyyrä, 1999, Holm et al., 1999). There are no studies with an emphasis on the role of the entitlement period length. In the late 1980s the Finnish UI system included a cut in the level of UI benefits. Within the period from April 1987 to July 1989, UI benefits were reduced by 12.5% after 40 weeks of unemployment (and finally expired after 100 weeks). Exploiting the removal of the cut in 1989, Uusitalo and Moisala (2003) find an increase of some 8% in the employment hazard when the cut kicks in, but this effect was not statistically significant. Empirical hazard rates for UI recipients aged 52 or less reported in Koskela and Uusitalo (2003) exhibit no peaks for transitions to employment, but a level shift at one year that is followed by an additional peak after two years for transitions to labour market programmes. These descriptive observations are in line with evidence from Sweden and Norway. The absence of significant increases in the employment hazard around the exhaustion of UI benefits in the Nordic countries is perhaps attributed to relatively long entitlement periods, the availability of welfare payments after UI benefit exhaustion, and easy access to active labour market programmes.

[10] A fully parametric Weibull model was adopted, so that all hazard functions were restricted to being constant, monotonically increasing, or monotonically decreasing.

3 The institutional framework in Finland

3.1 Unemployment compensation

The Finnish compensation system distinguishes between the basic unemployment allowance, earnings-related UI benefit, and labour market support. Earnings-related UI benefit is received by workers who have been working and contributing insurance payments to an unemployment fund for at least 10 months during the two years prior to unemployment.[11] Those fulfilling the employment criteria of having worked at least ten months but not belonging to an unemployment fund are eligible only for the basic allowance (which amounted to 115 euros per week in 2003). Workers not meeting the employment criteria or whose entitlement period has been exhausted can claim for labour market support, which is viewed as a minimum income for the long-term unemployed and those just entering the labour market. The maximum benefit level for labour market support equals the basic unemployment allowance, but it is means-tested against household income.

Basic unemployment allowance and earnings-related unemployment insurance can be received for a maximum of 500 working days (i.e. approximately two calendar years of continuous unemployment), but there is an exception for the older unemployed. The unemployed reaching the age of 57 (55 until 1.1.1997) before running out of benefits (i.e. reaching the maximum duration of 500 working days) are allowed to collect UI benefits until the age of 60, after which they are eligible for unemployment pension. Effectively this means that a newly unemployed individual must be at least 55 years and 1 month of age (53 years and 1 month prior to 1997) when becoming unemployed to be eligible for the extended UI benefits.

The compensation level of unemployment pension is determined by previous earnings but over a longer period than that used for determining UI benefits.[12] At the age of 65 unemployment pension is transformed into the conventional old-age pension. The combination of the extended UI entitlement period and unemployment pension has become known as the "unemployment tunnel". It secures income for the elderly unemployed until old-age retirement, covering periods that could be as long as 10 years.

3.2 Reforms of the unemployment tunnel scheme

The age threshold for the extended UI benefit duration was raised by two years, from 53 to 55, at the beginning of 1997. This was part of a larger reform package that appeared, first, as a white paper in September 1995, was officially proposed by the government in May 1996 and passed as a law by the Parliament in September 1996. Effectively the reform took place on January 1st 1997. However, according to a protection clause, the

[11] The unemployment funds are closely related to labour unions. The fund membership is voluntary and workers can join the fund without joining the union.

[12] Rantala (2003) finds that transitions from unemployment to unemployment pension were followed by an average increase of 16% in the gross compensation level in 1996 and 1997. For workers with a very steeply increasing earnings profile before job loss UI benefits can be higher than unemployment pension.

prior threshold was applied to individuals born before 1944 who were made redundant or resigned before June 1996 and were eligible for unemployment benefits on January 1st 1997 (or had received unemployment benefits for at least 100 days in 1996). As a result, individuals aged 53-54 becoming unemployed before June 1996 were eligible for the extension, but those becoming unemployed in 1997 no longer were. Because of the period of notice which can be 6 months at maximum, workers aged 53-54 who entered unemployment in 1996 between June and December may or may not be eligible for the extension, depending on the exact timing of resignation or dismissal.

The aim of this 1997 reform package was to cut unemployment expenditures, to improve employment incentives amongst the unemployed, and to close certain loopholes in the system. The unemployment pension scheme was perceived as a loophole, given that some companies had exploited the existing system when downsizing and it had turned into a somewhat generally acceptable early retirement scheme. Hence, the government wanted to phase the system out.[13] Since there is no reason to believe that the age threshold was raised in response to a change in relative labour market conditions for the older workers, our analysis should not be subject to endogenous policy bias.

We will also have to take into consideration another less obvious and smaller reform executed in 1996,[14] which cut the benefit levels for various early retirement schemes, including unemployment pensions.[15] Individuals becoming unemployed at January 1st 1996 or thereafter face the lower pension rates. However, an exemption was made for persons born before 1943 who were registered as unemployed precisely on the first working day of 1996.[16] They remain covered by the prior rules in case of early retirement (regardless of when in the future the event of retirement takes place).

3.3 Active labour market programmes

Although finding suitable employment for the unemployed is the main goal of policy makers we should bear in mind that in the Finnish case active labour market programmes are a significant route out of unemployment and, thus, cannot be ignored. However, the elderly unemployed were not specifically targeted with active measures during the 1990s, since unemployment amongst them was not foreseen as the most acute problem. In general, participation in ALMPs tends to be lowest in the oldest cohorts, ranging from some 25% of all unemployed in the youngest cohorts to some 17% in the eldest cohorts (Aho, 2005).

[13] The age threshold for the extended period of unemployment benefits was raised by two additional years in 2005. Over the period 2009-2014 the unemployment pension scheme will be gradually abolished and replaced with additional unemployment benefit days for the elderly until the age of 65.

[14] In what follows, we will refer to the reforms separately as the 1996 reform and the 1997 reform.

[15] The accrual rate for projected pensionable service was decreased from 1.5% per year to 1.2% per year while aged under 60 and to 0.8% per year while aged 60 to 65. (Projected pensionable service is the time period from the pension contingency to the age of 65. The projected pensionable service increases the amount of pension, since it is calculated as if the person had continued working until the age of 65. A pension accrues according to a lower accrual rate, however.)

[16] And it suffices to be registered as unemployed just on that given day.

The programmes are offered to the unemployed by the employment office, but an unemployed individual can also apply for these measures unprompted.[17] During the placement period in subsidised employment a participant receives the prevailing market wage set in collective agreements. Subsidised employment contracts in the private sector are expected to be continuous and may thus result in the unemployed individual renewing eligibility for earnings-related benefits by fulfilling the 10 months' time-at-work condition. Job placements in the public sector tend to be fixed-term and shorter by nature (often 6 months) and, thus, a single such period does not usually suffice to renew eligibility.[18] During participation in labour market training the participants receive a sum equalling their unemployment compensation together with a daily allowance for maintenance and possibly for accommodation. The training period neither consumes nor accumulates the UI entitlement period left.

While for some unemployed people ALMPs may truly be a way of obtaining contacts and skills leading to a secure job, there is also the possibility that others simply comply to participate in order to prolong the exhaustion of benefits or to regain foregone eligibility. If the latter should apply, then the individuals in our sample eligible for the extended UI benefits have little incentive to participate in these programmes, and we might also expect this to show up in our results.

[17] Should an unemployed individual refuse to participate in these measures without a valid reason if such a measure is offered to him, he will be deprived of unemployment compensation for a period of 60 days.

[18] According to Virjo et al. (2006) in 1996 and 1998 some 15 percent of all the unemployed were receiving earnings-related benefits owing at least in part to participation in the programmes.

4 Data and descriptive statistics

Our data were drawn from the records of the Employment Statistics (ES) database of Statistics Finland. Since 1987 the ES database has been updated regularly by merging information according to unique individual identity numbers from over 20 administrative registers. In brief, this database basically follows the entire Finnish population over time and across different labour market states. For research purposes the records of the ES database are currently available until the year 2000. In the subsequent analysis of unemployment durations, we focus on private sector workers aged 50-54 becoming unemployed between 1995 and 1998, having been continuously employed for at least one calendar year prior to the year of job loss, and being thus likely eligible for earnings-related UI benefits. We further restrict the sample by including only those with UI benefits higher than 14 euros per day,[19] which corresponds to the amount of basic unemployment allowance at the time.

4.1 Selection of treatment, comparison, and control groups

The age threshold for the newly unemployed to be eligible for the extended UI benefit was increased from 53 to 55 years in 1997. This effectively reduced the unlimited entitlement period to two years among the group aged 53-54 at the time of job loss, providing a quasi-experimental setting for studying the relationship between the length of the UI entitlement period and transitions out of unemployment. Workers aged 53-54 who entered unemployment in 1995 or 1996 are referred to as the *treatment* group since their behaviour is affected by eligibility for the extended UI benefits. Workers aged 53-54 who became unemployed in 1997 or 1998 serve as the *comparison* group. Apart from eligibility for the extended UI benefits and the timing of unemployment entry, this group is assumed to be highly similar to the treatment group. Workers aged 50-52 are labelled as the *control* group, as they are used to identify the business cycle effect. Due to the anticipatory effects of the reforms, we will exclude certain groups from our analysis.

First, in anticipation of the 1997 reform the flow into unemployment increased in the age group 53-54 at the end of 1996 (Kyyrä and Wilke, 2007). Hence, individuals aged 53-54 becoming unemployed between June 1996 and December 1996 are excluded from our analysis.[20] Second, anticipation of the 1996 reform led to an excess flow into unemployment at the end of 1995 among those benefiting from the protection clause, and thus individuals aged 53-54 becoming unemployed between October 1995 and January 5th 1996 are also excluded from this analysis. Third, due to evident seasonal variation in unemployment, we must also exclude individuals in our control group (aged 50-52) becoming unemployed during the same time periods.

[19] And less than 75 euros per day.

[20] Moreover, the eligibility status for these workers is not clear. While most of them are likely to be eligible for the extension due to the notice period, we cannot be sure, as the day of resignation or dismissal is not available in our data.

In addition, we exclude individuals born in 1942 or 1943 becoming unemployed in 1997 or 1998. These individuals had already turned 53 prior to the 1997 reform and thus could have got the extension in 1995 or 1996. The option was briefly taken away from them in 1997, but after turning 55 they were eligible for the extension again. Individuals who thought they were likely to wind up unemployed anyway may have profited from choosing to become unemployed sooner, i.e. prior to the 1997 reform as opposed to after it. Thus, those in these age cohorts becoming unemployed after the reform are likely to be a selected group and should be excluded.

Kyyrä and Wilke (2007), using the same data set as ours, describe in detail the post-unemployment destinations of the unemployed individuals by age, year, and month of unemployment entry. Their analysis clearly demonstrates that the reforms at the start of 1996 and 1997 did not affect the labour market behaviour of the 50-52 year olds, while the 53-54 year olds showed apparent divergence in their post-unemployment destination states before and after the reforms. Individuals in this age group becoming unemployed at the end of 1995 or 1996 are far more likely to retire from the labour market and far less likely to find employment than those becoming unemployed earlier the same year. This further corroborates our decision to exclude these individuals from the sample. By estimating various logit models for unemployment risk, Kyyrä and Wilke (2007) also formally test that the time intervals now included in this analysis are free from anticipatory effects.[21] This being said, we now feel we have a representative sample free from any anticipatory or otherwise distorting effects.

4.2 Descriptive statistics

For each worker in the data we observe the length of the unemployment spell (in days), exit destination, and a set of control variables. All unemployment spells that continue beyond the end of 2000 are treated as censored. Sample statistics by age group and by year of entering unemployment are shown in Table 1. The average duration of an unemployment spell is roughly one year for the 50 to 52-year-olds entering unemployment in 1995 – 1998, as well as for the 53 to 54-year-olds entering unemployment in 1997 and 1998, but it is much longer for the older cohort entering unemployment in 1995 and 1996. When considering only completed spells eventually ending in employment, differences between the groups diminish drastically. This, according to our view, indicates that a large fraction of the older age group entering unemployment in 1995 or 1996 is not actually searching for a job but is instead passively waiting for access to retirement.

Given the age structure of our sample, it is not surprising to find that most of the individuals are married, nor that only a small share have a dependent child in the family. Overall, the differences in mean individual characteristics among the groups remain small. The occupational distribution in both age groups is fairly similar across entering years.

[21] Our estimation sample is otherwise similar to that of Kyyrä and Wilke (2007) but we exclude individuals aged 55-57 as their behaviour is evidently also affected by other than business-cycle-related factors.

Table 1: Sample statistics by age and time of entering unemployment

	1995 Jan to Sep		1996 Jan to May		1997/1998 Jan to Dec	
	50-52	53-54	50-52	53-54	50-52	53-54
Unempl. duration, days	368	1252	290	968	284	299
	(382)	(866)	(350)	(767)	(291)	(265)
Duration to employment	190	215	151	144	151	155
	(214)	(279)	(175)	(187)	(176)	(164)
Female	.475	.553	.394	.455	.434	.476
Married	.687	.724	.700	.712	.685	.677
Female × married	.304	.388	.259	.301	.295	.314
Dependent child	.218	.099	.195	.098	.205	.100
Swedish-speaking	.061	.045	.051	.043	.040	.062
Post-secondary education	.083	.049	.061	.055	.059	.049
Occupation:						
Commercial	.183	.157	.131	.154	.156	.180
Technical	.083	.071	.081	.065	.078	.070
Humanist	.022	.021	.020	.015	.021	.029
Health care	.019	.014	.016	.017	.020	.020
Clerical	.188	.261	.149	.158	.151	.157
Agricultural	.026	.011	.076	.046	.016	.021
Transportation	.053	.057	.061	.060	.065	.061
Industrial	.324	.315	.382	.396	.395	.371
Services	.092	.091	.078	.077	.092	.085
Not classified	.009	.003	.005	.012	.006	.006
Firm size:						
50 employees or less	.435	.297	.474	.390	.470	.473
51-500 employees	.293	.313	.253	.293	.270	.271
Over 500 employees	.272	.390	.272	.317	.260	.256
Tenure $\geq$ 4 years	.460	.497	.437	.450	.437	.457
Unemployed in early 1990s	.383	.287	.570	.510	.549	.518
Past recall in early 1990s	.076	.062	.186	.158	.176	.157
Ln (UI benefits)	3.30	3.31	3.32	3.32	3.30	3.28
	(.33)	(.27)	(.34)	(.28)	(.33)	(.33)
Number of observations	935	1046	837	584	2939	657

Notes: Sample standard deviations for continuous variables in parentheses. The spells of those born in 1942 or 1943 for 53-54 years old excluded from 1997/1998 data.

Not surprisingly, the elderly unemployed are rather poorly educated on average. Roughly half of the sample were previously unemployed in the 1990s. This is a rather large fraction despite the fact that the early 1990s was a period of sky-high unemployment. On the other hand, almost half of the sample have stayed with the same employer at least for the previous four years. There is practically no difference in the average level of earnings-related unemployment benefits received by the different groups.

Table 2 shows the numbers of individuals at risk of exiting unemployment and observed exits to different destinations by each duration interval for our treatment group (those with extended UI benefits) and the pooled control and comparison group (all those with the fixed two-year entitlement period). Only 60% of the individuals in our treatment group exit unemployment during the observation period, while almost everyone from the control

Table 2: Risk set and observed exits by duration intervals

	Treatment group[1)]				Control and comparison groups[2)]			
	Exits to			Risk	Exits to			Risk
Interval	Emp	ALMP	Out	set	Emp	ALMP	Out	set
(0,2] months	150	10	23	1630	1041	117	100	5368
(2,4] months	121	7	20	1447	674	118	98	4110
(4,6] months	41	10	15	1299	387	103	64	3220
(6,8] months	25	16	11	1233	202	123	48	2666
(8,10] months	32	6	14	1181	129	97	49	2293
(10,12] months	9	9	10	1129	107	121	49	2018
(12,14] months	13	3	21	1101	55	160	43	1741
(14,16] months	9	6	10	1064	63	159	43	1483
(16,18] months	12	9	8	1039	51	128	31	1218
(18,20] months	11	5	9	1010	37	98	35	1008
(20,22] months	2	6	8	985	27	99	43	838
(22,24] months	4	4	6	969	28	86	37	669
(24,30] months	8	17	20	955	44	162	56	518
(30,36] months	2	4	14	910	8	31	20	221
(36,48] months	3	8	33	890	4	12	21	135
(48,60] months	2	4	34	846	1	4	3	49
(60,63] months	0	0	42	547	0	1	0	18
(63,∞] months	1	0	111	502	0	1	0	17
Sum	445	124	409		2858	1620	740	
(%)	(27.3)	(7.6)	(25.1)		(53.2)	(30.2)	(13.8)	

1) Individuals aged 53-54 becoming unemployed in 1995 or 1996. 2) Individuals aged 50-52 becoming unemployed in 1995-1998 and individuals aged 53-54 in 1997-1998.

and comparison groups do exit. Moreover, the distribution of end-states is quite different in the two groups. Exits to employment and to ALMPs are observed only for some 27% and 8% within our treatment group with extended UI benefits. Among those with limited benefit duration these exits are observed for 53% and 30%, respectively. Without the risk of benefit exhaustion, there appears to be considerably fewer people exiting both to employment and ALMPs. On the other hand, considerably more people, 25%, with unlimited UI benefits withdraw from the labour market altogether by the end of the observation period.

Given that some 40% of individuals in the treatment group are still unemployed at the end of the observation period of 5 to 6 years of length, it seems clear that a substantial fraction of workers with extended UI benefits are no longer active in the labour market. Ervasti (2003) studies preferences for job search using a survey of 970 unemployed workers.[22] According to their search activity he classifies the unemployed into four groups, where the most passive group consists of those who do not look for a new job, nor even wish to return to employment. Members of this group are rather old, i.e. entitled to extended UI benefits, receiving high UI benefits, having high reservation wages, and suffering from deteriorated health. Consequently, accounting for this considerably lower level

[22] The survey was subject to a large response bias.

of activity does appear necessary here for the treatment group. This will be accomplished by using the split population model described below.

5 Econometric methods

We need to be aware of several complications in the empirical analysis. First, the unemployment spells start at different points in time, and, in particular, workers aged 53-54 under different UI schemes enter unemployment in different years. We need to find a way to separate the changes in the hazard function owing to different UI schemes from the effects of changing macro-economic conditions. For purposes of identifying the effects of the business cycle, we adopt a difference-in-differences type of setting where workers aged 50-52 serve as a control group. Second, prior survey evidence and our descriptive analysis suggest that a notable fraction of workers with extended UI benefits are passive and effectively withdrawn from the labour market. This issue is taken into account by allowing the transition probabilities to employment and ALMPs to be zero for some individuals in this group.

5.1 The split population model

Consider a worker who loses his job and becomes unemployed at time τ. The worker is followed until the termination of the unemployment spell or the end of the observation period $\overline{\tau}$ (i.e. the last day of 2000). The duration of the unemployment spell T is continuous. If a transition out of unemployment occurs within the observation period, it will be followed by employment (e), participation in an ALMP (p), or withdrawal from the labour market (o). The unemployment spell is right-censored if it continues beyond the observation period, in which case we know only that $T > \overline{\tau} - \tau$. We allow for a possibility that the worker chooses to withdraw from active labour market behaviour, in which case he does not look for employment, nor is he willing to participate in ALMPs. Instead he is passively waiting for an opportunity to escape the labour force via some early retirement scheme. For simplicity, we assume this choice is made at the moment of entering unemployment. We denote $\varepsilon = 1$ if the worker is still active, and $\varepsilon = 0$ otherwise. The value of this latent choice variable is not directly observed. The time path of explanatory variables for the hazard functions from τ to $\tau + t$ is denoted by $\mathbf{X}(t, \tau)$. This set includes variables that are either fixed, change with spell duration t (unemployment benefits), or change with calendar time $\tau + t$ (calendar-time dummies and local unemployment rate). The subset of variables that also affect the probability distribution of ε is denoted with $\mathbf{z}$.

The conditional hazard rates at spell duration t to employment and ALMPs are multiplicative in ε :

$$\theta_k(t\,|\mathbf{X}(t,\tau),\varepsilon) = \theta_k(t\,|\mathbf{X}(t,\tau))\,\varepsilon, \qquad k \in \{e,p\}\,.$$

We assume that the worker is not active with probability $p(\mathbf{z}) = \Pr(\varepsilon = 0\,|\mathbf{z})$, in which case $\theta_k(t\,|\mathbf{X}(t,\tau), \varepsilon = 0) = 0$ for $k \in \{e,p\}$. The transition rate out of the labour force is independent of ε, so that

$$\theta_o(t\,|\mathbf{X}(t,\tau),\varepsilon) = \theta_o(t\,|\mathbf{X}(t,\tau))\,.$$

The survivor function conditional on $\mathbf{X}(t,\tau)$ and ε is given by

$$\begin{aligned} S(t\,|\mathbf{X}(t,\tau),\varepsilon) &= \exp\left\{-\varepsilon\int_0^t \theta_e(u\,|\mathbf{X}(u,\tau))\,du - \varepsilon\int_0^t \theta_p(u\,|\mathbf{X}(u,\tau))\,du \right. \\ &\quad \left. - \int_0^t \theta_o(u\,|\mathbf{X}(u,\tau))\,du\right\}. \end{aligned} \tag{1}$$

Since the value of ε is observed only in some cases, we cannot always condition on it. By taking the expected value of $S(t\,|\mathbf{X}(t,\tau),\varepsilon)$ with respect to ε, we obtain

$$\begin{aligned} S(t\,|\mathbf{X}(t,\tau)) &= [1-p(\mathbf{z})]\exp\left\{-\int_0^t \sum_j \theta_j(u\,|\mathbf{X}(u,\tau))\,du\right\} \\ &\quad + p(\mathbf{z})\exp\left\{-\int_0^t \theta_o(u\,|\mathbf{X}(u,\tau))\,du\right\}, \end{aligned} \tag{2}$$

where $\sum_j$ denotes the sum over all the three possible exit destinations.

If a transition to employment or to an ALMP occurred at spell duration t, the values of both ε and T are observed along with the destination state. The likelihood contribution in this case is given by

$$[1-p(\mathbf{z})]\,\theta_k(t\,|\mathbf{X}(t,\tau))\,S(t\,|\mathbf{X}(t,\tau),\varepsilon=1)\,, \qquad k\in\{e,p\}\,.$$

When the unemployment spell ended via withdrawal from the labour force at spell duration t, we observe the spell length but not the value of ε, and hence the likelihood contribution is

$$\theta_o(t\,|\mathbf{X}(t,\tau))\,S(t\,|\mathbf{X}(t,\tau))\,.$$

If the unemployment spell is still in progress at time $\overline{\tau}$, the value of ε remains unobserved and all we know is that $T>\overline{\tau}-\tau$. The likelihood contribution in this censored case is

$$S(\overline{\tau}-\tau\,|\mathbf{X}(t,\tau))\,.$$

Putting these pieces together, we obtain the log-likelihood of the model:

$$\begin{aligned} \mathcal{L} &= \sum_i\sum_j d_{ij}\ln\theta_j(t_i\,|\mathbf{X}_i(t_i,\tau_i)) + \sum_i (1-d_{ie}-d_{ip})\ln S(t_i\,|\mathbf{X}_i(t_i,\tau_i)) \\ &\quad + \sum_i (d_{ie}+d_{ip})\left[\ln[1-p(\mathbf{z}_i)] + \ln S(t_i\,|\mathbf{X}_i(t_i,\tau_i),\varepsilon_i=1)\right], \end{aligned} \tag{3}$$

where $\sum_i$ denotes the sum over all individuals in the sample; $\sum_j$ denotes the sum over all the three possible exit destinations; $d_{ij}=1$ if individual i exited to destination $j\in\{e,p,o\}$, and $d_{ij}=0$ otherwise; and for censored observations $t_i=\overline{\tau}-\tau_i$ and $\sum_j d_{ij}=0$. The log-likelihood function (3) is maximised with respect to the unknown determinants of $\theta_e(\cdot)$, $\theta_p(\cdot)$, $\theta_o(\cdot)$, and $p(\cdot)$.

5.2 Parametrization

We adopt a step-function approximation to the cause-specific hazard functions in a continuous time framework. Compared with the grouped data analysis, this specification is equally flexible but also allows us to exploit the variation in the observed durations between the spells that fall into the same time interval. The time axis for the length of unemployment spells is divided into M intervals as $(c_{m-1}, c_m]$, $m = 1, 2, ..., M$, with $c_0 \equiv 0$ and $c_M \equiv \infty$. Although any duration dependence can be approximated arbitrarily closely by increasing the number of time intervals, this is not possible in practice because of a finite sample size. We set the length of the first 12 intervals to two months. These are followed by two intervals of six months, two intervals of twelve months, one interval of three months, and the open-ended interval (see Table 2). We will impose some group-specific restrictions on the shape of hazard functions on the basis of observations available for estimation.

We assume the following hazard function for destination $k \in \{e, p, o\}$ at spell duration $t \in (c_{m-1}, c_m]$:

$$\theta_k(t\,|\mathbf{X}(t,\tau)) = \exp\left\{\lambda_k^m + \eta_k^m D + \alpha_k G_{50-52} + \gamma_k Y_{53-54\times 96} + \mathbf{x}_m'\boldsymbol{\beta}_k\right\},$$

where D is the dummy variable for the entitlement period of two years (i.e. 53-54 years old who entered unemployment in 1997-1998 and 50-52 years old in all entry years), G_{50-52} is the dummy variable for workers aged 50-52, $Y_{53-54\times 96}$ is the dummy variable for workers aged 53-54 whose unemployment started in 1996, and $\mathbf{x}_m$ is a vector of other covariate values for the mth duration interval. The hazard functions are constant within each interval but vary across intervals owing to time-varying parameters (λ_k^m and η_k^m) and time-varying covariates (included in $\mathbf{x}_m$). The time-varying covariates include unemployment compensation, local unemployment rate, and year and quarter dummies. Among individuals under the conventional UI scheme UI benefits lapse after 24 months of unemployment and are followed by labour market support. The local unemployment rate, year dummies, and quarter dummies are related to calendar time, not to the elapsed duration of unemployment. These variables control for changing labour demand conditions over time and across regions.[23]

For destination k the shape of the hazard function for the treatment group (53-54 years old in 1995 and 1996) is modelled with a set of dummy variables for the log baseline hazard function, λ_k^m, $m = 1, 2, ..., M$. Given a relatively small number of observed exits for this group (see Table 2), we need to impose some equality constraints on λ_k^m for subsequent intervals. Within the treatment group the hazard function is allowed to differ by a proportional factor of e^{γ_k} between workers who entered unemployment in 1995

[23] The local unemployment rate is computed as the average of monthly rates. The quarter/year dummy takes a value of one if the midpoint of the duration interval is located on that quarter/year. For the open-ended interval 3 months from the beginning of the period is used as a reference point instead of the midpoint, which is not well-defined.

and 1996. These two groups are covered by different rules for computing unemployment pension benefits. The pattern of duration dependence for the comparison group (53-54 years old who entered unemployment in 1997 and 1998) is very close to that of the control group (50-52 years old). This is not very surprising, given that the age difference is so small and they all are covered by the same UI scheme. Consequently, the hazard functions between these groups are allowed to differ only by a proportional shift factor of e^{α_k}.

The likelihood of choosing passivity, $p(\mathbf{z})$, varies with characteristics $\mathbf{z}$. On the basis of descriptive evidence, we assume that all workers with a fixed entitlement of two years are active, and hence $p(\mathbf{z}) = 0$ is imposed for them. Following Schmidt and Witte (1989) and Pudney and Thomas (1995) among others, we assume the logistic distribution for ε within the treatment group:

$$p(\mathbf{z}) = \begin{cases} 0, & \text{if } D = 1, \\ \dfrac{\exp\{\mathbf{z}'\boldsymbol{\delta}\}}{1 + \exp\{\mathbf{z}'\boldsymbol{\delta}\}}, & \text{if } D = 0, \end{cases}$$

where $\mathbf{z}$ includes a constant.

We will consider the standard competing risks model as a benchmark for our duration analysis. The model outlined above is reduced to the standard competing risks model with a piecewise-constant hazard if one imposes $p(\mathbf{z}) = 0$ for all individuals in the data. In this case the log-likelihood function (3) factorizes into separate components for the parameters of each destination, and hence the estimation may be done in three steps.

5.3 Cumulative incidence functions

In our model there is a variety of channels through which the extended UI benefits can affect the flows of workers out of unemployment to employment, ALMPs, and non-participation. The cause-specific hazard functions are affected by extended UI benefits in two ways. First, the cause-specific hazard functions of the treatment group can be of a different shape. The hazard rates to destination k in the mth interval differ by the proportional factor $e^{\eta_k^m}$ between workers who are otherwise identical but are covered by the two different UI schemes. Since η_k^m are allowed to vary freely across the intervals, this amounts to estimating separate baseline hazard functions for the treatment and comparison groups.[24] Second, unlike the other groups, workers aged 53-54 who entered unemployment in 1995 or 1996 do not lose their UI benefits after 24 months. This results

[24] The flexible specification for the effect of the entitlement period on the hazard functions is adopted because economic theory does not impose simple parametric restrictions. In the two-state search model of Mortensen (1977), for example, an increase in the maximum duration of UI benefits reduces the employment hazard over the forepart of the unemployment period but increases it close to and beyond the exhaustion point. In other words, the effect on the employment hazard is predicted to change over the course of the unemployment spell, potentially reversing its sign at some point. In the empirical analysis this possibility is sometimes ruled out a priori by imposing the restriction that changes in the length of the entitlement period may lead to level shifts in the underlying hazard function but cannot affect its shape (e.g. Hunt, 1995; and Lalive and Zweimüller, 2004). Moreover, economic theory provides little guidance on the expected responses in the hazard rates to ALMPs and non-participation.

in a different time pattern for the unemployment compensation variable that is included in $\mathbf{x}_m$, the effect of which comes on top of the difference in the baseline hazard functions for long-duration spells. It should be stressed that these entitlement period effects on the hazard functions are conditional on workers being active (i.e. conditional on $\varepsilon = 1$). There is an additional disincentive effect via the choice of labour market withdrawal: some of those with extended UI benefits may be discouraged and choose to withdraw from active labour market behaviour entirely. This effect can be heterogenous, and it is measured with $p(\mathbf{z})$.

To summarize all these potential effects in a coherent way, we will calculate the marginal cumulative incidence and distribution functions for different groups. The cumulative incidence function (CIF) for destination $k \in \{e, p, o\}$ at spell duration $t \in (c_{m-1}, c_m]$ conditional on $\mathbf{X}(t, \tau)$ and ε is defined as

$$F_k(t\,|\mathbf{X}(t,\tau), \varepsilon) = \int_0^t \theta_k(u\,|\mathbf{X}(u,\tau), \varepsilon)\, S(u\,|\mathbf{X}(u,\tau), \varepsilon)\, du, \tag{4}$$

and it equals the probability that the individual has entered destination k by spell duration t. Note that $F_k(t\,|\mathbf{X}(t,\tau), \varepsilon = 0) = 0$ for $k \in \{e, p\}$ and $F_o(t\,|\mathbf{X}(t,\tau), \varepsilon = 0) = 1 - S(t\,|\mathbf{X}(t,\tau), \varepsilon = 0)$.

The marginal CIF gives the predicted fraction of workers who have escaped from unemployment through a particular exit route by a given duration time. The marginal cumulative distribution function (CDF), which equals the sum of marginal CIFs for all exit destinations, gives the predicted share of those who have left unemployment by a given duration time for any reason. The estimated fractions of individuals who have escaped unemployment by spell duration t through employment and ALMPs are obtained as

$$\widehat{F}_k(t) = \frac{1}{N}\sum_{i=1}^{N} [1 - \widehat{p}(\mathbf{z}_i)]\, \widehat{F}_k(t\,|\mathbf{X}_i(t,\tau_i), \varepsilon_i = 1)\,, \qquad k \in \{e, p\}\,,$$

where i indexes individuals and N denotes the group size at time $t = 0$. The estimated fraction of the workers who have left the labour force by spell duration t is

$$\widehat{F}_o(t) = \frac{1}{N}\sum_{i=1}^{N} \Big([1 - \widehat{p}(\mathbf{z}_i)]\, \widehat{F}_o(t\,|\mathbf{X}_i(t,\tau_i), \varepsilon_i = 1) + \widehat{p}(\mathbf{z}_i) \left[1 - \widehat{S}(t\,|\mathbf{X}_i(t,\tau_i), \varepsilon_i = 0)\right] \Big)\,.$$

In other words, we first calculate the estimates of CIFs for each individual at various points in duration time, conditional on the time path of the individual's covariate values up until that point. Then we obtain the group-specific estimates of the marginal CIFs by taking averages over individual estimates within the groups. The marginal CDF at spell duration t can be estimated as $\widehat{F}(t) = \sum_k \widehat{F}_k(t) = 1 - \widehat{S}(t)$, where the marginal survivor function is

$$\widehat{S}\,(t) = \frac{1}{N}\sum_{i=1}^{N} \Big([1 - \widehat{p}(\mathbf{z}_i)]\, \widehat{S}(t\,|\mathbf{X}_i(t,\tau_i), \varepsilon_i = 1) + \widehat{p}(\mathbf{z}_i) \widehat{S}(t\,|\mathbf{X}_i(t,\tau_i), \varepsilon_i = 0) \Big)\,.$$

In the conventional competing risks specification we simply have $\widehat{p}(\mathbf{z}_i) = 0$ and $\varepsilon_i = 1$ for all individuals.

5.4 Discussion

One can think of the model outlined above as a special case of a multiplicative frailty (or unobserved heterogeneity) model (Sy and Taylor, 2000). As a frailty variable, ε has a probability distribution with two mass points at known values. Unlike in the standard frailty models, ε is not entirely unobservable, since it is observed for individuals who exited to employment or ALMPs, and its distribution can depend on the same observed characteristics as the hazard functions do. This class of duration models was introduced by Farewell (1977). The models are known as mixture or cure models in statistics, and split population or mover-stayer models in econometrics (Schmidt and Witte, 1989; and Abbring, 2002). Applications in labour economics include Yamaguchi (1992), Swaim and Podgursky (1994), Pudney and Thomas (1995), Addison and Portugal (2003), Ollikainen (2003), and Mavromaras and Orme (2004). Our specification departs from the existing literature in that the same ε enters multiplicatively in the two distinct cause-specific hazard functions and does so only for a particular subgroup of the population.

Farewell (1982) emphasizes that the split population models should not be applied indiscriminately. There must be strong evidence of a subgroup not at the risk of experiencing the event of interest. A feature of the model that the same covariates can affect the choice of withdrawal from labour market behaviour and exit rates out of unemployment allows additional flexibility in modelling, but opens up the possibility of over-parametrization (Sy and Taylor, 2000). Identifiability of the model generally requires a long observation period and a sufficiently large number of uncensored observations. In our application descriptive findings along with indirect survey results give convincing evidence of the existence of a group of the elderly unemployed who are no longer engaged in labour market activities. Having a follow-up period of 5-6 years, we should also be able to detect this group.

The focus of the econometric analysis of UI has been on detecting effects on the employment hazard or the overall hazard out of unemployment. However, the effect of UI on the employment hazard may be less relevant from a policy perspective than its effects on the likelihood of leaving unemployment via employment by a given time. The employment effect of a policy change in the UI system that affects the employment hazard may be reinforced or attenuated by changes in hazard rates out of the labour force and to ALMPs. As a consequence, a policy change with a strong effect on the employment hazard may have a negligible or even opposite effect on the probability that the unemployment spell will end with employment. Alternatively, a policy change with no effect on the employment hazard may still have a significant effect on the likelihood of finding a job due to indirect effects via competing hazards. Thus it may be hazardous to focus on estimating reform effects on the employment hazard only by treating individuals who exit to other destinations than employment as censored observations. We believe that an appropriate way of evaluating employment effects requires simultaneous account of all cause-specific hazards, and the cumulative incidence functions provide a useful way of summarizing the results of competing risks analysis in a coherent and policy-relevant way. Although the cumulative incidence approach has enjoyed popularity in the medical treatment literature (e.g. Pepe, 1991; Gaynor et al., 1993; and Satagopan et al., 2004),[25] it has attracted surprisingly little attention in economic applications.

[25] In clinical studies, it is common to compare nonparametric estimates of the cumulative incidence functions. In the presence of covariates, Fine and Gray (1999) and Fine (1999) propose semiparametric methods for directly estimating the cumulative incidence functions rather than combining estimates of the cause-specific hazard functions.

6 Estimation results

6.1 Hazard function estimates

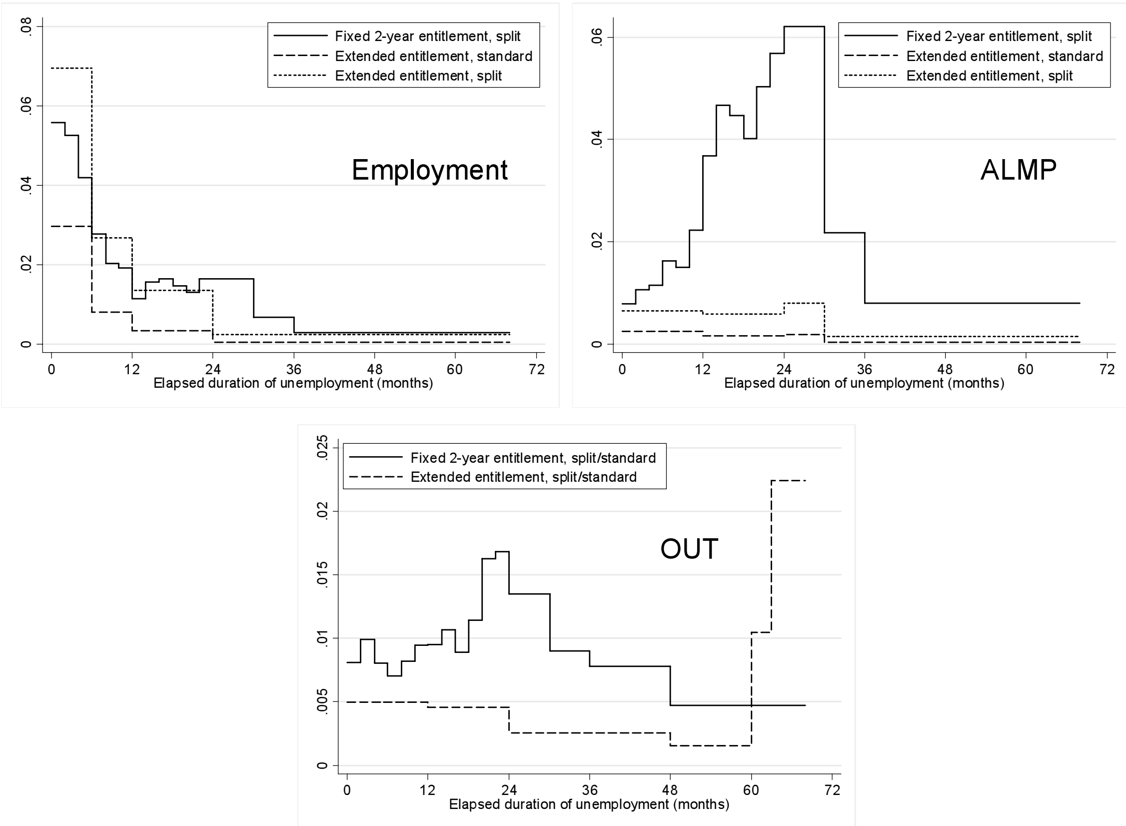
The hazard function estimates of the standard competing risks and split population specifications are presented in Figure 2 and Tables 3 and 4. The hazard functions are depicted for a reference person,[26] and the estimates obtained from the split population specification are conditional on searching (i.e. $\varepsilon = 1$). The baseline parameters (λ_k) not reported in the tables determine the shape of the cause-specific hazards for the treatment group with extended UI benefits in Figure 2. The time-varying coefficients (η_k) for individuals with the fixed two-year-entitlement period measure the difference between the hazard functions for the treatment and comparison groups, and the statistical significance thereof. The hazard functions for transitions to employment and ALMPs obtained from both specifications are plotted for a reference person with extended UI benefits. The corresponding hazard functions under the fixed entitlement period are roughly identical between the specifications, and hence only the hazard function estimate from the split population model is shown in Figure 2. Since both specifications produce identical estimates for the transition rate out of the labour force (see discussion in the Appendix and parameter estimates in Tables 3 and 4), the two hazard functions in the last graph are common to both models.

Under the two-year entitlement period the employment hazard exhibits considerable negative duration dependence. By the time the unemployment period has lasted for 6 months the employment hazard is more than halved. We find only a slight benefit-exhaustion-related upturn in the hazard for employment at 22-30 months, whereas most of the benefit-exhaustion-related exits are directed towards ALMPs. The hazard function for transitions to ALMPs peaks considerably both at one year and, even more so, at two years of continuous unemployment. These findings are in line with evidence from Norway (Bratberg and Vaage, 2000; and Røed and Zhang, 2003) and Sweden (Carling et al., 1996). The transition rate out of the labour force is fairly steady and low overall, peaking slightly around the time of benefit exhaustion.

The employment hazard of the worker with extended UI benefits obtained from the standard competing risks model declines with the elapsed duration, being significantly lower than the hazard of the worker with the fixed entitlement period (see time-varying coefficients and their standard errors in Table 3). However, the most dramatic discrepancies between the UI schemes are found for the transition rates to ALMPs. The hazard function of the treatment group for transitions to ALMPs obtained from the standard competing risks model lies flat at a very low level. The transition rate out of the labour force among workers with extended benefits is also low and flat until it peaks at the very

[26] The reference individual is married, lives in a region with the local unemployment rate of 20 per cent (slightly below the sample mean), and receives UI benefits equal to the mean benefits for the 53-54-year-olds (log of UI benefits = 3.3162). Other dummy variables are set to zero. Under the fixed entitlement period, the UI benefits of the reference individual are assumed to drop after 2 years of unemployment to the level of basic allowance (log of UI benefits = 2.6529).

Figure 2: Cause-specific hazard functions for a reference person under two different UI schemes



end of the observation period when the individuals start to exit via the unemployment pension scheme.

Our estimates of the labour market withdrawal probability from the split population model suggest that some half of the workers with extended UI benefits are inactive, having zero hazard rates to employment and ALMPs (see the next section). Hence, hazard estimates for the treatment group obtained from the standard competing risks specification are subject to bias, owing to a particular type of unobserved heterogeneity problem. The split population model takes this issue explicitly into account, and yields the hazard estimates for a reference worker with extended benefits who is still actively looking for a new job and considers ALMPs as a possible way of escaping unemployment. Not surprisingly, the hazard functions of the treatment group for transitions to employment and ALMPs conditional on being active are higher compared with the estimates from the conventional competing risks specification. The employment hazard for the treatment group exhibits negative duration dependence, being rather close to the estimated hazard for the comparison group. The difference between the employment hazard under fixed and extended entitlement is statistically significant at the conventional risk levels only between 4-6 and

Table 3: Results of the standard competing risks specification

	Hazard function for transitions to					
	Employment		ALMPs		Non-particip.	
Time-varying coefficient for 2-year entitlement period:						
Interval (0,2] months	0.634	(0.110)	1.150	(0.198)	0.491	(0.201)
Interval (2,4] months	0.580	(0.112)	1.457	(0.198)	0.691	(0.202)
Interval (4,6] months	0.361	(0.118)	1.544	(0.201)	0.481	(0.216)
Interval (6,8] months	1.258	(0.162)	1.891	(0.198)	0.347	(0.230)
Interval (8,10] months	0.946	(0.171)	1.808	(0.205)	0.495	(0.231)
Interval (10,12] months	0.889	(0.176)	2.207	(0.201)	0.639	(0.234)
Interval (12,14] months	1.235	(0.210)	3.177	(0.218)	0.730	(0.243)
Interval (14,16] months	1.569	(0.205)	3.418	(0.219)	0.849	(0.245)
Interval (16,18] months	1.609	(0.215)	3.372	(0.223)	0.663	(0.265)
Interval (18,20] months	1.508	(0.232)	3.267	(0.229)	0.916	(0.260)
Interval (20,22] months	1.372	(0.254)	3.494	(0.230)	1.270	(0.253)
Interval (22,24] months	1.610	(0.252)	3.614	(0.235)	1.305	(0.264)
Interval (24,30] months	3.120	(0.304)	5.025	(0.285)	1.847	(0.245)
Interval (30,36] months	2.231	(0.443)	5.738	(0.328)	1.442	(0.310)
Interval (36,48] months	1.387	(0.521)	4.734	(0.364)	1.300	(0.286)
Interval (48,60] months	1.387	(0.521)	4.734	(0.364)	1.300	(0.286)
Interval (60,63] months	1.387	(0.521)	4.734	(0.364)	−0.623	(0.329)
Interval (63,∞] months	1.387	(0.521)	4.734	(0.364)	−1.383	(0.306)
Age 50-52	0.117	(0.063)	−0.233	(0.077)	0.135	(0.119)
Female	−0.170	(0.071)	0.672	(0.093)	−0.006	(0.102)
Married	0.363	(0.050)	0.172	(0.086)	−0.109	(0.092)
Female $\times$ married	−0.368	(0.081)	0.021	(0.108)	0.149	(0.122)
Dependent child	0.155	(0.045)	0.062	(0.067)	−0.032	(0.091)
Swedish-speaking	0.204	(0.082)	−0.049	(0.122)	0.136	(0.141)
Tenure $\geq$ 4 years	−0.172	(0.037)	−0.126	(0.049)	0.064	(0.060)
Unemployed in early 1990s	0.312	(0.042)	0.146	(0.053)	0.065	(0.067)
Past recall in early 1990s	0.709	(0.047)	−0.291	(0.126)	0.089	(0.127)
Occupation: (ref. commercial)						
Technical	0.302	(0.081)	−0.131	(0.109)	−0.179	(0.147)
Humanist	0.071	(0.140)	−0.107	(0.161)	−0.163	(0.217)
Health care	0.708	(0.127)	−0.071	(0.200)	0.101	(0.253)
Clerical	−0.050	(0.074)	0.079	(0.073)	0.059	(0.096)
Agricultural	0.685	(0.102)	0.140	(0.275)	0.162	(0.282)
Transportation	0.239	(0.086)	−0.177	(0.136)	0.306	(0.130)
Industrial	0.383	(0.060)	−0.204	(0.073)	−0.092	(0.091)
Services	−0.037	(0.084)	−0.068	(0.090)	0.015	(0.114)
Not classified	−0.231	(0.264)	0.332	(0.295)	0.606	(0.299)
Year 1996 x age 53-54	0.043	(0.106)	−0.168	(0.199)	0.095	(0.136)
Time-varying covariates:						
Ln (UI benefits)	−0.735	(0.064)	2.155	(0.093)	0.271	(0.125)
Ln (local unemployment rate)	0.076	(0.080)	0.205	(0.096)	0.027	(0.119)
Quarter 2	0.268	(0.047)	−0.323	(0.072)	0.179	(0.091)
Quarter 3	−0.066	(0.056)	−0.197	(0.072)	0.079	(0.094)
Quarter 4	−0.201	(0.058)	0.012	(0.070)	0.206	(0.094)
Year 1996	0.183	(0.067)	0.117	(0.127)	−0.224	(0.132)
Year 1997	0.241	(0.070)	0.397	(0.129)	−0.213	(0.143)
Year 1998	0.210	(0.069)	0.306	(0.129)	−0.129	(0.141)
Year 1999	0.199	(0.087)	0.365	(0.138)	0.161	(0.158)
Year 2000	−0.235	(0.180)	0.395	(0.161)	0.667	(0.193)

Notes: Number of observations is 6,998. Standard errors in parentheses. Estimates of the baseline hazards not reported.

Table 4: Results of the split population specification

	Hazard function for transitions to					
	Employment		ALMPs		Non-particip.	
Time-varying coefficient for 2-year entitlement period:						
Interval (0,2] months	−0.220	(0.152)	0.174	(0.235)	0.491	(0.201)
Interval (2,4] months	−0.282	(0.154)	0.480	(0.235)	0.691	(0.202)
Interval (4,6] months	−0.508	(0.159)	0.565	(0.239)	0.481	(0.217)
Interval (6,8] months	0.035	(0.235)	0.910	(0.236)	0.347	(0.230)
Interval (8,10] months	−0.276	(0.239)	0.824	(0.242)	0.495	(0.232)
Interval (10,12] months	−0.335	(0.243)	1.221	(0.240)	0.639	(0.235)
Interval (12,14] months	−0.179	(0.294)	1.839	(0.288)	0.730	(0.243)
Interval (14,16] months	0.151	(0.293)	2.078	(0.289)	0.849	(0.245)
Interval (16,18] months	0.189	(0.300)	2.031	(0.292)	0.663	(0.265)
Interval (18,20] months	0.089	(0.314)	1.926	(0.296)	0.916	(0.260)
Interval (20,22] months	−0.044	(0.326)	2.151	(0.297)	1.270	(0.253)
Interval (22,24] months	0.196	(0.327)	2.273	(0.301)	1.305	(0.264)
Interval (24,30] months	1.481	(0.419)	3.440	(0.377)	1.847	(0.245)
Interval (30,36] months	0.595	(0.529)	4.039	(0.438)	1.442	(0.310)
Interval (36,48] months	−0.264	(0.597)	3.036	(0.465)	1.300	(0.286)
Interval (48,60] months	−0.264	(0.597)	3.036	(0.465)	1.300	(0.286)
Interval (60,63] months	−0.264	(0.597)	3.036	(0.465)	−0.623	(0.329)
Interval (63,∞] months	−0.264	(0.597)	3.036	(0.465)	−1.383	(0.306)
Age 50-52	0.124	(0.063)	−0.228	(0.077)	0.135	(0.119)
Female	−0.155	(0.074)	0.692	(0.095)	−0.006	(0.116)
Married	0.394	(0.052)	0.196	(0.088)	−0.109	(0.097)
Female $\times$ married	−0.350	(0.085)	0.010	(0.112)	0.149	(0.134)
Dependent child	0.135	(0.046)	0.055	(0.067)	−0.032	(0.091)
Swedish-speaking	0.198	(0.084)	−0.060	(0.124)	0.136	(0.141)
Tenure $\geq$ 4 years	−0.156	(0.039)	−0.124	(0.050)	0.064	(0.060)
Unemployed in early 1990s	0.291	(0.044)	0.143	(0.054)	0.065	(0.067)
Past recall in early 1990s	0.682	(0.049)	−0.312	(0.126)	0.089	(0.127)
Occupation: (ref. commercial)						
Technical	0.278	(0.084)	−0.140	(0.110)	−0.179	(0.148)
Humanist	0.009	(0.151)	−0.140	(0.163)	−0.163	(0.217)
Health care	0.666	(0.131)	−0.099	(0.202)	0.101	(0.253)
Clerical	−0.086	(0.077)	0.053	(0.074)	0.059	(0.096)
Agricultural	0.629	(0.106)	0.134	(0.277)	0.162	(0.283)
Transportation	0.264	(0.089)	−0.151	(0.137)	0.306	(0.131)
Industrial	0.381	(0.062)	−0.198	(0.074)	−0.092	(0.091)
Services	−0.082	(0.088)	−0.100	(0.091)	0.015	(0.115)
Not classified	−0.284	(0.272)	0.282	(0.298)	0.606	(0.299)
Year 1996 x age 53-54	−0.255	(0.171)	−0.379	(0.270)	0.095	(0.136)
Time-varying covariates:						
Ln (UI benefits)	−0.695	(0.064)	2.096	(0.093)	0.271	(0.126)
Ln (local unemployment rate)	0.094	(0.082)	0.217	(0.097)	0.027	(0.122)
Quarter 2	0.274	(0.047)	−0.321	(0.072)	0.179	(0.091)
Quarter 3	−0.038	(0.056)	−0.199	(0.072)	0.079	(0.094)
Quarter 4	−0.176	(0.058)	0.018	(0.071)	0.206	(0.094)
Year 1996	0.214	(0.069)	0.145	(0.129)	−0.224	(0.132)
Year 1997	0.267	(0.071)	0.430	(0.131)	−0.213	(0.143)
Year 1998	0.239	(0.070)	0.333	(0.131)	−0.129	(0.141)
Year 1999	0.241	(0.088)	0.395	(0.139)	0.161	(0.158)
Year 2000	−0.185	(0.181)	0.429	(0.162)	0.667	(0.193)

Notes: Number of observations is 6,998. Standard errors in parentheses. Estimates of the baseline hazards not reported.

24-30 months of unemployment (see Table 4). In other words, workers with extended UI benefits choosing to continue job search exit to employment at a similar rate as otherwise identical workers with a fixed period of UI benefits. By contrast, active individuals within the treatment group have very low transition rates to ALMPs compared with the comparison group. As a general remark, the transition rates out of unemployment under the two different UI schemes are of a different shape here. Restricting the effects of the length of the UI entitlement period to a proportional shift alone would, at least in our case, be far too constraining (for such a priori restrictions see e.g. Hunt, 1995, and Winter-Ebmer, 1998).

The coefficient estimates for the background characteristics in Tables 3 and 4 are in line, and hence we will only discuss the estimates of the split population specification. The findings are fairly conventional, indicating that women have a lower employment hazard and a higher hazard rate for transitions to active labour market programmes. Married people have higher transition rates both to employment and ALMPs. Individuals with dependent children also have a higher employment hazard, as do people speaking Swedish as their first language. The coefficients of occupational variables indicate differences in the transition rate to employment across different sectors.

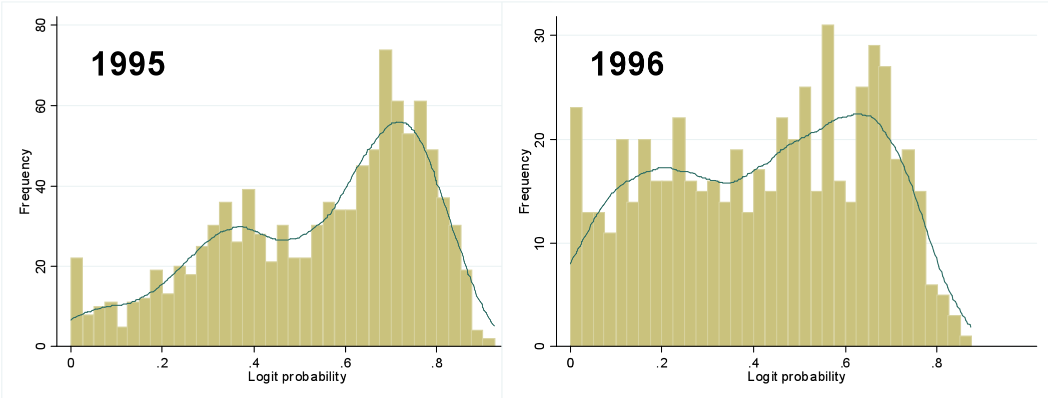
There is evidence of both quarterly and yearly variation in transition rates to all end-states. The time-dependent regional unemployment rate has a significant positive effect on the transition rate to ALMPs. This merely points out the strong regional aspect of the labour market policy practised in Finland. As discussed in Lilja (1992) and Ollikainen (2003), a higher than average proportion of individuals participating in active labour market programmes comes from regions with a high unemployment to vacancies ratio. As expected, the amount of unemployment benefits has a negative effect on the hazard rate for transitions to employment. This is generally viewed as a result of the individual's higher leeway in being selective about his employment. High unemployment benefits substantially increase the transition rate to ALMPs, as well.

6.2 Likelihood of labour market withdrawal

As expected on the basis of descriptive analysis, we find evidence of a large subgroup of workers with extended UI benefits who are no longer actively engaged in job search. Figure 3 shows the distributions of labour market withdrawal probabilities (i.e. the probability of $\varepsilon = 0$) across individuals with extended benefits by year of entry to unemployment. The mean probability of being inactive in our treatment group is 0.54 among those becoming unemployed in 1995 and 0.42 among those becoming unemployed in 1996. Thus, roughly half of the unemployed eligible for extended UI benefits are at no point interested in finding a way out of unemployment via employment or ALMPs, but instead are just passively waiting for retirement. The finding of such high overall withdrawal probabilities corroborates the necessity of implementing the split population model here.

Moreover, there is a lot of variation in the withdrawal probabilities across individuals.

Figure 3: Distributions of labour market withdrawal probabilities by year of entry to unemployment (Epanechnikov kernel)



In addition to the evident variation within each year, there might also be some difference in the activity levels of the groups between years. However, looking at the logit coefficients in Table 5 we find that, in fact, the dummy for 1996 is statistically insignificant, indicating that the observed yearly differences in the withdrawal probabilities are attributed to differences in the observed characteristics of the unemployed individuals. It appears that personal characteristics such as gender, being married or having dependent children do not significantly affect the decision of whether or not to search. Factors related to the individual's employment history turn out to be much more important. In all occupations individuals tend to be less likely discouraged than in the reference group, commercial work, although many of these differences are not statistically significant. The inflection point for the quadratic effect of unemployment benefits is 31 euros per day. Since two-thirds of individuals in the treatment group are receiving benefits lower than this, a small increase in UI benefits would discourage most of the people from searching for a new job.

As described in Kyyrä and Wilke (2007), the unemployment risk of those elderly workers who are eligible for extended UI benefits is much higher than that of other older workers in Finland, and particularly high in the case of larger firms with more than 50 employees. When forced to downsize, companies tend to target their dismissals on the elderly, some of whom may quite willingly retire via the unemployment tunnel scheme. In a tough situation this is the most easily approved line of action by the general public, and hence least damaging for the firms' reputation, as the income level for the elderly is well secured by the unemployment tunnel scheme. Moreover, firms with more than 50 employees are liable for a fraction of early retirement expenditures via partially experience-rated employer contributions. The largest firms with more than 500 employees are obliged to pay a higher cost share of disability pensions than of unemployment pensions. As a consequence, directing the elderly employees in the UT scheme may be economically rational, especially for large companies, as in this way they escape the potential disability

Table 5: Determinants of the labour market withdrawal probability in the split population model

	Logit coefficient		Marginal Effect
Year 1996	−0.295	(0.257)	−0.058
Female	0.254	(0.360)	0.050
Married	0.447	(0.305)	0.087
Female × married	0.103	(0.388)	0.020
Dependent child	−0.475	(0.265)	−0.093
Swedish-speaking	−0.101	(0.402)	−0.020
Tenure $\geq$ 4 years	0.224	(0.163)	0.044
Unemployed in early 1990s	−0.170	(0.188)	−0.033
Past recall in early 1990s	−0.796	(0.308)	−0.156
Occupation: (ref. commercial)			
Technical	−0.848	(0.384)	−0.166
Humanist	−0.947	(0.733)	−0.185
Health care	−1.304	(0.642)	−0.255
Clerical	−0.680	(0.289)	−0.133
Agricultural	−1.955	(0.655)	−0.382
Transportation	−0.008	(0.414)	−0.002
Industrial	−0.628	(0.261)	−0.123
Services	−0.523	(0.413)	−0.102
Not classified	−1.347	(1.564)	−0.263
Firm size: (ref. $\leq$ 50 employees)			
51-500 employees	1.031	(0.236)	0.202
Over 500 employees	1.264	(0.246)	0.247
Ln (local unemployment rate)	−0.042	(0.409)	−0.008
(UI benefits) / 10	5.064	(1.035)	0.990
(UI benefits)2 / 100	−0.817	(0.175)	−0.160
Constant	−7.963	(1.782)	

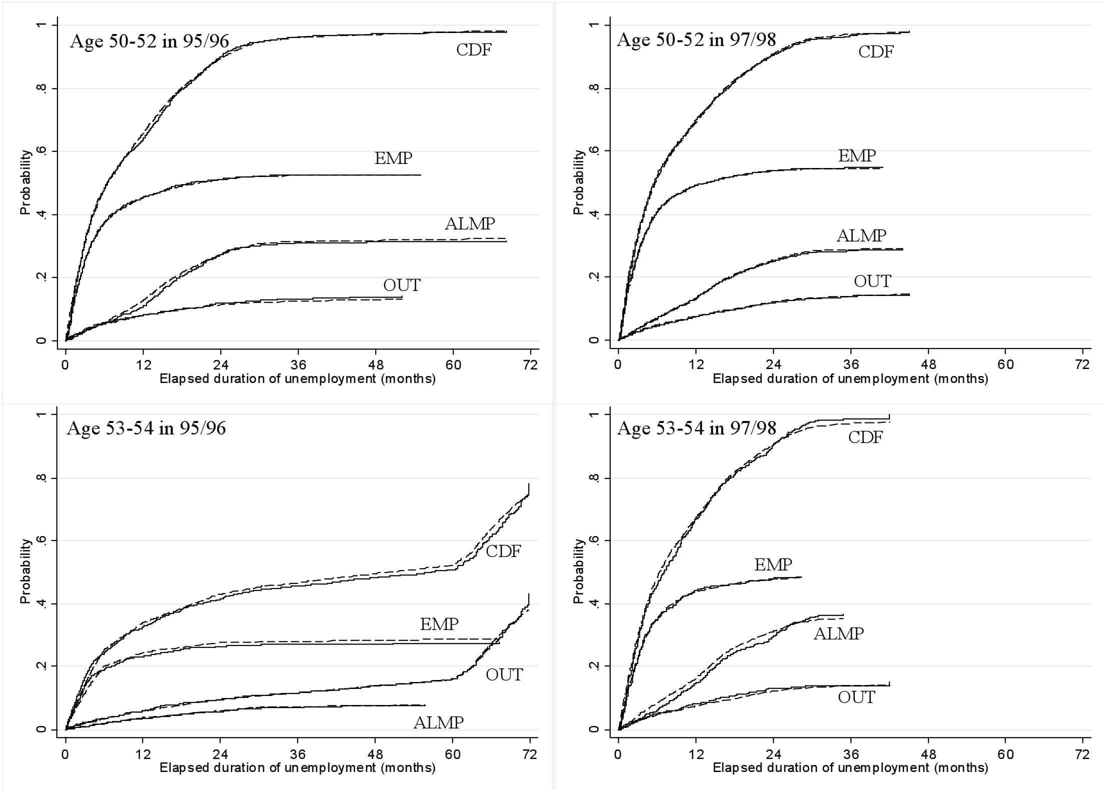
Notes: Standard errors of logit coefficients in parentheses. Marginal effects are computed by taking the average of individual-specific effects.

pension expenditures of these employees. Of course, the employers have an incentive to get rid of workers with the highest risk of disability in the first place. Such workers may also be more likely to become passive in the case of job loss. This is consistent with our result that individuals laid off from medium-sized or large firms are much more likely to withdraw from job search than those laid off from small firms with no more than 50 employees. Moreover, the estimated withdrawal probability is highest for the employees of the largest firms.

6.3 Cumulative incidence functions

We found evidence of a large fraction of inactive workers within the treatment group and notable differences in the cause-specific hazard functions between the groups. To summarize these results, we calculate the marginal cumulative incidence and distribution functions for different groups defined by age category and time of entry to unemployment. In Figures 4 and 5 the solid lines depict the nonparametric estimates given by the data and the dashed ones depict the estimates of our models. A comparison of the predictions

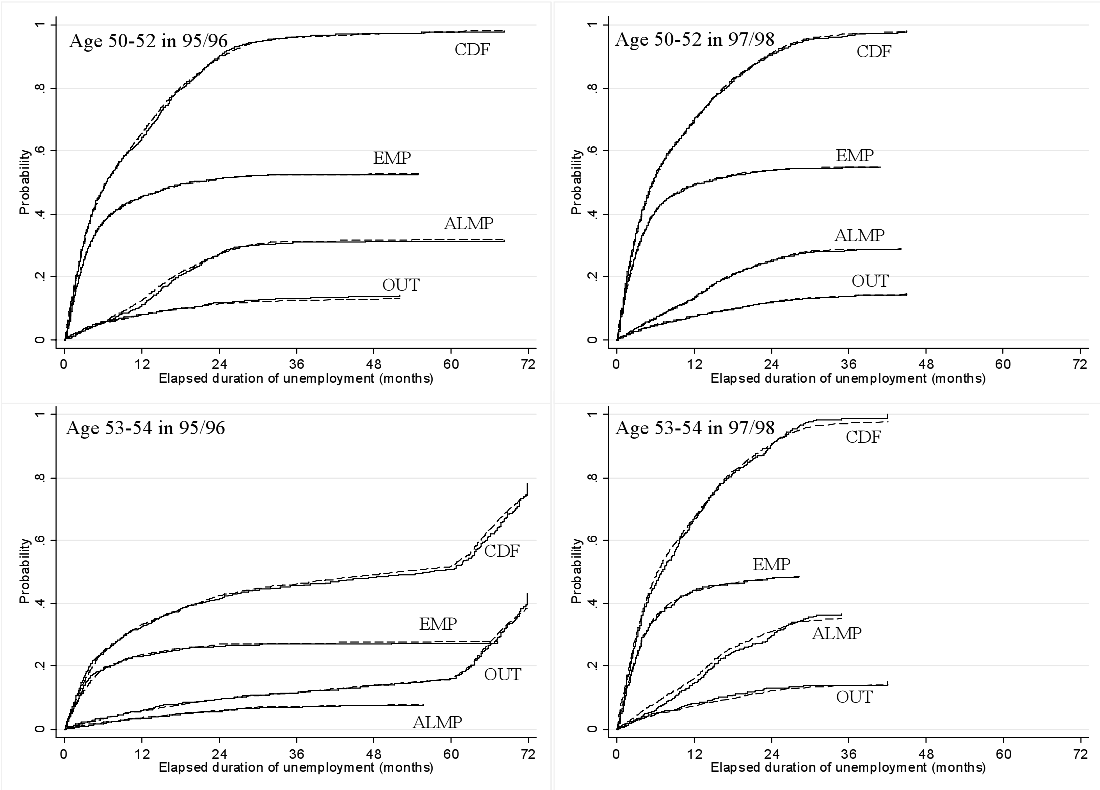
Figure 4: Cumulative incidence and distribution functions from standard duration model by age and entry year of unemployment (Note: Solid lines are nonparametric estimates, and the dashed ones are the fitted curves from the model)



obtained from our model with the corresponding nonparametric estimates provides a simple procedure for assessing the goodness of fit. The cumulative incidence and distribution function estimates from the standard competing risks specification and from the split population specification are roughly identical in all the other groups except for the treatment group. Even there, the fit of the split population model is only marginally better compared to the standard duration model. Hence, the contribution of our paper lies not in claiming the split population model to be superbly better in obtaining an accurate fit in case of this particular data set, but in stating that with the split population model we can dig deeper into the reasons behind the phenomenon we are observing.

We find that the treatment group, as such, is very different from the other three groups, with considerably lower cumulative probabilities of exiting to employment and ALMPs. The cumulative incidence of employment by 12 months is some 45 to 50% in the control and comparison groups, but only 25% in the treatment group. In the control groups the overall probability of employment eventually converges up to 55% and in the comparison group to 50%, while in the treatment group this figure is less than 30%. Also, very few people in the treatment group eventually escape unemployment through ALMPs, while the cumulative

Figure 5: Cumulative incidence and distribution functions from split population model by age and entry year of unemployment (Note: Solid lines are nonparametric estimates, and the dashed ones are the fitted curves from the model)

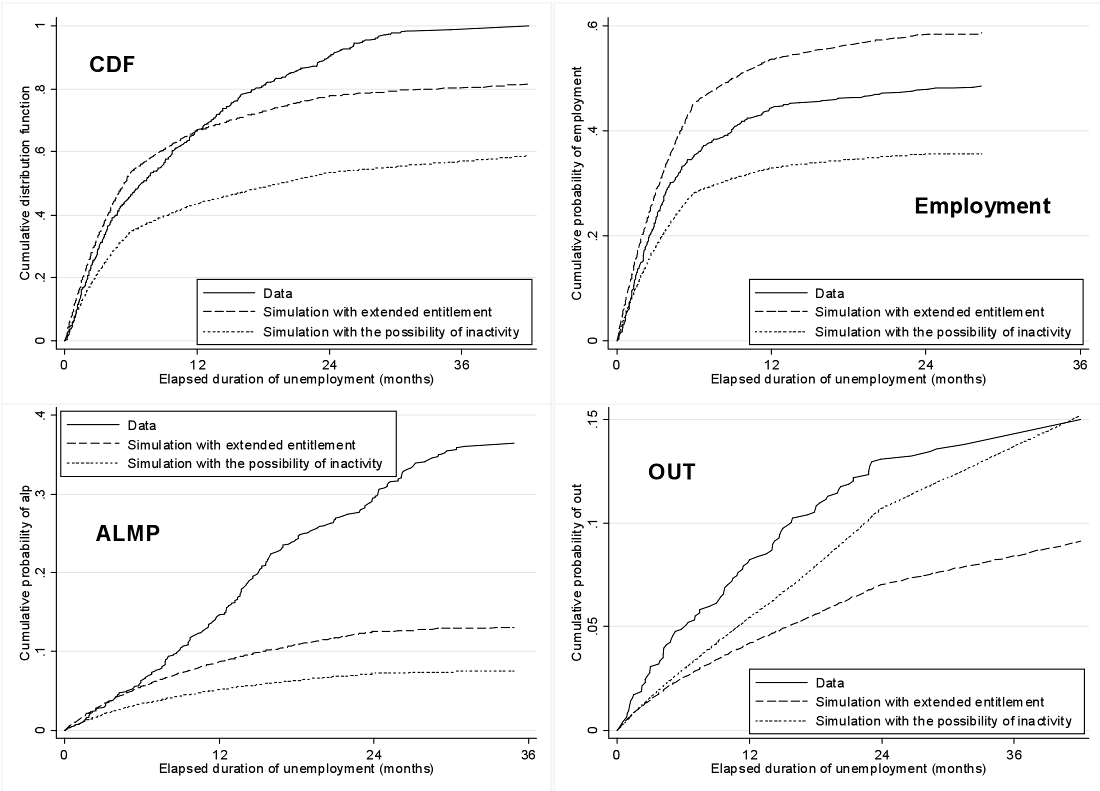


probability of participation in such programmes in the control and comparison groups converges even up to 35%. As discussed in Section 3.3, if the unemployed participate in ALMPs simply in order to prolong the exhaustion of benefits or to regain eligibility, then the individuals in our treatment group have little incentive to participate, which appears to be the case here. Overall, the percentage of the unemployed eventually leaving unemployment in the group with extended benefit duration converges to only some 50% by duration of 60 months, after which it converges further to 75% due to retirement.

6.4 Simulating the effects of indefinite UI benefits

Eligibility for extended UI benefits was removed from workers aged 53 and 54 in 1997. As we have already seen, such a sharp change in the UI scheme has had a strong effect on the unemployment experiences of those who were affected. In this section we address this issue further by simulating the hypothetical effects of extended UI benefits on workers aged 53-54 who entered unemployment in 1997 or in 1998. In other words, we ask how this group of the unemployed would have behaved, had they been eligible for the extended benefits. We emphasize that our simulation exercise does not tell much about the overall

Figure 6: Simulated marginal cumulative distribution and incidence functions under different UI schemes



effect of the 1997 reform, since it also had a strong impact on the inflow to unemployment (Kyyrä and Wilke, 2007). In the absence of the reform there would have been more 53-54-year-old employees becoming unemployed between 1997 and 1998. That is, our results apply only to the group of workers who lost their jobs and became unemployed in the world where the reform actually took place. Using the parameter estimates from the split population model we are able to separately identify the effects of extended UI benefits on the participation choice and conditional cumulative probabilities of exits to employment, ALMPs, and non-participation.

In Figure 6 we illustrate the results of this experiment by plotting the cumulative distribution and incidence functions, first, for the observed data, i.e. individuals aged 53-54 entering unemployment in 1997 or 1998. (These are simply the nonparametric estimates.) Second, for the same group in the hypothetical case of being eligible for the extended UI entitlement, but assuming that all individuals remain active (i.e. by setting $\varepsilon = 1$ for all), and third, in this same hypothetical situation, but allowing individuals to withdraw from the labour market with the estimated probabilities. These simulations were performed separately for each individual in the data, and the averages of fitted curves over all individuals are plotted in Figure 6.

We begin with a comparison of actual outcomes and hypothetical outcomes when all workers with extended UI benefits are assumed to remain active. In the real-world data the cumulative distribution function converges to 1, but among the active population with extended entitlement it converges to some 0.8, i.e. 20% of the active unemployed are predicted to remain unemployed secured by the extension. The probability of employment in the actual sample with fixed 2-year UI entitlement converges to 50%, but allowing for the extended entitlement it, in fact, approaches 60% when all individuals are assumed to remain active. This results from the drastically low transition rates to ALMPs and non-participation among the extension-entitled population. In the real-world data exits to ALMPs account for a notable part of all exits, as some 36% of all unemployment spells eventually end with a transition to such programmes. Adding extended entitlement, it would only be about 13%. The probability of exiting the labour force altogether converges to 15% in the real-world data, but if we should allow for extended entitlement the incidence of exiting would decline to 9%. Hence, among the active unemployed the entitlement extension effectively increases the probability of employment by 10 percentage points, but due to the otherwise low transitions rates, the probability of remaining unemployed is also higher for them.

If we consider the population with the extended entitlement and allow for the possibility of inactivity, the predicted cumulative probability of exiting from any cause by the end of the observation period is only some 60%. The probability of employment decreases by 15 percentage points compared with the real-world case with fixed entitlement, converging eventually to 35%. Allowing for the possibility of inactivity for the extension-entitled population results in the probability of participation in ALMPs decreasing further to 7%, while the probability of withdrawing from the labour market returns to the level of the real-world case.

Overall, our simulation exercise illustrates that consideration of the end-state specific hazard functions does not give a full picture of the underlying phenomena. In this particular case, the employment hazard functions reveal hardly any statistically significant difference in the case of the fixed two-year UI entitlement and the extended entitlement among the active population (Figure 2). This might falsely lead us to conclude that, after conditioning on remaining active, the extension of the entitlement period has no effect on the probability of exiting to employment. Looking at our simulated cumulative incidence functions we do, however, find that among the active extension-entitled population the cumulative probability of employment is actually *higher* due to the lower hazards for transitions to ALMPs and non-participation. This result is reversed once we allow some of those with extended UI benefits to choose inactivity. To summarize, via our simulation exercise we are able to say, that had this UI entitlement extension from two years to indefinite been applied to the same cohort becoming unemployed in 1997 or 1998, it would have resulted in a 15% decrease in their employment probability.

7 Concluding remarks

In this paper we have investigated flows out of unemployment among older workers under indefinite UI duration and a fixed entitlement period of two years. We took advantage of the Finnish UI reform to identify the impact of the fixed UI duration within the difference-in-differences framework. The possibility that some older unemployed people may be discouraged from labour market activities was accounted for by using the competing risks specification of the split population duration model.

We found no evidence of large increases in the employment hazard around the time of benefit exhaustion for those workers with the entitlement period of two years. By contrast, the hazard rates for labour market programmes and non-participation exhibit large increases as the time of benefit exhaustion approaches. These findings are in line with evidence from other Nordic countries. The combination of extended UI benefits and an early retirement scheme serves as a popular pathway to labour market withdrawal several years prior to the normal old-age pension. Our results suggest that as many as half of the elderly unemployed entitled to extended UI benefits choose to withdraw from the labour market, remaining passive until early retirement. The likelihood of labour market withdrawal varies with occupation, the level of UI benefits, and the size of the past employer. There are no notable discrepancies in the employment hazards between active workers with extended UI benefits and those with the entitlement period of two years. However, active workers with extended UI benefits have much lower transition rates to labour market programmes and non-participation (prior to access to early retirement). As a consequence, compared with those who will lose their benefits after two years of unemployment, active workers entitled to extended UI benefits are more likely to enter employment but also more likely to still be unemployed 36 months after entry to unemployment.

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A Technical details

A.1 Nonparametric estimators

Let $t_1 < t_2 < \cdots < t_M$ be the observed durations of completed unemployment spells in the data. Denote the number of individuals who exit unemployment to destination $k \in \{e, p, o\}$ at spell duration t_j with d_{kj}. The Kaplan-Meier estimate of the survivor function is

$$\widehat{S}(t) = \prod_{j|t_j \le t} \left(1 - \frac{d_j}{n_j}\right),$$

where $d_j = d_{ej} + d_{pj} + d_{oj}$ is the total number of exits and n_j is the number of individuals at risk at spell duration t_j. The estimate of the cumulative incidence function for destination k is given by

$$\widehat{F}_k(t) = \prod_{j|t_j \le t} \frac{d_{kj}}{n_j} \widehat{S}(t_{j-1}),$$

see for example Gaynor et al. (1993). It can be verified that $\widehat{F}_e(t) + \widehat{F}_p(t) + \widehat{F}_o(t) = 1 - \widehat{S}(t)$.

A.2 Survivor functions

For ease of exposition, we denote the hazard rate for destination k at spell duration $t \in (c_{m-1}, c_m]$ as $\theta_k(\mathbf{x}_m) \equiv \theta_k(t\,|\mathbf{X}(t,\tau))$, where $\mathbf{x}_m$ is a vector of covariate values in the mth duration interval. The overall hazard rate is denoted with $\theta(\mathbf{x}_m) = \theta_e(\mathbf{x}_m) + \theta_p(\mathbf{x}_m) + \theta_o(\mathbf{x}_m)$.

The survivor function at spell duration $t \in (c_{m-1}, c_m]$ is given by

$$S(t\,|\mathbf{X}(t,\tau)) = [1 - p(\mathbf{z})]\, S(t\,|\mathbf{X}(t,\tau), \varepsilon = 1) + p(\mathbf{z}) S(t\,|\mathbf{X}(t,\tau), \varepsilon = 0),$$

where

$$\begin{aligned}
S(t\,|\mathbf{X}(t,\tau), \varepsilon = 1) &= \exp\left\{ -\int_0^t \sum\nolimits_k \theta_k(u\,|\mathbf{X}(u,\tau))\, du \right\} \\
&= \exp\left\{ -\sum_{j=1}^{m-1} \int_{c_{j-1}}^{c_j} \theta(\mathbf{x}_j) du - \int_{c_{m-1}}^{t} \theta(\mathbf{x}_m) du \right\} \\
&= \left[\prod_{j=1}^{m-1} \exp\left\{-\theta(\mathbf{x}_j)\,(c_j - c_{j-1})\right\} \right] \exp\left\{-\theta(\mathbf{x}_m)\,(t - c_{m-1})\right\}.
\end{aligned}$$

and

$$S(t\,|\mathbf{X}(t,\tau), \varepsilon = 0) = \left[\prod_{j=1}^{m-1} \exp\left\{-\theta_o(\mathbf{x}_j)\,(c_j - c_{j-1})\right\} \right] \exp\left\{-\theta_o(\mathbf{x}_m)\,(t - c_{m-1})\right\}.$$

A.3 Cumulative incidence functions

The cumulative incidence function for destination k at spell duration $t \in (c_{m-1}, c_m]$

$$F_k(t\,|\mathbf{X}(t,\tau)) = [1 - p(\mathbf{z})]\, F_k(t\,|\mathbf{X}(t,\tau), \varepsilon = 1) + p(\mathbf{z}) F_k(t\,|\mathbf{X}(t,\tau), \varepsilon = 0),$$

where

$$
\begin{aligned}
F_k(t\,|\mathbf{X}(t,\tau),\varepsilon=1) &= \int_0^t \theta_k(u\,|\mathbf{X}(u,\tau))\,S(u\,|\mathbf{X}(u,\tau),\varepsilon=1)\,du \\
&= \sum_{j=1}^{m-1}\int_{c_{j-1}}^{c_j}\theta_k(\mathbf{x}_j)S(u\,|\cdot)\,du + \int_{c_{m-1}}^{t}\theta_k(\mathbf{x}_m)S(u\,|\cdot)\,du \\
&= \sum_{j=1}^{m-1}\theta_k(\mathbf{x}_j)S(c_{j-1}\,|\cdot)\int_{c_{j-1}}^{c_j}\exp\{-\theta(\mathbf{x}_j)\,(u-c_{j-1})\}\,du \\
&\quad +\theta_k(\mathbf{x}_m)S(c_{m-1}\,|\cdot)\int_{c_{m-1}}^{t}\exp\{-\theta(\mathbf{x}_m)\,(u-c_{m-1})\}\,du \\
&= \sum_{j=1}^{m-1}\frac{\theta_k(\mathbf{x}_j)S(c_{j-1}\,|\cdot)}{\theta(\mathbf{x}_j)}\left[1-\exp\{-\theta(\mathbf{x}_j)\,(c_j-c_{j-1})\}\right] \\
&\quad +\frac{\theta_k(\mathbf{x}_m)S(c_{m-1}\,|\cdot)}{\theta(\mathbf{x}_m)}\left[1-\exp\{-\theta(\mathbf{x}_m)\,(t-c_{m-1})\}\right] \\
&= \sum_{j=1}^{m-1}\frac{\theta_k(\mathbf{x}_j)}{\theta(\mathbf{x}_j)}\left[S(c_{j-1}\,|\cdot)-S(c_j\,|\cdot)\right]+\frac{\theta_k(\mathbf{x}_m)}{\theta(\mathbf{x}_m)}\left[S(c_{m-1}\,|\cdot)-S(t\,|\cdot)\right],
\end{aligned}
$$

with $S(c_0\,|\cdot)\equiv 1$, and

$$F_k(t\,|\mathbf{X}(t,\tau),\varepsilon=0)=0, \qquad k\in\{e,p\}\,,$$

and

$$F_o(t\,|\mathbf{X}(t,\tau),\varepsilon=0)=1-S(t\,|\mathbf{X}(t,\tau),\varepsilon=0)\,.$$

A.4 The log-likelihood function

For ease of exposition, the conditioning covariates, $\mathbf{X}(t,\tau)$, are suppressed from the following expressions. Denote the cumulative hazard function for destination $k\in\{e,p,o\}$ at spell duration $t\in(c_{m-1},c_m]$ with

$$\Lambda_k(t)\equiv-\sum_{j=1}^{m-1}\theta_k(\mathbf{x}_j)\,(c_j-c_{j-1})-\theta_k(\mathbf{x}_m)\,(t-c_{m-1})\,.$$

We can write

$$\ln S(t\,|\varepsilon=1)=-\sum_j\Lambda_j(t),$$

where $\sum_j$ denotes the sum over all the three exit destinations. For an individual in the treatment group with $p(\mathbf{z})>0$ we have

$$
\begin{aligned}
\ln S(t) &= \ln\left([1-p(\mathbf{z})]\,e^{-\sum_j\Lambda_j(t)}+p(\mathbf{z})e^{-\Lambda_o(t)}\right) \\
&= \ln\left(e^{-\sum_j\Lambda_j(t)}+e^{\mathbf{z}'\boldsymbol{\delta}-\Lambda_o(t)}\right)-\ln\left(1+e^{\mathbf{z}'\boldsymbol{\delta}}\right) \\
&= \ln\left(e^{-\Lambda_e(t)-\Lambda_p(t)}+e^{\mathbf{z}'\boldsymbol{\delta}}\right)-\Lambda_o(t)-\ln\left(1+e^{\mathbf{z}'\boldsymbol{\delta}}\right).
\end{aligned}
$$

By substituting these expressions into (3), we find that the contribution of individual i to the log-likelihood function is

$$\begin{aligned}
\mathcal{L}_i &= \sum_j d_{ij} \ln \theta_j(t_i) - (d_{ie} + d_{ip}) \left[\ln \left(1 + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) + \sum\nolimits_j \Lambda_j(t) \right] \\
&\quad + (1 - d_{ie} - d_{ip}) \left[\ln \left(e^{-\Lambda_e(t_i) - \Lambda_p(t_i)} + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) - \Lambda_o(t_i) - \ln \left(1 + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) \right] \\
&= \sum_j d_{ij} \ln \theta_j(t_i) - \Lambda_o(t_i) - \ln \left(1 + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) - (d_{ie} + d_{ip}) \left[\Lambda_e(t_i) + \Lambda_p(t_i) \right] \\
&\quad + (1 - d_{ie} - d_{ip}) \ln \left(e^{-\Lambda_e(t_i) - \Lambda_p(t_i)} + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right),
\end{aligned}$$

if he belongs to the treatment group, and

$$\mathcal{L}_i = \sum_j d_{ij} \ln \theta_j(t_i) - \sum\nolimits_j \Lambda_j(t)$$

otherwise. These expressions can be combined:

$$\begin{aligned}
\mathcal{L}_i &= \sum_j d_{ij} \ln \theta_j(t_i) - \Lambda_o(t_i) - \left[1 - q_i + q_i \left(d_{ie} + d_{ip} \right) \right] \left[\Lambda_e(t_i) + \Lambda_p(t_i) \right] \\
&\quad + q_i \left[(1 - d_{ie} - d_{ip}) \ln \left(e^{-\Lambda_e(t_i) - \Lambda_p(t_i)} + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) - \ln \left(1 + e^{\mathbf{z}_i' \boldsymbol{\delta}} \right) \right],
\end{aligned}$$

where $q_i = 1$ if individual i belongs to the treatment group, and $q_i = 0$ otherwise.

The structure of the log-likelihood function implies that the ML estimators of the parameters of the non-participation hazard are statistically independent of all other parameters of the model (since the matrix of the second derivatives of the log-likelihood function is block diagonal). As a consequence, the estimation of the conventional competing risks and split population specifications will lead to identical parameter estimates for the hazard function for transitions out of the labour force.

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