

# Early retirement policy in the presence of competing exit pathways: Evidence from policy reforms in Finland\*

— PRELIMINARY VERSION —

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## Abstract

A majority of older Finns withdraw from employment via early retirement schemes years before the old-age pension. A series of policy reforms aimed at reducing the widespread use of early exit pathways have caused exogenous differences in eligibility rules between people born in different years. By exploiting this variation for identification, the paper estimates the effects of changes in the eligibility age thresholds for unemployment and part-time pension schemes and tightening medical screening for disability pension benefits. The findings imply that these reforms have jointly raised the average age at which older workers leave employment by 3.8 months. This effect results from a sharp drop in the disability enrolment rate among workers aged 58 and above, and from a lower incidence of unemployment at younger groups. The policy effects are found to be heterogenous, so that different subgroups have been affected differently by different reforms.

**Keywords:** Early retirement, policy reform, disability, unemployment.

**JEL-codes:** J14, J26.

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# 1 Introduction

A tendency of older workers to retire several years before the official retire age is a common problem in Europe. In Finland, a majority of older people withdraw from regular employment through some early retirement scheme. As a result, the effective retirement age is below 60, over five years less than the general retirement age. Most early leavers enter either disability pension or unemployment-related benefits, which are typically collected until the age of 65 when the entitlement to an old-age pension begins. A disability pension is generally payable to all people whose working capacity has decreased substantially, but earlier cohorts had been able to qualify for disability pension benefits under less strict medical criteria. A high incidence of long-term unemployment among older people can be attributed to the combination of extended unemployment benefits for workers above a given age threshold and an unemployment pension payable to the aged long-term unemployed. Moreover, a gradual withdrawal from employment is possible through a part-time pension, which can be awarded to a worker above a certain age threshold who switches from full-time to part-time work.

The widespread use of the early retirement schemes has raised concerns about the financial sustainability of the pension system. This is because the old-age dependency ratio – the ratio of the population aged 65 and over to the population aged 15-64 – has been estimated to rise from the current level of 26% to 45% by 2030 (Statistics Finland, 2009), when Finland is expected to have one of the highest dependency ratios among the OECD countries. There is a consensus among labour unions, employers and political parties that people must be induced to work at older ages in order to cope with the financial pressure resulting from the aging society. During the past two decades, a number of policy reforms have been already implemented, but additional measures are called for.

This study provides a comprehensive evaluation of the effects of a series of reforms that altered the eligibility age thresholds of two early retirement schemes and tightened medical criteria for disability pension benefits. The first reform took place in 1997 when the age threshold for the extended unemployment benefits was raised by two years. In the next year the age threshold for part-time pension was reduced by two years, but it was subsequently increased back to its original level five years later. In 2003 the possibility of qualifying for disability pension benefits under lenient medical criteria was abolished. A common feature of these reforms is that the new rules were applied only to people born after a given year. Since different cohorts were affected by the different reforms, a series of the Finnish reforms provides an exceptionally good quasi-experimental setting to identify the causal effects of various early retirement options and their interactions.

In most existing studies of early retirement, there has been no exogenous variation in eligibility criteria or benefit levels to permit credible identification of causal effects (e.g. Riphahn, 1997, Lindeboom, 1998, Kerkhofs *et al.*, 1999, and Dahl *et al.*, 2000) or identification hinges entirely on functional forms imposed by some theoretical model (e.g. Heyma, 2004). Krueger and Pischke (1992), Røed and Haugen (2003), and Brathberg *et al.* (2004) do exploit law changes for identification, but the focus of these studies is on very special pension reforms. In particular, they do not examine the role of the age thresholds, which will be the focal point of the present analysis.

The empirical analysis is based on a large sample drawn from a database of Statistics Finland, which comprises longitudinal information on the entire Finnish population from over 20 administrative registers. The effects of the strictness of medical criteria for disability pension benefits are first

examined in the difference-in-differences setting where the labour market outcomes of two groups of workers that are subject to different medical criteria are compared. The identical approach is also applied to study the role of different eligibility ages for the part-time pension. In both cases, all the workers included in the analysis were born within six months from each others, but only those born after the last day of a certain year were affected by the reform. This kind of analysis can be performed without imposing much parametric restrictions on modelling, but the results are not very well suitable for out-of-sample predictions or for assessing the joint effect of several different reforms.

Therefore, the paper proceeds to model conditional probabilities (i.e. hazard rates) for transitions out of employment into unemployment, disability retirement and nonparticipation as a function of age. This competing risks approach provides an alternative way of estimating the impact of being eligible for a certain early retirement pathway. In addition to the strictness of medical criteria and part-time pension eligibility, the impact of eligibility for the extended unemployment benefits is considered in this setting. The eligibility effects are allowed to vary with individual characteristics, so that different subgroups can respond differently to different policy changes. As before, the eligibility effects are identified by exploiting exogenous variation in the eligibility rules resulting from the policy reforms. The hazard estimates are then used to evaluate the average exit age of employment and outflows to competing exit destinations under different policy regimes. This kind of counterfactual analysis also identifies subgroups that were most strongly affected by a given reform.

According to the results, altogether the three reforms have raised the average age at which older people leave employment by 3.8 months. Much of this increase is attributed to a decline in the incidence of disability, caused by the adoption of stricter medical criteria for disability pension benefits. Also the increase in the age threshold for extended unemployment benefits has postponed the average exit age by 1.3 months. It turns out that being eligible for the part-time pension scheme has only a moderate effect on transitions out of employment. However, the scheme induces full-time workers to switch part-time, and may therefore reduce effective labour supply. It noteworthy that different subgroups have been affected by different reforms. Tightening disability criteria affected particularly strongly educated women in the public sector, postponing their exits from employment by some six months, whereas less educated men in the manufacturing sector were affected by the reform of the unemployment scheme.

The paper proceeds as follows. The next section gives details for the retirement schemes and their changes over time. It also reviews some previous studies and discusses the likely effects of different reforms in the light of existing evidence for Finland and other countries. Section 3 describes the data. This is followed by a section describing the macroeconomic environment and reporting evidence on aggregate labour market outcomes before and after the reforms. Section 5 contains the difference-in-difference analyses of the two policy reforms. Section 6 reports the hazard estimates and the associated counterfactual analysis of a series of the policy reforms. The final section concludes.

## **2 Institutional framework for Finland**

This section describes the criteria for granting various pension benefits in force from the early 1990s up to 2004, which is the last year covered by our data. The Finnish pension system is two-tiered.

The first tier is a statutory employment pension scheme. It is administrated by various pension institutions and includes several public- and private-sector pension schemes. Employment pension accrues separately from each employment relationship, and hence the benefit level is determined by the length of employment history and the amount of past earnings. The second tier is a flat-rate national pension scheme, provided by the Social Insurance Institution. National pension is paid in proportion to other pension income. It follows that national pension is only granted to individuals whose other pension income is sufficiently low due to poor earnings history. Together the employment and national pension schemes guarantee a pension benefit that is at least equal to a full national pension.

Before the official retirement age, a pension can be granted in form of early old-age, disability, part-time or unemployment pension provided that required eligibility conditions are met.<sup>1</sup> These benefits can be received up to the age when the entitlement to an ordinary old-age pension begins. The eligibility criteria for the early retirement schemes have changed several times from the mid-1990s to 2003. Because of these changes, individuals born in different years have been able to apply for a given type of pension benefits at different ages. This is illustrated in figure 1, which shows the eligibility status as a function of age for 1941 to 1948 birth cohorts.

## 2.1 Old-age pension

An *ordinary old-age pension* is payable to people who have attained the general retirement age of 65. However, a lower general retirement age exists in some public-sector employment contracts and in certain professions. Moreover, one can claim the old-age pension in the form of an *early old-age pension* before the general retirement age. This option becomes available at age 60 for the private-sector employees and at age 58 for the public-sector employees. Alternatively, one may postpone the receipt of old-age pension benefits beyond the general retirement age. If the old-age pension is claimed early (postponed), the benefit level will be permanently reduced (increased).

The old-age pension system did not change during the period under investigation. Nevertheless, the reforms of other schemes may have had spillover effects on transitions to old-age pension benefits. For example, a competing risks analysis of Kerkhofs *et al.* (1999) suggests that eligibility for an early retirement scheme lowers transition rates to disability and unemployment insurance benefits in the Netherlands. This suggests that the Finnish reforms of the unemployment and disability schemes may have induced some workers to take up early old-age pensions.

## 2.2 Disability pensions

An *ordinary disability (OD) pension* is payable to individuals aged 16 to 64 whose working capacity has decreased at least by 60% (by 40% for a partial pension). When determining eligibility, individual's chances to support himself or herself by regular work, age, education, occupation, and place of residence are taken into account along with the medical assessment. The OD pension can be granted either indefinitely (if return to work is not likely) or for a specific period. In the latter case the OD pension is also referred to as a rehabilitation subsidy, and its receipt is conditional on a rehabilitation plan.

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<sup>1</sup>A part-time pension can be received only under the employment pension scheme. There are also some pension benefits that are paid under the special acts for farmers, widows/widowers, and war veterans.

Figure 1: Age thresholds for early retirement schemes by year of birth.

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004
1941	51	52	<b>53</b>	<b>*54*</b>	<b>*55*</b>	<b>56</b>	<b>57</b>	<b>*58*</b>	<b>59</b>	<b>60</b>	<b>61</b>	<b>62</b>	<b>63</b>
1942		51	52	<b>*53*</b>	<b>*54*</b>	<b>55</b>	<b>56</b>	<b>*57*</b>	<b>58</b>	<b>59</b>	<b>60</b>	<b>61</b>	<b>62</b>
1943			51	52	<b>*53*</b>	54	<b>55</b>	<b>*56*</b>	<b>57</b>	<b>58</b>	<b>59</b>	<b>60</b>	<b>61</b>
1944				51	52	53	54	<b>*55*</b>	<b>56</b>	<b>57</b>	<b>58</b>	<b>59</b>	<b>60</b>
1945	<b>*Anticipation*</b>				51	52	53	54	<b>55</b>	<b>56</b>	<b>57</b>	<b>58</b>	<b>59</b>
1946	<b>Unemployment tunnel</b>					51	52	53	54	<b>55</b>	<b>56</b>	<b>57</b>	<b>58</b>
1947	Part-time pension						51	52	53	54	<b>55</b>	<b>56</b>	<b>57</b>
1948	Individual early retirement							51	52	53	54	<b>55</b>	<b>56</b>

The *individual early retirement (IER) pension* is another disability pension which is available for workers who have a long working career and who are unable to continue in their current job because of deteriorated health. Compared to the OD pension, eligibility for the IER pension is subject to more lenient medical criteria. The minimum degree of working incapacity is not defined and other factors, like the length of service and working conditions, have a greater weight. But, unlike the OD pension, the IER pension is payable only to workers above a certain age threshold. The age threshold was 58 until 2000 when it was raised to 60 and subsequently abolished entirely in 2004. These changes in the age threshold were applied only to people born in 1944 or later, whereas the older birth cohorts maintained their eligibility (see figure 1). Essentially, the OD and IER pension schemes are identical pathways to disability pension benefits except the latter is subject to less strict medical screening. As such, abolition of the IER pension effectively made medical screening for disability pension benefits stricter for the younger cohorts. For these reasons, OD and IER pensions are viewed as alternative routes to a general disability pension, and workers who cannot qualify for disability benefits under the IER scheme are simply subject to stricter medical criteria.

In the light of existing evidence, the strictness of eligibility criteria is expected to play an important role in determining the incidence of disability retirement among older people. Börsch-Supan (2007) analyzed cross-country differences in the disability enrolment rates of people aged 50 to 65 years, using harmonized survey data for 12 European countries and similar survey data for the UK and US. He found very little explanatory power for demographic and health-related differences across the countries, whereas three quarters of the cross-country variation were explained by institutional variables describing the generosity and the ease of access of disability insurance. The most influential institutional factor turned out to be the strictness by which vocational considerations are applied when determining eligibility.

Autor and Duggan (2003, 2006) argue that a dramatic expansion of disability insurance enrolment during the past two decades in the US cannot be explained by a true increase in the incidence of disabling illness, but the reduced stringency of the screening for disability benefits has played an important role. This development has also reduced the unemployment rate because the US disability system has begun "function much like a long-term unemployment insurance program for the unemployable" (Autor and Duggan 2006, p. 74). Karström *et al.* (2008) examined the effect of a reform in Sweden where the special rules for people aged 60-64 that allowed them to qualify for disability benefits under more lenient medical requirements compared to the younger groups were abolished. They found the reform led to increases in the unemployment and sickpay insurance reciprocity rates in the affected age group, rather than an increased employment rate. The Swedish

reform was very similar to the Finnish disability pension reform evaluated in this paper.

### 2.3 Unemployment tunnel

The *unemployment pension* is payable to people aged 60-64 who have been unemployed for at least two years. While the ordinary entitlement period of unemployment benefits is two years, workers above a certain age threshold at the time of unemployment entry can collect unemployment benefits until they turn 60 and become eligible for the unemployment pension benefit.<sup>2</sup> The combination of extended unemployment benefits and unemployment pension is known as the 'unemployment tunnel' (UT) scheme. In 1997, the age threshold for the extended unemployment benefits was raised by two years from 53 to 55. This reform was passed as a law by the parliament in September 1996, and it came into effect on January 1, 1997. However, according to the protection clause, the former age threshold was applied to workers born before 1944 who either were resigned from their job or were made redundant before June 1996 and were unemployed on January 1, 1997 (or had received unemployment benefits at least for 100 days in 1996). As a consequence of the reform, workers aged 53 or 54 at the beginning of their unemployment spells who resigned or were made redundant in June 1996 or later lost their eligibility for the UT scheme. In anticipation of the forthcoming increase in the age threshold, large numbers of people born in 1943 or earlier entered unemployment at the end of 1996 (see Kyyrä and Wilke, 2007), which must be kept in mind when interpreting our empirical figures. It should be emphasized that most people who qualify for extended benefits do not return employment but remain unemployed until retirement. Kyyrä and Ollikainen (2008) estimate that roughly one-half of workers entitled to extended benefits withdraw from job search altogether, and passively wait for early retirement.

### 2.4 Part-time pension

A gradual withdrawal from the labor market is possible through *part-time pension*, which can be awarded to a worker above a certain age threshold but less than 65 years of age who is working on a part-time basis. This requires that both the employer and employee agree on the arrangement. The part-time pension benefit is half of the difference between full-time and part-time earnings. Being on part-time pension did not reduce the level of future old-age pension before 2003. It follows that the scheme effectively subsidizes reductions in working hours rather generously.

On July 1, 1998 the age threshold of part-time pensions was temporarily lowered from 58 to 56. This change was meant to be in force until 2000, but was later postponed. In 2003, the age threshold was changed back to 58 and the old-age pension following receipt of a part-time pension was cut for those born after 1946. These changes were passed as a law by the parliament in July 2002 and they came into effect on January 1, 2003.

The objective of the part-time pension scheme is to provide an alternative for older people with reduced work capacity or work motivation who might otherwise withdraw from employment entirely. The part-time pensions are supposed to reduce particularly disability pension enrolment. If so, the scheme can provide a means of postponing retirement decisions and increasing the employment rates of older groups. On the other hand, the part-time pension scheme can be a costly measure. It provides a rather generous subsidy for a reduction in working time. It is likely that at least some

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<sup>2</sup>In practice, the employment authorities do not offer active labour market programmes to workers on extended benefits, nor monitor their job-search efforts.

people reduce their working time to gain from the subsidy without staying employed any longer. In order to the scheme be economically beneficial from the society's point of view, taking up a part-time pension should, on average, postpone notably full-time retirement.

## 2.5 Other changes in pension schemes requiring attention

Although our focus lies on changes in the eligibility age thresholds, we cannot ignore some reforms cutting early pension benefit levels. This is because these reforms were associated with protection clauses, which led to notable anticipatory behavior just before the law changes became in effect. The first one ('1996 reform') cut benefit levels for various early retirement schemes, including unemployment pensions. The reduced benefit levels apply to workers who start collecting early pension benefits in 1996 or later. This law was enacted by the parliament in September 1995. According to the protection clause, all those born before 1943 who were unemployed on January 1, 1996 remain covered by the old rules in case of early retirement (regardless of the day the early retirement event takes places in the future). Anticipation of the law change caused an excess inflow to unemployment at the end of 1995 among older employees who benefited from the protection clause (see Kyyrä and Wilke, 2007). Although this reform is not of our interest, we need to take it into account in our research design.

Another reform ('2000 reform') cut unemployment pension benefits. This cut did not apply to those born before 1945 who were unemployed on January 1, 2000 if they have either resigned from their job or were made redundant before August 1999. Hence, one may expect to see an excess inflow to unemployment in 1999 for these cohorts.

In 2005, there was a massive reform ('2005 reform') of the old-age pension scheme. It led to changes in the pension accrual rates and the length of the earnings history accounted for when determining the level of the pension benefit. In addition, the upper bound for the old-age pension benefit was abolished, and a more flexible scheme for old-age retirement, allowing individuals to freely choose their retirement age between 63 and 68, was introduced. Although these changes came into effect after our observation period, the reform may have induced anticipatory behavior in 2004. This is because for those who retire between 2005 and 2011 the pension benefits are calculated according to both the old and new rules, and the higher benefit is granted. As a consequence, some people who would have retired in 2004 in the absence of the reform may have postponed their entry into old-age pension benefits, which should be kept in mind when interpreting our results.

## 3 Data

Data for this study come from the Finnish Longitudinal Employer-Employee Database (FLEED). The worker records of FLEED have been obtained by merging information from over 20 administrative registers through the use of unique personality identity numbers. This database effectively covers all people with permanent residence in Finland, and its information content is extensive. Along with standard socio-demographic background variables, the database includes detailed information on annual income (from the tax authorities), employment (from the pension institutes), unemployment and participation in labour market programmes (from the employment offices). Importantly, there is also information on what kind of pension benefits, if any, a person is received at the end of each year.

With this source of data one can follow the entire Finnish population over time and across different labour market states. At the time of this study, a one-third random sample of all people in the database was readily accessible. These workers are observed over a period from 1990 to 2004. In practice, only certain older cohorts of the sample are needed when analyzing retirement behavior.

## 4 Economic conditions

Before turning to the econometric analysis of policy reforms, it is useful to take a brief look at labour market outcomes before and after a series of the reforms and consider changes in the economic environment. One should bear in mind that the period under investigation covers very turbulent economic times. Finland experienced a severe depression in the early 1990s. The economy contracted three years in a row (1991-1993), leading to a drop of over 10% in the GDP and pushing the unemployment rate above 16%. The depression years were followed by a period of strong economic growth that lasted until 2008 when the global economic crisis, triggered by the US subprime mortgage collapse, hit the Finnish economy.

Figure 2 shows the population shares of 40-65-year-old people occupying different labour market states at the end of 1996 and 2004. The first period offers a snapshot of the time just before the first policy reform under investigation came into effect (i.e. the two-year increase in the age threshold of extended unemployment benefits in 1997). The later one describes the time when all the reforms had been in force at least for a while. As seen in figure 2d, almost all people were on old-age pension benefits at age 65 in both years. Only 6-7% of people have postponed their entry into old-age retirement beyond the official retirement age. In 1996 a higher share of people was on old-age pension at all ages, but the difference compared to 2004 is noteworthy only at ages 61 and 63. Among those not yet awarded an old-age pension, there are striking differences in the incidence of employment, unemployment and disability retirement between the two periods.<sup>3</sup>

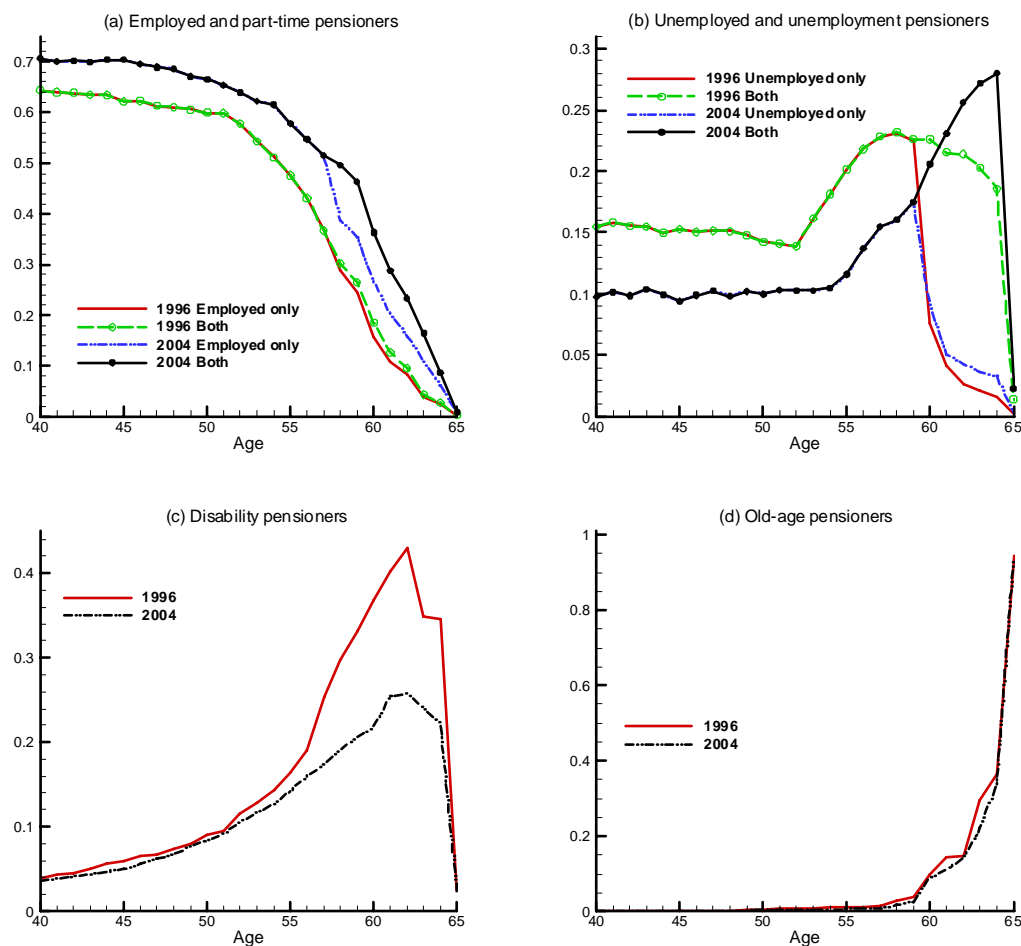
The employment rate for people in their 40s is some 5 percentage points higher in 2004 than in 1996, whereas the difference in the unemployment rates is of the same size but in the opposite direction, indicating a strong positive employment effect for the improved economic environment in 2004. Among older groups, employment has improved even much more; the employment rates of people in their late 50s and early 60s are at least some 10 percentage points higher in 2004. When also part-time pension recipients are counted as employed, the improvement in the employment level is even bigger for these groups. In that case, the employment rate in 2004 exceeds the 1996 level by some 15 to 20 percentage points. The employment gains seem to be associated with the sharp drop in disability pension enrolment rates (figure 2c), perhaps resulting from the abolition of the IER scheme.

Despite the dramatic improvement in the economic environment from the mid-1990s to 2004, the incidence of being on unemployment-related benefits has increased for people above 60 years of age. At the end of 2004, one in four of people aged 62 to 64 were granted an unemployment pension or received unemployment benefits (figure 2b). At a glance, these numbers seem puzzling, as one might have expected to find lower unemployment rates for the oldest groups in 2004 as a result of

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<sup>3</sup>Only salary and wage earners are counted as employed in figure 2, because their behaviour will be analyzed. Adding entrepreneurs to the numbers would increase the employment levels, without altering the shape of the curves.

Figure 2: Population shares in different labour market states at the end of 1996 and 2004 by age (Source: Author's computations from FLEED)



improved economic conditions and the reforms of the UT scheme in the late 1990s and 2000 that tightened the eligibility criteria and cut benefit levels. Furthermore, it is unlikely that the reformed disability pension scheme plays a role here. The abolition of the IER scheme affected the cohorts that were 60 years of age or less at the end of 2004 (see figure 1), so that the stricter eligibility criteria for disability pension benefits cannot explain the large shift from disability benefits towards unemployment-related benefits in the oldest group. The astonishing levels of unemployment benefit receipt for the oldest people in 2004 arises, at least partly, from anticipation behavior associated with the implementation of the UT scheme reforms. As seen in figure 1, people aged 60 or more at the end of 2004 have been covered by the protection clause of at least one reform affecting the UT scheme. Consequently, many of these people entered unemployment in 1995, 1996 or 1999 in anticipation of the next year's reform, ending up to collect unemployment benefits for several years until qualifying for an unemployment pension at age 60. See Kyyrä and Wilke (2007) for evidence of excess unemployment inflows in 1995 and 1996.

To summarize, the employment rates of people in their late 50s and early 60s have increased substantially from the mid-1990s to 2004. While the flourishing economic environment since the mid-1990s explains part of the improvement, it is not the whole story. The sharp decline in the disability enrolment rates together with the increasing number of the oldest people on unemployment pension benefits point to an important role for the pension reforms. In what follows, the paper aims at separating the causal effects of distinct policy changes from the business cycle effects, while making a distinction between the anticipatory and permanent policy effects, that is, between transitory versus long-term effects.

## 5 Difference-in-differences analysis

### 5.1 Abolishing the individual early retirement scheme

As part of a larger reform package, the individual early retirement scheme was removed in 2000 from all workers born in 1944 or later. The oldest affected workers were 56 at the time when the change became into effect, and hence below the pre-reform age threshold of 58, ruling out anticipation behavior towards exits into IER benefits. At a glance, this seems to provide a sound regression discontinuity setting for evaluating the impact of the reform, where the labour market outcomes of workers born just before and after the last day of 1943 would have been compared at age 58 and beyond. But, as seen below, there were anticipatory behavior toward exits to other labour market states in 1999 and an asymmetric long-term anticipatory effect of the past UT scheme reform taking place in 1996. To deal with these issues, one needs to consider the labour market states occupied at younger ages as well.

Specifically, two worker groups are considered: those born in the last quarter of 1943 ('comparison group') and those born in the first quarter of 1944 ('treatment group'). These people are roughly of the same age but the IER pension option was taken away only from those born in 1944. By comparing these two groups over time, one can assess to what extent eligibility to the IER pension scheme affects various labour market outcomes.<sup>4</sup> From the last quarter of 2001 onwards, when the 1943 cohort attained the age threshold of 58 for the IER pension (see figure 1), the groups have been in different positions with respect to medical requirements for disability pension eligibility.

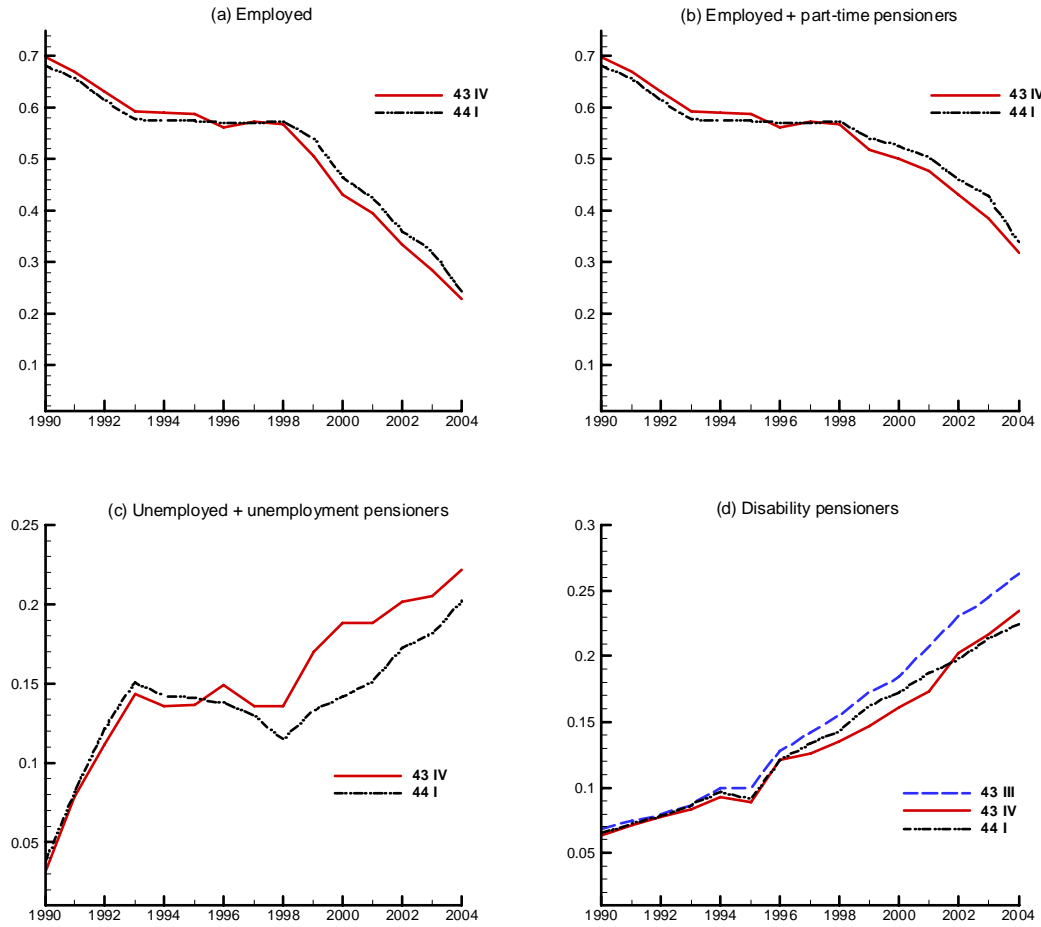
As seen in figure 3a, the employment rates of the two cohorts were very similar until 1998, but since then the employment rate of the 1944 cohort has been at a higher level. Adding part-time pensioners to the employed stock does not change the overall picture (figure 3b). It is noteworthy that the gap in employment is roughly constant over the period 1999–2004, even though the workers born in 1943 did not qualify for the IER scheme before 2001. So, the higher employment rate of the 1944 cohort must arise from some sort of anticipatory behavior taking place already in 1999.

Unemployment patterns in figure 3c illustrate the importance of asymmetric anticipatory effects. The unemployment rates were rather similar until 1997 with the exception of a temporary increase in 1996 for the 1943 cohort. Unlike the marginally younger cohort, workers in the 1943 cohort reached the age of 53 at the last quarter of 1996 and hence were covered by the protection close of the 1997 reform of the UT scheme, which led to the excess flow into unemployment at the end

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<sup>4</sup>By 'eligibility' it is meant that the worker may qualify for IER pension benefits because of his or her age. As discussed previously, other criteria about medical and occupational conditions must be met before an IER pension can be granted.

Figure 3: Labour market state at the end of year by birth quarter



1996. Interestingly, the unemployment level of this group rebounded notably by the end of the next year, but still remained above the level of the 1944 group. The difference in the unemployment rates increased from 1998 to 2000, but then decreased over the next years. Recall that the workers born in 1943 regained their eligibility for extended unemployment benefits in 1998, which explains the increase in that year, but the additional increase in 1999 must be related to the differential anticipatory effects associated with the 2000 reform. The larger growth rate of unemployment from 2000 onwards in the 1944 group suggests that those who cannot apply for IER benefits are subject to a higher risk of ending up unemployed.

In figure 3d no differences in the disability enrolment rates existed before 1997.<sup>5</sup> It is surprising to find that the incidence of disability pension receipt is lower over the period from 1997 to 2001 for the 1943 cohort than for the marginally younger group. Moreover, this appears to be true only for the chosen comparison group born in the last quarter of 1943, not for those who were born earlier

<sup>5</sup>Statistics Finland changed its procedure of merging register data on pension benefits in 1995, which led to the transitory drop in the number of disability pension recipients for that year.

in the same year. This is illustrated in figure 3d by showing the disability pension enrolment rate also for workers born in the third quarter of 1943. The 1943 cohort reached the age threshold for the IER scheme in 2001, which led to an acceleration of their disability enrolment rate compared to the non-eligible 1944 cohort. This suggests that eligibility to the IER scheme has increased the entry rate to disability pension benefits at age 58 and thereafter. Obviously, the same conclusion is drawn taking either the third or last quarter of 1943 as a comparison group.

To assess the statistical significance of the between-group differences in the labour market outcomes, the following simple linear probability model was estimated:

$$Y_{it} = \mathbf{X}_{it}\boldsymbol{\beta} + \eta D_i + \delta_t (D_i \times R_t) + \lambda_t + \varepsilon_{it}, \quad (1)$$

where  $i$  indexes individuals and  $t$  is the time period or, equivalently, age.  $Y_{it} = 1$  if the worker occupies the labour market state of interest (employment, unemployment or disability) at the end of year  $t$ , and  $Y_{it} = 0$  otherwise.  $D_i = 1$  if the worker was born in the first quarter of 1944, and  $D_i = 0$  if in the last quarter of 1943.  $R_t = 0$  for the reference period prior to the policy reform, and  $R_t = 1$  for all the subsequent periods.  $\mathbf{X}_{it}$  is a vector of control variables for marital status, own house, education level, and living region. The general age/time pattern common to all workers is captured by  $\lambda_t$ 's. The difference between the two groups in the reference year is measured by  $\eta$ . Under the assumption that this gap would have remained constant over time without the policy change,  $\delta_t$  measures the causal effect of abolishing the IER scheme on  $E(Y_{it} | \mathbf{X}_{it})$ , the probability of being in the state of interest at the end of year  $t$  given  $\mathbf{X}_{it}$ .

The model was estimated for two periods: 1997–2004 and 2000–2004. In both cases, the first year of the time interval serves as the reference period (i.e.  $R_{1997} = 0$  or  $R_{2000} = 0$ ). The results obtained from the different time intervals should be interpreted in different ways, however. At the end of 1997 neither group was eligible for extended unemployment benefits, nor for part-time pension, and no anticipation behavior related to the 2000 reform did occur so early. Thus, when the 1997–2004 period is analyzed,  $\eta$  captures the effect of the small between-group age difference plus the difference due to anticipation behavior of the 1943 cohort taking place in 1996 in response to the 1997 reform. It follows that  $\delta_t$  describes the effect of abolishing the IER scheme as part of the 2000 reform. To be specific, the counterfactual outcomes reflect all the other features of the 2000 reform, including anticipatory behavior triggered by the protection clause associated with the reduction in the unemployment pension benefits for those who become unemployed after January 1, 2000. The difference-in-differences (DiD) estimates of  $\delta_t$  obtained from the 1997–2004 period should therefore be interpreted as *transitory* effects that apply only to the narrow birth cohorts potentially affected by the protection clause of the 2000 reform. It should be emphasized that both the treatment and comparison group were covered by the protection clause, but these groups may have reacted differently due to different outside options in the future.

When estimated from the shorter time period, the model aims to capture the *persistent* effect of abolishing the IER scheme in absence of anticipatory behavior. At the end of 2000, both the birth cohorts were in similar position with respect to their eligibility for extended unemployment benefits and part-time pension, while those born in 1943 were 57 years old and thereby not yet entitled to the IER pension. So, when the reference period is 2000,  $\eta$  captures also (possibly asymmetric) anticipatory effects associated with the 2000 reform taking place in 1999 in addition to the same source of the between-group differences as in the case of the longer time period.

Table 1: DiD estimates of the effects of abolishing the IER scheme scheme

Period	Age of treated	Transitory effect			Persistent effect		
		coeff.	ste	t	coeff.	ste	t
<b>A. Employed</b>							
1998	54	0.009	0.006	1.60			
1999	55	<b>0.037</b>	0.008	4.71			
2000	56	<b>0.035</b>	0.010	3.56			
2001	57	<b>0.030</b>	0.011	2.78	-0.006	0.006	-0.88
2002	58	<b>0.029</b>	0.011	2.53	-0.007	0.008	-0.81
2003	59	<b>0.038</b>	0.012	3.21	0.003	0.009	0.30
2004	60	0.015	0.012	1.23	-0.020	0.010	-1.91
<b>B. Employed + part-time pensioners</b>							
1998	54	0.009	0.006	1.61			
1999	55	<b>0.023</b>	0.008	2.93			
2000	56	<b>0.028</b>	0.009	3.18			
2001	57	<b>0.027</b>	0.010	2.77	-0.001	0.006	-0.23
2002	58	<b>0.031</b>	0.010	2.98	0.003	0.008	0.42
2003	59	<b>0.045</b>	0.011	4.02	0.018	0.009	1.92
2004	60	0.021	0.012	1.70	-0.007	0.011	-0.68
<b>C. Unemployed</b>							
1998	54	<b>-0.015</b>	0.006	-2.63			
1999	55	<b>-0.030</b>	0.007	-4.13			
2000	56	<b>-0.040</b>	0.008	-5.08			
2001	57	<b>-0.030</b>	0.008	-3.52	0.010	0.005	1.92
2002	58	<b>-0.023</b>	0.009	-2.51	<b>0.017</b>	0.006	2.77
2003	59	-0.016	0.009	-1.76	<b>0.023</b>	0.007	3.32
2004	60	-0.012	0.010	-1.22	<b>0.028</b>	0.008	3.68
<b>D. Disability retirement</b>							
1998	54	0.000	0.003	-0.03			
1999	55	0.008	0.004	1.73			
2000	56	0.004	0.005	0.67			
2001	57	0.006	0.006	1.05	0.003	0.003	0.83
2002	58	-0.013	0.007	-1.84	<b>-0.016</b>	0.005	-3.17
2003	59	-0.011	0.008	-1.41	<b>-0.014</b>	0.006	-2.31
2004	60	<b>-0.018</b>	0.008	-2.14	<b>-0.021</b>	0.007	-2.98
Reference period			1997			2000	
# of individuals			7781			7626	
# of observations			60,711			37,505	

Notes: The unemployed also includes unemployment pension recipients. Disability retirement refers to receipt of either OD and IER pension benefits. The control variables include dummies for marital status, own house, educational level (4 categories), and leaving region (20 NUTS3 regions). Reported standard errors and t-values are robust to intragroup correlation of repeated observations on the same individuals. The point estimates that are significantly different from zero at the 95 percent confidence level are in bold.

The estimates of  $\delta_t$  obtained from the two time periods are reported in table 1.<sup>6</sup> Consider the transitory effects first. As seen in panels A and B, abolition of the IER scheme increased the employment probability between 1999 and 2003 when the affected workers were 55 to 59 years old. This is mainly driven by the higher incidence of unemployment of the 1943 cohort between 1998 and 2002 (see figure 3c). The employment effect drops and loses its statistical significance – though only at the 95% confidence level – in 2004, which is driven by a sharp decline in the employment rate of the 1944 cohort (see figure 3a). These workers turned 60 and hence those who were covered by a private-sector pension scheme qualified for an early old-age pension in the first quarter of 2004. Given that there is no evidence of a similar drop in employment for the 1943 cohort in 2003, when they turned 60, nor in the next year, it seems that ineligibility to an IER pension increases the take-up rate of the early old-age pension.

There is a weak transitory effect on the incidence of disability retirement at the end of the observation period (see panel D). Namely, abolition of the IER scheme reduced the disability enrolment rate by some two percentage points in 2004. This effect is almost of the same size than the persistent effect in 2004. Note the abolition of the IER scheme has a statistically significant persistent effect on disability retirement participation also at ages 58 and 59. There is no evidence of persistent employment effects, given that none of the persistent effects in panels A or B differ statistically from zero at the 95% confidence level (although the persistent effect at 59 is significant positive at the 90% confidence level). In contrast to the transitory unemployment effects, the persistent effects on unemployment are positive and statistically significant from age 58 onwards. Overall, the DiD estimates of the persistent effects lend support to the claim that exits to unemployment and disability pension schemes are substitutes. Specifically, being unable to apply for an IER pension decreases the disability enrolment rate, but also increases the unemployment rate. As a result of the similar magnitude of these two effects, there is no effect on the likelihood of being employed.

To summarize, during the period 2001–2004, the workers born in 1943 had lower transition rates into disability pension benefits and a higher incidence of unemployment than the marginally older group, whereas the evolution of the employment rate was rather similar for the 1943 and 1944 groups. That is, a lack of the IER scheme seems to lower the disability risk at the expense of the increased unemployment risk, without affecting notably the employment probability. These are the long-run effects that should be relevant also for the cohorts born after 1944. In addition, the implementation of the 2000 reform led to anticipatory behavior that had substantial short-term effects on labour market outcomes for the certain cohorts.

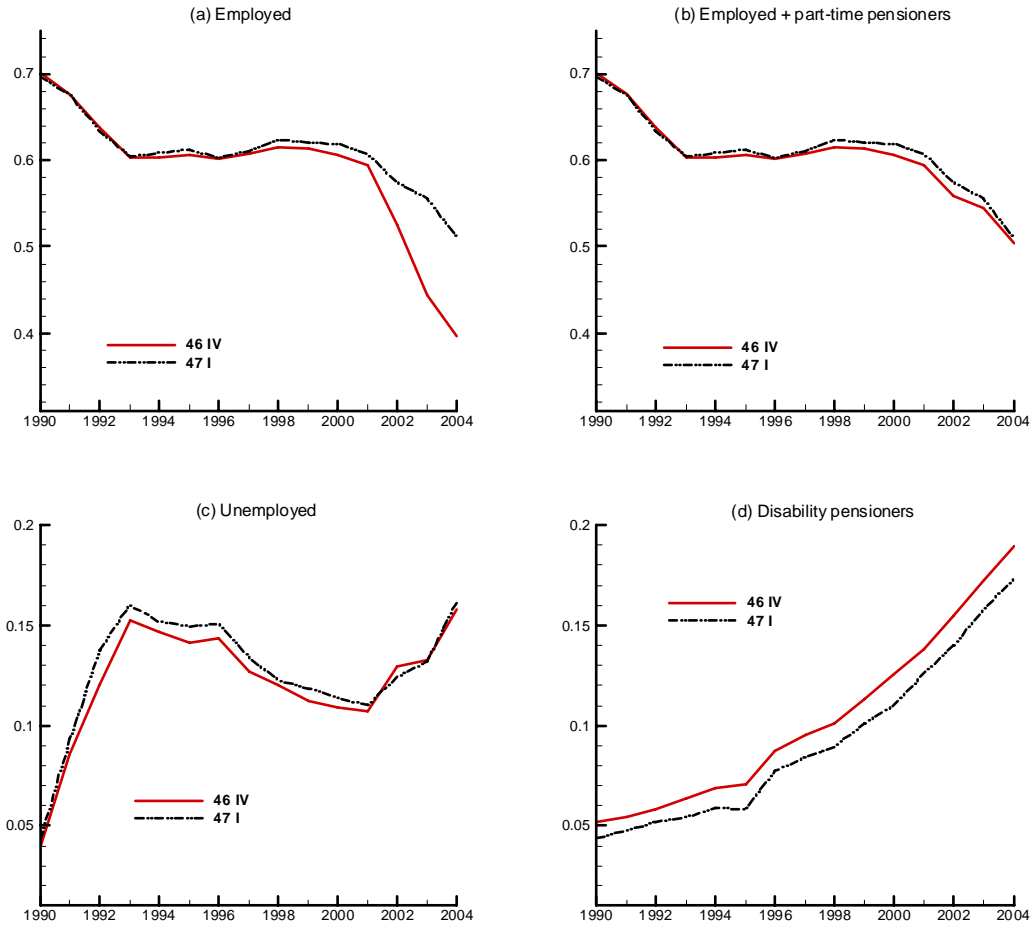
## 5.2 Different age thresholds for part-time pension

In July 2003 the age threshold for part-time pension was increased from 56 to 58 for those born after 1946. When the change came into effect, the affected cohorts were 55 years old or younger, and thereby unable to advance their entry into part-time retirement in anticipation of the reform. Hence, one can study the effect of the increase in the age threshold by comparing labour market outcomes between workers born in the last quarter of 1946 ('comparison group') and those born in the first quarter of 1947 ('treatment group'). It should be stressed that also an old-age pension following receipt of a part-time pension was reduced for those born after 1946. By implication,

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<sup>6</sup>Controlling for  $\mathbf{X}_{it}$  has hardly any effect on the point estimates, nor on the standard errors

Figure 4: Labour market state at the end of year by birth quarter



the difference-in-differences setting identifies the joint effect of the two-year increase in the age threshold and somewhat lower economic incentives for part-time pension. Unlike in the case of the abolition of the IER scheme, there are no other reforms that would have affected the 1946 and 1947 cohorts asymmetrically. So there should not be pre-reform differences between the two groups calling for particular attention, nor the issue about the transitory and persistent effects.

As shown in figure 4a, the two groups had very similar employment rates until 2001. In the next year the 1946 cohort turned 56, and thereby became eligible for a part-time pension. As a result, their employment rate began to decline sharply, being only 40% at the end of 2004, when 51% of the 1947 cohort were still employed. When also part-time pensioners are counted employed, the share of those still working at the end of the observation period is 50% for the 1946 group (see figure 4b), which is only one percentage point below the level of the treatment group. In figure 4c, there is no evidence that eligibility for part-time pension benefits would have reduced the incidence of unemployment. By contrast, the unemployment rate of the 1946 cohort increased at a marginally faster rate in the last years of the observation period, when they were already attained

Table 2: DiD estimates of the effects of the 2-year increase in part-time pension eligibility

Period	Age of treated	Persistent effect		
		coeff.	ste	t
<b>A. Employed</b>				
2001	54	0.000	0.005	0.05
2002	55	<b>0.036</b>	0.007	5.26
2003	56	<b>0.098</b>	0.008	12.13
2004	57	<b>0.101</b>	0.009	11.85
<b>B. Employed + part-time pensioners</b>				
2001	54	0.000	0.005	0.06
2002	55	0.002	0.007	0.32
2003	56	-0.002	0.007	-0.31
2004	57	-0.004	0.008	-0.55
<b>C. Unemployed</b>				
2001	54	-0.001	0.005	-0.22
2002	55	-0.010	0.006	-1.69
2003	56	-0.005	0.007	-0.81
2004	57	-0.001	0.007	-0.11
<b>D. Disability retirement</b>				
2001	54	0.003	0.003	1.12
2002	55	0.001	0.004	0.20
2003	56	0.000	0.005	0.02
2004	57	-0.002	0.005	-0.39
Reference period		2000		
# of individuals		10,834		
# of observations		53,484		

Notes: The unemployed also includes unemployment pension recipients. Disability retirement refers to receipt of either OD and IER pension benefits. The control variables include dummies for marital status, own house, educational level (4 categories), and leaving region (20 NUTS3 regions). Reported standard errors and t-values are robust to intragroup correlation of repeated observations on the same individuals. The point estimates that are significantly different from zero at the 95 percent confidence level are in bold.

the age threshold for part-time pension. The rate of disability pension receipt is somewhat higher for the 1947 cohort in all years, which is surprising as they are younger, albeit only a few months. But as this difference does not change over the last years, it cannot be related to the differing age thresholds for part-time pensions.

Testing statistical significance can be done by applying a difference-in-differences approach similar to one applied in evaluating the effect of abolishing the IER scheme. Specifically, the model outlined in (1) was estimated using data on individuals who were born either in the last quarter of 1946 ( $D_i = 0$ ) or in the first quarter of 1947 ( $D_i = 1$ ). The observation period used covers years 2000–2004. The reference period is 2000 (i.e.  $R_{2000} = 0$ ) when both the groups were still covered by the conventional unemployment benefits rules. The same set of control variables was used as previously. The DiD estimates of the policy change – the two-year increase in the eligibility age of part-time pension benefits plus a cut in the future old-age pension benefits – on the likelihood of different labour market outcomes,  $\delta_t$ 's, are reported in table 2.

Only statistically significant effects can be found in panel A, suggesting that a higher rate of full-time employment is the only effect of the policy change. In particular, there is no evidence of significant effects on the incidence of unemployment or disability retirement. These estimates should be interpreted with keeping in mind that the affected workers in the analysis were 57 years old at the end of 2004. It is possible that being on a part-time pension has an effect on unemployment or disability but only at the older ages, in which case the effect is simply not recovered in the present setting.

## 6 Competing risks analysis of exits from employment

In the DiD analysis above, a single reform was considered at a time and the outcome variable was the likelihood of occupying a given labour market state as opposed to occupying any other state. In this section, the analysis is completed by modelling the eligibility effects of different policy schemes on transition rates out of employment to different destinations simultaneously. The transition probabilities conditional on surviving employed up to a given age takes into account that the occurrence of one event (e.g. a dismissal) removes the individual from the risk of the other events (e.g. exit from work to a disability pension or to an early old-age pension). In this way, the selection issues resulting from anticipation behavior induced by the protection clauses of some reforms can be dealt with. Finally, the estimates of a structural parametric model can be used to evaluate expected counterfactual labour market outcomes associated with various combinations of early retirement schemes.

### 6.1 Sample and descriptive statistics

The target population includes people born between 1941 and 1948 who worked for the local or central government, or held a private-sector job covered by the Employees' Pension Act (TEL) at the end of the year when they were 51 years old (see figure 1).<sup>7</sup> In addition, it is required that the worker was employed at least for the past two years, had worked for the current employer at least one year, had annual earnings no less than €6000 and did not receive any pension benefits in the year when he or she turned 51. The resulting sample obviously represents people with a strong labor market attachment in their 50's, which should be kept in mind when interpreting the results.

To reduce computational burden, a 30% random sample of the original sample was drawn to be used for estimating the hazard models. These workers are followed until the end of the observation period or until they leave employment for one of the competing reasons: unemployment, disability retirement or nonparticipation. The labour market states are observed at the end of each year. Part-time pensioners are classified to be employed. Nonparticipation includes all other states outside the labour force except being on a disability pension. Although some workers died or emigrated, the most common reason for leaving the labour force without a disability pension at older ages is being awarded an (early) old-age pension benefit. Among workers aged 58 and above, 84% of exits to nonparticipation is attributed to receipt of an old-age pension benefit.

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<sup>7</sup>These sample restrictions drop entrepreneurs and those who work for Church, Bank of Finland, the Social Insurance Institution of Finland or the Åland Parliament, as well as a few small private-sector employee groups who are covered by the Temporary Employees' Pension Act (LEL) or by the Pension Act for Performing Artists and Certain Groups of Employees (TaEL).

Table 3: Mean characteristics at age 51 by exit destination

	Exit destination:				
	All (1)	Unempl. (2)	Disability (3)	Nonpart. (4)	Censored (5)
Female	0.537	0.512	0.549	0.521	0.546
Married	0.747	0.725	0.739	0.764	0.751
Own house	0.846	0.848	0.841	0.859	0.844
Living region:					
Uusimaa	0.326	0.255	0.282	0.380	0.345
Southern Finland	0.372	0.448	0.370	0.344	0.355
Eastern Finland	0.103	0.107	0.132	0.095	0.098
Mid-Finland	0.117	0.110	0.121	0.098	0.122
Northern Finland	0.081	0.079	0.094	0.083	0.079
Tenure < 2 years	0.057	0.064	0.046	0.054	0.057
Sector:					
Public sector	0.386	0.132	0.449	0.521	0.423
Manufacturing	0.256	0.438	0.253	0.155	0.222
Trade	0.123	0.174	0.100	0.096	0.118
Transportation	0.054	0.075	0.063	0.042	0.049
Services	0.090	0.111	0.059	0.094	0.089
Other	0.090	0.070	0.076	0.091	0.099
Unemployment rate	0.201	0.213	0.211	0.207	0.194
Schooling years	11.290	10.480	10.720	11.636	11.576
Earnings, €/year	29,261	26,450	26,119	32,363	30,081
Taxable wealth, €/year	19,792	17,659	17,755	20,968	20,598
# of individuals	19,674	3493	2259	2408	11,514

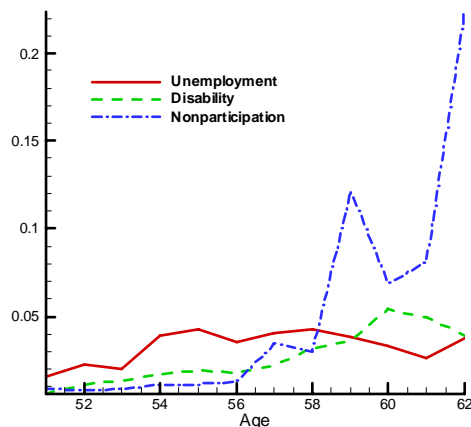
Notes: Mean values are computed at the age of 51. Workers with more than one employment spells are classified according to the exit destination of the last spell. Censored observations include those who were employed at the end of 2004. Southern Finland also includes Åland.

For the age groups analyzed here the transition out of employment is typically an absorbing state. However, a small group of workers (3.5%) in the sample exited from employment but then returned employment during the observation period. For these individuals more one employment episode is recorded and used in the analysis. Overall, the sample used in the estimations contains 131,162 year-specific observations on 19,674 individuals.

Table 3 reports the sample means of some background variables for all the workers included in the analysis (column 1) and for subgroups leaving employment through different exit pathways (columns 2 to 4). The last column reports the sample means for those whose exit destination was not observed during the observation period, as they were still employed at the end of 2004. The variables are measured at the first year of the observation period, when the workers were 51 years old.

From the first rows one can see that there are no differences in exit pathways between women and men, nor between married and non-married people. Entries to disability pension may be slightly more common in Eastern Finland than in other parts of the country, whereas workers living in Uusimaa are less likely to end up unemployment. Overall these regional differences are relatively small. There are more notable discrepancies in the exit pathways between individuals holding

Figure 5: Empirical hazard rates



jobs in different sectors. Public-sector employees are much less likely to be laid off (or quit for unemployment), but more likely to be granted a disability pension or moving to nonparticipation compared to other workers. By contrast, manufacturing workers have the highest risk of becoming unemployed and a relatively low risk of leaving work for nonparticipation. A job tenure less than 2 years may increase the risk of unemployment and lower the likelihood of being granted a disability pension. Compared to those entering unemployment or a disability pension, workers who survived employed until the end of the observation period or those who left work for nonparticipation are, on average, somewhat better educated, have higher earnings and higher taxable wealth.

As seen in the last row of the table, over a half of the individuals did not leave employment by the end of the observation period. Such a high rate of censoring is not surprising given that the youngest cohorts in the data were less than 60 years old in 2004. The most popular exit destination is unemployment, whereas exits to disability pension and nonparticipation are almost equally common. It is worth emphasizing that the truncation in the observation period implies that many exits to old-age pension (i.e. to nonparticipation) are not observed. This is evident in figure 5, where raw hazard rates for the three competing destinations are depicted.<sup>8</sup> By comparing the hazard rates, it clear that a transition to nonparticipation is the by far most likely outcome for workers who remain employed until age 62. In fact, exits to nonparticipation are very rare until age 56, but then there is a level shift in the nonparticipation hazard at age 57. This is because the public-sector employees become eligible for early old-age pension during the next year. The much larger group of private-sector employees qualifies for the early old-age pension two years later, which explains the sharp peak in the nonparticipation hazard at age 59.

The hazard rate for exits to disability pension increases until age 60 when it starts to decline. 57-year-old workers born before 1944 reach the age threshold for the IER scheme during the next year, which probably explains a sharply increasing pattern of the disability hazard rate between ages 57 and 60. The declining pattern of the hazard thereafter may indicate that the early old-age

<sup>8</sup>The hazard rate for exit destination  $j$  at age  $a$  is defined as the fraction of people employed at the end of the year when they were  $a$  years old who occupied state  $j$  by the end of the next year.

pension, which is available for all workers aged 60 or more, provides an alternative exit pathway for those with reduced working capacity. The hazard rate to unemployment doubles at age 54, as most people become eligible for extended benefits during the next year for the first time. It is noteworthy that unemployment is the most common reason for leaving employment until age 58.

The hazard rates in figure 5 show the magnitude of different transition probabilities at different ages. But they do not tell much about the importance of the early retirement options. Firstly, in the pooled sample, some of people employed at a given age can qualify for a certain pension scheme while others cannot, depending on their birth year. Secondly, in order to be at risk of exiting from employment at a given age, one has to remain employed until that point (or to be re-employed if exited at a younger age). Because this dynamic selection process is driven by observed and unobserved characteristics of individuals, the composition of the employed at older ages can be very different from the initial sample employed at age 51. In order to recover the causal effects of pension scheme eligibility, the dynamic selection issues must be dealt with.

## 6.2 Mixed logit model for transition probabilities

Consider a worker  $i$  who is  $a$  years old and employed at the end of year  $t$ . One year later the worker can either be unemployed ( $j = U$ ), on a disability pension ( $j = D$ ), outside the labour force for another reason ( $j = O$ ), or be still employed ( $j = E$ ). The log odds of leaving employment for state  $j$  by age  $a + 1$  (i.e. by the end of year  $t + 1$ ) rather than staying employed for one additional year is given by

$$\ln \left( \frac{h_{ij}(a | \mathbf{v}_i)}{h_{iE}(a | \mathbf{v}_i)} \right) = \alpha_{aj} + \mathbf{X}_{ia}\boldsymbol{\beta}_j + \varphi_j \text{Early}_{ia} + \eta_j \text{IER}_{ia} + \delta_j \text{PTP}_{ia} + \lambda_j \text{UT}_{ia} + v_{ij}, \quad (2)$$

where  $h_{ij}(a | \mathbf{v}_i) \equiv \Pr(a < A_i \leq a + 1, J = j | A_i > a, \mathbf{X}_{ia}, \text{IER}_{ia}, \text{PTP}_{ia}, \text{UT}_{ia}, \mathbf{v}_i)$  is the probability of exit to state  $j$  by age  $a + 1$  for an employed worker with covariates  $\mathbf{X}_{ia}$ , retirement options  $\text{IER}_{ia}$ ,  $\text{PTP}_{ia}$  and  $\text{UT}_{ia}$ , and unobserved factors  $\mathbf{v}_i = (v_{iU}, v_{iD}, v_{iO})$ . The reference event is staying employed until the end of the next year (i.e. until age  $a + 1$ ), which probability is  $h_{iE}(a | \mathbf{v}_i)$ . The hazard function for exit to state  $j \in \{U, D, O\}$  at age  $a$  is

$$h_{ij}(a | \mathbf{v}_i) = \frac{\exp(\alpha_{aj} + \mathbf{X}_{ia}\boldsymbol{\beta}_j + \varphi_j \text{Early}_{ia} + \eta_j \text{IER}_{ia} + \delta_j \text{PTP}_{ia} + \lambda_j \text{UT}_{ia} + v_{ij})}{1 + \sum_k \exp(\alpha_{ak} + \mathbf{X}_{ia}\boldsymbol{\beta}_k + \varphi_k \text{Early}_{ia} + \eta_k \text{IER}_{ia} + \delta_k \text{PTP}_{ia} + \lambda_k \text{UT}_{ia} + v_{ik})}, \quad (3)$$

and that for staying employed until the end of the next year is  $h_{iE}(a | \mathbf{v}_i) = 1 - \sum_k h_{ik}(a | \mathbf{v}_i)$ . Essentially, the model is a discrete-time competing risks model for the length of the employment career beyond age 51, and  $h_{ij}$ ,  $j = \{U, D, O\}$  are the parametrized versions of the raw hazard rates shown in figure 5. This modelling approach is convenient because it allows us to deal with the competing exit routes, time-varying regressors (e.g. age and eligibility status for different retirement schemes) and censored observations (persons still employed at the end of 2004).

The covariates in  $\mathbf{X}_{ia}$  include a regional-specific unemployment rate, gender, marital status, years of schooling, earnings, a dummy for a short employment relationship, sector dummies, year fixed effects, and 'anticipation dummies', defined as interaction terms for workers born between 1941 and 1942 in 1995, for those born between 1941 and 1943 in 1996, and for those born between 1941 and 1944 in 1999 (see the cells marked with stars in figure 1). The earnings are interpreted as a proxy variable for earnings potential. The lagged values are used because receipt of disability

pension benefits typically precedes periods of sick leave, which can reduce earnings while still employed. The anticipation dummies allow for a temporary increase in the likelihood of the exit rate one year before the reforms of 1996, 1997 and 1999 for those workers who were covered by the protection clauses. The year fixed effects capture the effects of economic environment, attitudes towards early retirement and general improvement in health that change over time in the same way. Some variables reported in table 3 did not have statistically significant effects, and were thereby excluded from the covariate set.

The age-specific fixed effects ( $\alpha_{aj}$ 's) allow for fully flexible age profiles for the hazard rates.  $Early_{ia}$ ,  $IER_{ia}$ ,  $PTP_{ia}$  and  $UT_{ia}$  are dummy variables for eligibility to apply for an early old-age pension, for an IER pension, for a part-time pension and for extended unemployment benefits during the next year, respectively. Their coefficients are the parameters of primary interest, capturing the eligibility effects of various early retirement schemes. Given that all the eligibility variables enter all the odds ratios, the substitution effects are explicitly allowed for. As an example, eligibility for the IER scheme can directly affect the odds ratios between unemployment and employment (as measured by  $\eta_U$ ), and between nonparticipation and employment (as measured by  $\eta_O$ ).

Just like in the DiD analyses before, identification of the effects of  $IER_{ia}$ ,  $PTP_{ia}$  and  $UT_{ia}$  hinges on variation caused by the pension reforms. The eligibility effects and unrestricted age effects are separately identified because of variation in the eligibility age thresholds across individuals born in different years (see figure 1). Since eligibility for different early retirement schemes are determined by age and birth year only, the eligibility variables are necessarily independent of the unobservables, and hence their coefficients have a causal interpretation.<sup>9</sup> Although the early old-age pension scheme did not change during the observation period, the different age thresholds for private- and public-sector employees allow the estimation of the coefficient of  $Early_{ia}$ . However, the causal interpretation is questionable in this case, because workers are not randomly allocated in the different sectors.

Variation stemming from the reforms does not eliminate the need to control for unmeasured differences between individuals, which can arise for a variety of reasons. For example, people are likely to differ in their valuation of leisure. Some may also have permanent physical or mental defects that affect their possibilities to continue working at older ages. In general, workers with lower employability for unobserved reasons leave employment at a higher rate at each age, and hence their proportion among those still employed at a given age decreases with age. As a result, failure to control for unobserved heterogeneity tend to lead to a downward bias in the estimates of the hazard rates at older ages. This is serious cause for concern in our case, because the early retirement schemes are not readily available at age 51, but one has to survive in employment for some years before qualifying for such schemes. For these reasons, the model incorporates three heterogeneity terms, which are normally distributed and potentially correlated. As a consequence, the model avoids the Independence of Irrelevant Alternative assumption, which has been seen as a major weakness of the multinomial logit specification. For example, a less motivated worker may consider all alternatives to escape employment, suggesting a positive correlation between the

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<sup>9</sup>This also explains why our model includes a dummy variable for eligibility for the part-time pension scheme, but not for actual receipt of a part-time pension benefits. Although the eligibility is an exogenous covariate, the eligible individual's choice to exploit the part-time pension option is certainly not. To estimate the causal effect of the receipt of part-time pension benefits, one has to deal with the self-selection problem, which calls for a much more complicated econometric approach. While such an effect of obvious interest, its estimating is left for future work.

unobservable factors.

It is worth noting that the parameters of the mixed logit model should be interpreted with some caution. The coefficients of the covariates have a straightforward interpretation in terms of the effects on the odds ratios. Consider, for example, the effect of being eligible for extended UI benefits.  $e^{\lambda_U}$  gives the proportional effect of UT eligibility on the odds ratio between unemployment and employment, but it does not tell how the hazard rate for exit to unemployment is affected by the UT scheme. In addition to  $\lambda_U$ , the latter effect depends also on  $\lambda_D$  and  $\lambda_O$ , and more generally, on the values of all the conditioning covariates. If UT eligibility also increases transitions to disability retirement and/or nonparticipation, it is possible that the odds ratio of becoming unemployed vs. staying employed increases but the conditional probability (the hazard rate) of becoming unemployed decreases. This may be the case for some people but not for all, depending on the value of other covariates. Nor there is a single parameter that would tell us how the probability of remaining employed until a given age (the survival rate) are the probability of eventually exiting from employment to unemployment are affected by a change in the UT scheme. For these reasons, the mixed logit estimates will be summarized by computing the probabilities of interest under different policy schemes. Before proceeding to such an analysis, the parameters of the mode must be estimated, which is done by the method of (simulated) maximum likelihood.

### 6.3 Estimation

Suppose that worker  $i$  leaves employment for state  $j$  between ages  $a$  and  $a + 1$ . The likelihood contribution, conditional on observed characteristics and unknown parameters, is

$$f(a_i, j_i) = \int S_i(a | \mathbf{v}_i) h_{ij}(a | \mathbf{v}_i) \phi(\mathbf{v}_i) d\mathbf{v}_i, \quad (4)$$

where the survivor function

$$S_i(a | \mathbf{v}_i) = \prod_{l=51}^{a-1} h_{iE}(l | \mathbf{v}_i),$$

gives the probability of staying employed until age  $a$ . If the worker was still employed at the end of 2004 when he or she was  $a + 1$  years old, the last term  $h_{ij}(a | \mathbf{v}) = h_{iE}(a | \mathbf{v})$ , and hence the likelihood contribution equals  $S_i(a + 1)$ , the probability of surviving in employment until age  $a + 1$ .

The model was estimated under different assumptions about the random effect structure. As a benchmark for more complex specifications, the issue of unobserved heterogeneity was first ignored by imposing  $\mathbf{v}_i = \mathbf{0}$  for all  $i$ . This was followed by a specification where the random effects were assumed be independent normal variants, with zero means and unknown standard deviations. This specification does not allow for correlation between the competing exit channels. The third specification assumed the cross-nesting structure, where  $v_{iU} = e_{iUD} + e_{iUO}$ ,  $v_{iD} = e_{iUD} + e_{iDO}$  and  $v_{iO} = e_{iDO} + e_{iUO}$ , with  $e_i$ 's normal variants with zero means and unknown standard deviations. In this case, a large estimate of the standard deviation of  $e_{iUD}$ , for example, is an indication of the correlation between exits via unemployment and disability retirement. In the most general specification  $\mathbf{v}_i$  was assumed to be multivariate normal with the unrestricted covariance structure. Finally, the perfect correlation between the random effects was imposed but allowed the importance of the random effect to vary across the odds ratios. Namely, this specification imposes  $v_{ij} = \phi_j e_i$ , where  $e_i$  is the standard normal variate, and  $\phi_j$ 's are unknown scale factors to be estimated.

The specification without unobserved heterogeneity reduces to a simple pooled multinomial logit model for hazard rates. In other cases, the estimation procedure is complicated by the fact that the integral in (4) does not have a closed-form solution. To overcome this difficulty, the worker  $i$ 's likelihood contribution is approximated with the simulated probability

$$\tilde{f}(a_i, j_i) = \frac{1}{R} \sum_{r=1}^R S_i(a | \mathbf{v}_i^r) h_{ij}(a | \mathbf{v}_i^r), \quad (5)$$

where  $\mathbf{v}_i^r$  is the  $r$ -th Halton draw from density  $\phi$ . The unknown parameters of the model, including those that characterize the distribution of random effects, are then obtained by maximizing the simulated log-likelihood function based on (5).

#### 6.4 Parameter estimates

The results from the specification with homogeneous policy effects are reported in table 4. To ease interpretation of the age effects, the model includes an intercept for each  $j$  (i.e.  $\mathbf{X}_{ia}$  includes a constant) and  $\alpha_{51j}$  is restricted to 0, so that  $e^{\alpha_{aj}}$  describes the odds ratio of leaving employment for state  $j$  for a worker aged  $a$  compared to the same worker at the reference age 51. As expected, there is strong age dependency towards all exit routes. From age 51 to age 60 the odds of becoming unemployed becomes eightfold ( $e^{2.097} \approx 8.14$ ), whereas the odds of being granted a disability pension becomes tenfold. The odds of leaving employment for some other reason increases even much more, with a large discrete change at age 59. As expected, there are anticipation effects one year before the 1996, 1997 and 2000 reforms became into effect. The odds of becoming unemployed increased by 77%, 98% and by 51% in 1995, 1996 and 1999 among those covered by the protection clauses. There is no evidence of statistically significant anticipation behavior towards other exit channels.

Being eligible to apply for IER pension benefits increases the odds of being granted a disability pension by 128%, but has no other effects on the other odds ratios. Eligibility for the UT scheme has no effect on the odds of becoming a disability pension recipient. The odds of becoming unemployed increase by 134% when the worker exceeds the age threshold for extended unemployment benefits. Although we did not find any effects for being entitled to apply for a part-time pension benefit (except for substituting full-time work by subsidized part-time work) in our DiD analysis, part-time pension eligibility decreases the odds of leaving employment for unemployment by 20%.

It appears that the odds ratios vary strongly with education, income level and sector. For example, the odds of becoming unemployed is 88% lower for the civil servant than for the otherwise similar worker in manufacturing ( $e^{-2.095} \approx 0.123$ ). This raises a doubt whether the eligibility effects are the same for all people. As such we estimated a number of models where the eligibility dummies were interacted with various sets of background variables. Many such interactions do not enter statistically significantly and, hence, should not be added to the model. Table 5 presents the results of the specification with interactions that have at least one statistically significant coefficient. In this specification, the coefficients of the eligibility dummies describe the policy effects on the odds ratios for the reference person for whom all the interactions terms are zero. The coefficients of part-time pension eligibility, for example, describes the effects for a man with 11 years of schooling who holds a private-sector job that pays about € 22,000 per year. It turns out that being eligible to apply for a part-time pension reduces the odds of nonparticipation by 53% ( $e^{-0.752} \approx 0.471$ ) for

Table 4: Hazard parameters from a model with homogeneous policy effects

	Unemployment		Disability		Nonparticipation	
	coeff.	ste	coeff.	ste	coeff.	ste
Intercept	<b>-1.799</b>	0.190	<b>-6.033</b>	0.505	<b>-5.177</b>	0.301
age 52	0.148	0.113	<b>0.711</b>	0.158	0.024	0.143
age 53	0.218	0.115	<b>1.017</b>	0.169	0.182	0.143
age 54	<b>0.568</b>	0.172	<b>1.409</b>	0.280	<b>0.632</b>	0.287
age 55	<b>1.086</b>	0.199	<b>1.606</b>	0.322	<b>0.834</b>	0.321
age 56	<b>1.071</b>	0.214	<b>1.558</b>	0.348	<b>0.981</b>	0.334
age 57	<b>1.446</b>	0.244	<b>1.430</b>	0.387	<b>1.581</b>	0.369
age 58	<b>1.714</b>	0.260	<b>1.846</b>	0.407	<b>1.526</b>	0.383
age 59	<b>2.232</b>	0.315	<b>2.032</b>	0.448	<b>2.804</b>	0.408
age 60	<b>2.097</b>	0.345	<b>2.305</b>	0.492	<b>2.295</b>	0.440
age 61	<b>2.027</b>	0.382	<b>2.282</b>	0.528	<b>2.616</b>	0.462
age 62	<b>2.692</b>	0.423	<b>2.194</b>	0.586	<b>4.041</b>	0.499
1995 x born 1941-42	<b>0.570</b>	0.274	-0.119	0.405	0.394	0.379
1996 x born 1941-43	<b>0.682</b>	0.236	0.372	0.353	-0.327	0.317
1999 x born 1941-44	<b>0.411</b>	0.141	0.001	0.170	0.057	0.181
log (unemployment rate)	<b>0.844</b>	0.084	<b>0.439</b>	0.105	0.012	0.086
Female	-0.045	0.050	<b>-0.168</b>	0.067	<b>-0.140</b>	0.066
Married	<b>-0.182</b>	0.048	-0.085	0.063	0.019	0.055
Schooling years	<b>-0.081</b>	0.013	<b>-0.153</b>	0.019	<b>-0.040</b>	0.015
log (earnings)	<b>-0.824</b>	0.070	<b>-0.487</b>	0.098	<b>0.348</b>	0.108
Tenure < 2 years	<b>0.674</b>	0.058	0.148	0.090	<b>0.404</b>	0.072
Public sector	<b>-2.095</b>	0.076	<b>0.206</b>	0.087	<b>0.379</b>	0.085
Trade	<b>-0.290</b>	0.065	<b>-0.290</b>	0.107	<b>0.281</b>	0.103
Transportation	<b>-0.304</b>	0.084	<b>0.395</b>	0.124	<b>0.453</b>	0.118
Services	<b>-0.227</b>	0.075	-0.159	0.124	<b>0.568</b>	0.099
Other sector	<b>-0.847</b>	0.081	0.049	0.114	<b>0.315</b>	0.102
Early O-A eligibility	<b>-0.440</b>	0.139	0.098	0.119	<b>0.931</b>	0.107
IER eligibility	0.074	0.107	<b>0.823</b>	0.141	0.172	0.114
PTP eligibility	<b>-0.227</b>	0.101	0.060	0.137	-0.167	0.145
UT eligibility	<b>0.850</b>	0.139	-0.067	0.204	-0.081	0.246
Log-likelihood			-37,866.5			

Notes: The model also includes 11 year fixed effects. The reference worker is a non-married 51-year-old man who holds a manufacturing job with the duration of no less than two years. Schooling years are measured in deviation from 11, which corresponds to the secondary level of education. Earnings refer to the average earnings in years t-1 and t-2; and log earnings is measured in deviation from 10, which corresponds to about 22,000 euros a year. The parameter estimates that are significantly different from zero at the 95 percent confidence level are in bold. Standard errors are based on the robust 'sandwich' covariance matrix.

Table 5: Hazard parameters from a model with heterogeneous policy effects

	Unemployment		Disability		Nonparticipation	
	coeff.	ste	coeff.	ste	coeff.	ste
Intercept	<b>-1.858</b>	0.195	<b>-5.881</b>	0.508	<b>-4.901</b>	0.292
age 52	0.185	0.115	<b>0.702</b>	0.156	0.041	0.144
age 53	<b>0.266</b>	0.117	<b>0.998</b>	0.164	0.206	0.145
age 54	<b>0.607</b>	0.174	<b>1.383</b>	0.274	<b>0.658</b>	0.290
age 55	<b>1.124</b>	0.201	<b>1.577</b>	0.314	<b>0.856</b>	0.324
age 56	<b>1.117</b>	0.217	<b>1.525</b>	0.339	<b>1.004</b>	0.338
age 57	<b>1.431</b>	0.247	<b>1.342</b>	0.381	<b>1.692</b>	0.374
age 58	<b>1.706</b>	0.263	<b>1.761</b>	0.400	<b>1.631</b>	0.388
age 59	<b>1.691</b>	0.332	<b>1.871</b>	0.448	<b>3.087</b>	0.415
age 60	<b>1.571</b>	0.359	<b>2.142</b>	0.491	<b>2.605</b>	0.446
age 61	<b>1.488</b>	0.397	<b>2.117</b>	0.524	<b>2.950</b>	0.469
age 62	<b>2.111</b>	0.437	<b>2.033</b>	0.578	<b>4.427</b>	0.504
1995 x born 1941-42	<b>0.593</b>	0.277	-0.104	0.400	0.387	0.381
1996 x born 1941-43	<b>0.685</b>	0.237	0.378	0.350	-0.328	0.318
1999 x born 1941-44	<b>0.412</b>	0.142	0.007	0.169	0.058	0.183
log (unemployment rate)	<b>0.868</b>	0.086	<b>0.442</b>	0.102	-0.039	0.088
Female	-0.012	0.054	-0.131	0.072	<b>-0.608</b>	0.078
Married	<b>-0.195</b>	0.049	-0.088	0.061	0.046	0.057
Schooling years	<b>-0.065</b>	0.016	<b>-0.144</b>	0.022	<b>0.079</b>	0.021
log (earnings)	<b>-1.435</b>	0.145	<b>-0.528</b>	0.177	<b>-0.475</b>	0.198
Tenure < 2 years	<b>0.683</b>	0.058	0.139	0.089	<b>0.397</b>	0.073
Public sector	<b>-1.894</b>	0.082	<b>0.249</b>	0.094	-0.050	0.107
Trade	<b>-0.308</b>	0.066	<b>-0.287</b>	0.105	<b>0.313</b>	0.103
Transportation	<b>-0.309</b>	0.086	<b>0.386</b>	0.121	<b>0.445</b>	0.118
Services	<b>-0.231</b>	0.077	-0.161	0.121	<b>0.602</b>	0.099
Other sector	<b>-0.868</b>	0.084	0.020	0.111	<b>0.391</b>	0.101
Early O-A eligibility	0.251	0.180	0.226	0.161	<b>0.402</b>	0.142
x female	-0.220	0.145	0.013	0.156	<b>0.672</b>	0.142
IER eligibility	0.149	0.127	<b>1.028</b>	0.178	-0.277	0.155
x public sector	-0.232	0.177	<b>-0.347</b>	0.147	<b>0.366</b>	0.131
x female	-0.097	0.131	-0.108	0.160	<b>0.426</b>	0.149
PTP eligibility	-0.147	0.107	0.088	0.152	<b>-0.752</b>	0.176
x public sector	<b>-0.669</b>	0.141	-0.024	0.131	<b>0.702</b>	0.139
x schooling years	-0.037	0.021	-0.014	0.028	<b>-0.185</b>	0.025
x log (earnings)	0.009	0.117	-0.137	0.172	<b>1.290</b>	0.231
UT eligibility	<b>0.792</b>	0.139	-0.092	0.204	-0.098	0.257
x log (earnings)	<b>0.665</b>	0.153	0.171	0.197	0.062	0.238
Log-likelihood	-37,701.7					

Notes: The model also includes 11 year fixed effects. The reference worker is a non-married 51-year-old man who holds a manufacturing job with the duration of no less than two years. Schooling years are measured in deviation from 11, which corresponds to the secondary level of education. Earnings refer to the average earnings in years t-1 and t-2; and log earnings is measured in deviation from 10, which corresponds to about 22,000 euros a year. The parameter estimates that are significantly different from zero at the 95 percent confidence level are in bold. Standard errors are based on the robust 'sandwich' covariance matrix.

such a person. But note there is no such effect for otherwise similar women. Furthermore, there is evidence that part-time pension eligibility tend to reduce the odds of becoming unemployed for those working in the public sector. This effect has no much impact in practice, as the overall risk of unemployment is so low for public-sector employees. This degree of heterogeneity implies that the part-time pension scheme has very different implications for different people.

The effect of IER eligibility on the odds of being granted a disability pension for a private sector employee is significantly lower than the effect for an otherwise similar public servant. Allowing for heterogeneity in the policy effects does not alter our conclusion that the odds of becoming unemployed does not depend on the IER scheme. The overall effect of abolishing the IER scheme on employment careers is non-trivial. Its magnitude and sign probably vary between workers.

The effect of being eligible for extended unemployment benefits is rather similar for different people, varying only with earnings. A stronger effect on the odds of entering unemployment for workers with high earnings may indicate that some of those who will be entitled to higher earnings-related unemployment insurance benefits select themselves into unemployment. The UT scheme does not affect the odds of being granted a disability pension or transitions to nonparticipation, which is consistent with the estimates in table 4. Hence, it seems that the two-year increase in the age threshold of the UT scheme in 1997 has led to an increase in the expected employment careers.

## 6.5 Effects of policy reforms

The eligibility effects on the odds of hazard rates reported above are informative about the importance of the availability of different early retirement schemes, but their magnitudes are difficult to interpret in terms of quantities of interest from the policy point of view.<sup>10</sup> So it is not easy to say whether a given policy reform has been economically important or not based on the estimated changes in the odds of hazard rates. The policy effects are studied by comparing cumulative exit probabilities and the expected duration of the remaining employment career associated with different counterfactual policy designs for workers born in 1947 or 1948 who were included in the estimation sample (6324 individuals). These cohorts were chosen because they have been affected by all the reforms investigated in this study. The question addressed is how this group would have behaved, had they been covered by a given combination of the early retirement rules. The starting point is the eligibility rules that were in force in the mid-1990s. The reforms are then introduced one by one, leading to the set of the eligibility rules that these individuals actually faced. More precisely, the following counterfactual policy regimes are considered:

0. The old rules that were in force in 1996: the age thresholds of 53, 58 and 56 for extended unemployment benefits, IER pension benefits and part-time pension benefits, respectively.
1. The rules 0 but the eligibility age for the extended unemployment benefits raised by two years from 53 to 55 (the 1997 reform).
2. The rules 1 but the IER scheme abolished (the 2000 reform).

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<sup>10</sup>For example, the effect of a given covariate on the odds ratio between the hazard of being granted a disability pension and that of remaining employed does not provide direct information about the covariate's effect on the hazard function of entering a disability pension benefits, nor on its effect on the probability that the employment spell will be eventually followed by a disability pension. This is because the sign and magnitude of the latter effects depend on all the parameters of the model. They also vary across individuals with different background characteristics.

3. The rules 2 but the eligibility age for the part-time pension raised by two years from 56 to 58 (the 2003 reform). These are the rules the workers born in 1947 and 1948 actually faced.

The analysis is based on the logit specification with the heterogeneous policy effects. For a worker  $i$  who has been employed from age 49 to 51, the probability of being still employed at age  $a \in [52, 63]$  under policy  $p \in \{0, 1, 2, 3\}$  is given by the survival function

$$S_i^p(a) = \int S_i^p(a|\mathbf{v}_i)\phi(\mathbf{v}_i) d\mathbf{v}_i. \quad (6)$$

In the computations worker  $i$ 's covariates  $\mathbf{X}_{ia}$  are fixed at their initial values at age 51 (except a dummy for short job tenure), whereas the age patterns of the eligibility dummies,  $Early_{ia}$ ,  $IER_{ia}$ ,  $PTP_{ia}$  and  $UT_{ia}$ , and their interactions with the covariates are chosen to be consistent with a given policy regime. There are no anticipation effects, so that the results describe the permanent policy effects. The integral is approximated using the same simulation technique as was applied in the simulated maximum likelihood estimation.

The expected age of exit from employment under policy  $p$  for worker  $i$  is computed as

$$E^p(A_i) = \int \left( \sum_{l=51}^{65} (l + 1/2) [1 - h_{iE}^p(l|\mathbf{v}_i)] S_i^p(l|\mathbf{v}_i) \right) \phi(\mathbf{v}_i) d\mathbf{v}_i. \quad (7)$$

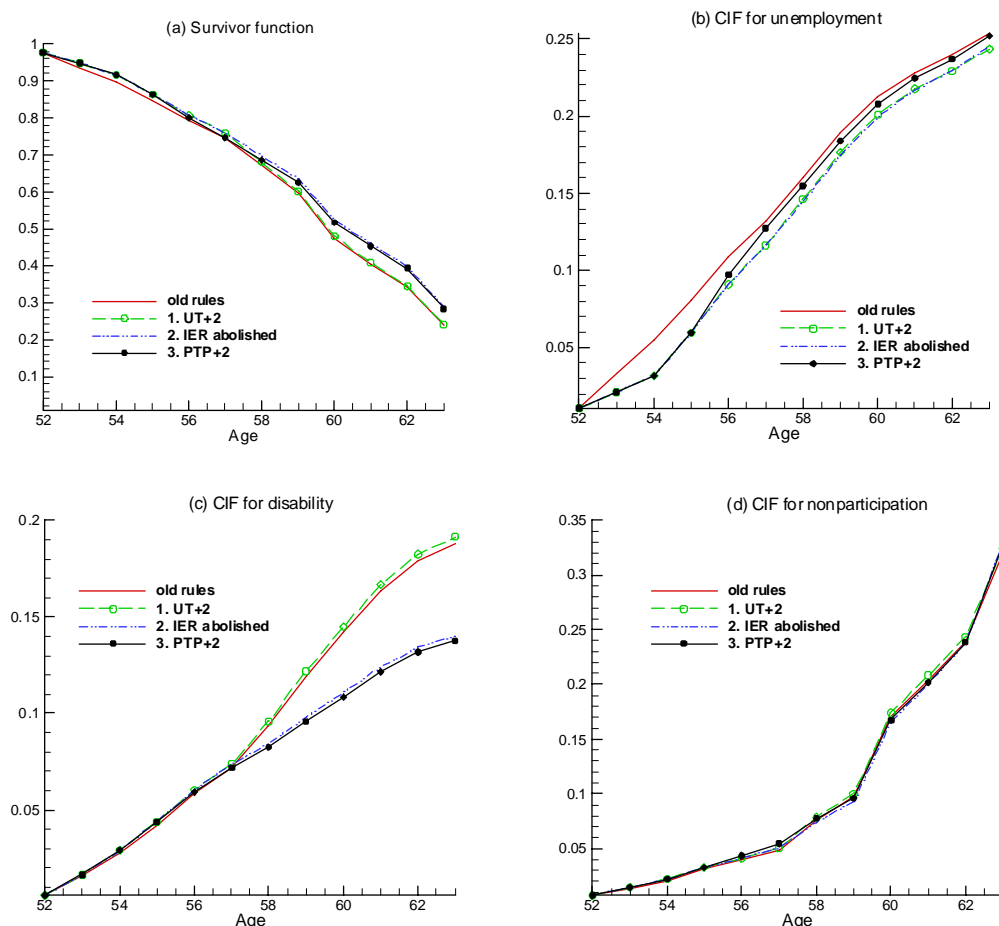
The logit hazard gives the probability that the worker exits from employment by  $a + 1$  conditional on being employed at age  $a$ . It follows that the worker can be either  $a$  or  $a + 1$  years old at the day when she left her job, and thereby the retirement age in the equation above is set to  $a + 1/2$  when exit occurs between ages  $a$  and  $a + 1$ . Moreover, since the hazard estimates are available only up to age 62, it is assumed that  $h_{iE}^p(63|\mathbf{v}_i) = h_{iE}^p(62|\mathbf{v}_i)$  and  $h_{iE}^p(64|\mathbf{v}_i) = 0$ . The first assumption may overestimate the likelihood that a 63-year-old employee stays employed for one additional year, whereas the latter ignores the possibility of postponing the old-age pension beyond age 65. Given that the same assumptions are imposed in all cases, they should not affect notably differences in the expected exit ages associated with different policy regimes

The effect of policy reform  $p \geq 1$  compared to the counterfactual under policy  $p - 1$  for worker  $i$  can be measured as  $S_i^p(a) - S_i^{p-1}(a)$  and  $E^p(A_i) - E^{p-1}(A_i)$ . The sum of these effects over reforms 1, 2 and 3 gives the joint effect of all the three reforms compared to the counterfactual case of the mid-1990s rules. It should be stressed that the survival function and the expected exit age are the two sides of the same coin in the sense they both describe the expected remaining time in employment. In particular, they both are solely determined by the hazard of staying employed or, equivalently, the *sum* of the hazards for exits to unemployment, disability and nonparticipation, and thereby do not contain information on the relative importance of different exit pathways. Therefore, cumulative incidence functions are used to study changes in the use of alternative exit pathways. The cumulative incidence function (CIF) for state  $j$  under policy  $p$  is

$$F_{ij}^p(a) = \int \left( \sum_{l=51}^{a-1} h_{ij}^p(l|\mathbf{v}_i) S_i^p(l|\mathbf{v}_i) \right) \phi(\mathbf{v}_i) d\mathbf{v}_i, \quad (8)$$

and it gives the probability of having exited from employment to state  $j$  by age  $a \in [52, 63]$ . Because  $S_i^p(a) = 1 - \sum_k F_{ik}^p(a)$ , the effects of a policy reform  $p$  in terms of changes in CIFs,  $F_{ij}^p(a) - F_{ij}^{p-1}(a)$ , provide a useful decomposition for the overall effect on the survival probability.

Figure 6: Marginal survivor and cumulative incidence functions for 1947 and 1948 cohorts under different pension schemes



Using the estimated parameters from the hazard model, the cumulative incidence and survival rates as well as the expected exit ages were computed for all individuals born in 1947 or 1948 under the different policy regimes. The average values of individual-specific survival and cumulative incidence functions (i.e. the marginal curves) are depicted in figure 6. The marginal CIFs give the predicted shares of people escaping from employment via different pathways by a given age, whereas the marginal survival rate gives the predicted share that is expected to be still employed by a given age. Discrepancies in the curves associated with the different policy regimes describe the policy effects on the cumulative flows out of employment. Tables 6 and 7 provide additional information. The first column of table 6 shows the policy effects on the expected exit age. The other columns show the policy effects on the survival rate and CIFs at age 63. Their average effects correspond to the differences in the last values of the marginal curves in figure 6. To illustrate heterogeneity in the policy effects across individuals, also the 10th and 90th percentiles along with the median effects are reported in table 6. Furthermore, table 7 reports the mean characteristics of individuals at the lower and upper end of the distributions of the policy effects on the expected

exit age. The first decile include workers whose exit age was subject to the smallest change due to a given reform and the tenth decile those who were subject to the largest change.

The hazard estimates indicate that the two-year increase in the age threshold for the UT scheme in 1997 reduced notably the odds of becoming unemployed at ages 52 and 53. Consequently, the cumulative probability of entering unemployment under policy regime 1 lies clearly at a lower level than in the pre-reform case in figure 6b. The difference also declines with age, and hence the effect on the cumulative incidence of unemployment at age 63 is only  $-0.01$  (see panel A in table 6). Moreover, the overall effect on the likelihood of surviving employed until that age is practically zero (0.002) due to the slightly increased cumulative risks of disability retirement and nonparticipation. Since the hazard estimates did not indicate significant spillover effects on other exit rates for UT eligibility (see table 5), these increases are largely driven by the dynamic selection process: when the worker's likelihood of becoming unemployed decreases at a given age, he is more likely to be employed at older ages and, hence, at the higher risk of entering other states than unemployment. As seen in panel A of table 6, the mean duration of the employment career is estimated to be increased by 1.3 months ( $= 12 \times 0.109$ ) due to the two-year increase in the age threshold of the UT scheme. This effect also varies across individuals, as the effect increases from .029 to .201 from the 10th to 90th percentile. The reform has hardly any effect on educated women in the public sector who dominate the first decile of the effect distribution in table 7, whereas the employment careers of male manufacturing workers are subject to the largest increases. In particular, 97% of individuals in the tenth decile of the effect distribution holds a job in the manufacturing sector. This stems at least part from the highest overall risk of unemployment for manufacturing workers (see table 5).

Turning to the next reform, abolishing the IER scheme reduces the odds of begin granted a disability pension after age 57, resulting in a substantial reduction in the cumulative incidence of disability retirement at older ages (figure 6c). The likelihood of having left employment for disability pension benefits by age 63 reduces by 0.051 compared to the counterfactual level with the IER scheme available at age 58. In relative terms this reduction is as much as one-fourth. The overall effect on the likelihood of staying employed is almost equally large, being 0.046 at age 63. This results an increase of 3.4 months ( $= 12 \times 0.284$ ) in the expected exit age on average. This average effect masks a large degree of heterogeneity between individuals. Abolishing the IER scheme lengthens the expected employment career 6.2 months or more for one-tenth of people, whereas the effect is 1.7 months or less for another one-tenth (panel B in table 6). This heterogeneity stems from different responses towards transitions to unemployment and nonparticipation, as the 10th and 90th percentiles of the changes in the cumulative incidences of unemployment and nonparticipation at age 63 are of the opposite signs. As seen in table 7, this reform had the strongest impact on educated women in the public sector, having very little effect on educated men in the private sector.

The last reform – the two-year increase in part-time pension eligibility – increases the likelihood of entering unemployment, but decreases the likelihood of being granted a disability pension and that of exiting to nonparticipation. All these effects are very small, being close to zero on average. As a result, the average survival rate is hardly affected, and the average drop in the expected exit age is less than one month (panel C in table 6). The effects of this reform do not exhibit much variation across individuals either. For the one-tenth of the population most strongly affected, the expected age of exit was reduced by 1.5 months or more. A common characteristic for this group is high education. Overall, the two-year increase in the part-time pension eligibility had relatively

Table 6: Distributions of policy effects

	$E(A)$	$S(63)$	$F_u(63)$	$F_d(63)$	$F_o(63)$
	(1)	(2)	(3)	(4)	(5)
<b>Reference level (regime 0)</b>	<b>59.595</b>	<b>0.239</b>	<b>0.253</b>	<b>0.188</b>	<b>0.320</b>
<b>A. 2-year increase in UT age limit (from regime 0 to 1)</b>					
Mean	<b>0.109</b>	<b>0.002</b>	<b>-0.010</b>	<b>0.003</b>	<b>0.004</b>
10% percentile	0.029	0.001	-0.014	0.001	0.003
Median	0.111	0.002	-0.010	0.003	0.005
90% percentile	0.201	0.004	-0.005	0.006	0.006
<b>B. Abolition of IER scheme (from regime 1 to 2)</b>					
Mean	<b>0.284</b>	<b>0.046</b>	<b>0.002</b>	<b>-0.051</b>	<b>0.004</b>
10th percentile	0.138	0.015	-0.014	-0.071	-0.072
Median	0.239	0.038	0.005	-0.053	0.008
90th percentile	0.517	0.094	0.013	-0.030	0.066
<b>C. 2-year increase in PTP age limit (from regime 2 to 3)</b>					
Mean	<b>-0.069</b>	<b>-0.004</b>	<b>0.007</b>	<b>-0.003</b>	<b>-0.001</b>
10th percentile	-0.130	-0.007	0.002	-0.005	-0.007
Median	-0.070	-0.003	0.005	-0.003	-0.001
90th percentile	-0.013	-0.001	0.015	0.000	0.006
<b>D. Joint effect of 3 reforms (from regime 0 to 3)</b>					
Mean	<b>0.324</b>	<b>0.045</b>	<b>-0.002</b>	<b>-0.051</b>	<b>0.007</b>
10th percentile	0.152	0.015	-0.022	-0.069	-0.074
Median	0.342	0.038	-0.002	-0.052	0.014
90th percentile	0.478	0.089	0.020	-0.030	0.072

Notes: Reference level is the level under old rules in force in 1996.

moderate effect on all individuals. This finding is in accordance with the result of the DiD analysis that the part-time pension scheme mainly induces part-time working at the expense of full-time working, without notable effects on the transitions to other labour market states.

The joint effect of all the three reforms in panel D in table 6 suggests that the expected age at which workers leave employment increased by some 3.9 months ( $= 12 \times 0.324$ ). Since the 10th percentile of the joint effect is 1.8 months, the series of the policy reforms have induced basically all groups of workers to work longer. For one-tenth of the most strongly affected workers the increase is about a half year or more. This group consists solely of female public servants. It is evident that abolition of the IER scheme is responsible for much of the overall increase in the length of the employment careers. The likelihood of being granted a disability pension by age 63 has decreased by 0.051 compared to the counterfactual case of the mid-1990s rules, which is a substantial drop. Also, the two-year increase in the age threshold for the UT scheme has played an important role in postponing exits from employment. While the cumulative incidence of unemployment at age 63 is almost at the same level it would have been under the mid-1990s rules, those who enter unemployment do so at the later point in their employment career. Also, it should be noted that the present focus on exits from employment ignores the increase in the transition rate out of unemployment for 53-54-year-old job seekers caused by the 1997 reform (see Kyyrä and Wilke,

Table 7: Mean characteristics of workers in the 1st and 10th deciles of the distributions of policy effects on the expected age of exit from employment

	2-yr increase in UT age limit		Abolition of IER scheme		2-yr increase in PTP scheme		Joint effect of 3 reforms	
	1st	10th	1st	10th	1st	10th	1st	10th
log (unemployment rate)	-2.099	-1.525	-1.965	-1.977	-1.768	-1.879	-1.950	-1.924
Female	0.647	0.289	0.002	1.000	0.515	0.137	0.037	1.000
Married	0.895	0.571	0.857	0.791	0.707	0.788	0.862	0.756
Schooling years	1.884	-0.823	2.264	1.736	3.388	-0.855	3.841	0.646
log (earnings)	0.118	0.338	0.777	0.264	0.168	0.633	0.593	0.234
Tenure < 2years	0.060	0.023	0.077	0.031	0.056	0.063	0.069	0.026
Public sector	0.936	0.000	0.024	1.000	0.319	0.244	0.258	0.990
Manufacturing	0.016	0.971	0.219	0.000	0.165	0.329	0.130	0.010
Trade	0.010	0.008	0.235	0.000	0.110	0.119	0.150	0.000
Transportation	0.006	0.003	0.042	0.000	0.053	0.093	0.042	0.000
Services	0.008	0.018	0.259	0.000	0.205	0.088	0.217	0.000
Other sector	0.024	0.000	0.221	0.000	0.148	0.127	0.203	0.000
Mean effect on $E(A)$	0.020	0.218	0.053	0.534	-0.162	0.017	0.059	0.510

Notes: XXX

2007, or Kyyrä and Ollikainen, 2008), and hence the present study somewhat understates the overall effect of the increase in the UT age threshold on unemployment.

When interpreting the effect of the increase in the age threshold for the part-time pension, one should bear in mind that part-time pensioners are counted employed. Although the reform slightly lowered the average age of exit from employment, it has also reduced transitions from full-time to part-time work among workers aged 56 and 57, which does not show up here. The results of the DiD analysis in Section 5.2 suggest that the latter effect is very strong, and thereby the part-time pension reform has almost surely increased the total number of hours worked.

## 7 Concluding remarks

Altogether the three reforms have raised the average age at which older people leave employment by 3.8 months. It is noteworthy that the expected exit age has increased among all workers due to these reforms, albeit the size of the effect varies across different workers. The most effective reform was the abolition of the IER scheme, which has reduced remarkable the disability enrolment rates at age 58 and beyond. The average increase of 3.4 months in the exit age due to this reform masks a large degree of impact heterogeneity. Among the educated women in the public sector, the exit age was postponed by a half year. This raises a question why this group responded so strongly. Overall, the strong effect for the increased stringency of medical screening for disability pension benefits is in line with the results from other countries, including Autor and Duggan (2003, 2006), Karström *et al.* (2006) and Börsch-Supan (2007).

The two-year increase in the age threshold for extended unemployment benefits had the strong effect on the incidence of unemployment among those in their early 50s but the cumulative effect deciles with age. This reform has postponed entries to unemployment, and thereby led to an

increase of 1.3 months in the average age of exit from employment. Forthcoming reforms will increase the age threshold by one year for those born between 1950 and 1954 and by two years for the later cohorts. These changes may have a stronger effect on the average exit age than the 1997 reform had, because the unemployment hazard increases with age and because the affected cohorts are less likely to withdraw via the disability pension scheme due to the stricter medical screening. It should be kept in mind that this kind of reforms are effective mainly in the manufacturing sector, which relative importance will decrease in the long run.

The DiD analysis showed the part-time pension scheme induces full-time workers to switch part-time, without affecting notably the incidence of unemployment or disability retirement. The hazard analysis also suggested that eligibility for the part-time pension scheme has not strong effects transitions out of employment. It follows that the two-year increase in the age threshold has probably increased effective labour supply as measured by hours worked, despite a small negative effect on the average exit age. Having said this, the forthcoming one-year increase in the age threshold for those born in 1953 and later seems to an appropriate policy change.

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