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Pensions, Retirement, and the Financial Position of the Elderly

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In my beginning is my end. In my end is my beginning.
T.S. Eliot, "East Coker," Four Quartets

Preface

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Jim Been, Leiden, January 2015

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1 Introduction

This thesis collects five studies that are related to *pensions, retirement and the financial position of the elderly*. The five studies each aim to contribute to the understanding of pensions, retirement and the financial position of the elderly. The chapters in the thesis can be read independently. This introductory chapter provides the motivation for this thesis' topics (section 1.1) followed by the research questions underlying each of the chapters (section 1.2) and a summary of the main findings of each chapter (section 1.3).

1.1 Motivation

Long-term trends, such as the aging of the population and the increased life-expectancy, and the consequences of the recent financial crisis, have raised concerns about the sustainability of pension systems. The aging of the population increases the old-age dependency ratio thereby increasing the pressure on the working population. This effect on the dependency ratio is enforced by the low and decreasing fertility rates. Whereas this dependency ratio was 29 in 2010, it is expected to increase to 55 in 2040, meaning that there are 55 persons over the age of 65 for every 100 persons of age 15-64. The increased life-expectancy implies that people, on average, receive pension benefits for a longer period. The recent financial crisis has had a large impact on the capital reserves of pension funds which resulted in pension benefits cuts accordingly. Current pension systems are not well-prepared to face the long-term trends and the short-term volatility.

Consequently, many OECD countries have proposed and implemented reforms to alleviate the pension system from the pressure of demographic aging and to create sustainable pension systems for the future (OECD 2011b). Many of the reforms implemented to alleviate the pension systems from the consequences of aging populations are related to increasing both the statutory and effective retirement age, making pension benefits less generous and increasing contributions. Increasing the retirement age aims to increase the labor supply at older ages such that older people postpone receiving pension benefits and contribute to the pension system for a longer period. One of the most prominent measures taken to induce labor supply at older ages is making the options to retire earlier than the statutory retirement age less available and less generous. Decreasing the benefits and increasing the contributions of pensions means that the accumulation of pensions per contributory year decreases.

These reforms imply that public- and private pensions have become less generous. As a consequence, the proposed and implemented reforms have raised a lot of discussion about the financial position of current and future retirees. What are the effects of the reforms on current retirees? Do we expect future retirees to have an adequate pension for consumption needs? Do people need to retire later while working more hours during their career? What are the consequences of making early retirement less available and less generous? What jobs are needed to retire later? This thesis aims to understand the effects of aging on people's retirement behavior and the adequacy of their (future) pensions.

1.2 Research questions

Reforms are complicated by institutional path dependency and long-term contracts. Therefore, reforms implemented to combat the negative consequences of aging to society, such as increases in the statutory retirement age, are often politically controversial and difficult to implement. Hence, the approach that many countries have initially taken is to privatize pensions which entails a shift from the relative importance of public- to private pensions and from defined benefit (DB) to defined contribution (DC) sys-

tems (Barr and Diamond 2009, OECD 2009b, Orenstein 2011). This shift reduces the pressure of demographic aging on the public finances, as mandatory public pensions are usually financed by a PAYG system, but it may increase inequality of income of the elderly.

Chapter 2 addresses the research question *Do shifts from public- to private pension provision lead to higher levels of income inequality or poverty among older people?* The chapter explains that private pension plans are generally less redistributive than public pension plans. Therefore, it is expected that the observed shifts from public- to private pension plans has led to higher levels of income inequality among the elderly. Chapter 2 empirically analyzes the distributional effects of shifts from public- to private pension provision by relating the relative importance of the public- and private pensions to indicators of income inequality and poverty among elderly in 15 European countries for the period 1995-2007. This approach allows us to draw conclusions about the consequences of reforming the pension system for the financial well-being of the elderly.

To assess the financial well-being of the elderly, *Chapter 3* focuses on the questions *Do Dutch households save adequately for retirement? Which pension components are important, and what are the vulnerable groups?* The chapter argues that it is important to assess the adequacy of pensions based on analyzing microdata which allows us to investigate the heterogeneity and taking into account the pension that people accumulate in the current system. Prior assessments of the adequacy of pension systems calculated replacement rates for a fictitious person who earns a median income during his career and who is assumed to accumulate a pension during a full career (Mercer 2013, OECD 2013c). According to these studies, the Netherlands are ranked high among pension systems that give an adequate income during retirement. However, the question arises whether adequacy is still that high when the adequacy of retirement income is based on microdata and taking into account the pension that people actually accumulate in the current system.

Chapter 3 develops a much needed integral method to assess the adequacy of retirement savings of Dutch households based on actual

accumulation in the current system by taking into account private wealth[1] next to wealth accumulated in public (1st pension pillar) and private pensions (2nd and 3rd pension pillar). Taking into account private wealth is important since different forms of assets may act as substitutes for one another. Solely focussing on public and private pensions may therefore underestimate the income available at retirement as (housing) wealth may be annuitized for consumption purposes during retirement. Analysis of the adequacy of retirement savings is decomposed for different generations and potentially vulnerable groups such as the self-employed.

The analysis identifies the mandatory (occupational) private pensions (2nd pension pillar) as one of the main components of retirement income. These mandatory private pensions are accumulated over the life-cycle according to an earnings related pension formula in the Netherlands. Life-cycles wage profiles are therefore crucial in determining income available at retirement. This line of reasoning also holds for other countries where earnings related pension formulas also (partly) determine the accumulation of wealth in public pensions (1st pension pillar).[2] Wage profiles are a central component in life-cycle models of consumption and savings since income uncertainty is derived from the deterministic component in wages over the life-cycle (Gourinchas and Parker 2002, Scholz et al. 2006) and so for the analysis of the adequacy of income during retirement. Based on a life-cycle model, Scholz et al. (2006) conclude that the extent and magnitude of undersaving is generally small in the US, although younger cohorts tend to be less likely to have saved sufficiently for retirement.

The conclusions of life-cycle models depend on the correct specification of the life-cycle wage-profile. However, life-cycle models do not consider selection effects into work. Wages are likely to be observed non-randomly over the life-cycle, e.g. wages are observed for people who are working. These individuals may earn a different potential wage than the individuals who are not working. Neglecting this non-random selection into work may bias estimated wage-profiles (Casanova 2010, 2013). For the analysis of the adequacy of retirement savings in the Netherlands, this is important

[1] Housing wealth and other private savings such as net savings account, securities, stocks substantial shareholders and business assets.

[2] For example, Norway, France, the UK and the US.

because accumulation in mandatory private pensions is based on these life-cycle earnings. Since only people in paid-employment earn a wage and accumulate mandatory private pensions,[3] a biased estimation of the wage-profile leads to incorrect assumptions regarding the future accumulation in mandatory private pensions.

Chapter 4 proposes a new panel data sample selection estimator correcting for simultaneous decisions in participation and working hours decisions. Whereas prior panel data sample selection estimators only include a selection correction for work versus non-work (Dustmann and Rochina-Barrachina 2007, Kyriazidou 1997, Rochina-Barrachina 1999, Semykina and Wooldridge 2010, 2011, Wooldridge 1995), we add additional information about part-time and full-time work. Adding information regarding part-time versus full-time work decisions provides extra information regarding the individuals' unobserved characteristics that may influence their wages (e.g. tastes, preferences, ability, effort). This is especially relevant when analyzing life-cycle wages as part-time employment is often chosen among women raising children (Booth and Van Ours 2008, Gregory and Connolly 2008). Men, on the other hand, often use part-time employment as partial retirement mechanism (Cahill et al. 2006, Ruhm 2006) with substantial decreases in wages as a consequence (Aaronson and French 2004, Casanova 2013). The proposed new estimator is used to estimate wage-profiles over the complete life-cycle for both men and women and to show the existence of selection effects in life-cycle wage profiles.

Existence of selection over the life-cycle would suggest that earlier estimates of wage profiles, without selection correction[4] or with binary selection correction can be improved. The new estimator can be applied to estimate selection corrected life-cycle wage profiles to assess the savings adequacy for retirement in both life-cycle models (Scholz et al. 2006) and microsimulation models (Borella 2004). The new panel data sample

[3] Although it depends on the pension fund to what extent unemployed and disabled persons can also accumulate private pensions.

[4] A huge strand of literature uses prime-aged men to estimate wage profiles and assume that non-random selection in work is negligible among this group. However, the assumption is arguable and the consequence of this approach is that the results of the wage-profile can not be generalized to women and old-age men.

selection model forms the basis for the Dutch contribution to the *OECD Retirement Savings Adequacy* project (OECD 2014).

Whereas Chapter 4 analyzes selection effects into part-time employment and full-time employment and the consequences for life-cycle wages, *Chapter 5* focuses on the extent to which non-standard employment, such as part-time employment, may postpone early retirement. The research question that is dealt with in Chapter 5 is *Did the rise in non-standard employment, such as self-employment and part-time employment, contribute to the increased labor force participation observed among older workers across Europe?* Early labor market withdrawal has contributed substantially to the decreasing working lives observed in Europe (Brugiavini and Peracchi 2005). Pestieau (2003) argues that the financial sustainability of pension systems is substantially affected by these low participation rates of older workers. Barr (2006) argues that *'The problem [of the financial sustainability of pension systems] is not that people are living longer, but that they retire too early.'* Longer working lives would alleviate some of the pressure of population aging on the sustainability of pension systems (Maestas and Zissimopoulos 2010). Inducing longer working lives can be attained by either increasing the statutory retirement or decreasing early labor market withdrawal (or both). Longer working lives have two effects on the sustainability of pensions: 1) increases the number of years to accumulate pensions[5] and 2) reduces the number of years to pay benefits. In relation to the adequacy of retirement savings as analyzed in Chapter 3, the results of Angelini et al. (2009) suggest that increasing participation rates among older workers may also increase the adequacy of pensions to finance consumption in retirement.

Gruber and Wise (1998) argue that the low participation rates of older workers have been a consequence of large disincentives to work at older ages. Formal early retirement programs as well as social insurance programs, that were often used to smooth transitions from work to retirement in practice, are prominent explanations of the relatively low participation rates among older persons. Since the 1990s, many governments have

[5]Increasing the statutory retirement age by a year does, however, not imply an increase in the effective retirement age of one year because of early retirement possibilities.

started to reform formal and informal early retirement possibilities (Casey et al. 2003). Hence, the participation rates of older workers have been rising for both men and women in many European countries, although the participation rates are still low compared to those of prime age workers.

Simultaneously, rapid increases in non-standard forms of employment, such as part-time employment and self-employment, have been observed since then and especially among people aged 50-64 (Chen et al. 2013). Both forms of non-standard employment provide flexibility in working hours and may therefore be used as gradual retirement mechanisms to bridge the period between full-time employment and full retirement (Bruce et al. 2000, Cahill et al. 2006, Ruhm 2006). Morris and Mallier (2003) show that the high and increasing importance of such non-standard employment opportunities among this age-group in European countries can be related to the countries' patterns in labor force participation at older ages. Chapter 5 captures this idea into a formal retirement model and estimates a reduced form empirical model at the country-level so to analyze the effect of part-time employment rates and self-employment rates on early retirement measured at both the extensive (e.g. participation decision) and intensive margin (e.g. hours decision).

Insight in this relationship between non-standard employment and postponement of retirement is relevant for policy-makers because of two main reasons. Firstly, policy makers can increase the possibilities of non-standard work for older workers to increase the sustainability of the pension system. Secondly, non-standard work may be used to supplement pension income at older ages. Regarding this second reason, the financial well-being of elderly may not only be evaluated by pension wealth (as in Chapter 2 and 3) and private wealth (as in Chapter 3), but also by the ability to postpone retirement and to supplement the income streams from pension wealth and private wealth. Next to private (housing) wealth and informal care by children and grandchildren, the ability to postpone full retirement may contribute to the financial well-being of elderly besides the income streams from public and private pensions.

Thus far, it is argued that the increase in non-standard employment is related to voluntary decisions to decrease working hours prior to retirement. However, the increase in non-standard employment is also likely to

be in anticipation to declining opportunities in both (full-time) employment (Dorn and Sousa-Poza 2010) and early retirement possibilities such as disability- and unemployment insurance (Casey et al. 2003). Reforms in early retirement schemes may have shifted persons towards non-standard employment as early retirement schemes may no longer be available or as generous as prior to the reforms. Chapter 5's analysis takes into account such spill-over effects between early retirement and non-standard employment, but does not try to explain the nature of the increases in non-standard employment.

Chapter 6, on the other hand, does try to shed some light on the reasons of choosing non-standard employment, more specifically self-employment, at older ages by answering the question *Is there evidence for necessity-driven self-employment at the end of working life?* Self-employment is found to be relatively important among the 50+ population, compared to younger age groups (Hurd 1996, Karoly and Zissimopoulos 2004, Zissimopoulos and Karoly 2007). One of the main explanations for the relatively high importance of self-employment among older persons is that older persons, who tend to have stronger preferences for leisure (Kantarci and Van Soest 2008), primarily choose a period of self-employment before full retirement as a gradual retirement route in which they are able to decumulate the number of hours worked.[6] However, the 50+ population also faces difficulties finding a new job once unemployed (Chan and Stevens 2001, Maestas and Li 2006). Chapter 6 sheds a new light on the often assumed gradual retirement function of self-employment by constructing three testable hypotheses related to the labor market dynamics of older workers in order to analyze the importance of necessity-driven self-employment at older ages. The three testable hypotheses all relate to ending or avoiding unemployment at older ages which is often associated with the necessity of self-employment.[7]

[6]This is suggested by Bruce et al. (2000), Fuchs (1982), Giandrea et al. (2008), Gu (2009), Hurd (1996), Morris and Mallier (2003), Zissimopoulos and Karoly (2007).

[7]Such as suggested by Earle and Sakova (2000), Glocker and Steiner (2007), Kellard et al. (2002), Kuhn and Schuetze (2001), Reize (2000), Rissman (2003), Taylor (1999).

Main findings 1.3

This section provides the answers to the questions raised in section 1.2.

Chapter 2 hypothesizes that the relative shifts from public to private pensions in order to relieve the pressure on the public finances from the aging population has led to higher levels of income inequality among the elderly. *A priori* shifts from public to private pensions are expected to have such an effect as private social security plans are generally less redistributive than public social security. Using macroeconomic information on 15 European countries over the period 1995-2007 does not confirm this hypothesis. No evidence is found that shifts from public to private pension provision are associated with higher levels of income inequality and poverty among elderly people. Intriguingly, this is not in line with findings in the literature on pension reform and income inequality (Arza 2008, Fukawa 2006, Hughes and Stewart 2004, Milligan 2008, Oshio and Shimizutani 2005, Weller 2004) and with literature on the redistributive effects of public and private social security in general (Caminada and Goudswaard 2005, Goudswaard and Caminada 2010).

Nevertheless, this chapter is the first to analyze the relationship between public- and private pensions and income inequality using pooled cross-section and time-series data which should provide more information on the true relationship between the public/private pension mix and income inequality than analyses solely based on cross-sectional (Brown and Prus 2004, Fukawa 2006, Weller 2004) or time-series data (Milligan 2008, Myles 2000, Oshio and Shimizutani 2005, Schirle 2009).[8] Chapter 2 discusses a number of tentative explanations that are conceivable for not finding a positive effect of shifts from public to private pension provisions on income inequality among elderly using our analysis.

Chapter 3 is concerned with the adequacy of retirement savings of Dutch households. To investigate the adequacy of retirement savings, this chapter analyzes not only public and occupational private pension rights, but also

[8] In a cross-sectional approach, the effects of pension reform cannot be analyzed over time. Studies taking a time-series approach in one country have difficulties examining whether the findings also hold for other comparable pension reforms in other countries.

annuity insurances, housing wealth and private savings. The adequacy of the total of available pension income at retirement is determined by replacement rates and absolute levels of pension annuities. Taking into account the total of pension annuities and summing over all age- and socioeconomic groups gives a median gross replacement rate of 83% and a median net replacement rate of 101%. Average and median total resources available at retirement are 33,000 and 27,000 euros per year respectively. By only taking into account public- and private pensions, the median gross replacement rates would be 71%.

Regarding the relative importance of each pension component, public and occupational pensions each account for more than 35% of total pension annuities. Private non-housing assets account for 14% and imputed rental income from net housing wealth accounts for about 10%. Assuming that households deplete their housing wealth during retirement would increase the median gross replacement rate by another 5%-points. The relative importance of the pension components differs between generations: younger generations tend to have higher accumulated occupational pensions while the older generations tend to have a relatively larger share of their wealth in private savings and housing.

In the literature, it is often assumed that a 70% gross replacement rate of previous earnings is the norm for an adequate pension (Haveman et al. 2007). Using a replacement rate of 70% of current gross income, 31% of all households face an inadequate income during retirement when taking into account the total of pension annuities. Based on relatively low replacement rates and low absolute levels of pension income available at retirement, the chapter identified several potentially vulnerable groups in terms of retirement savings adequacy such as the self-employed, first-generation immigrants, single women and recipients of social assistance-, unemployment- and disability benefits.

Aforementioned results regarding the adequacy of retirement savings of Dutch households are sensitive to the assumptions made for the future. Deviating from an intermediate scenario, young generations would benefit most from an optimistic future scenario but also suffer more from a pessimistic scenario compared to older generations.

A different approach to assess the adequacy of retirement income is taken by Scholz et al. (2006) who base their analysis on the estimation of a life-cycle model of consumption and savings. Like in all life-cycle models, conclusions regarding consumption and saving depend on the correct specification of the wage-profile as the earnings process is the most important determinant of income over the life-cycle. However, most life-cycle models do not correct for non-randomness in observed wages. As a consequence, corrections for non-random selection into work might are necessary to get unbiased estimates of the life-cycle wage profile (Casanova 2013). Unbiased estimates of life-cycle wages are not only important in life-cycle models (Gourinchas and Parker 2002, Scholz et al. 2006) but also in other models that depend on life-cycle earnings processes such as earnings inequality models (Cappellari 2004) and microsimulation models of future pension accumulation (Borella 2004).

Chapter 4 argues that the proposed new estimator improves estimates of life-cycle wage-profiles compared to models without selection or with binary selection as it allows to take into account more unobserved heterogeneity from working hours decisions. Applying the proposed method to Dutch administrative data showed that conclusions regarding selection into work over the life-cycle are different from applying the method suggested by Rochina-Barrachina (1999). Whereas the binary selection model of Rochina-Barrachina (1999) suggests negative selection into work for both men and women, the proposed ordered selection model finds positive selection among men and less pronounced negative selection among women. Positive selection suggests that persons with more affluent observed and unobserved characteristics work. The difference indicates that it is important to take into account both participation and hours decisions to account for non-randomness in observed wages.

Analyzing selection effects into part-time and full-time employment and decomposing the analysis for educational levels shows that the importance of the selection may differ between part-time and full-time employment and across educational levels, but the direction of the selection effects is generally positive. Persons with beneficial characteristics select themselves in part-time and full-time employment. This applies to both low-

and high-educated persons. The existence of such selection effects should be corrected for in models that depend on estimating wage processes.

Furthermore, the application of the model indicates that career breaks have a substantial negative effect on life-cycle wages with an average effect of 11% (men) and 7% (women) of the first year which increases up to 21% (men) and 17% (women) from the third year. Also, the model finds substantial part-time wage penalties in life-cycle wages of about 30% for women. For men, we do not find such a part-time wage penalty.

Causal effects between the rise in non-standard work and the decline in early retirement across Europe are analyzed in *Chapter 5*. The analysis finds that part-time employment decreases the labor market withdrawal of older men. More specifically, an increase in the part-time employment rate of 1%-point leads to a reduction in early retirement of 1.7% according to our definition of early retirement in the baseline specification. The results suggest that this is mainly because of the possibility to reduce working hours as we find that specifically voluntary part-time employment induces labor force participation at older ages. However, we do not define whether this reduction in working hours is due to phased retirement (e.g. a reduction in working hours in the same job) or bridge-jobs (e.g. a change in a less demaning job often associated with fewer working hours and a lower wage). Part-time employment as retirement mechnism is also found by Cahill et al. (2006), Gustman and Steinmeier (1984), Kim and DeVaney (2005), Quinn and Kozy (1996), Ruhm (1990, 2006).

Among women, the role of part-time employment in early retirement is somewhat more ambiguous. Effects of part-time employment on early retirement are smaller if significant at all. This finding is in line with the gender differences in gradual retirement found by Peracchi and Welch (1994) and the fact that part-time employment serves a different role over the life-cycle among women (Booth and Van Ours 2008, Gregory and Connolly 2008).

The results at the intensive margin follow a similar pattern as the results at the extensive margin: part-time employment significantly increases the average hours worked of persons aged 55-64 but only for men. Hours effects of part-time employment are not found for women. According to

these results, labor force participation of older men can be increased by creating opportunities for gradual retirement. One such opportunity may be part-time pensions, although it does not increase the labor supply of the elderly per se (Wadensjo 2006). The analysis does not suggest that variation in self-employment has a statistically significant effect on early retirement despite the fact the self-employment is often regarded as a gradual retirement mechanism (Bruce et al. 2000, Fuchs 1982, Giandrea et al. 2008, Gu 2009, Hurd 1996, Morris and Mallier 2003, Zissimopoulos and Karoly 2007).

Chapter 6 rejects the often assumed hypothesis of self-employment being a gradual retirement mechanism by testing whether there is evidence for necessity-driven self-employment prior to retirement. For this purpose Chapter 6 analyzes three testable necessity-hypotheses: 1) whether transitions from unemployment to self-employment are relatively important and increase with age, 2) whether high unemployment rates push workers from paid-employment to self-employment, and 3) whether the introduction of job search requirements for unemployed older workers increases self-employment.

Regarding the first necessity-hypothesis, unemployed and inactive individuals have a higher probability to enter self-employment at the end of working life than those in paid-employment. The probability of moving from paid-employment to self-employment is relatively low and does not increase with age (as would be the case when self-employment would be chosen out of opportunity to reduce working hours at the end of working life). Testing the second necessity-hypothesis suggests that the unemployment rate has a positive effect on transitions from paid-employment to self-employment among men, which suggests that men in paid-employment become self-employed at older ages in order to avoid a period of unemployment. For women, on the other hand, we find a negative effect of the unemployment rate on transitions from paid-employment to self-employment. The third necessity-hypothesis suggests that the introduction of job search requirements at the end of working life have stimulated people to exit unemployment and discouraged people to enter unemployment. This reform, however, did not increase necessity or

opportunity driven self-employment. Individuals that are confronted with search requirements are partly able to find a job, but there are also large substitution effects between unemployment and inactivity (mostly early retirement) which suggests that these options are still more attractive than using self-employment as (voluntary) retirement mechanism.

Based on the three testable hypotheses, the findings in Chapter 6 suggest that at the end of working life individuals with a relatively weak labor market position are more likely to switch to self-employment. The results do not suggest that self-employment is used as a gradual retirement route in the Netherlands despite the many evidence found for the US (Bruce et al. 2000, Fuchs 1982, Giandrea et al. 2008, Gu 2009, Hurd 1996, Zissimopoulos and Karoly 2007). It is likely that these different findings can be explained by the fact that part-time employment is generally more likely in Europe than in the US (OECD 2014). As a consequence, bridge jobs are much less important in European countries (Brunello and Langella 2012) which may explain the higher incidence of transitioning to self-employment at older ages in the US.

Together with Chapter 6, Chapter 5 suggests that part-time (paid) employment possibilities may especially be used as gradual retirement route (for men) in the Netherlands. As long as such part-time work at older ages induces labor supply (as suggested by Chapter 5) part-time employment possibilities may complement the income streams available at old age, thereby increasing the adequacy of resources available in retirement (as analyzed in Chapter 3) as well as the resilience of financial well-being to reforms in the pension system (as mentioned in Chapter 2). However, Chapter 4 indicates that individuals with more affluent characteristics self-select themselves into part-time employment. This means that individuals with less affluent characteristics, who are likely to be more vulnerable to a low income during retirement, have fewer opportunities and/or do not choose to exploit the advantages of part-time employment.

2 Pension reform and income inequality among the elderly in 15 European countries

Abstract

The ageing of populations and hampering economic growth increase pressure on public finances in many advanced capitalist societies. Consequently, governments have adopted pension reforms in order to relieve pressure on public finances. These reforms have contributed to a relative shift from public to private pension schemes. Since private social security plans are generally less redistributive than public social security, it can be hypothesized that the privatization of pension plans has led to higher levels of income inequality among the elderly. This study contributes to the income inequality and pension literature by empirically analysing the distributional effects of shifts from public to private pension provision in 15 European countries for the period 1995-2007. We do not find evidence that shifts from public to private pension provision lead to higher levels of income inequality or poverty among elderly people. The results appear to be robust for a wide range of econometric specifications.

A journal version of this chapter is published in the *International Journal of Social Welfare* (Van Vliet et al. 2012). Different versions of this chapter are published as Goudswaard et al. (2012), Van Vliet et al. (2011). The chapter is co-authored by Olaf van Vliet, Koen Caminada and Kees Goudswaard and was presented at the 18th International Research Seminar of the Foundation for International Studies on Social Security, Sigtuna, Sweden, 8-10 June, 2011 and the NAKE Researchers Day 2011. We would like to thank the participants, in particular Bjørn Hvinden, Tim van Rie, Frank Vandenbroucke and Peter Whiteford, for their helpful suggestions. In addition, thanks to Marike Knoef and Ben van Velthoven for their useful comments.

2.1 Introduction

The ageing of populations has triggered pension reform in many industrialized countries over the past decades. Plans to alleviate the pressure of ageing on public finances have resulted in a trend towards more private pension provision (OECD 2009b, Orenstein 2011). In the pension literature, remarkably little attention has been paid to the distributive effects of these reforms for the elderly. Since private social security arrangements generally entail less income redistribution than public social security (Goudswaard and Caminada 2010), it could be expected that shifts from public to private in the pension provision lead to higher levels of income inequality and poverty among elderly people (Arza 2008). This would imply a trade-off between alleviating the pressure on public finances on the one hand and income inequality among the elderly on the other. The empirical literature in this field exists mainly of either cross-national studies at one moment in time (for example Smeeding and Williamson 2001) or descriptive analyses for a single country (for example Milligan 2008). As a consequence, relatively little insight has been gained about how pension reforms have influenced income inequality and poverty among the elderly in advanced capitalist countries over the past decades.

The aim of the present study was to examine the relationship between the developments in pension systems and the variation in income inequality and poverty among the elderly across countries and over time. First, we analysed to what extent reforms have resulted in a trend towards relatively more private pension provision across advanced capitalist countries. To that end, we used the most recent release of the OECD Social Expenditure database (OECD 2010b). Indeed, in the pension systems of many countries, there have been shifts from public to private in the period 1995-2007, but there is substantial variation across countries. Subsequently, we examined to what extent these shifts have influenced income inequality levels and poverty rates among the elderly, based on a number of pooled time series cross-section regression analyses. Our focus on annual macro data for a relatively short period implies that this article does not contain an integral income redistribution analysis based on the discounted values of lifetime income, contributions paid and benefits received. The main result of our

analysis is that a relatively higher private share of pension provision in a country is not associated with higher levels of income inequality or poverty in that country. With respect to inequality and poverty, the analysis has relied mainly on EU ECHP/SILC data (Eurostat 2011b), but the results appear to be robust for other data and for a wide range of econometric specifications.

The article is structured as follows. In section 2.2, the relationship between pension reform and income inequality among elderly people is introduced. Section 2.3 describes the data, measures and method used in the study. Descriptive statistics, the results of the regression analyses and the sensitivity analyses are presented in section 2.4. In the fifth section, some explanations for our findings are discussed and the article closes with the conclusions in the sixth and last section.

Pension reforms 2.2

Public and private pensions 2.2.1

In an era of ageing populations, relieving public finances is one of the most important drivers of pension reform. An increase in the number of pensioners relative to the labour force leads to increasing budgetary pressure. Budgetary problems as a consequence of cyclical shocks, such as the recession of 2008-2009, may increase the pressure to reform (public) pension systems even further. However, even though the pressure to reform pensions is high, pension reforms are in reality often unruly. Since pensions are based on long-term contracts, reforms are complicated by institutional path dependency (Myles and Pierson 2001). Changes as higher statutory pension ages or reductions of pension benefits are often controversial from a political viewpoint and therefore difficult to realize. Hence, many countries have chosen for a different approach to pension reform. This approach, which is often labelled pension privatization, entails shifts from public pension provision to a mix of public and private pension provisions and a change from the defined benefit to the defined

contribution system (OECD 2009b, Orenstein 2011).[1] In Germany, for instance, the 2001 pension reform consisted of a reduction of the public pension levels and the introduction of state-subsidised, voluntary private pension schemes (Natali and Rhodes 2008). Based on a comparative case study, Arza (2008) showed that this is the type of pension reform that has also been opted for in Italy, Sweden, Poland and the United Kingdom.

The relevant question here is how relative shifts in the pension provision affect the income distribution among elderly people. Public pension plans are generally based on income-related funding and flat rate benefits, which relatively strongly benefit lower income groups. Therefore, public pensions are expected to generate a more equal income distribution and less poverty among the elderly. In a number of OECD countries, the level of public pension benefits is such that a relatively small percentage of pensioners has fallen below the poverty line. Private pension plans, in contrast, are based on a link between contributions paid and benefits received and therefore are not expected to contain elements of (ex ante) income redistribution. A private pension insurance is actuarially fair as a rule. This means that each individual is provided with benefits whose actuarial value is equal to his contributions, given the chance of the insured event occurring. This is the case for individual private pension insurances that have a defined contribution character. However, private earnings-related pension schemes (in the second pillar) may not be actuarially fair and may contain elements of redistribution. This is often the case when (supplementary) private schemes are negotiated by social partners in collective labour contracts. These schemes are mandatory for (a group of) workers. Defined benefit pension schemes, for example, generally redistribute resources both within generations (for instance through redistributive elements such as thresholds or ceilings) and across generations (risk sharing, back service). Also, tax advantages (to households or to employers) can be used to stimulate the provision of private pensions. This is often the case in supplementary pension programs, where contributions are tax

[1] A shift from public to private pension provision alleviates the pressure on public finances, but it does not solve financing problems of the pension system. If a deficit of the pension system is considered as unsustainable, the only solution to make it sustainable is reducing benefits, increasing contributions, or both, either publicly or privately (Barr and Diamond 2009).

exempt (Yoo and De Serres 2004). The fiscal advantages related to, for example, supplementary private pension plans are positively related to income levels in most countries and thus favour the rich (Goudswaard and Caminada 2010). In general, as Ferrarini and Nelson (2003) showed, social insurance is less equalising after taxation in all countries.

In summary, it seems plausible that private pension schemes will generate less income redistribution from rich to poor than public programs, although at this stage their distributional impact in a cross-country analysis is not fully clear. In other words, there are good reasons to expect that relative shifts from public to private pensions lead to higher income inequality among the elderly.

Earlier findings 2.2.2

Much literature analyses the relationship between social security and income inequality in general. Based on a cross-national study at the macro level Smeeding and Williamson (2001) concluded that high levels of public social spending are associated with low levels of income inequality and poverty. Caminada and Goudswaard (2005) and Goudswaard and Caminada (2010) compared the redistributive effects of public and private social security. Taking a broad definition of social security and based on an international comparative analysis, they concluded that the redistributive effect of private social security is smaller than that of public social security. However, Caminada et al. (2012) found no significant effects of private social expenditure on poverty rates.

Interestingly, the findings for pensions seem to be less ambiguous than for social security in general. A number of cross-sectional studies indicate that income inequality among elderly people is lower as larger shares of the income of the elderly consist of public pension benefits (Brown and Prus 2004, Fukawa 2006, Weller 2004). The number of studies on the income effects of private pensions is considerably smaller, but Schirle (2009) found for Canada that a larger private share in the pension provision is associated with an increasing income inequality among elderly people. Combining the results of the studies on public and private pensions, it

seems plausible that a shift from public to private leads to more income inequality among the elderly.

Comparable effects of shifts in the public/private-mix of pensions were found for poverty among elderly people. Based on country-specific analyses over time, Oshio and Shimizutani (2005) and Milligan (2008) concluded that a larger public share in the pension provision is related to less poverty among elderly people. Hughes and Stewart (2004) found that increases in the private share are associated with an increase in the poverty rate among elderly people.

From a methodological perspective, the empirical literature on pension reform and income inequality can be divided into two types of studies. The first type consists of cross-sectional studies, comparing a number of countries in a certain year (Brown and Prus 2004, Fukawa 2006, Weller 2004). In these studies, the effects of pension reform cannot be analysed over time. The second type of studies is mainly focused on developments over a longer period but, for a single country (Milligan 2008, Myles 2000, Oshio and Shimizutani 2005, Schirle 2009). In these studies, it is quite difficult to examine whether the findings also hold for other comparable pension reforms in other countries. Therefore, in this study the dimension time will be incorporated into a cross-sectional analysis.

2.3 Data, measures and method

2.3.1 Public and private pension expenditure

Most comparative studies on welfare states rely on social expenditures as indicator to analyse welfare state reforms across different countries. To examine changes in the public/private-mix of pension provision, we used data from the most recent OECD Social Expenditure Database (OECD 2010b). This database contains social expenditure data on both public and private pension schemes. In this database, programmes are classified as social when two conditions are simultaneously satisfied (Adema 2010, Adema and Ladaique 2009). First, they have to be intended to serve a social purpose, such as old-age. Other policy areas with a social purpose are:

survivors, incapacity related benefits, health, family, active labour market policies, unemployment, housing and a category of other social security areas. Second, they have to involve either interpersonal redistribution or compulsory participation. Hence, purely private old-age plans which are the result of direct market transactions by individual people are not included. The distinction between public and private social security is based on the institution that controls the financial flows, namely public agencies or private bodies.

Our study analysed public and private social pension expenditure, both expressed as percentage of GDP and as millions of U.S. dollars (constant (2000) prices, ppp) per pensioner. A relevant measure is the share of private social pension expenditure as percentage of total social pension expenditure. This measure provides a good indication of shifts in the public/private-mix. The measure for private social pension expenditure indicates the total of expenditures on mandatory and voluntary pension schemes.[2] Furthermore, the indicators include expenditures on incomes of people who retired at the statutory retirement age and of early retirees.[3] Expenditures on survivor pensions are not included in the indicators.

In a cross-national analysis at the macro-level, expenditure indicators have some limitations (Van Vliet 2010). First, differences in expenditure patterns may be driven by differences in demographic trends across countries. When increases in pension expenditure fall short of increases in the number of retirees, this may have negative consequences for the incomes of elderly people and for the income inequality among the elderly. To control for the ageing of populations, we included a control variable measuring the percentage of population aged 65 and above. For this measure,we used data from Eurostat (2011a). Second, expenditures do not indicate institutional differences in pension systems, such as a pay-as-you-go ver-

[2]The OECD Social Expenditure Database also provides the possibility to present expenditures on mandatory and voluntary private pension separately. However, since the classification of private pension spending into mandatory and voluntary pension schemes is not unequivocal, we mainly use the total of these two categories. As a robustness check, we also analysed the income effects of mandatory and voluntary pension schemes separately, which did not alter the results (reported below). Note that the classification into public, mandatory private and voluntary private pensions coincides with the classification into first, second and third pillar pensions respectively.

[3]Expenditures on public pensions also include spending on some other services for the elderly (see Vandenbroucke and Vleminckx 2011).

sus a funded system, or a defined benefit versus a defined contribution system. Third, the variation in the tax treatment of contributions and benefits across countries was not taken into account. Ideally, we would use net expenditure on pensions, after tax, but international standardised data for such an indicator are unfortunately not available for a longer period. Despite these limitations, pension expenditures can give an indication of shifts from public to private pensions.

2.3.2 Income inequality and poverty among the elderly

For income inequality and poverty among the elderly, the study relied on two indicators provided by Eurostat (2011b). Income inequality among the elderly was measured by the S80/S20 ratio of people aged 65 and over. This indicator is constructed by dividing the total disposable income of the top 20 percent incomes of elderly by the total equivalized disposable income of the bottom 20 percent incomes of people aged 65 and over. A higher value of this indicator implies a higher inequality among the elderly. Although this indicator gives a good indication of income inequality at the extremes of the distribution, it neglects shifts between other quintiles. Therefore, the *Gini-coefficient* and the *Atkinson index* may be preferred measures of income inequality. However, data on income inequality among elderly measured by either the Gini coefficient or the Atkinson index are not available for a reasonable number of years.

Poverty among the elderly is measured by the percentage of people aged 65 and over who live below the poverty line of 60 percent of median equivalized (disposable) income of the total population. This poverty line of 60 percent is also officially used as poverty measure by the European Union. A higher value of this indicator implies a higher rate of poverty among elderly. Note that this indicator is a relative poverty line and can therefore be seen as a detailed representation of income inequality for the lower part of the income distribution.

Method 2.3.3

To examine the relationship between pension reform and income inequality among the elderly, we run a number of pooled time series cross-section regression analyses. Our estimations take the following form:

$$Q_{it} = \alpha + \beta X_{it} + \delta Z_{it} + \mu_i + \lambda_t + \epsilon_{it} \quad (2.1)$$

In equation (2.1), Q represents the dependent variables of income inequality (S80/S20) or poverty (PL 60) among the elderly. Variables describing the pension system, public pension expenditures, private pension expenditures and total pension expenditures (all as a percentage of GDP) and the private share of pension expenditures (private pension expenditures as percentage of total pension expenditures), are represented in vector X. The control variables ageing (share of people aged 65 and over relative to total population) and real GDP per capita (constant (2000) prices ppp) are represented in vector Z. For the latter variable, we used data from the OECD (OECD 2011a). Recognising that the variation in income inequality and poverty among the elderly may be related to unobserved country- and year-specific effects, country (i) and year (t) dummies are modelled by μ and λ, respectively. The error-term ϵ follows an *AR(1)*-process to correct for autocorrelation. In addition, we used panel-corrected standard errors to correct for panel-heteroskedasticity and simultaneous spatial correlation (Beck and Katz 1995).

Constrained by data availability, the study covered 15 European countries (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Sweden and the United Kingdom) for the years 1995 up till 2007. In the dataset, a number of observations were missing. This is in particular the case for the income inequality and poverty data and especially for Scandinavian countries. However, all countries were included in the regression analyses. Several sensitivity tests, which are discussed below, indicate that results did not suffer from the missing data.

2.4 Empirical analysis

2.4.1 Descriptive statistics

Table 2.1 illustrates the developments in pension expenditures for the included countries from 1995 up till 2007. On average, social expenditures as a percentage of GDP on both public and private pensions have been increased. Hence, total pension expenditure as a percentage of GDP has increased too. Furthermore, the data show an increase in the private share of total pension expenditure, albeit to a limited extent. Private spending as a share of total pension expenditure rose on average from 14.3 percent in 1995 to 14.9 percent in 2007. This indicates a relative shift from public to private in the pension provision. More interestingly, there is substantial variation in the developments of private pension expenditure as a share of total pension expenditure across countries. In Belgium for instance, social expenditures on private pensions increased more than expenditures on public pensions. This has resulted in a shift from public to private in the total pension expenditure. This trend fits well with the trend that is reported in Peeters et al. (2003), which is based on data from national sources. In other countries, such as Denmark, Ireland, Italy, the Netherlands, Sweden and the UK, shifts in the public/private-mix are the result of opposing trends in public and private pension expenditure. Indeed, there is a negative correlation between yearly changes in private and public pension expenditure of -0.22 which is significant at the 1 percent level.

Table 2.2 shows a general trend towards less income inequality and less poverty among the elderly in the period 1995-2007.[4] In 2007, the average income inequality among elderly (mean 12 countries) has decreased by

[4]These results should be interpreted with caution, because there is a disruption in the time series of inequality and poverty indicators presented in table 2.2. Until 2001, data were provided by the European Community Household Panel survey (ECHP). Since 2005 all EU-15 countries provide data from the new European Union Statistics on Income and Living Conditions (EU-SILC). During the transitional period poverty indicators were provided by national sources which were harmonized ex-post as closely as possible with EU-SILC definitions by Eurostat. Despite the fact that most EU-SILC variables are defined in the same way as the corresponding ECHP variables, some differences arise. The transition from ECHP to EU-SILC possibly explains the large number of missing observations in this period. See for more details Eurostat (2005).

almost 18 percent compared to 1995. A decreasing trend over time is also shown by the poverty rate among the elderly which has decreased by almost 9 percent on average. There is some variation in trends between countries still. Greece and Portugal, for example, have shown a huge decline in poverty rates among elderly over time. However, Finland and Ireland have faced a relatively large increase in poverty among elderly in the same period. These trends are robust with respect to the poverty lines applied (50, 60 or 70% of median equivalized income). Nevertheless, different patterns of poverty can be seen within countries. Germany and the Netherlands, for example, have shown a decrease in poverty rates among elderly when using poverty line of 50 percent, while both countries have shown an increase in poverty rates in the same period when using a poverty line of 60 and 70 percent. These observations imply that relatively more elderly live at risk of poverty in 2007 compared with 1995, but less elderly find themselves at the absolute bottom of the income distribution.

Table 2.1: Pension expenditures in European countries, 1995-2007

	Public pension expenditure (%GDP)			Private pension expenditure (%GDP)			Total pension expenditure (%GDP)			Private pension expenditure (%total)		
	1995	2007	Change 95-07	1995	2007	Change 95-07	1995	2007	Change 95-07	1995	2007	Change 95-07
Austria	10	10.7	0.7	0.4	0.5	0.1	10.4	11.3	0.8	3.7	4.7	0.9
Belgium	7	7.1	0.1	1.3	2.8	1.5	8.3	9.9	1.6	15.5	28.4	12.8
Denmark	8.4	7.3	-1.1	1.8	2.2	0.3	10.2	9.5	-0.7	18	23	5
Finland	8.5	8.4	-0.1	0.3	0.2	-0.1	8.8	8.6	-0.2	3.4	2.2	-1.1
France	10.6	11.1	0.5	0.1	0.2	0.1	10.7	11.2	0.5	1	1.5	0.5
Germany	8	8.7	0.7	0.7	0.7	0.1	8.6	9.4	0.8	7.6	7.9	0.3
Greece	9.2	10	0.9	0.4	0.4	0	9.5	10.4	0.9	3.8	3.5	-0.3
Ireland	2.9	3.1	0.2	1.1	0.9	-0.2	3.9	4	0.1	26.8	22.4	-4.4
Italy	9.3	11.7	2.4	2.8	1.3	-1.5	12.2	13.1	0.9	23.4	10.1	-13.3
Luxembourg	8.2	4.8	-3.4	0.6	0.3	-0.3	8.8	5.2	-3.7	6.8	6.6	-0.2
Netherlands	5.5	5.3	-0.3	2.6	3.5	0.9	8.1	8.8	0.7	31.9	40.2	8.4
Norway	7.1	6.2	-0.8	0.6	0.6	-0.1	7.7	6.8	-0.9	8.2	8.5	0.3
Portugal	6	9.2	3.2	0.2	0.2	0	6.2	9.4	3.2	2.7	1.8	-0.8
Sweden	9.8	9	-0.8	1.9	2.1	0.2	11.7	11.1	-0.7	16.3	19	2.7
United Kingdom	5.5	5.8	0.3	4.7	4.5	-0.2	10.2	10.3	0	46.2	43.8	-2.5
Mean	7.7	7.9	0.2	1.3	1.4	0.1	9	9.3	0.2	14.3	14.9	0.6

Note: Figures for Luxembourg (1995) are based on linear extrapolation. Source: OECD Social Expenditure Database (OECD, 2010) and own calculations.

Regression results 2.4.2

The social outcomes presented above suggest that there is no evidence that an increasing share of private pensions leads to higher income inequality and poverty among elderly. In Belgium, for instance, the country with the largest relative shift from public to private, income inequality and poverty among the elderly decreased. In Italy, the country with the largest relative shift from private to public, an increase in income inequality and poverty rates among the elderly can be observed. In order to take our analysis beyond the descriptive statistics, we continue with regression analyses on the 15 European countries over the years 1995-2007.

The results of the regression analyses are presented in table 2.3 and table 2.4. The effects of public pension expenditure as percentage of GDP on income inequality among the elderly are negative, but not significant. Model 7 indicates that public pension expenditure as percentage of GDP is negatively and significantly related to poverty among the elderly. Consistent with our expectations based on the literature, this suggests that higher social spending on public pensions is associated with lower poverty rates among the elderly. However, the results in Models 9 and 11 indicate that there is only weak evidence for this relationship.

With respect to private pension expenditure as percentage of GDP, the results do not indicate a positive effect of private pension expenditure on income inequality. In contrast, the negative coefficients suggest that private pension expenditure as a percentage of GDP is negatively related to income inequality among the elderly. Model 2 indicates also a negative coefficient for private pension expenditure as percentage of total pension expenditure, but the Models 3 and 5 show a positive effect for the private share of the pension provision, albeit not significant. However, when pension expenditure is expressed in dollars per pensioner, to exclude any denominator effect of GDP, the results indicate a negative effect for the private pension expenditure as a share of total pension expenditure again. This implies that higher spending on private pensions in general, and a shift from public to private pensions in particular, are not associated with higher income inequality among the elderly. Turning to poverty, all

Table 2.2: Trends in social outcomes among elderly people, 1995-2007

	Income inequality among the elderly (S80/S20)			Poverty among the elderly (PL 60)		
	1995	2007	Change 95-07	1995	2007	Change 95-07
Austria	4.0	3.2	-0.8	20.0	14.0	-6.0
Belgium	4.9	3.4	-1.5	25.0	23.0	-2.0
Denmark	-	2.7	-	-	17.7	-
Finland	3.0	2.9	-0.1	12.0	21.6	9.6
France	4.8	4.0	-0.8	19.0	13.1	-5.9
Germany	4.9	4.2	-0.7	15.0	16.2	1.2
Greece	7.6	4.8	-2.8	35.0	22.9	-12.1
Ireland	3.9	3.4	-0.5	19.0	28.3	9.3
Italy	4.6	4.7	0.1	18.0	21.9	3.9
Luxembourg	4.1	3.2	-0.9	12.0	7.2	-4.8
Netherlands	4.2	3.2	-1.0	8.0	9.5	1.5
Norway	-	2.8	-	-	14.1	-
Portugal	6.6	6.0	-0.6	38.0	25.5	-12.5
Sweden	-	2.8	-	-	9.9	-
United Kingdom	4.9	4.4	-0.5	32.0	27.6	-4.4
Mean (all countries)	4.8	3.7	-1.1	21.1	18.2	-2.9
Mean (12 countries)	5.2	4.3	-0.9	23.0	21.0	-2.0

Note: Mean 12 countries excluding Denmark, Norway and Sweden. Eurostat SILC-database (Eurostat, 2011a) and own calculations.

measures for private pension provision are not significantly related to poverty among the elderly.

The results for total pension expenditure are comparable to the case of public pension expenditures. Total pension expenditure as a percentage of GDP, which consists of the sum of public and private pension expenditure, is negatively and significantly correlated with poverty among the elderly, while no significant correlation can be observed between total pension expenditures and income inequality among the elderly.

As to graying populations, the results indicate that the effect of graying on income inequality and poverty among the elderly is limited. It seems that the percentage of the population aged 65 and over is slightly negatively correlated with income inequality among the elderly, while no correlation can be observed between this variable and poverty among the elderly. The results suggest that there is no clear linkage between GDP per capita and income inequality among the elderly. However, GDP per capita is positively and significantly associated with poverty rates among the elderly.

In summary, the results of the regression analyses suggest that higher private expenditure for pensions as a percentage of GDP, per pensioner and as a share of total pension expenditure are not associated with higher levels of income inequality among the population aged 65 and above. Furthermore, the regression analyses indicate a poor linkage between private provisions of pension schemes and poverty rates among the elderly. Taken together, these results do not provide evidence for the expectation that shifts from public to private pension provision are associated with higher levels of income inequality.

Table 2.3: Panel data regressions of pension expenditures and income inequality (s80/s20) among the elderly (65+)

	Model 1		Model 2		Model 3		Model 4		Model 5		Model 6	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.13	0.12			-0.07	0.12			-0.19	0.13		
Private pension exp. (%GDP)	-0.45***	0.17			-1.22**	0.52			-0.87**	0.42		
Private share (%total)			-3.41**	1.52	9.2	5.85			3.58	4.77	-6.82***	2.52
Total pension exp. (%GDP)							-0.2	0.13				
Public pension exp. p.p (/100)											-0.01*	0.01
Private pension exp. p.p. (/100)											-0.01	0.01
Population aged 65 and over (%total)	-0.22**	0.11	-0.22*	0.12	-0.26**	0.11	-0.18	0.11	-0.34***	0.13		
GDP per capita (/1000)									-0.09***	0.03	0.00	0.03
Constant	8.96***	2.14	7.62***	1.78	8.99***	2.02	9.03***	2.23	13.74***	2.92	7.12***	1.12
Observations	135		135		135		135		135		135	
Adj. R-squared	0.84		0.84		0.84		0.84		0.85		0.83	
Rho	0.41		0.42		0.35		0.48		0.36		0.39	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level. Exp. p.p. is expenditures per pensioner. Each regression also includes country and year dummies (not shown here). Countries included: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Sweden, and the United Kingdom. Data sources: Income inequality (Eurostat, 2011); Pension expenditure: OECD Social Expenditure Database (2010).

Table 2.4: Panel data regressions of pension expenditures and poverty (PL 60) among the elderly (65+)

	Model 7		Model 8		Model 9		Model 10		Model 11		Model 12	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-1.50*	0.83			-1.13	0.91			-0.29	1.14		
Private pension exp. (%GDP)					-4.04	3.31			-5.23	3.84		
Private share (%total)			2.61	11.96	40.38	48.43			65.32	56.73	-0.95	39.82
Total pension exp. (%GDP)							-1.34*	0.70				
Public pension exp. p.p (/100)											-0.11	0.07
Private pension exp. p.p. (/100)											-0.01	0.15
Population aged 65 and over (%total)	-0.54	0.80	-0.78	0.76	-0.70	0.80	0.61	0.78	-0.02	0.64		
GDP per capita (/1000)									0.81**	0.34	1.08***	0.30
Constant	42.75***	11.58	30.42***	11.49	41.05***	11.46	42.37***	11.67	1.38	16.85	12.50	15.06
Observations	154		154		154		154		154		154	
Adj. R-squared	0.79		0.78		0.79		0.79		0.80		0.80	
Rho	0.63		0.66		0.63		0.64		0.63		0.59	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level. Exp. p.p. is expenditures per pensioner. Each regression also includes country and year dummies (not shown here). Countries included: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Sweden, and the United Kingdom. Data sources: Income inequality (Eurostat, 2011); Pension expenditure: OECD Social Expenditure Database (2010).

2.4.3 Sensitivity analyses

Since the results are not in line with our expectations based on both theoretical and empirical literature on pension reform and income inequality, we perform a variety of robustness checks. First, we examine the dependence of the results on different specifications of the empirical model. Estimations without correction for autocorrelation or panel-corrected standard errors do not alter the result that shifts towards more private pensions are not correlated with higher income inequality or poverty levels among the elderly. With respect to the most important independent variable, the share of private pension expenditure as percentage of total pension expenditure, it should be noted that the variation within countries over time is rather small. In combination with country fixed-effects, this reduces in itself the chance to find any significant effects for this variable. Therefore, we ran the analyses also without country fixed-effects, which did not alter the results. Other specifications that we applied, such as first differences, log-transformations, lagged variables or the exclusion of year fixed-effects did not change the results. Neither do the results change if an independent variable as graying is excluded or if a measure for a country's wealth such as GDP per capita is included.

To further probe the robustness of our results, we took into account that our analyses were based on unbalanced panels owing to a number of missing observations. This is especially the case for Scandinavian countries with regard to the inequality and poverty indicators. This could lead to biased results, since Denmark, Finland, Norway and Sweden have below average poverty and inequality levels. Therefore, we also ran regressions in which the number of observations is extended. We employed regression analyses in which both the dependent and the independent variables are linearly inter- and extrapolated and analyses in which only the dependent variables are inter- and extrapolated. Since extrapolation is associated with more uncertainty than interpolation, we also ran regressions for only interpolated dependent variables. In addition, we used more sophisticated techniques such as cubic- and cubic-spline interpolation.[5] All analyses

[5]We use several inter- and extrapolation techniques such as linear inter- and extrapolation, cubic interpolation and cubic spline interpolation. For applications of linear interpolation, see for example L'horty and Rault (2003), Clarke et al. (2008), Stern (2005)

indicate that our results are not biased by the missing observations. This is also confirmed by the results of regression analyses in which the group of Scandinavian countries is omitted. These results are in line with the findings of Gustafsson and Johansson (1999), who found that the group of Scandinavian countries do not influence the results of regression analyses on income inequality and social expenditure very strongly. Moreover, the results are neither affected by excluding the countries one by one in the regression analyses. We also tested to what extent the results are driven by the countries with the largest shifts in the public/private-mix, namely Belgium, Italy and the Netherlands. Regressions without these three countries yielded similar results.

Finally, we examined the sensitivity of the results for the use of alternative indicators and data sources. With regard to poverty, the results of estimation of poverty lines among the elderly of 50 and 70 percent are comparable to the results of the poverty line of 60 percent. Subsequently, we ran our main empirical specification for four different indicators for income inequality among the elderly based on data from the OECD (OECD 2008).[6] Our results with respect to the linkage between the share of private pension expenditure and income inequality among the elderly appear to be robust for Gini coefficients before and after taxes and transfers, the standard coefficient of variation and the mean log deviation. Additionally, the replication of the results presented in table 2.3 and 2.4 with Gini coefficients and poverty lines[7] among the elderly from the Luxembourg Income Study (LIS 2011) confirm our empirical results based on Eurostat data. As to the independent variable, the measures for private pension expenditure can be disaggregated into mandatory and voluntary private pension expenditure. The results of the regression analyses with the dis-

and Toroj (2008). An example of cubic-spline interpolation in economics, which is based on polynomial instead of linear methods, can be found in Nanda and Ross (2008).

[6]The main advantage of these OECD data is the availability of more sophisticated income inequality indicators. However, the most important disadvantage of these data is that at most 6 data points per country are available in the waves from mid 1970s till mid 2000s. Another advantage is that these data are available for a longer period and a larger group of countries than the Eurostat data (Eurostat 2011b). Our results also hold for this larger country group and longer period.

[7]The Gini coefficient among the elderly is provided by Wang and Caminada (2011) who constructed this indicator from the micro data. The 40, 50 and 60 percent poverty lines among the elderly are taken from the LIS Key Data (LIS 2011).

aggregated measure do not differ from the results with the aggregated measures. In summary, the combined evidence of these robustness checks suggest that our results are robust with respect to different specifications, variables and data sources.

2.5 Discussion

A number of tentative explanations is conceivable for our main finding that shifts towards relatively more private pensions are not related to higher levels of income inequality among the elderly. The level of supplementary pension benefits is often strongly related to the income level during working life. A more private pension provision therefore leads to a higher supplementary pension for higher incomes than for lower incomes. But it could be possible that even though the absolute increase in private pension benefits is smaller for lower incomes than for higher incomes, the relative increase for lower incomes is much larger than for higher incomes. This is illustrated by Myles and Pierson (2001) in a study on pension reform in Canada in the beginning of the 1990s. Burtless (2006) also stated that the effects of changes in the public/private-mix of pensions on replacement rates, the income from pensions relative to income from work in the past, vary along the income distribution. A possible scenario is that the coverage of private pensions has increased and that this is mainly the case for lower income groups. This could be an explanation for the fact that we did not find a relationship between shifts in the public/private-mix of pensions and income inequality and poverty among the elderly. Hence, further research at the macro-level could be focused on specifying the effects of pension reform for different quintiles of the income distribution.

In addition, it should be noted that the analyses in this study do not account for determinants of income at the individual level. On the one hand, this concerns general personal characteristics which determine income such as education. On the other hand, current individual pension benefits are determined by long-term effects such as previous wages, contributions paid and macro-economic conditions in the past. It is hardly

possible to capture this time dimension in a macro-level analysis. Another factor that might influence pension benefits and incomes of elderly people is the prevalence of deficiencies in contributions paid in the past (Esping-Anderson and Myles 2006). Future empirical research based on micro-data, in which it is possible to control for individual characteristics, may provide more insight into the relationship between pension reform and income inequality.

Finally, the use of pension expenditure data at the macro-level implies some restrictions. Much information can be lost in classifying pension programmes into pillars (Whitehouse 2002). Moreover, as mentioned above, shifts in pension expenditure can only give a rough indication of changes in institutional characteristics of pension systems.

Conclusion 2.6

In many industrialized countries, public pension systems have been reformed in order to alleviate the pressure on public finances resulting from ageing populations. This has often led to shifts in the pension provision from public to private. The average magnitude of these shifts remains limited, but in a number of countries there have been substantial changes. Since private pensions are probably less redistributive than public pensions, these shifts could be hypothesized to lead to more income inequality among retirees. This study contributes to the income inequality and pension literature by empirically analyzing the income effects of shifts in the public/private-mix of pensions in 15 European countries for the period 1995-2007, using pooled time series cross-section regression analyses. The most important finding is that shifts in the pension provision from public to private do not (yet) seem to entail higher levels of income inequality or poverty among people aged 65 and older. Intriguingly, this finding is not in line with expectations in the literature on pension reform and income inequality (Arza 2008, Fukawa 2006, Hughes and Stewart 2004, Milligan 2008, Oshio and Shimizutani 2005, Weller 2004) and with literature on the redistributive effects of public and private social security in general (Caminada and Goudswaard 2005, Goudswaard and Caminada

2010). A tentative explanation for this finding is that more people in lower income deciles have been covered by private pension plans. As a result, the increases in the pension benefits of people with lower incomes were relatively larger than for people with higher incomes.

The policy implication of our findings seems to be that the pressure of the pension expenditures on public finances can be alleviated without serious consequences for income inequality or poverty among elderly people. However, this policy implication should be taken with much caution, even though the results are robust for other data sources and a broad range of alternative econometric specifications. As suggested before, our results could be explained by increases in the coverage of private supplementary pension schemes rather than policy reforms. A higher coverage of private programs also causes a shift from public to private, but will probably have a rather different distributional impact compared to cutting public pension benefits. In addition, empirical research at the macro-level goes along with a number of limitations with respect to institutional characteristics of pension systems and individual characteristics of pensioners.

Finally, it should be noted that our analysis does not include the years after 2007. This implies that we have no prospect of the income effects of the pension reforms which are triggered by the credit crisis at the beginning of the 21st century. The results of this study provide no reason to expect that recent reforms in many European countries will lead to more income inequality and higher poverty rates among the elderly. Future research should provide more insight into the answer to this question.

2.A Sensitivity analyses

Tables 2.3 and 2.4 of the main text present the results of panel data regressions of pension expenditures and social outcomes among the elderly (65+). Tables 2.5 to 2.15 below present the result of several robustness checks:

Table 2.5: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **excluding graying** (population aged 65 and over, % total)

Table 2.6: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **including GDP per capita**
Table 2.7: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly inter- and extrapolated dependent and independent variables**
Table 2.8: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly inter- and extrapolated independent variables**
Table 2.9: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly interpolated independent variables**
Table 2.10: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with cubically interpolated independent variables**
Table 2.11: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with cubically spline interpolated independent variables**
Table 2.12: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **excluding Scandinavian countries**
Table 2.13: Panel data regressions of pension expenditures and **different poverty lines (PL 50, PL 70)** among the elderly (65+)
Table 2.14: Panel data regressions of pension expenditures and income inequality among the elderly (65+) **using income distribution variables of OECD (2010) instead of Eurostat (2011)**
Table 2.15: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **using a decomposition of private pension expenditures into mandatory and voluntary expenditures**

Table 2.5: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **excluding graying**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.18	0.14					-1.63**	0.79				
Private pension exp. (%GDP)	-0.41**	0.18					-1.02	0.84				
Private share (%total)			-2.99**	1.25					3.93	11.05		
Total pension exp. (%GDP)					-0.23	0.14					-1.43**	-0.70
Population aged 65 and over (%total)												
Constant	6.25***	1.48	4.37***	0.11	34.23***	7.50	36.13***	8.24	18.86***	1.46	34.23***	-7.50
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	135		135		135		154		154		154	
Adj. R-squared	0.84		0.83		0.84		0.79		0.78		0.79	
Rho	0.42		0.42		0.47		0.62		0.65		0.64	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.6: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **including GDP per capita**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.22*	0.12					-0.94	1.01				
Private pension exp. (%GDP)	-0.58***	0.18					-0.50	0.79				
Private share (%total)			-4.39***	1.51					7.98	14.05		
Total pension exp. (%GDP)					-0.28**	0.14					-0.78	0.87
Population aged 65 and over (%total)	-0.33***	0.12	-0.32***	0.13	-0.27**	0.12	0.13	0.65	0.11	0.61	0.04	0.67
GDP per capita (/1000)	-0.10***	0.03	-0.08**	0.03	-0.08***	0.03	0.71***	0.36	0.81***	0.29	0.70**	0.36
Constant	14.03***	2.95	11.05***	2.35	13.32***	3.06	9.03	16.89	-3.42***	9.30	8.98	16.52
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	135		135		135		154		154		154	
Adj. R-squared	0.85		0.84		0.85		0.79		0.79		0.79	
Rho	0.36		0.34		0.46		0.64		0.65		0.65	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.7: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly inter- and extrapolated dependent and independent variables**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	0.08	0.07					-0.06	0.48				
Private pension exp. (%GDP)	-0.34***	0.12					-0.03	0.77				
Private share (%total)			-5.31***	1.64					1.42	12.04		
Total pension exp. (%GDP)					0.07	0.09					0.03	0.55
Population aged 65 and over (%total)	0.00	0.10	0.03	0.10	0.03	0.09	1.64***	0.55	1.58***	0.48	1.57***	0.51
Constant	3.77***	1.46	4.14***	1.52	3.16*	1.83	-3.69	6.74	-3.53	7.54	-3.69	7.64
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	195		195		195		195		195		195	
Adj. R-squared	0.75		0.76		0.75		0.72		0.71		0.71	
Rho	0.55		0.58		0.61		0.70		0.74		0.73	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.8: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly inter- and extrapolated independent variables**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	0.18	0.10					-0.02	0.58				
Private pension exp. (%GDP)	-0.25***	0.13					0.03	0.72				
Private share (%total)			-5.20***	1.57					-1.70	9.47		
Total pension exp. (%GDP)					0.07	0.09					0.00	0.52
Population aged 65 and over (%total)	-0.01	0.12	0.04	0.12	0.05	0.11	1.68***	0.44	1.67***	0.36	1.66***	0.43
Constant	2.76	2.06	3.97**	0.76	2.9	2.03	-4.75	5.85	-4.82	5.61	-4.66	5.90
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	189		189		189		189		189		189	
Adj. R-squared	0.76		0.76		0.75		0.72		0.71		0.71	
Rho	0.60		0.60		0.61		0.71		0.72		0.71	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.9: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with linearly interpolated independent variables**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	0.09	0.11					-1.38	0.64				
Private pension exp. (%GDP)	-0.35***	0.13					-1.05	0.79				
Private share (%total)			-2.90**	1.44					-1.33	10.64		
Total pension exp. (%GDP)					-0.16	0.11					-1.22**	0.57
Population aged 65 and over (%total)	-0.21**	0.10	-0.21*	0.11	-0.18*	0.10	-0.45	0.57	-0.63	0.54	-0.50	0.56
Constant	8.43***	1.91	7.47***	1.63	8.50***	1.94	40.28***	0.55	28.4***	8.35	39.54***	8.64
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	164		164		164		169		169		169	
Adj. R-squared	0.83		0.83		0.83		0.75		0.73		0.74	
Rho	0.43		0.45		0.49		0.62		0.66		0.64	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.10: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with cubically interpolated independent variables**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.09	0.11					-1.39*	0.64				
Private pension exp. (%GDP)	-0.33**	0.14					-1.02	0.78				
Private share (%total)			-2.73**	1.47					1.58	10.51		
Total pension exp. (%GDP)					-0.15	0.11					-1.23**	0.56
Population aged 65 and over (%total)	-0.20**	0.10	-0.21*	0.11	-0.18*	0.10	-0.45	0.52	-0.64	0.50	-0.51	0.52
Constant	8.37***	1.97	7.43***	1.67	8.46***	2.00	40.41***	8.02	28.58***	7.70	39.73***	8.09
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	164		164		164		168		168		168	
Adj. R-squared	0.83		0.82		0.82		0.75		0.74		0.75	
Rho	0.48		0.48		0.53		0.62		0.65		0.63	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.11: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **with cubically spline interpolated independent variables**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.09	0.11					-1.40*	0.73				
Private pension exp. (%GDP)	-0.32**	0.14					-1.02	0.85				
Private share (%total)			-2.67*	1.48					2.83	12.21		
Total pension exp. (%GDP)					-0.15	0.11					-1.26*	0.65
Population aged 65 and over (%total)	-0.21*	0.11	-0.21*	0.11	-0.18*	0.11	-0.47	0.66	-0.69	0.64	-0.55	0.63
Constant	8.39***	2.00	7.47***	1.71	8.51***	2.07	40.95***	9.8	29.09***	9.80	40.45***	9.88
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	160		160		160		162		162		162	
Adj. R-squared	0.83		0.83		0.83		0.77		0.76		0.77	
Rho	0.43		0.43		0.54		0.64		0.67		0.65	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.12: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **excluding Scandinavian countries**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	0.00	0.11					-1.58*	0.84				
Private pension exp. (%GDP)	-0.31**	0.16					-0.90	0.81				
Private share (%total)			-2.77*	1.56					7.10	12.12		
Total pension exp. (%GDP)					-0.06	0.11					-1.30*	0.69
Population aged 65 and over (%total)	-0.24*	0.11	-0.22**	0.11	-0.21*	0.11	-0.94	0.75	-1.18	0.73	-1.06	0.74
Constant	7.96***	1.81	7.7***	1.67	8.00**	1.85	49.86***	11.40	36.79***	10.92	49.06***	11.46
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	112		112		112		123		123		123	
Adj. R-squared	0.85		0.86		0.86		0.79		0.78		0.79	
Rho	0.45		0.48		0.54		0.61		0.65		0.63	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.13: Panel data regressions of pension expenditures and **different poverty lines (PL 50, PL 70)** among the elderly (65+)

	Poverty (PL 50) among the elderly (65+)						*Poverty (PL 70)* among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-1.98***	0.59					-1.77**	0.86				
Private pension exp. (%GDP)	-1.88**	0.92					-1.40	1.19				
Private share (%total)			-6.62	11.39					-5.35	12.93		
Total pension exp. (%GDP)					-1.91***	0.53					-1.66**	0.68
Population aged 65 and over (%total)	-0.87*	0.51	-1.16**	0.47	-0.90*	0.49	0.18	0.53	-0.11	0.51	-0.12	0.49
Constant	43.26***	6.22	26.92***	7.16	42.99***	6.38	44.08***	8.88	30.05***	7.93	43.9***	9.03
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	128		128		128		128		128		128	
Adj. R-squared	0.79		0.78		0.79		0.88		0.88		0.88	
Rho	0.42		0.38		0.43		0.60		0.59		0.60	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

Table 2.14: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **using income distribution variables of OECD (2010) instead of Eurostat (2011)**

	Gini before taxes and transfers		Gini after taxes and transfers		Standard coefficient of variation		Mean log deviation		Gini before taxes and transfers		Gini after taxes and transfers		Standard coefficient of variation		Mean log deviation	
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**		**Model 7**		**Model 8**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Private share (%total)	-0.09	0.20	-0.07***	0.02	-0.03	0.04	-0.13***	0.04	-0.10	0.24	-0.13***	0.04	-0.01***	0.00	-0.22***	0.07
Constant	0.76***	0.08	0.32***	0.01	1.83***	1.40	0.2***	0.02	0.81***	0.01	0.31***	0.00	1.02***	0.20	0.23***	0.03
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	No		No		No		No		No		No		No		No	
Year restriction	No		No		No		No		Yes		Yes		Yes		Yes	
Country restriction	No		No		No		No		Yes		Yes		Yes		Yes	
Observations	85		108		96		96		37		45		42		42	
Adj. R-squared	0.72		0.95		0.45		0.97		0.88		0.93		0.79		0.89	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level. The coefficients and standard errors are multiplied by a factor of 100 in the columns (1), (2), (4), (5), (6) and (8). The full time span covers the years 1985-2005. The restricted period covers the years 1995-2005 to make the time span comparable to our main analyses based on the data from Eurostat (2011a). The full country group includes the 15 European countries and Australia, Canada, Czech Republic, Hungary, Iceland, Japan, Korea, Mexico, New Zealand, Poland, Slovak Republic, Switzerland, Turkey and the United States. The countries are restricted to the 15 European countries from the Eurostat (2011a) data set to make the analyses comparable.

Table 2.15: Panel data regressions of pension expenditures and social outcomes among the elderly (65+) **using a decomposition of private pension expenditures into mandatory and voluntary expenditures**

	Income inequality (s80/s20) among the elderly (65+)						Poverty (PL 60) among the elderly (65+)					
	Model 1		**Model 2**		**Model 3**		**Model 4**		**Model 5**		**Model 6**	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Public pension exp. (%GDP)	-0.43*	0.23					-1.66	2.08				
Mandatory pension exp. (%GDP)	-0.37	0.30					-1.40	2.50				
Voluntary pension expenditures (%GDP)	-0.76***	0.13					-0.66	0.84				
Voluntary share (% total pension exp.)			-3.31**	1.39	-7.55*	1.14			12.61	11.95	-1.83	10.23
Mandatory share (% total pension exp.)					1.57	1.90					1.71	17.89
Population aged 65 and over (%total)	0.17**	0.08	-0.19*	0.11	0.18**	0.09	1.65	1.13	-0.84	0.75	1.45	1.13
Constant	omitted		4.39***	0.11	1.30	1.70	10.02	29.26	30.99***	11.25	-6.74	21.34
Country dummies	Yes		Yes		Yes		Yes		Yes		Yes	
Time dummies	Yes		Yes		Yes		Yes		Yes		Yes	
AR(1) disturbances	Yes		Yes		Yes		Yes		Yes		Yes	
Observations	49		135		49		60		154		60	
Adj. R-squared	0.67		0.84		0.67		0.82		0.78		0.82	
Rho	-0.02		0.41		-0.04		0.44		0.65		0.46	

OLS regressions; unstandardized coefficients; panel-corrected standard errors; Prais-Winsten transformation (AR (1) disturbances). * Significant at the .10 level; ** at the .05 level; *** at the .01 level.

3 Measuring Retirement Savings Adequacy; Developing a multi-pillar approach in the Netherlands

Abstract

The Dutch pension system is highly ranked on adequacy. These rankings, however, are based on fictitious replacement rates for median income earners. This paper investigates whether the Dutch pension adequacy is still high when we take into account the resources that people really accumulate, using a large administrative data set. A comprehensive approach is followed: not only public and private pension rights, but also private savings and housing wealth are taken into account. Summed over all age- and socioeconomic groups we find a median gross replacement rate of 83% and a net replacement rate of 101%. At retirement age, 31% of all households face a gross replacement rate that is lower than 70% of current income. Public and occupational pensions each account for more than 35% of total pension annuities. Private non-housing assets account for 14% and imputed rental income from net housing wealth accounts for about 10%. Some vulnerable groups, such as the self-employed, have

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below average replacement rates. Results are fairly similar to results found in the UK, indicating that we should be careful in evaluating the adequacy of pensions systems on the basis of fictitious replacement rates.

3.1 Adequate retirement savings

In many Western countries, pension systems are affected by demographic aging (OECD 2013c) and reforms are needed to keep the system sustainable and adequate. A good pension system protects people against poverty and smooths people's income over their life-cycle. To achieve these goals countries organize their pension system in very different ways. Considerable effort has been made to compare pension systems across countries and to identify strengths and weaknesses of different systems (Allianz 2011, EC 2012, Mercer 2013, OECD 2013c). In these comparisons the current Dutch pension system compares very favorably with regard to pension adequacy.[1] This is due to a relatively high flat-rate public pension, but also to a high replacement rate for a fictitious person who earns a median income during his whole career and accumulates a pension for 45 years.[2] In practice, however, there are few Dutch people who actually accumulate a pension for the full 45 years. The question arises whether adequacy is still that high when we base adequacy on pension rights that people actually accumulate in the current system. This paper therefore examines the public and private pension rights that households have accumulated. We also investigate the role of private savings and housing

[1] In Mercer (2013) the Netherlands achieves the first place with regard to adequacy. The EC (2012) shows that the Netherlands have one of the highest replacement rates and lowest poverty rates for future retirees compared to other European countries. According the Allianz (2011), the Netherlands ranks fifth, just after Australia, Sweden, Denmark and New Zealand, in terms of the Pension Sustainability Index. This index does not take into account the adequacy of pensions.

[2] OECD (2013c) produces replacement rates for fictitious persons who earn a median income in several OECD countries. The Netherlands achieves the first place with a gross replacement rate of 91.4% and a net replacement rate of 103.8%. These kind of numbers are being used in pension system indicators such as the Melbourne Mercer Global Pension Index (Mercer 2013).

wealth could play during retirement. This gives us necessary integrated results regarding the available resources to finance retirement.[3]

To be able to evaluate adequacy we need to know how much resources retirees need. A variety of standards can be chosen against which to judge adequacy. The Life Cycle Hypothesis (LCH) is the main theoretical framework for assessing the adequacy of savings (Banks et al. 1998). In this model, consumption is not determined by current income, but by (expected) lifetime resources. According to the LCH, it is optimal for individuals or households to save (or borrow) to the extent that, after discounting, the marginal utility of consumption is smoothed over the life cycle. However, the model does not provide straightforward predictions on how much people save in various stages of their lives. Households with identical lifetime incomes might choose different levels of savings for a number of reasons - including uncertainty over future incomes and over future needs, different degrees of risk aversion, variations in time preference rates, the possible existence of liquidity constraints and bequest motives. The model is therefore consistent with a substantial degree of inequality in saving.

Several studies have used the life cycle model to analyze retirement readiness. Bernheim et al. (2001) found a decline in consumption at retirement that is highly correlated with the household income replacement ratio. Households appear to discover that their resources after retirement are insufficient to maintain their standards of living, and adjust their consumption downward accordingly. In other words, people do not save enough to smooth their consumption. Engen et al. (1999) argued that smoothing marginal utility of consumption may have a different impact on saving behavior than smoothing consumption as such. They developed a stochastic life cycle model in which people save both for retirement and for precautionary reasons, including uncertain lifespan. They conclude that savings are adequate for a majority of households. Scholz et al. (2006) developed an extensive stochastic life cycle model that also incorporates government transfers and taxes, as well as medical expenses. They found

[3]Knoef et al. (2013a) simulate household income of the elderly in 2020. Compared to Knoef et al. (2013a), we add analyses about private savings, housing wealth and the composition of pension entitlements. Knoef et al. (2013b) preluded this project.

that the model provides a good representation of households' savings behavior in the US. Fewer than 20% of the households save less than their optimal levels, and the extent of under-saving is generally small.

A common element of these approaches is that people should save enough to maintain their living standards after retirement. This does not imply that consumption after retirement should be equal to consumption before retirement. Consumption requirements are likely to fall when people retire (Scholz et al. 2006). The most commonly used measure of relative well-being after retirement is the income replacement rate. This is the ratio of some post-retirement income (from pensions, annuitized wealth holdings and so forth) to some pre-retirement income (such as earnings during the years preceding retirement, or average earnings during the career). Replacement rates are an important indicator of pension systems. (OECD 2013c), for example, shows replacement rates for fictitious persons in several countries with median earnings throughout their working life. Boskin and Shoven (1987) argued that a replacement rate of less than unity is consistent with the life cycle theory. Haveman et al. (2007) indicated that a widely accepted standard in the literature is having a retirement income equal to or greater than 70% of previous earnings. This is regarded as the income necessary to maintain preretirement consumption. Binswanger and Schunk (2012) investigated minimum acceptable income replacement rates using surveys in the US and the Netherlands, and found that these rates range from 95% to 45% across income quintiles in the US, and from 75% to 60% across income quintiles in the Netherlands. In this study we also use (expected) replacement rates as a key indicator of savings adequacy and retirement readiness. The standard is set at 70%, but we can also show the results for alternative replacement rates. A second approach is to set a social standard for adequacy. In this approach, retirement income is considered adequate when it is equal to or greater than poverty levels of income (Haveman et al. 2007). There are three ways of setting the poverty line: an absolute standard, a relative standard and a subjective standard (Caminada et al. 2012). The US poverty line is based on an absolute standard, which remains fixed over time in real terms. The EU-agreed relative poverty line is set at a fixed percentage of the median income in each country. The at-risk-of-poverty rate is defined as

the share of persons with an equivalized disposable income below 60% of the national median equivalized disposable income. In several OECD studies the poverty line is set at 50% of the median equivalized disposable income. The subjective poverty line is based on respondents' answers to questions regarding what they consider to be an adequate standard of living. Walker (1987) introduced the consensual budget standards method, where members of the public together with some experts reach agreement (consensus) about what people need as a minimum and then draw up budgets to meet those needs. Hoff et al. (2009) applied this method for the Netherlands and found, for example, that in 2008 a single man of age 75 needed about 800 euro per month. De Bresser and Knoef (2014), on the other hand, show that half of the respondents in a representative Dutch household panel expect that they would need between 1.095 en 1.825 euros per month to meet their own minimal expenditure needs (in 2008 euros and equivalized to a single person household).

Another issue is that resource adequacy at the time of retirement does not necessarily mean that incomes are adequate throughout a person's remaining lifetime. Resources may increase during retirement - due to additional asset accumulation, bequests and so forth. But resources may also deteriorate during retirement - due to cuts in pension benefits, for example, or bad investments or increasing uncovered health costs. Haveman et al. (2007) therefore examined the resource adequacy at two points in time: at the time of retirement and ten years later. VanDerhei and Copeland (2010) also measured retirement readiness at several points in time. They argued that replacement rate measures are useful, but that it is difficult to accurately integrate the concepts of longevity risks, post retirement investment risks and uninsured healthcare risks. They follow an approach in which a household is considered to run short of money if its resources are not sufficient to meet minimum retirement expenses plus uncovered expenses from the nursing home and healthcare. Expenses are derived from the Consumer Expenditure Survey, based on actual observed expenditure of the elderly for different family sizes and income levels. This approach has its advantages, but the problem is that observed consumption patterns of retirees are constrained by their resources. If their resources fall short, their observed expenses may not reflect their real

needs in retirement. Consequently, the method used by VanDerhei and Copeland cannot provide the (only) benchmark against which to judge the adequacy of resources.

This paper uses a large administrative data set to scrutinize the resources that households of different generations have accumulated to finance retirement. This is in stark contrast with the approach taken in OECD (2013c) where a fictitious person is analyzed. To develop a more comprehensive view on pension adequacy not only public and private pensions, but also private savings and housing wealth are taken in to account. Private savings and housing wealth are annuitized, taking into account household age, age differences between household members, and economies of scale. Furthermore, because of the large administrative data set, we can draw credible conclusions for specific vulnerable groups. To investigate the bandwidth of the results, we investigate several scenarios as to what will happen from the time of observation until retirement. This multi-pillar approach is highly applicable to other countries, although the implementation of the approach may be limited by the availability of country-specific data.

When we only consider public and occupational pension income we find a median gross and net replacement rate of 71% and 84%, respectively. Private savings and housing wealth can play a substantial role to increase adequacy, but even when these are taken into account about 31% of the households do not reach a gross replacement rate of 70%. The results are fairly comparable to the results of Crawford and O'Dea (2012), who perform a comparable type of analysis for the UK. So, although the pension system of the UK achieves a much lower international rank on adequacy than the Dutch pensions system,[4] the results are fairly comparable when we consider adequacy on the basis of real pension savings. This shows that that we should be cautious in drawing conclusions about the performance of pension systems across countries on the basis of indices that use fictitious replacement rates.

[4]The UK achieves the 8th place in the Melbourne Mercer Global Pension Index and has a gross replacement rate of 37.9% and a net replacement rate of 48.0% for a median earner. The Netherlands achieves the 1st place with a gross replacement rate of 91.4% and a net replacement rate of 103.8% for a median earner.

The structure of the paper is as follows. Section 3.2 describes the Dutch pension system and section 3.3 introduces the data. Section 3.4 shows descriptive statistics of income and wealth in 2008 and section 3.5 describes our method. Section 3.6 predicts financial resources during retirement and replacement rates. Section 3.7 focuses on several vulnerable groups, such as self-employed households, immigrants and households on social assistance. Section 3.8 analyzes the sensitivity of the results with regard to assumptions about indexation, real rates of return, housing prices and the depletion of housing wealth. Finally, section 3.9 concludes.

3.2 The Dutch Pension System

As in many European countries, the Dutch pension system consists of three pillars. The first is a pay-as-you-go system and involves a flat-rate public pension benefit for all residents as from the statutory retirement age of 65 onwards. The level of the public pension is linked to the net minimum wage and depends on the number of years that a person has resided in the Netherlands. Each pensioner living in a couple household who has lived in the Netherlands between the ages of 15 and 65 receives 50% of the minimum wage, and single pensioners receive 70% of the minimum wage. For people with a low pension income and almost no wealth, the first pillar is topped up with social assistance to guarantee a social minimum.

Several OECD countries have recently increased their statutory pension age, or will do so in the coming decades (OECD 2013c). In the Netherlands, the statutory retirement age increased by one month as of January 2013, and will gradually increase to 66 in 2019 and 67 in 2023. It has been proposed to increase the statutory retirement age more rapidly: to 66 in 2018 and 67 in 2021.

The Dutch second pillar consists of capital-funded occupational pensions, of which the primary responsibility lies with employers and employees. Mandatory occupational pension accumulation is agreed upon in collective labor market agreements in the Netherlands. As a consequence, 90% of all employees have a pension scheme with their employer.

Occupational pensions mainly consist of defined-benefit pension plans. Until the beginning of the 21st century, most pension plans aimed to pay a pension income of 70% of final gross wage from the age of 65 onwards if an employee had worked fulltime for at least 40 years. From 2003 onwards, pension funds have lowered their ambition, and they now aim to pay 70% of the average career salary, instead of 70% of the final gross salary (including public pension benefits). The recent financial crisis has shown that the Dutch pension system is vulnerable to shocks in financial markets. Many pension funds have had difficulties achieving their indexation ambitions, and several funds recently were even compelled to cut nominal pension rights. Also, annual tax-favored pension accruals have been reduced from 2.25% to 2.15% and will be reduced further to 1.875%. This means that the percentage by which pensions are built up each year is reduced and that one has to work more years to achieve the same pension income. Furthermore, the age that forms the basis for the determination of the pension premiums increased from 65 to 67 as of 2014. Early retirement will consequently become financially less attractive, and the pension income of future retirees is likely to become less generous.

The third pillar is formed by private individual pension products (such as life annuities) and other private savings. Until a major tax reform in 2001, everyone could buy life annuities at tax beneficial terms up to a certain limit (e.g. premiums up to 2,808 euro were fiscally attractive in the year 2000). After the tax reform, this limit was reduced in 2002 to 1,069 euro, and only the self-employed and individuals with a gap in their pension entitlements were allowed to buy life annuities at fiscally attractive terms up to higher amounts. Other pillars are housing wealth or an extension of working life on a part-time or fulltime basis. People who have amortized part of their mortgage benefit from lower housing costs during retirement. Although not commonly done by the current generation of elderly, people may move or use reverse mortgages to deplete housing wealth.

Data 3.3

To estimate the extent of financial resources available to the current labor force upon entering retirement, we combine administrative data with assumptions as to what will happen from the time of observation until the day of retirement. This section describes the data that are used. We combine as many wealth components as possible in evaluating the retirement readiness of the Dutch population: public pension rights (PAYG), occupational pension rights, individual annuity insurances, housing wealth and private savings. The most recent data about occupational pension rights come from 2008. Therefore, a representative sample of households in 2008 forms the basis of all of our data.

To assess the pension rights accumulated in public old-age pensions, we take administrative data from the 2008 'Dutch statistics on public pension entitlements' (in Dutch: Algemene Ouderdomswet aanspraken totaal, AOWA). These data contain information about the public pension entitlements that have been built up by people between the ages of 15 and 64.

Concerning occupational pensions, we use of the 2008 'Dutch statistics on occupational pension entitlements' (PA). These data provide information about the occupational pension entitlements that have been built up by people between the ages of 15 and 64. This information is gathered by Statistics Netherlands from occupational pension funds in the Netherlands. Pension funds deliver data to Statistics Netherlands about the annuity that participants would receive in case they remain employed in their current job with their current wage rate until the statutory retirement age of 65. Not all pension funds have provided data to Statistics Netherlands, but the aggregate amount of pension entitlement in the Netherlands is available from the Dutch Central Bank (DNB), and Statistics Netherlands used this information, together with employment data, to correct the individual pension entitlements (Eenkhoorn and Zijlmans 2010). After a divorce, occupational pension benefits are often partly paid out to the ex-partner.[5]

[5]Either an ex-partner receives part of the occupational pension benefits when the ex-husband or ex-wife becomes 65, or entitlements are converted directly after the divorce into two separate entitlements for both members of the divorced couple. Then, for example, the benefits can start at different moments in time. Conversions are included

To assess information about income and other wealth assets, administrative data were taken from the 2008 Dutch Income Panel data (IPO), with wealth information from the tax office, banks and social security administrations. Banks have to deliver data about savings accounts that exceed 500 euro or yield interest of more than 15 euro a year. Checking accounts are not included. Furthermore, the data contain information on stocks, bonds and wealth from an own business. With regard to housing, the data include information about the value of the house and the mortgage, the value of secondary houses and some moveable properties such as houseboats.

Whereas the AOWA and PA data set contain information about the entire Dutch population, IPO contains a representative sample of Dutch households that are followed over time. We therefore merge AOWA and PA to the IPO sample. Major advantages of these administrative data are a very low attrition rate and a high level of representativeness. Attrition takes place only because of immigration or death. Another advantage of administrative data is that the observed variables are measured with a high degree of accuracy. In this progress report we merged only the 2008 data, since this is the most recent year for which AOWA and PA are available.

The data have some shortcomings. They do not, for example, provide information about assets accumulated in personal defined-contribution pension plans (third pillar). Data is available, however, regarding contributions made to third pillar pension plans as from 1989, which provides information about the wealth accumulated in third pillar pension plans (Caminada 2000). Furthermore, young generations in the Netherlands often seek to avoid taxes through an endowment mortgage or an investment-based mortgage. This means that the mortgage is not paid off during the term of the mortgage. Instead, money is paid to an insurance company or a bank, such that (part of) the mortgage can be paid off at the end of the term. The money accumulated at the insurance company or at the bank is not observed by the tax office, and is not available in the data.[6] Also, we

in the data but there is no information regarding pensions that are partly paid out to ex-partners when the participant becomes 65.

[6] In 2008, about 30% of the mortgages were endowment- or investment-based mortgages (Dijkhuizen 2013).

do not know which households own an endowment or investment-based mortgage.

Descriptive analysis 3.4

This section describes current income (3.4.1) and wealth (3.4.2) observed for several age groups.

Income 3.4.1

Table 3.1 describes gross equivalized household income and the proportion of households receiving income from the various income sources in 2008. Income is measured in 2010 euros using the consumer price index. In order to standardize household income to a single-person household, we use the equivalence scale provided by Statistics Netherlands (Siermann et al. 2004), which assumes that two adults need 37% more income than a single adult to achieve the same welfare level.[7] The households' key person, who is randomly drawn from the Dutch population and who is followed over time in the IPO data set, determines the age category of the household.

As expected, labor income is the most important income component and is highest for people between the ages of 50 and 54. Average disability and unemployment benefits increase until the statutory retirement age of 65. This growth is a combination of age, period and cohort effects, which cannot be distinguished. Furthermore, older individuals have longer unemployment durations on average, which lead to higher unemployment benefits per year. Public pension benefits are received as from the age of 65, so before the age of 65 we only observe public pension benefits of household members that are 65 or older (e.g. partners or parents in the same household). In the age group 60-64, early retirement income becomes important and seems to replace labor income at least partly. Non-labor income includes interest received from bank accounts, dividends from

[7] Kalmijn and Alessie (2008) found that the modified OECD scale and the equivalence scale of Statistics Netherlands yield very similar results.

Table 3.1: Household income, 2008[a]

Age group	35-49	50-54	55-59	60-64	65-69	70+	All
Average income							
Labor income	32,332	35,776	31,767	15,992	4,388	1,507	22,908
Disability insurance	696	1,534	2,121	2,769	611	92	1,106
Unemployment insurance	294	382	667	843	204	26	353
Public pension (AOW)	206	220	365	1,545	10,853	12,545	3,495
Private pension	451	1,264	3,623	13,342	13,276	9,263	5,027
Non-labor income	-2,242	-642	176	1,344	2,092	2,657	-129
Profit from business	3,974	4,091	3,816	3,392	1,435	306	3,028
Social assistance	538	582	595	606	192	149	458
Child- and study allowances	732	487	185	49	27	13	376
Other transfers[b]	268	287	239	277	280	461	304
Gross income	37,249	43,980	43,854	40,160	33,357	27,019	36,926
Disposable income	24,968	28,892	28,623	26,965	25,194	21,788	25,502
Median income							
Labor income	29,808	33,603	28,661	4,286	0	0	18,019
Disability insurance	0	0	0	0	0	0	0
Unemployment insurance	0	0	0	0	0	0	0
Public pension (AOW)	0	0	0	0	13,001	13,033	0
Private pension	0	0	0	6,377	8,699	5,333	0
Non-labor income	-1,823	-578	-25	11	289	469	0
Profit from business	0	0	0	0	0	0	0
Social assistance	0	0	0	0	0	0	0
Child- and study allowances	718	0	0	0	0	0	0
Other transfers[b]	0	0	0	0	0	0	0
Gross income	32,208	38,332	37,472	32,581	26,361	21,288	30,769
Disposable income	22,542	26,169	25,484	22,918	21,360	18,465	22,349
Proportion of households receiving various income components							
Labor income	0.91	0.90	0.84	0.60	0.30	0.11	0.67
Disability insurance	0.10	0.17	0.21	0.23	0.09	0.01	0.12
Unemployment insurance	0.07	0.08	0.09	0.08	0.03	0.01	0.06
Public pension (AOW)	0.03	0.03	0.05	0.18	1.00	1.00	0.30
Private pension	0.09	0.16	0.29	0.67	0.93	0.87	0.40
Non-labor income	0.92	0.92	0.92	0.92	0.91	0.88	0.91
Profit from business	0.17	0.16	0.14	0.12	0.07	0.02	0.13
Social assistance	0.07	0.07	0.07	0.06	0.04	0.03	0.06
Child- and study allowances	0.67	0.41	0.16	0.04	0.02	0.01	0.34
Other transfers[b]	0.34	0.28	0.14	0.11	0.15	0.24	0.25
Gross income	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Disposable income	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Observations	22,245	6,645	6,277	6,479	4,620	10,299	56,565

[a] Equivalized household income in 2010 euros. The age of the key person in the households determines the age category of the household.
[b] Rental house allowance, home owner grant, alimony and study costs allowance.

stocks, income from bonds, imputed rent, mortgage interest, and income from other property such as second houses. By using imputed rent (as defined in IPO) and mortgage interest, we take into account that those who paid off their mortgage take advantage of low housing expenses. Mortgage interest explains the negative values for non-labor income in the young age groups. Among the 70+ population, total gross income is relatively low, which may be related to cohort effects. Finally, since the income distribution is positively skewed, mean income is higher than median income, which means that higher deciles earn a proportionally larger share of total income.

Wealth 3.4.2

Table 3.2 presents average household wealth, median household wealth and the proportion of households owning various wealth components in 2008. Wealth is measured in 2010 euros using the consumer price index and is not equivalized to a one-person household.

The results indicate that wealth in savings accounts increases with age, at least until the age of 70. Debts other than mortgage are owned by somewhat more than 10% of the sample and are highest in the 55-59 age category. Stocks from a substantial holding are relatively high, but only owned by not more than 1% of the sample. Securities, however, are owned by more than 25% of the sample, and increase on average from about 11,638 euro in the age category 35-49 to 25,641 euro in the 70+ age category.

Property is owned by 65% of the sample. Most of them (78%) also have a mortgage. The proportion of homeowners with a mortgage is high in the 35-49 age category (67/72=93%), but also in the 70+ category 41% of the homeowners still have a mortgage. Net housing wealth (property value minus the mortgage) is substantial and varies over age categories: it is lowest in the 35-49 age category and highest in the 60-64 age category, with an average of 206,066 and a median of 142,433 euro. The relatively high levels of net housing wealth among older generations can be explained by amortization of mortgages but also by home price increases before 2008. Between the beginning of the 1990s and 2008 home prices increased

Table 3.2: Household wealth, 2008[ab]

Age group	35-49	50-54	55-59	60-64	65-69	70+	All
Average wealth							
Savings account	33,836	41,911	51,498	60,199	60,441	57,215	46,194
Debt other than mortgage	16,036	17,729	21,777	18,992	16,905	8,065	15,830
Stocks substantial shareholders	22,764	22,523	33,633	41,561	22,072	11,526	23,992
Securities	11,638	19,802	22,294	23,933	24,427	25,641	18,782
Mortgage	144,295	103,716	85,120	64,963	42,896	14,676	91,992
Property	248,294	270,271	272,856	271,029	248,363	160,313	240,192
Business assets	5,468	6,849	6,017	5,670	5,163	989	4,874
Net housing wealth	103,999	166,555	187,736	206,066	205,467	145,637	148,200
Mortgage to property ratio	0.63	0.43	0.37	0.27	0.20	0.10	0.43
Total wealth	161,669	239,910	279,401	318,437	300,664	232,943	226,211
Median wealth							
Savings account	9,378	11,452	17,922	21,232	24,115	24,175	14,987
Debt other than mortgage	0	0	0	0	0	0	0
Stocks substantial shareholders	0	0	0	0	0	0	0
Securities	0	0	0	0	0	0	0
Mortgage	117,420	64,049	39,760	8,100	0	0	23,251
Property	227,579	234,051	229,736	222,186	201,693	0	211,401
Business assets	0	0	0	0	0	0	0
Net housing wealth	40,604	108,895	125,952	142,433	136,797	0	66,220
Mortgage to property ratio	0.63	0.37	0.28	0.21	0.13	0.00	0.36
Total wealth	70,826	136,023	155,577	181,251	173,095	68,523	105,828
Proportion of households owning various wealth components							
Savings account	0.87	0.89	0.91	0.92	0.94	0.94	0.90
Debt other than mortgage	0.11	0.13	0.14	0.13	0.11	0.06	0.11
Stocks substantial shareholders	0.02	0.02	0.02	0.02	0.01	0.01	0.02
Securities	0.27	0.28	0.28	0.28	0.26	0.20	0.26
Mortgage	0.67	0.63	0.58	0.51	0.41	0.18	0.53
Property	0.72	0.72	0.70	0.68	0.62	0.44	0.65
Business assets	0.15	0.13	0.11	0.09	0.05	0.01	0.10
Total wealth	0.92	0.92	0.94	0.94	0.96	0.95	0.93
Observations	22,245	6,645	6,277	6,479	4,620	10,299	56,565

[a] Household wealth in 2010 euros. The age of the key person in the households determines the age category of the household.

[b] 7% of the households do not have any wealth according to the IPO data. These households may only own checking accounts (with unlimited amounts of money) and/or savings accounts that do not exceed 500 euro (or yield interest of more than 15 euro).

substantially, with an increase of about 180% between 1995 and 2008. This was at least partly due to decreasing mortgage interest rates and reduced borrowing constraints (before the 1990s, income from second earners was only taken into account for five years, and this became 30 years). The share of homeowners decreases after the age of 65; therefore, also average net property value decreases after the age of 70. A possible explanation for this is that people's health or the death of a partner forces them to move to a nursing home or a smaller house. In addition, cohort effects may play a role (homeownership is relatively low in old cohorts). Due to fiscally attractive mortgage constructions, described in section 3, we underestimate housing wealth. Housing wealth is rather illiquid, however, and is therefore often excluded in empirical studies on savings adequacy (Venti and Wise 1991). People in the Netherlands strongly prefer to stay in their own home as long as possible (De Graaf and Rouwendal 2012). Reverse mortgages could be used to access a portion of home equity, but are still rare in the Netherlands. Nevertheless, housing wealth is very important in saving for retirement. Persons owning a house, given that they have repaid part of the loan on the house, need less income to finance their necessary expenses than persons who live in a rental house.

Method and assumptions 3.5

This section explains the method and assumptions that we use to predict financial resources during retirement for future generations of retirees. Households may deplete wealth to finance their retirement. In view of this we first describe how we annuitize household wealth. Secondly, we describe the assumptions that we make for the pension components.

Annuitizing household wealth 3.5.1

Whereas pension rights and annuity insurances are observed at the individual level, private savings and housing wealth are observed at the household level. We do not know how the members of a couple divide their wealth over each other. Therefore, to determine pension savings

adequacy we assume that couples smooth their wealth over time and over each other. In the annuitization process we take into account that members of a couple are often of different age and do not have the same life expectancy. Furthermore, we take into account economies of scale to reckon that when one of the partners dies, the remaining widow(er) needs to deplete relatively more wealth to be equally well off as before, since he or she loses economies of scale.

To investigate pension savings adequacy we project financial resources as from the age of 65.[8] To take both economies of scale and the age difference between members of a couple into account, we distinguish between the period where only the oldest member of the couple is 65 years or older and the period where both members are 65 or older. When the man is older than the woman we compute the annuity as follows:

$$A = K/\left[\sum_{n=max(65-a_m,1)}^{64-a_f}\left((1-{}_np_{af})\,{}_nq_{am}+0.5\cdot E\cdot {}_np_{af}\,{}_nq_{am}\right)\frac{1}{(1+r)^n}+\right.$$
$$\left.\sum_{n=65-a_f}^{T-a_f}\left({}_np_{af}(1-{}_nq_{am})+(1-{}_np_{af})\,{}_nq_{am}+E\cdot {}_np_{af}\,{}_nq_{am}\right)\frac{1}{(1+r)^n}\right] \tag{3.1}$$

where K is the amount of capital needed for annuity A as from the age of 65. a_m is the age of the man, a_f is the age of the woman, ${}_np_{af}$ is the probability that a woman of age a is still alive after n years and ${}_nq_{am}$ is the probability that a man of age a is still alive after n years. T is the maximum life expectancy and E reflects the equivalence scale (how much extra income a two-person household needs to be as well-off as a one-person household). We standardize the annuity to a one-person household. The first term of equation (3.1) reflects the period in which the man already reached the age of 65 and the woman is younger than 65. In case the woman is no longer alive, the man needs an annuity A; in case the woman is still alive, we assume the man needs $0.5 \times E$ of an

[8] The baseline scenario analyzes all pension components as from the age of 65. Section 3.8 shows the results when all components are computed as from the age of 64 and 67. We do not differentiate the retirement age between cohorts, although young cohorts may be better equipped to work longer than older cohorts are.

annuity, because of the economies of scale. The second term of equation (3.2) reflects the period in which both the man and the woman are of age 65 or older. In case only the man or only the woman is alive, the household needs annuity A. In case both are alive they need $E \times A$. The other way around, when the woman is older than the man, we use

$$A = K/\left[\sum_{n=max(65-a_f,1)}^{64-a_m}\left((1-{}_nq_{am})\,{}_np_{af}+0.5\cdot E\cdot {}_np_{af}\,{}_nq_{am}\right)\frac{1}{(1+r)^n}+\right.$$
$$\left.\sum_{n=65-a_m}^{T-a_m}\left({}_nq_{am}(1-{}_np_{af})+(1-{}_nq_{am})\,{}_np_{af}+E\cdot {}_np_{af}\,{}_nq_{am}\right)\frac{1}{(1+r)^n}\right] \tag{3.2}$$

When both men and women have the same age we only keep the second term of equation (3.1) or (3.2), because there is no period in time where one of the members is 65 or older and the other member has not yet reached the age of 65 in this situation.

Assumptions 3.5.2

This section describes the assumptions for each pension component. With regard to the first pillar we assume that people stay in the Netherlands as from 2008 until the age of 65. To compute the public pension benefit that households receive we use the full gross public pension benefit level of 2008, measured in 2010 euros (13,033 euro per year for singles and 17,993 euro per year for couples). We include social assistance benefits that are used to guarantee a social minimum (e.g. for immigrants), and we assume that public pension benefits will be indexed. Finally, the public pension eligibility age is higher for future generations of retirees. To be able to compare public pensions across generations, we compute public pension benefits for everyone as if they are received as from the age of 65, using an actuarially neutral reduction rate for young generations that have a public pension eligibility age higher than 65.[9]

[9]The Dutch public pension system has no flexible public pension retirement age. However, since in this paper we analyze all pension components as from the age of 65,

Regarding occupational pensions, we use the data of Statistics Netherlands about occupational pension rights, which assume that people remain employed in their current job with their current wage rate until the age of 65. In future research we will test how robust the results are with respect to this assumption by estimating and simulating wage profiles and labor market transitions, taking into account part-time work and stochastic non-employment spells. For the moment, we have to bear in mind that we do not take into account wage growth for younger workers and that we do not take into account unemployment and early retirement for this group (not all people will be working until the age of 65). In general, occupational pension entitlements in the Netherlands are nominal rights with price indexation conditional on the financial situation of the pension fund. Because of the poor financial situation of most pension funds in the Netherlands in recent years, pension funds have been unable to make inflation corrections. For the future we assume that 50% of the inflation will be corrected and that inflation amounts to 2% per year.[10] Furthermore, we make the rather optimistic assumption that no pension cuts take place. For 65+ individuals we do not observe second pillar pension entitlements, but we do observe the amount of second and third pillar pension benefits that they receive.

To approximate wealth accumulated in third pillar pension plans, we use the yearly contributions made to third pillar pension products as from 1989 and add a fictitious real return of 1% (after tax) per year. For the future we assume that, until the age of 65, people deposit the same amount into the pension product every year as they did on average during 2006-2010 (in real terms). We assume a future real rate of return of 1% (after tax) per year.[11]

we also compute public pensions as from the age of 65, as if households can borrow against their future public pension income. In this way, all results are based on the same retirement age.

[10] We assume an indexation of 50% in the baseline scenario. Section 3.8 shows calculations with no indexation (pessimistic scenario) and full indexation (optimistic scenario).

[11] We assume a yearly real rate of return of 1% in the baseline scenario. Section 3.8 shows the calculations with 0% (pessimistic scenario) and 2% (optimistic scenario).

For the annuitization of private savings we use an annual real rate of return of 1% after tax,[12] and the most recent mortality rates per cohort predicted by Statistics Netherlands (December 17th 2010).[13] Mortality differences between men and women and between cohorts are taken into account. We do not consider differential mortality by income (Kalwij et al. 2013) and we assume that the remaining lifetimes of couples are independent. As for the future, we assume that no additional private savings are being made to finance retirement.

With regard to housing, we assume an average yearly drop in real property prices of 1%. This means that an individual of age 40 in 2008 experiences a drop in the real value of his house of 22% between now and the age of 65. The average drop in housing prices was already 20% (in real terms) between the year 2008 and 2013. So, for this person real house prices should stay more or less constant after 2013 for this assumption to be true. Homeowners who have amortized part of their mortgage have relatively low housing costs. We take this into account by a small percentage (4%) of the net capital accrued in property (imputed rent). With an inflation of 2% we have an imputed rent in real terms of 2% (4%-2%). Until the age of 65, imputed rental income increases net housing wealth (e.g. by amortizing the mortgage). It can be seen as a return on housing wealth.

We assume that no additional private savings and mortgage amortizations will be made between 2008 and the year in which people reach the age of 65 to finance retirement. Thus, for the present we look only at current savings to determine pension savings sufficiency, and we compare current savings with current income. Also, we assume that retirement is the only savings motive for households, although other motives may exist, such as bequests (Van Gilst et al. 2008). We also assume that children have left the household at the time the key person of the household reaches the age of 65. Furthermore, we allow for widowhood, but assume that couples stay together and singles remain single. To standardize household

[12] We vary this annual real rate of return over different scenarios in section 3.8. An annual real interest rate of 0% is used in the pessimistic scenario and 2% in the optimistic scenario.

[13] We assume that remaining lifetimes of couples are independent.

income we use the equivalence scale provided by Statistics Netherlands (Siermann et al. 2004), described above.

3.6 Results

Section 3.6.1 shows the results for future retirement income. Next, we compare these with current gross and net income (3.6.2 and 3.6.3), and investigate poverty during retirement as an indicator of how well households are prepared for their retirement (section 3.6.4).

3.6.1 Future retirement income

Table 3.3 shows equivalized pension annuities. We see that - for most households - public pension benefits and occupational pensions are the most important sources of income after retirement. These components together provide 65% of the average total annuitized wealth. Despite the relatively high percentage of households that have voluntary pension products, the holdings in these accounts are small and therefore contribute to the total pension wealth only marginally. Second pillar pension entitlements are highest in the two youngest age categories. There are several reasons for this. First, the pension coverage for young cohorts is higher than for old cohorts (especially among women). Second, we assume that people keep their current job until the age of 65, while older people have had more time to run into a gap, caused by a period of part-time employment or unemployment. Also, they may already have retired early, which decreases the occupational pension they receive as from the age of 65. If we would take into account the possibility that young cohorts will also run into unemployment, disability and/or early retirement, then their occupational pension would also be lower. Finally, a reduction of tax favored pension accruals will especially influence younger cohorts.

The mean and median occupational pension benefits show that the distribution of occupational pension entitlements is skewed to the right (private pensions are distributed unequally such that high deciles receive a proportionally larger share of total private pensions). The distribution of

public pension entitlements, however, is evenly distributed (most people receive a full state pension that consists of a flat rate).

Annuitized wealth from net savings accounts and securities is relatively high among the 70+ population because their remaining life expectancy is relatively low. Note that in this paper we assess whether current savings are adequate. We make no predictions about the extent of resources available to individuals at age 65, but estimate how much they would have in light of their current resources. We have to keep in mind that younger generations have more time to supplement their private savings. Also, private savings are probably higher especially for those persons who have an occupational pension gap.

Figure 3.1 shows the average annuitized wealth components over the income distribution, where households are sorted from low- to high gross income. As expected, public pensions are flat over the entire income distribution. All other wealth components increase with gross income, with a large peak at the higher end of the income distribution. The importance of net savings accounts and securities increases at the higher end of the income distribution.

Gross replacement rates 3.6.2

As a first measure of pension savings adequacy, we divide predicted retirement income by gross current income. This gives a replacement rate for households, using their current income, their current wealth, and the assumption that people keep their current job with their current wage and do not build up more capital (other than first and second pillar entitlements). Basically, we indicate to what extent current savings can replace current income conditional on the current job and wage.

Table 3.4 shows three different replacement rates. The first replacement rate only takes into account public and occupational pension benefits. The second includes voluntary pension products and other financial wealth, and the third also includes the imputed rental income of net housing wealth. These three replacement rates indicate the importance of different wealth components and provide insight into the replacement rates when

Table 3.3: Predicted yearly retirement income (annuitized wealth)[a]

Age group	35-49	50-54	55-59	60-64	65-69	70+	All
Wealth: average annuity							
Public pension[b]	11,141	11,233	12,107	12,817	12,533	12,955	11,895
Occupational pension[c]	14,431	13,474	12,107	8,806	669	73	9,678
Voluntary pension products[d]	779	915	917	752	47	5	606
Private pension benefits 65+ [e]	22	150	375	2,159	12,685	9,254	3,036
Net savings account	873	1,120	1,235	1,692	1,973	5,578	1,982
Stocks SH	1,048	1,021	1,373	1,649	952	820	1,101
Securities	547	955	983	962	1,086	2,905	1,164
Business assets	255	299	250	223	211	72	219
Imputed rent	2,584	3,789	4,061	4,252	4,037	2,601	3,202
Total pension annuity	31,680	32,955	33,408	33,313	34,193	34,263	32,884
Wealth: median annuity							
Public pension[b]	11,426	11,426	12,384	13,033	13,075	13,033	11,426
Occupational pension[c]	12,485	11,026	9,306	5,377	0	0	6,333
Voluntary pension products[d]	0	0	0	0	0	0	0
Private pension benefits 65+ [e]	0	0	0	0	8,293	5,352	0
Net savings account	379	421	626	719	973	2,020	656
Stocks SH	0	0	0	0	0	0	0
Securities	0	0	0	0	0	0	0
Business assets	0	0	0	0	0	0	0
Imputed rent	1,035	2,472	2,735	2,947	2,679	0	1,558
Total pension annuity	27,926	28,511	28,109	26,990	27,442	24,416	27,275
Proportion of households with entitlements from various pension arrangements							
Public pension[b]	1.00	1.00	1.00	1.00	1.00	1.00	1.00
Occupational pension[c]	0.97	0.95	0.94	0.87	0.16	0.02	0.71
Voluntary pension products[d]	0.44	0.49	0.49	0.41	0.04	0.00	0.33
Private pension benefits 65+ [e]	0.00	0.01	0.03	0.02	0.92	0.86	0.26

[a] Equivalized household income in 2010 euros. The age of the key person in the households determines the age category of the household.
[b] We assume that persons reside in the Netherlands at least until the age of 65.
[c] For persons younger than 65 we observe the annuity that participants would receive in case they remain employed in their current job with their current income until the age of 65 (no career/income developments).
[d] Pension rights accumulated in the third pillar are approximated using yearly contributions as from 1989.
[e] For persons of age 65 and older we do not observe occupational pension rights and the amount of wealth accumulated in voluntary pension products, but we do observe the sum of actual private pension income.

Figure 3.1: Composition of pension annuities over the income distribution

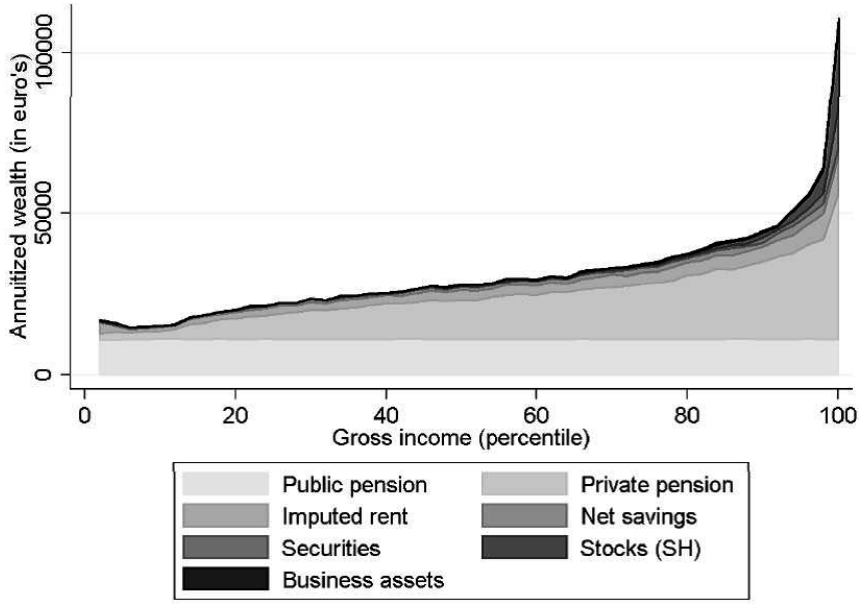


Table 3.4: Gross replacement rates, 2008[ab]

Age group	35-49	50-54	55-59	60-64	All
1st and 2nd pillars					
ratio p25	0.61	0.50	0.49	0.50	0.55
ratio p50	0.76	0.63	0.64	0.68	0.71
ratio p75	0.91	0.77	0.78	0.87	0.87
1st and 2nd pillars and private wealth					
ratio p25	0.66	0.55	0.55	0.59	0.61
ratio p50	0.82	0.69	0.70	0.77	0.77
ratio p75	0.98	0.84	0.85	0.96	0.94
Total pension annuity					
ratio p25	0.70	0.60	0.62	0.66	0.66
ratio p50	0.87	0.76	0.77	0.84	0.83
ratio p75	1.06	0.93	0.95	1.08	1.03

[a] The three replacement rates in this table give an impression of the importance of 1st and 2nd pillar pensions, privately saved wealth and property to finance retirement. However, it should be noted that the ratios cannot be compared mutually, because of the rearranging of the quartiles with respect to the wealth components that are taken into account in calculating the replacement rates.

[b] The table reports three quartiles (p25, p50 and p75) of the distribution of the replacement rates (replacing current gross income). At the bottom, 25% of the households have a replacement rate below the first quartile (p25). p50 indicates the median replacement rate. At the top, 25% of the households have a replacement rate higher than p75.

households do or do not deplete financial wealth.[14] Table 3.4 shows that the total median gross replacement rate is 83% (p50). Half of the sample has a total gross replacement rate between 66% and 103% (p25 and p75, respectively). The ratio is relatively high for the youngest age category as well as for the category 60-64. This can partially be explained by the fact that current income is relatively low among these households, as observed in table 3.1. If we only take into account wealth in the first and second pension pillars, replacement rates become substantially lower in all age categories. This substantial contribution of non-pension wealth, including housing, to retirement income is also found by Crawford and O'Dea (2012), who performed a comparable analysis for the UK.

Table 3.5 shows that when account is taken only of public and occupational pensions, a considerable share of the households (49%) has a

[14] Net housing wealth can also be depleted by moving to a smaller or rental house or by a reverse mortgage. Among current retirees this is not very common but it may become more common in the future. Section 3.8.2 describes the scenario in which households deplete housing wealth.

Table 3.5: Share of households below 70% and 100% gross replacement rates, 2008

Age group	35-49	50-54	55-59	60-64	All
Share below 70% gross replacement rate					
1st and 2nd pillars	0.40	0.63	0.61	0.52	0.49
Idem, including private wealth	0.31	0.52	0.50	0.41	0.39
Total pension annuity	0.26	0.41	0.38	0.30	0.31
Share below 100% gross replacement rate					
1st and 2nd pillars	0.85	0.94	0.93	0.85	0.87
Idem, including private wealth	0.77	0.89	0.88	0.78	0.81
Total pension annuity	0.68	0.81	0.80	0.68	0.72

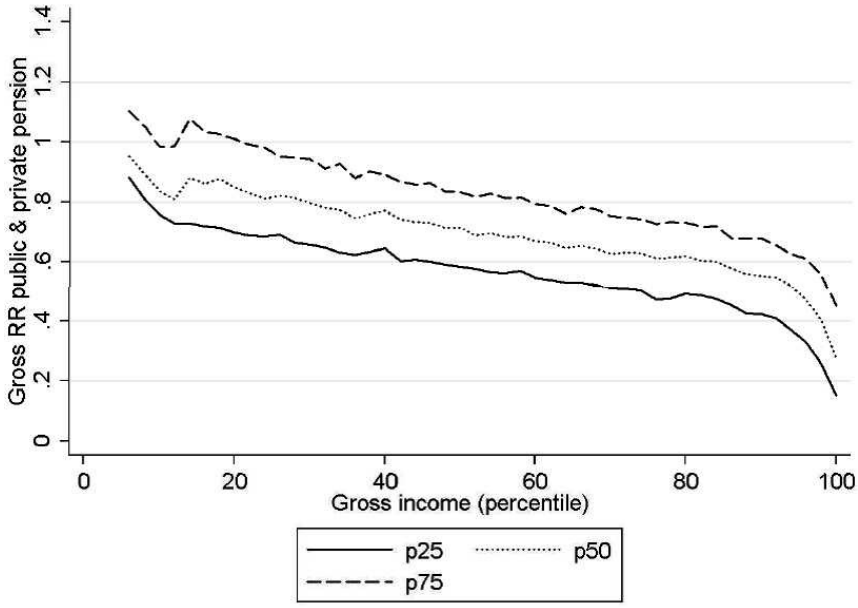
gross replacement rate below 70%. When account is taken of the third pillar, private wealth and imputed rental income from net housing, this percentage decreases to 31%. On the other hand, 28% (=100%-72%) of the households can replace at least their current income using the total of their pension annuities.

Figure 3.2 presents the development and variation of the gross replacement rate over the income distribution. Figure 3.2a focuses on the ratio of public and private pensions to gross current income. Here, it should be noted that high replacement rates of about 100% for low gross incomes are institutionally determined with the 'social minimum'.

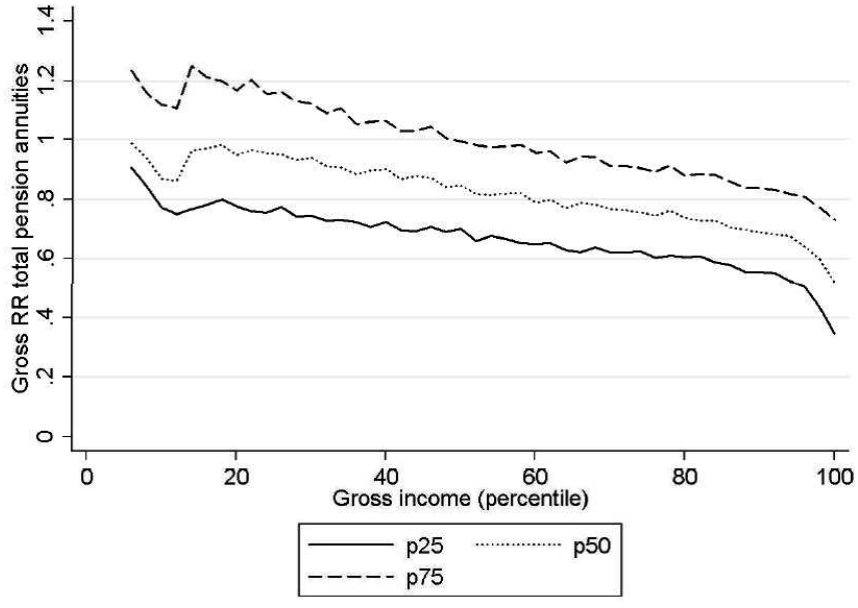
The replacement rate declines over the income distribution from a median replacement rate of 95% at the lower end of the income distribution to a median replacement rate of 34% at the top of the income distribution. Figure 3.2b shows the ratio of the total pension annuity compared to current gross income. Comparison of figures 3.2a and 3.2b reveals that wealth from voluntary pension products, private savings and property has a substantial positive effect on replacement rates; as from the 25th income percentile, replacement rates are approximately 15%-points higher when taking into account voluntary pension products, private savings and property. The replacement rate even increases by about 25%-points for the top quartile of the replacement rate (the dashed line) because of including private savings and housing wealth. The increase in the replacement rate is less substantial (about 8%) for the bottom quartile (the solid line). Finally, the decline in replacement rates over the income percentiles is lower when

Figure 3.2: Replacement rates over the income distribution

(a) First and second pillar



(b) Total pension annuity



we take into account private wealth and housing. As might be expected, the replacement rates that include the total pension annuity show a larger variation than the replacement rates that only take into account first- and second pension pillars.

Net replacement rates 3.6.3

The analysis thus far has focused on gross income and gross replacement rates, and has not considered the Dutch labor income tax and benefits system. However, net replacement rates that take into account the taxes and benefits system may give a better indication of the extent to which households are adequately prepared for their retirement. Individuals above the statutory retirement age face lower marginal tax rates in the first two brackets of the income tax system and do not pay premiums for social insurance and social security. This means that net replacement rates are in general higher than gross replacement rates.

Total disposable income (the denominator of the net replacement rate) is easily determined by summing primary income minus taxes plus transfers in the IPO database. To compute net pension annuities (the numerator), we compute the average tax burden of 65+ singles and couples in different income deciles. We distinguish homeowners and renters, since mortgage interests are tax deductible. The appendix describes the tax burdens found in IPO, which are reasonably comparable to those found in Microtax (a model that simulates Dutch taxes, CPB (2008)). We apply these tax burdens to the sum of predicted pension annuities in the first, second and third pillar. We do not tax annuities from financial wealth and housing wealth (actually, they were already taxed at the moment they were received as income). Wealth taxation is 1.2% of the financial wealth above the threshold of 20,000 euro per person. We do not take into account wealth taxation explicitly, but implicitly: when we use a real rate of return of 1% we assume that this is net of taxes.

Table 3.6 shows net replacement rates. Whereas the median replacement rate of first and second pillar pensions was 71% in gross terms (table 3.4), this is 84% in net terms, indicating that the majority of households are able to replace 84% of their current disposable income with net

Table 3.6: Net replacement rates, 2008[ab]

Age group	35-49	50-54	55-59	60-64	All
1st and 2nd pillars					
ratio p25	0.73	0.62	0.62	0.63	0.67
ratio p50	0.90	0.76	0.78	0.82	0.84
ratio p75	1.06	0.91	0.92	1.00	1.01
1st and 2nd pillars and private wealth					
ratio p25	0.79	0.68	0.70	0.73	0.75
ratio p50	0.97	0.84	0.86	0.92	0.92
ratio p75	1.15	1.01	1.02	1.12	1.11
Total pension annuity					
ratio p25	0.84	0.76	0.79	0.83	0.81
ratio p50	1.04	0.94	0.96	1.03	1.01
ratio p75	1.27	1.15	1.16	1.28	1.24

public and occupational pension benefits. The median net replacement rate increases to 92% when we take into account voluntary third pillar pensions and private wealth, and to 101% when we also add the imputed rental income of net housing.

Table 3.7 shows that only 24% of the households face a net replacement rate that is lower than 80% when all pension annuities are taken into account. When we only take into account first and second pillar pensions, this percentage is substantially higher (43%). Furthermore, table 3.7 indicates that 51% of all households are able to fully replace current disposable income with net pension annuities.

For international comparison, Crawford and O'Dea (2012) find that 53% of the individuals have a replacement rate below 80% in the UK in 2008, taking into account pension income alone. If the authors take into account all sources of wealth, only 21% of UK individuals fall below a replacement rate of 80%. Based on households, for the Netherlands we find that 43% (first and second pillars) and 24% (total pension annuity) fall below the 80% replacement rate in the Netherlands in 2008. Both calculations are based on a nominal interest rate of 3%.

Table 3.7: Share of households below 80% and 100% net replacement rates, 2008

Age group	35-49	50-54	55-59	60-64	All
Share below 80% gross replacement rate					
1st and 2nd pillars	0.35	0.57	0.53	0.47	0.43
Idem, including private wealth	0.26	0.43	0.39	0.33	0.32
Total pension annuity	0.21	0.31	0.26	0.22	0.24
Share below 100% gross replacement rate					
1st and 2nd pillars	0.67	0.85	0.84	0.75	0.74
Idem, including private wealth	0.55	0.74	0.73	0.61	0.61
Total pension annuity	0.45	0.59	0.56	0.46	0.49

3.6.4 Poverty

Whereas Sections 3.6.2 and 3.6.3 focused on replacement rates, high replacement rates do not necessarily reflect high incomes during retirement. For example, low-income households may face relatively high replacement rates because public pensions and social security benefits provide almost everyone with a social minimum. On the other hand, relative poverty may be high among them.

Unlike the relative poverty thresholds used by the EU, the Netherlands uses an absolute poverty line as official poverty indicator. The official poverty line in the Netherlands is the absolute social minimum proposed by the Netherlands Institute for Social Research (SCP). The social minimum implies that a single person aged 65 or over is in poverty if the person's income is lower than 928 euro (net, excluding holiday allowance) a month in 2008. Pensioners have a low probability to fall in official poverty, since a full public pension equals the social minimum. An individual who has not lived in the Netherlands all of the years between age 15 and 65 may have an incomplete public pension, but in the event that household income is lower than the social minimum and household wealth is lower than 5,325 euro per person, he is supplemented with social assistance benefits up to the social minimum (homeowners may own an additional amount of wealth of 44,950 euro). Due to these supplements up to the social minimum there are almost no elderly households living in absolute poverty. Only if someone has not lived in the Netherlands all of the years

between the age of 15 and 65 and his income or wealth is above these thresholds, or if someone does not possess the Dutch nationality or if that person is in detention he will not receive social assistance to supplement income to the social minimum. Furthermore, the take-up rate of these social assistance supplements is not 100%. Those who do not take-up the social assistance supplement live in poverty.

Our predictions of retirement income indicate that approximately 4% of all households that are currently in the age group 35-64 will need social assistance when they are retired, in order to top up public pension benefits to the social minimum. Among first-generation immigrant households, about 34% will need social assistance. Those who do not take up social assistance will live in poverty. Furthermore, low-income but wealthy households who do not receive a complete public pension may live in poverty according to the social minimum income definition, because they do not receive a social assistance supplement. Considering their wealth, however, these households may not really be said to live in poverty. Finally, some low-income households with relatively high mortgage rents and low imputed rents may fall into poverty according to the social minimum income definition.

Official poverty lines of the EU are based on 60%, 50% and 40% of median equivalized household income (MEI). For the Netherlands, these EU indicators of poverty imply that households fall into poverty when they have a yearly income lower than 12,003, 10,003 or 8,003 euro[15] for a poverty line based on 60%, 50% and 40% of MEI, respectively. These EU poverty thresholds are lower than or about the same as the social minimum. This means that only the above-mentioned households who do not take up social assistance, who have a low income but a high wealth level, or those with relatively high mortgage rents, may fall into poverty according to the EU definitions.

[15] In 2010 euros. Non-deflated poverty lines for 2008 as reported by EU-SILC are 11,713, 9,761 and 7,809 euro respectively.

Vulnerable groups 3.7

This section focuses on several potentially vulnerable groups. We study households with self-employment, since self-employed individuals do not have to participate in a pension fund (in contrast to most Dutch paid workers), and the vulnerability of immigrants, single women, renters, and households that faced unemployment or disability for at least two years between 1989 and 2008. First-generation immigrants may be vulnerable since they have not fully accumulated public pension entitlements. Furthermore, single women may be potentially vulnerable because of small or non-existent occupational pensions due to part-time work and providing care to their children. Recipients of unemployment insurance or social assistance may be vulnerable, since in general they do not accumulate occupational pension rights. For persons in disability insurance this is different. In nearly all pension funds individuals in disability insurance build up occupational pension rights as if the person still works in his previous job, with a dispensation from paying occupational pension premiums. Finally, renters are in general low-income households and they do not build up housing wealth (relatively tax beneficial).

To construct a robust indicator of a vulnerable group we use not only information of the year 2008, but also the years 1989 to 2007. Year-to-year movements in and out of social insurance, for example, are substantial.[16]

Clearly, there is a large overlap between these groups. For example, 36% of the households with at least one first-generation immigrant and 14% of the single women received social assistance for at least one year between 1989 and 2008. This percentage is even higher for households with a single female first-generation immigrant (44%).

Table 3.8 presents median pension annuities and gross median replacement rates of the potentially vulnerable groups. The first column of the table indicates that most of the potentially vulnerable groups have a relatively low annuity from pensions, private wealth and housing. The lowest retirement income is observed among households that experienced at least one year of social assistance. Among these households, retirement income

[16]We do not present net replacement rates in this section. Especially the self-employed have extensive tax facilities. Compared to the wage employed they have a relatively low tax pressure.

is generally not much higher than the basic public pension. The table also indicates that second pillar pensions are substantially lower among self-employed households than among all working age households. Renter have a relatively low median pension annuity and this difference is not completely due to the nonexistence of housing wealth.

The second column shows median gross replacement rates. We see that all potentially vulnerable groups except the self-employed have a replacement rate that is close to or above 70%, indicating that current income can to a large extent be maintained after retirement. Note, however, that high replacement rates among these groups are caused by relatively low current income levels, such that receiving a public pension may already be sufficient to replace current income. This seems to be especially the case for households that received at least one year of social assistance. By comparing three types of replacement rates we find that first-generation immigrants and households on social assistance have barely accumulated non-pension wealth, while those in unemployment or disability insurance have accumulated non-pension wealth.

The median self-employed household is expected to replace only 50% of current income when taking into account just first and second pillar pensions (this is 71% for all working age households). Adding third pillar pensions, private wealth and imputed rental income from net housing reduces the gap. Adding these components allows the median self-employed household to replace 74% of their current income after retirement (compared to 83% for all working age households). The spread around this median replacement rate is larger for the self-employed than for the general population.

Column 3 shows the percentage of households that fall below a replacement rate of 70%. Assuming that a replacement rate of 70% is sufficient, we observe that about 40% of the households in the potentially vulnerable groups do not reach a sufficient replacement rate. This is about 10%-points more than for all working age households. Single women and households on social assistance perform relatively well. For households on social assistance this is due to a construct of the social insurance system in which social assistance benefits are equal to the state pension. Self-employed households, on the other hand, are more often confronted with

a gross replacement rate below 70%. 46% of the self-employed households have a gross replacement rate lower than 70%, when taking into account all wealth components (31% for all working age households). So, the self-employed are less likely to maintain their standard of living. Also, due to extensive tax facilities for the self-employed, the replacement rate will not increase that much when moving from gross to net replacement rates. Note, however, that current income is on average substantially higher among self-employed households than among all working age households.

Scenario analyses 3.8

This section analyzes the sensitivity of pension savings adequacy to the assumptions made. Section 3.8.1 presents an optimistic and a pessimistic scenario in which we vary the indexation of occupational pensions, the real rate of return on non-housing wealth, the real return on property, and the retirement age. In the optimistic scenario, all factors are set optimistically. The opposite is the case in the pessimistic scenario. Hence, the two scenarios provide upper- and lower bounds on the resources available at retirement that actual outcomes are likely to fall into.

Section 3.8.2 uses the parameters of the baseline scenario again and shows the effect of housing wealth depletion after retirement (instead of only taking into account the imputed rental income from net housing).

Optimistic and pessimistic scenarios 3.8.1

Table 3.9 shows the parameters of the baseline, pessimistic and optimistic scenarios. Several assumptions remain constant across the scenarios. In all three scenarios we assume an inflation rate of 2%, a nominal imputed rent of 4% and real return on past third pillar payments of 1% after tax.

Other assumptions vary by scenario. The pessimistic scenario assumes no indexation of occupational pensions, such that the real value of occupational pension rights declines 2% every year due to inflation. The baseline

Table 3.8: Median pension annuity (PA) and gross replacement rates (GRR) of potentially vulnerable groups, 2008[a]

Group (share of all households, 35-64)	Median PA	Median GRR	GRR< 70%
Self-employed (12%)			
1st and 2nd pillar	18,488	0.50	0.73
Idem, including private wealth	24,689	0.62	0.58
Total pension annuity	30,016	0.74	0.46
First generation immigrants (8%)			
1st and 2nd pillar	13,818	0.72	0.46
Idem, including private wealth	14,190	0.74	0.43
Total pension annuity	14,524	0.76	0.41
Persons with no homeownership[b](35%)			
1st and 2nd pillar	16,622	0.72	0.46
Idem, including private wealth	17,410	0.75	0.41
Total pension annuity	17,453	0.75	0.41
Single women (16%)			
1st and 2nd pillar	15,209	0.73	0.44
Idem, including private wealth	16,471	0.77	0.37
Total pension annuity	17,540	0.80	0.32
At least two years experience of unemployment (5%)			
1st and 2nd pillar	20,180	0.67	0.56
Idem, including private wealth	21,732	0.72	0.47
Total pension annuity	24,105	0.78	0.39
At least two years experience of disability (11%)			
1st and 2nd pillar	18,168	0.64	0.61
Idem, including private wealth	19,872	0.69	0.52
Total pension annuity	22,138	0.75	0.42
At least one year experience of social assistance (2%)			
1st and 2nd pillar	12,048	0.81	0.33
Idem, including private wealth	12,164	0.81	0.32
Total pension annuity	12,185	0.82	0.31
All households, 35-64 (100%)			
1st and 2nd pillar	22,699	0.71	0.49
Idem, including private wealth	25,006	0.77	0.39
Total pension annuity	27,905	0.83	0.31

[a] Equivalized household income in 2010 euros.
[b] Renters do not reiceve income from imputed rent; some renters, however, own real estate (holiday homes or a houseboat).

scenario assumes 50% indexation. Real occupational pension rights are not reduced in the optimistic scenario, where full indexation takes place.

We assume a real rate of return of 0%, 1% and 2% in the pessimistic-, baseline- and optimistic scenarios, respectively. Furthermore, the average real rate of return on property from 2008 until retirement is -2%, -1% and 0% in the pessimistic-, baseline- and optimistic scenarios, respectively. This means that in the pessimistic scenario, an individual of age 40 in 2008 experiences a drop in the real value of his house of almost 40% between now and the age of 65; in the optimistic scenario, the drop is 0% (which entails a positive average real rate of return on property as from 2012 until the age of 65, since real housing prices have decreased between 2008 and 2012).

Finally, we assume different retirement ages in the three scenarios. A relatively low retirement age has a negative effect on retirement income and is, therefore, assumed in the pessimistic scenario. On the other hand, a relatively high retirement age has a positive effect on retirement income, and this is assumed in the optimistic scenario. People stop working and start using their pension annuity as from the age of 64 in the pessimistic scenario, 65 in the baseline scenario and 67 in the optimistic scenario. We adjust accumulated pension rights in an actuarially neutral way, using the factors of CPB (2009). This means that we cut occupational pension rights by 8% when the retirement age is 64, and increase occupational pension rights by 2 x 8=16% when the retirement age is 67. For public pensions we use an actuarially fair adjustment rate of 6.5% per year, and private savings are annuitized at age 64 in the pessimistic scenario and at age 67 in the optimistic scenario.

Table 3.10 shows median pension annuities in the pessimistic and optimistic scenarios. The table shows that the different assumptions have the highest impact on occupational pensions and imputed rent (induced by the indexation assumption, the retirement age, and the assumed development of housing prices).

The pessimistic scenario is most harmful to the young cohorts, since they have a longer period without indexation and with decreasing housing prices until they reach retirement. Older cohorts, who are closer to retirement, are relatively well-off in the pessimistic scenario compared to the

Table 3.9: Assumptions in the pessimistic, baseline and optimistic scenario

Scenarios	Pessimistic	Baseline	Optimistic
Inflation	2%	2%	2%
Indexation	0%	50%	100%
Real return assets (after tax)	0%	1%	2%
Real return property (after tax)	-2%	-1%	0%
Imputed rent	4%	4%	4%
Past real return 3rd pension pillar	1%	1%	1%
Future real return 3rd pension pillar	0%	1%	2%
Retirement age	64	65	67

households in the 35-49 age category. On the other hand, in the optimistic scenario, young cohorts have a relatively long period until retirement in which they can benefit from returns on investments and housing wealth.

Compared to the baseline scenario presented in table 3.3, older cohorts perform relatively well in the pessimistic scenario, young cohorts perform relatively well in the optimistic scenario, while the baseline scenario is slightly in favor of the younger age groups. This relatively good position of the younger age group can primarily be explained by relatively high occupational pensions. We may, however, overestimate the occupational pension accumulation of the young due to the assumption that individuals remain in their current job until the age of 65 (64 or 67 in the other two scenarios). In practice it is observed that persons tend to reduce working hours as from the age of 50 (women) or 55 (men), and to retire or become unemployed before the age of 65.

Table 3.11 presents gross replacement rates in the pessimistic and optimistic scenarios, and shows similar patterns as table 3.4. As expected, median replacement rates including all components are lower in the pessimistic scenario (0.70) and higher in the optimistic scenario (1.04) relative to the baseline scenario (0.83) presented in table 3.4. Replacement rates are lower for young cohorts compared to the older cohorts in the pessimistic scenario, while the reverse is true for the baseline and optimistic scenarios. All in all, we can conclude that results regarding retirement savings adequacy are sensitive to different future scenarios. Young generations benefit

Table 3.10: Median pension annuities in the pessimistic- and optimistic scenarios[a]

Age group	35-49	50-54	55-59	60-64	All
Pessimistic					
Public pension	10,573	10,573	11,530	12,186	10,573
Occupational pension	9,027	8,826	7,748	4,690	8,246
Voluntary pension products	0	0	0	0	0
Private pension benefits 65+	0	0	0	0	0
Net savings account	256	313	492	594	332
Stocks substantial shareholders	0	0	0	0	0
Securities	0	0	0	0	0
Business assets	0	0	0	0	0
Imputed rent	812	2,178	2,519	2,849	1,456
Total pension annuity	22,508	24,293	24,829	24,727	23,380
Optimistic					
Public pension	13,134	13,134	14,091	14,728	13,134
Occupational pension	18,530	14,831	11,945	6,598	14,954
Voluntary pension products	0	0	0	0	0
Private pension benefits 65+	0	0	0	0	0
Net savings account	592	599	850	926	676
Stocks substantial shareholders	0	0	0	0	0
Securities	0	0	0	0	0
Business assets	0	0	0	0	0
Imputed rent	1,360	2,930	3,067	3,158	2,147
Total pension annuity	37,170	35,410	33,712	31,059	35,432

[a] Equivalized household income in 2010 euros.

most from an optimistic scenario but also suffer more from a pessimistic scenario, compared to older generations.

Table 3.11: Gross replacement rates in the pessimistic- and optimistic scenarios

Age group	35-49	50-54	55-59	60-64	All
Pessimistic					
1st and 2nd pillars					
ratio p25	0.50	0.43	0.44	0.46	0.47
ratio p50	0.62	0.54	0.56	0.63	0.60
ratio p75	0.75	0.67	0.70	0.80	0.74
1st and 2nd pillars and private wealth					
ratio p25	0.54	0.47	0.48	0.53	0.51
ratio p50	0.66	0.58	0.61	0.69	0.65
ratio p75	0.79	0.72	0.75	0.88	0.79
Total pension annuity					
ratio p25	0.57	0.52	0.55	0.60	0.56
ratio p50	0.71	0.65	0.68	0.77	0.70
ratio p75	0.86	0.80	0.84	0.99	0.86
Optimistic					
1st and 2nd pillars					
ratio p25	0.79	0.62	0.58	0.58	0.68
ratio p50	0.99	0.78	0.76	0.80	0.89
ratio p75	1.21	0.95	0.93	1.00	1.11
1st and 2nd pillars and private wealth					
ratio p25	0.86	0.68	0.67	0.68	0.76
ratio p50	1.08	0.86	0.85	0.89	0.97
ratio p75	1.31	1.05	1.02	1.11	1.21
Total pension annuity					
ratio p25	0.91	0.75	0.74	0.76	0.82
ratio p50	1.15	0.94	0.92	0.98	1.04
ratio p75	1.42	1.16	1.13	1.24	1.32

3.8.2 Depletion of housing wealth

Throughout the paper the assumption has been made that households receive an imputed rental income on net housing wealth, but that households do not deplete housing wealth. So, households neither move to a smaller house or rental house, nor use reverse mortgages to finance retirement with housing wealth. This section assumes that net housing

wealth will be depleted after retirement. At the retirement age people buy an annuity from their net housing wealth, in the same way as we assumed for private wealth (explained in section 3.5.1). Thus, households still receive an imputed rental income, but also 'eat up' their housing wealth. Compared to table 3.3, table 3.12 shows that the total median pension annuity is substantially higher when net housing wealth will be depleted. This holds especially for older generations, who have a relatively high net housing wealth. For the age category 60-64, the median total pension annuity increases more than 3,000 euro per year.

Table 3.12: Median pension annuities and gross replacement rates when housing wealth will be depleted[a]

Age group	35-49	50-54	55-59	60-64	All
Net housing annuity	1,976	4,795	5,278	5,729	3,339
Total pension annuity	29,825	31,637	31,270	30,337	30,340
Gross RR, 25th percentile	0.70	0.60	0.62	0.66	0.66
Gross RR, median	0.87	0.76	0.77	0.84	0.83
Gross RR, 75th percentile	1.06	0.93	0.95	1.08	1.03

[a] Equivalized household income in 2010 euros.

The higher total pension annuity due to the depletion of housing wealth also translate into higher gross replacement rates. Taking into account the depletion of housing wealth instead of only taking into account the imputed rent increases the median replacement rate from 0.83 to 0.88, indicating that the median household can replace about 90% of current gross income during retirement if one takes into account pensions, private wealth and the depletion of housing wealth. Half of the households have a gross replacement rate between 69% and 112%.

Summary and conclusions 3.9

The performance of pension systems on adequacy is often evaluated on the basis of fictitious replacement rates for median earners. This study indicates that, although the Dutch pension system is very highly ranked on adequacy, results on adequacy are somewhat less promising when we use microdata to examine the pension that people actually accumulate

in the current system. Results on projected replacement rates are fairly comparable to the UK, despite the fact that the UK pension system has a much lower rank on pension savings adequacy. This suggests that we should be careful in evaluating the adequacy of pensions systems on the basis of fictitious replacement rates.

This study examines the extent of the resources people have available for retirement. Our results show that equivalized financial resources during retirement are on average about 33,000 euro per year and have a median of 27,000 euro per year. Young generations own relatively more occupational pension rights, whereas older generations have accumulated more private wealth and housing wealth. Private wealth and housing wealth raise median replacement rates substantially. Whereas the median gross replacement rate from public and private pensions is 71%, this increases to 83% when account is taken of all pension annuities.

The large-scale administrative data used in this study make it possible to focus on several vulnerable groups, such as households with self-employment. Self-employed households have relatively low occupational pension rights, but relatively high voluntary pensions, private savings and net housing wealth. The total pension annuity has a median of about 30,000 euros. This is somewhat higher than the pension annuities in the total population, however, which also includes inactive households. Replacement rates of the self-employed are relatively low, with a median of 74% for all pension components together. Other vulnerable groups include first-generation immigrants, single women and households that have faced unemployment, disability and/or social assistance. Whereas households with first-generation immigrants and households with social assistance rely almost fully on public pensions (and potentially a supplement from social assistance), households with unemployment or disability often own private wealth, which increases their median replacement rate by more than 10%-points.

Assumptions about indexation, housing prices and the retirement age influence the results. Occupational pension rights decrease dramatically when no indexation takes place between now and retirement, and developments in housing prices influence the imputed rental value of households' net housing wealth. The median total pension annuity varies from 23,000

euro in our pessimistic scenario to 35,000 euro in our optimistic scenario. Associated median replacement rates vary between 70% and 104%. If people were to deplete net housing wealth, the median pension annuity in the baseline scenario would increase by about 3,000 euro per year, which implies an increase in the median gross replacement rate of about 5%-points.

This study represents a first step in the assessment of retirement savings adequacy on the basis of microdata for the Netherlands. There are several important issues to bear in mind when interpreting the results. First, we use the data of Statistics Netherlands about occupational pension rights, which assume that people stay employed in their current job until a fixed retirement age. A natural next step involves taking into account how future wages and labor force participation will evolve. In the current analysis, we are likely to overestimate the occupational pension rights of the young generation since it is questionable whether they will work until the age of 65 and tax favored pension accruals decrease.

Second, we currently assume that no additional private savings will be made. In reality, private savings may increase, especially when households know that the second pillar will become less generous. For example, Alessie et al. (2013) suggested that social security wealth and pension wealth partly displace private savings, and Jia and Zhu (2012) found that this displacement is higher among high-income households than among low-income households. So, cuts in occupational pensions will partly be compensated by private savings, especially among high-income households. Structural models can be estimated to explain private saving behavior of Dutch households, and these models can be used in policy simulations.

Third, this study focuses on retirement income at the retirement age. Pension cuts after that age are not taken into account.

Fourth, we do not take into account that life expectancies differ substantially among income classes. Remaining life expectancy at age 65 is on average 2.5 years shorter for a low-income individual compared to a high-income individual (Kalwij et al. 2013). This means that on average annuitized private savings will be higher for low-income groups and lower for high-income groups.

Finally, whereas current simulations show deterministic outcomes, they are surrounded by uncertainty. In addition to existing uncertainty in future earnings, uncertainty in second pillar pension benefits will increase because risk of return and increases in life expectancy (macro longevity risk) will be deferred to participants of second pillar pension schemes. This is an interesting track for future research.

Given the pension- and long-term care reforms still to come, we argue that it is important to extend this research to convincingly evaluate the effect of several policy-relevant scenarios on a wide variety of households.

3.A Taxes

Table 3.13 shows the median tax burden of 65+ singles and couples in different income deciles. We also distinguish homeowners and renters, since mortgage interests are tax deductible. The median tax burden varies between 10 percent for the lower income deciles and 36 percent for the highest income decile.

Table 3.13: Median tax pressure over income deciles for four types of 65+ households, 2008[a]

Household income deciles	1	2	3	4	5	6	7	8	9	10
Income level (max)	15,514	19,846	25,111	30,228	35,362	40,915	47,150	55,544	70,014	-
Single 65+ households (house)	0.12	0.11	0.14	0.17	0.20	0.23	0.25	0.27	0.30	0.34
Single 65+ households (no house)	0.12	0.10	0.14	0.18	0.21	0.25	0.26	0.28	0.30	0.36
Couples 65+ households (house)	0.12	0.11	0.13	0.17	0.20	0.23	0.25	0.27	0.29	0.34
Couples 65+ households (no house)	0.11	0.11	0.14	0.18	0.21	0.24	0.27	0.27	0.29	0.32

[a] Equivalized household income in 2010 euros.

4 Estimating a panel data sample selection model with part-time employment: Selection issues in wages over the life-cycle

Abstract

This paper proposes a new panel data sample selection model for estimating wages over the life-cycle. The new estimator is an extension of the work of Rochina-Barrachina (1999) who proposed an estimator for panel data selection models where both the selection and the wage equation contain individual effects allowed to be correlated with the observable variables. Instead of solely correcting for systematic differences between those who work and those who do not work (*binary* selection), we extend the model by taking into account part-time and full-time work (*ordered* selection). Since part-time employment decisions provide additional information about unobserved characteristics. Our proposed method is likely to estimate improved wage profiles compared to models that use a binary selection indicator. The newly proposed estimator is applied to a large administrative data set based on Dutch tax records (2001-2011). The application allows us to analyze selection effects in part-time and full-time

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employment as well as the part-time wage penalty over the life-cycle. Education-specific life-cycle wage profiles show the existence of positive selection. For the average man, we do not find a part-time wage penalty. For the average low- and high-educated woman, we find part-time wage penalties of about 30%.

4.1 Introduction

Aging of the population confronts society with a growing number of dilemmas regarding the sustainability of public finances and collective arrangements. In OECD countries, pension systems are affected by demographic aging (OECD 2011b) and many countries have implemented or proposed reforms to alleviate the stress on the sustainability of the system primarily by increasing the statutory retirement age, making pension benefits less generous and increasing contribution rates. Forecasting the resources that people have available for post-retirement consumption is crucial when evaluating the impact of these reforms on government finances and financial well-being of retirees.

Since most pension formulas for pension accumulation are based on earnings during working life, life-cycle wage profiles are crucial in determining income available at retirement. Wages and wage processes are therefore a central component in life-cycle models. Especially, wage uncertainty and the persistence of income shocks play an important role in life-cycle models of consumption- and savings behavior that are used to evaluate retirement savings adequacy (Scholz et al. 2006). A life-cycle earnings model can also be used to simulate future (occupational) pension accumulation and the consequences of proposed reforms for such future pension accumulation (Borella 2004). A good understanding of the life-cycle wage profile is vital in this literature because deviations from the estimated deterministic component of the life-cycle wage profile are supposed to be the result of shocks.

Life-cycle models can be used to analyze retirement savings adequacy (Scholz et al. 2006). The conclusions of such analyses depend on the correct specification of the life-cycle wage-profile. However, the wage

profile estimated in life-cycle models in general does not consider selection effects into work. Wages are likely to be observed non-randomly over the life-cycle, e.g. wages are only observed for people who are working. These same individuals may earn a different potential wage than the individuals who are not working. Neglecting this non-random selection into work may bias estimated wages (Heckman 1979) and wage-profiles (Casanova 2013).

The main objective of this paper is to estimate life-cycle wage profiles of persons in wage employment. We do not distinguish other labor market statuses like self-employment, unemployment, disability, early retirement and other inactivity.[1] We estimate life-cycle wage profiles by using panel data sample selection models with special attention given to selection into full-time and part-time employment. The incorporation of part-time employment is important as part-time employment plays an important role throughout the life-cycle for both men and women. Women tend to prefer part-time employment jobs in general because of the possibility to combine work and care (Booth and Van Ours 2008, Gregory and Connolly 2008). Such part-time employment is often associated with a lower wage than full-time wages among women (Manning and Petrongolo 2008). Among men, part-time employment is often preferred at older ages (Kantarci and Van Soest 2008) as a way to reduce working hours prior to full retirement (e.g. Ruhm (2006), Cahill et al. (2006)). Such end-of-career transitions often imply substantial drops in wages (Aaronson and French 2004, Casanova 2013, Hurd 1996, Johnson and Neumark 1996).

To estimate life-cycle wage profiles using a panel data sample selection model with part-time employment, we propose a new estimator that extends the work of Rochina-Barrachina (1999). Rochina-Barrachina (1999) proposed an estimator for panel data selection models where both the selection and the wage equation contain individual effects allowed to be correlated with the observable variables.[2] Compared to Rochina-Barrachina

[1] Introducing self-employment as a separate state would include another endogenous decision. Also, we do not have information on the number of hours worked by the self-employed.

[2] Other studies dealing with the estimation of panel data sample selection models are Wooldridge (1995) and Kyriazidou (1997). Dustmann and Rochina-Barrachina (2007) provide a comparison of these three aforementioned estimation methods. Endogeneity

(1999), who uses a binary selection rule in the selection equation, we implement an ordered selection rule. By using an ordered indicator instead of a binary selection indicator we are able to take into account extra information regarding unobserved individual characteristics, such as ability and preferences, from selection into part-time and full-time work that may influence wages. Instead of only correcting for systematic differences between those who work and those who do not work, we also take into account unobserved differences between those who are employed part-time and full-time in a panel data sample selection model.

Like Rochina-Barrachina (1999) we eliminate individual specific effects from the equation of interest by taking first- and higher order differences. Furthermore, a conditional mean independence assumption (Wooldridge 1995) is made to deal with the possible correlations between the unobserved individual specific effects and the explanatory variables in the selection equation. In the literature, discrete choice models have been used to analyze part-time and full-time wages, amongst others, by Ermisch and Wright (1993), Dustmann and Schmidt (2000). In contrast to these papers we use a combination of a bivariate ordered probit selection model and a wage equation in differences in order to eliminate individual specific unobserved effects nonparametrically in the second stage. The advantage of using differences in the wage equation is that it allows for an unknown conditional mean of the individual effects.

To estimate the model we use administrative data from the Dutch tax office for the years 2001-2011, which are more representative and reliable than survey data which are often used for the estimation of wage profiles.[3] Our proposed estimator allows us to analyze selection effects, selection into full-time and part-time employment, the part-time wage penalty and the effect of career breaks on wages over the life-cycle.

Earlier contributions to selection into work over the life-cycle shows a diverse picture. Ejrnaes and Kunze (2011) show the existence of negative selection in reentering full-time work after birth among German women.

issues and dynamic panel data sample selection models are dealt with in Semykina and Wooldridge (2010) and Semykina and Wooldridge (2011) respectively.

[3]Most studies analyzing life-cycle wage profiles rely on survey data from PSID. A number of shortcomings of the PSID for analyzing earnings dynamics are mentioned in Pischke (1995).

However, whether selection is positive or negative is found to possibly change over time among women (Mulligan and Rubinstein 2008). Myck (2010) finds that British men approaching the retirement age and who maintain their employment status are more likely to be the lower wage individuals (e.g. negative selection), whereas German men with higher wages are more likely to remain employed (e.g. positive selection). For the US, Casanova (2010) finds negative selection for older men. Using different selection terms Casanova (2013) does not find evidence for selection effects among older men (50+) at all.

For men, the results of applying our two-step estimator suggest the existence of positive selection into work over the life-cycle. This is in contrast with the results we obtain when using the binary selection correction proposed by Rochina-Barrachina (1999). Applying a binary selection indicator suggests negative selection into work. However, adding extra information using an ordered selection indicator changes the sign of selection. We also find positive selection into part-time employment and full-time employment among both men and women as well among low-educated and high-educated groups. Actual selection corrected life-cycle wage profiles however differ between these groups. Estimating education-specific models, we find no part-time wage penalties for the average low- and high educated man respectively. For the average woman, we find part-time wage penalties of 30% and 34% for low- and high-educated women respectively. This wage differential between part-time and full-time work may be a compensation for the ability to combine work with care and a consequence of less experience being accumulated (Boeri and Van Ours 2008). Career breaks have a significant downward effect on life-cycle wages for both men and women although the effect is somewhat more pronounced among men.

The proposed two-step estimator in this paper is likely to be useful in all applications of life-cycle earnings models as the second-stage wage equation is likely to give better estimates of the coefficients of wages over the life-cycle than wage profiles estimated without correction for

selection[4] or with binary selection correction.[5] Applications of the model can vary from estimating life-cycle models (Gourinchas and Parker 2002, Scholz et al. 2006), analyzing earnings inequality (Baker and Solon 2003, Cappellari 2004, Haider 2001) to microsimulation exercises (Borella 2004).

The paper proceeds as follows. First, we describe the administrative data, the selection of the sample, and we provide a descriptive analysis of observed full-time and part-time wages over the life-cycle for men and women (section 4.2). Second, section 4.3 describes the basic model and explains the empirical specification. Section 4.4 shows the estimation results. Education-specific estimates are shown in section 4.4.3. Finally, section 4.5 concludes to what extent it is important to correct life-cycle wage profiles for selection into work and hours.

4.2 Data

The data in this study are taken from the 2001-2011 Income Panel Study from the Netherlands (IPO, CBS 2009a), the 2001-2011 Data on working hours (Baanprsjaarbedragtab, CBS 2010a) and the 2001-2011 data on the highest level of education (Hoogsteopltab, CBS 2010b). All three data sets are gathered by Statistics Netherlands. The IPO, a representative sample from the Dutch population, consists of an administrative panel dataset of, on average, 95,000 selected individuals per year who are followed longitudinally. Sampling is based on individuals' national security number, and the selected individuals are followed for as long as they are residing in the Netherlands on December 31 of the sample year. Individuals born

[4]To bypass possible selection papers focused on prime-aged males who are generally assumed to work to estimate wage profiles. MaCurdy (1982), Abowd and Card (1989), Baker (1997), Lillard and Reville (1999), Haider (2001), Meghir and Pistaferri (2004), Heathcote et al. (2010), Storesletten et al. (2004), Moffitt and Gottschalk (2012), Guvenen (2009), Altonji et al. (2009), Gottschalk and Zhang (2010), Ziliak et al. (2011) and Moffitt and Gottschalk (2011) (US). Dickens (2000), Ramos (2003) and Kalwij and Alessie (2007) (UK). Cappellari (2004) and Borella (2004) (Italy). Baker and Solon (2003) (Canada). Bonke et al. (2011) (Germany). Santos and Souza (2007) (Brazil). Magnac et al. (2011) (France). Sologon and O'Donoghue (2009) (Europe). As a consequence, the results of these models can not be generalized to women and persons approaching the retirement age (Kassi 2013).

[5]Such as Casanova (2010, 2013), Ejrnaes and Kunze (2011), Hanoch and Honig (1985), Johnson and Neumark (1996), Mulligan and Rubinstein (2008), Myck (2010).

in the Netherlands enter the panel for the first time in the year of their birth, and immigrants to the Netherlands in the year of their arrival. The main advantages of using this administrative dataset compared to using survey data for our analysis are, the large sample size, the long panel aspect of the data, the accuracy of tax data compared to survey questions, and representativeness. *Baanprsjaarbedragtab* contains information about working hours (the number of hours worked in proportion to a yearly full time job) for the whole Dutch population. *Hoogsteopltab* provides information on the highest level of education for a subsample of the Dutch population. We merge this information with the internationally standardized *ISCED3* measures of educational levels. The three data sets are merged based on the individuals' personal identifier.

Variable definitions and data selection 4.2.1

The dependent variables in our analysis is the real full-time equivalent wage expressed in 2010 euros. To construct wages, we divide yearly earnings by the proportion of hours worked relative to a full time job. This leaves us with a yearly full-time equivalent wage. Inevitably, we do not observe wages for people that do not work.

In this study we select individuals between the ages of 24 and 64 (309,025 observations for men and 305,678 observations for women). In the estimates, we only use information of persons born no later than 1980. Disentangling age, period and cohort effects works better when an individual is observed over a long time-span. Persons born later than 1980 are only available in the years 2006-2011 at relatively young ages.

Subsequently, we drop some outliers. First, we drop persons who worked less than one-twelfth of a full-time year. We argue they work to little to calculate a reliable wage. Second, we drop observations where the wage rate is higher than the 99th-percentile[6]. In this way we correct for possible measurement error in either earnings or the full-time employment factor leading to a very high wage. Third, we delete observations where the wage is lower than the minimum wage since the minimum wage is

[6] 149,681 euros for men and 89,930 euros for women on average.

legally binding (we take into account yearly differences in the minimum wage level). Fourth, observations are dropped if a year-to-year-change in the wage rate is lower than -50% or higher than 80%. It is highly unlikely that these persons face a year-to-year change in their wage that is due to promotion or demotion. It is more likely that such big changes in year-to-year wages are a consequence of measurement error in the part-time employment factor. Finally, since a lot of people retire during the year observations about the last year of work before retirement are sensitive to mistakes in the number of hours worked in that year. Therefore, we drop observations for which the wage rate dropped more than 30% or increased more than 80% in the last year before retirement.

For the analysis that differentiates between education levels, we end up with 87,401 men and 84,757 women for whom the education level is known. We use population weights to make the sample representative with respect to age, gender, marital status, province, household size and the age of the head of the household.

4.2.2 Descriptive statistics

Table 4.1 shows the development of earnings and wages in our period of observation (2001-2011). The table shows that labor income (including zeros for non-workers) is over time for men and increased for women. Also the average and median wages rates (second column) are slightly increasing over time for men and women. For women, we observe that the average part-time employment factor (which is equal to one if full-time employed throughout the year) increased substantially over the years 2001-2011 from 0.39 to 0.47. For men, the table indicates that average wages are quite stable over time while median wages seem to have increased over time.

Table 4.1 solely focuses on trends over time. To gain insight in wage-differences over the life-cycle and between cohorts we construct age-cohort figures. Figure 4.1 presents average earnings for men and women (including those who do not work). For men, average earnings are about 20,000 euros per year at the age of 25 and grow up to about 35,000 euros per year around the age of 50. After the age of 50, we observe a decline

Table 4.1: Descriptives of real earnings and wage rates

Year	Average earnings[a]	Average wage	Median wage	S.D.	Part-time factor	Obs.[b]
Men						
2001	30,266	43,212	37,948	19,249	0.72	26,142
2002	30,047	43,160	37,974	19,025	0.72	25,764
2003	29,947	43,793	38,511	19,402	0.71	25,891
2004	29,933	44,311	39,076	19,883	0.70	25,717
2005	29,710	44,357	39,005	20,140	0.70	25,686
2006	29,856	44,474	39,196	20,153	0.70	25,823
2007	29,978	44,199	38,906	19,989	0.70	25,954
2008	30,213	44,299	39,062	20,097	0.71	25,820
2009	30,196	44,973	39,752	20,300	0.70	25,913
2010	29,703	44,913	39,692	20,479	0.70	25,831
2011	29,813	44,774	39,376	20,928	0.70	25,552
Women						
2001	13,007	33,564	31,149	11,879	0.39	24,118
2002	13,398	33,908	31,574	11,987	0.40	23,926
2003	13,573	34,265	31,911	12,008	0.40	24,177
2004	13,745	34,634	32,198	12,309	0.40	24,127
2005	13,900	34,688	32,003	12,749	0.40	24,290
2006	14,283	33,329	31,070	11,647	0.42	25,927
2007	14,927	33,483	31,213	11,843	0.44	25,085
2008	15,471	33,671	31,296	11,981	0.45	25,172
2009	15,907	34,489	32,008	12,247	0.46	25,460
2010	16,176	34,876	32,428	12,496	0.46	25,260
2011	16,370	34,678	32,073	12,559	0.47	25,139

[a] Average earnings include observations with earnings equal to zero. Wage rates are only observed for workers.
[b] Total number of observations, including observations with earnings equal to zero.

in average yearly earnings with a huge drop in earnings around the age of 60. The decline in average earnings among men may be explained by several phenomena: 1) early retirement, 2) drops in hours worked (partial retirement), 3) older people receive lower wages and 4) cohort effects. Profound cohort differences are observed among women, because of the increased female labor force participation in the last decades. We observe that a 25 year-old female earns about 17,000 euros per year on average. Around the age of 35 (when most women raise their children) earnings are relatively low, probably because of a drop in the labor force participation and/or the number of hours work. Thereafter, earnings increase and as from the age of 50 earnings decrease again.

Unemployment and part-time employment shape the earnings profile as shown in figure 4.1.[7] Figure 4.2 therefore shows the percentage of men in full-time and part-time employment over the life-cycle for different cohorts. About 70% of all men in all cohorts seem to work full-time until the age 55.[8] However, between 2001 and 2011 it seems in all cohorts about 10% of the men moved from a full-time to a part-time job. Most men seem to end up in unemployment at older ages defined as everyone not in paid-employment. About 30% is unemployed at the age of 55 and this increases to about 90% at the age of 64 for the oldest cohort. As expected, younger cohorts of men retire later.

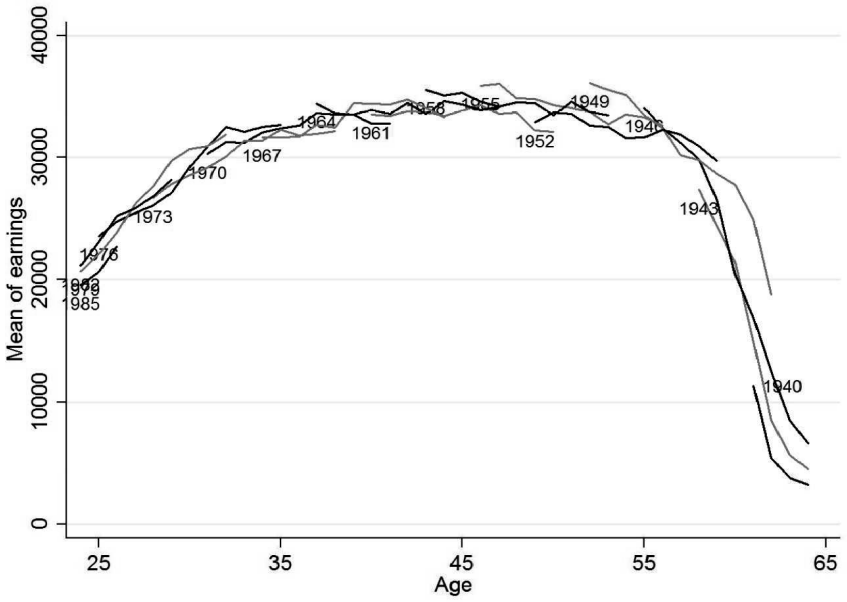
Figure 4.3 presents the percentage of women in full-time and part-time employment over the life-cycle for different cohorts. The figure indicates a substantial drop in full-time employment around the age at which women raise children. Before the age of 30 about 30-40% of women work full-time and this drops to less than 15% at the age of 40, after which it stays constant which is in line with the findings of Bosch et al. (2010). Part-time jobs, on the other hand, increase between the age of 30 and 40 from about

[7] In this paper we define people to be unemployed when they do not earn labor income from paid employment.

[8] We assume persons to be working full-time if the part-time employment factor is equal to one. Every person with a part-time employment factor of smaller than one is considered to be working part-time or unemployed. The effect of considering people with a part-time employment factor of 0.9 or bigger would be marginal as only 5.3% of men and 3.7% of women have a part-time employment factor of larger than 0.9 but smaller than one.

Figure 4.1: Life-cycle earnings of men (a) and women (b)

(a) Earnings men



(b) Earnings women

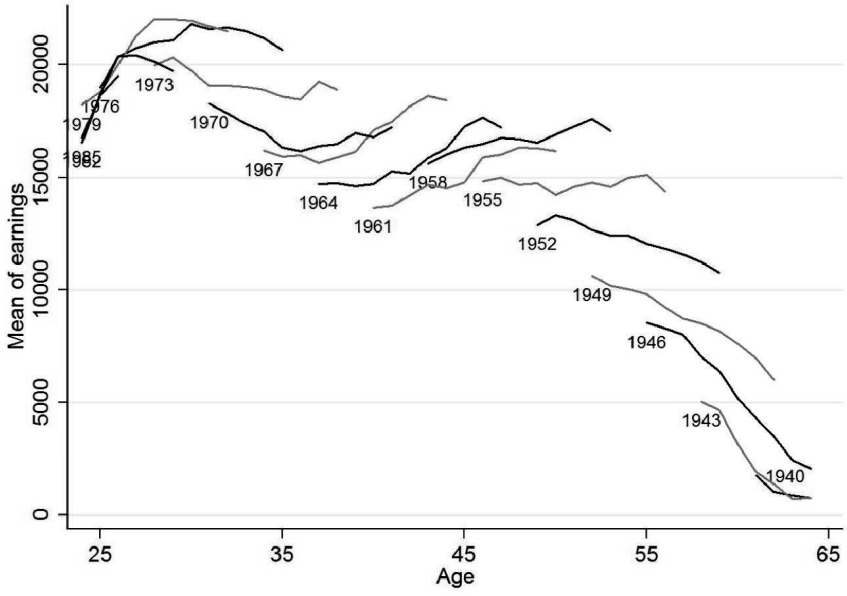
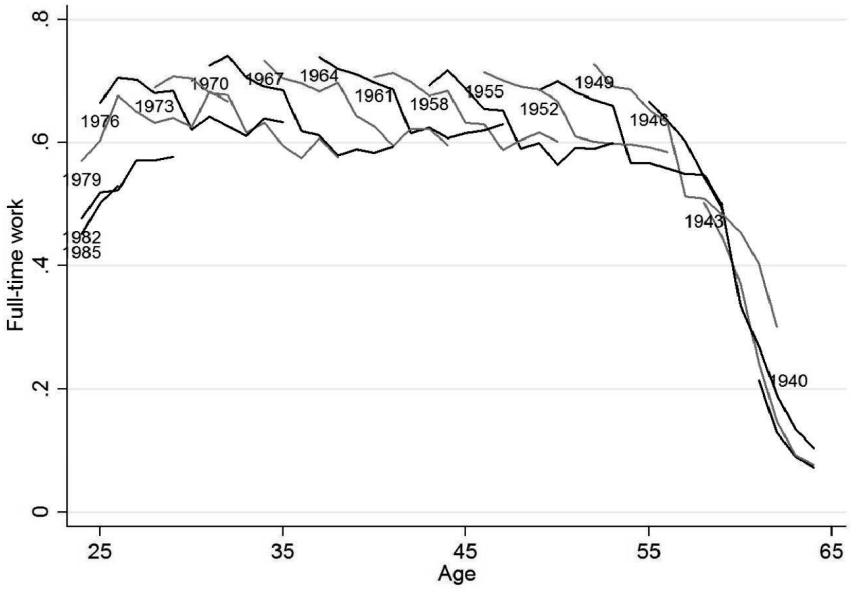
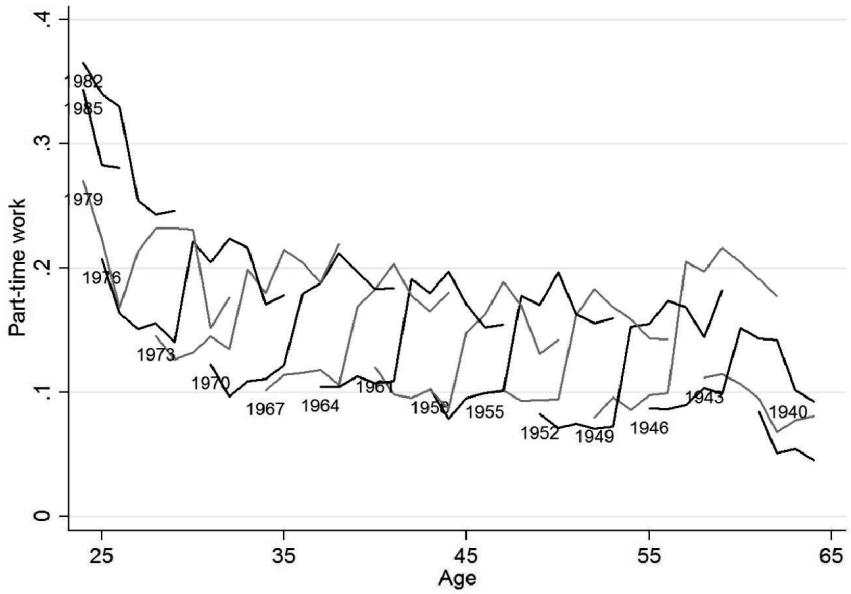


Figure 4.2: Percentage of men in full-time employment (a) and part-time employment (b)

(a) Full-time employment (%)



(b) Part-time employment (%)



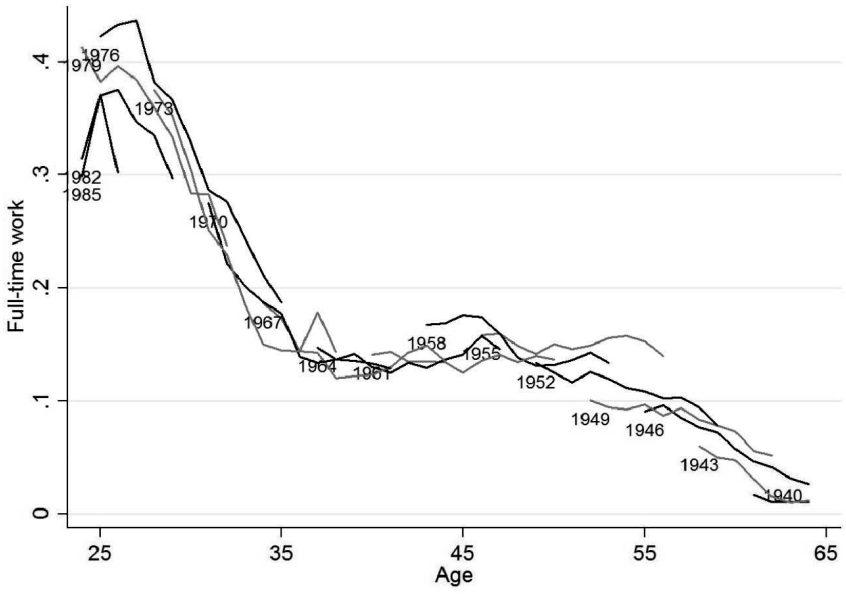
7 to 15%. Unemployment is much lower for younger cohorts than for older cohorts of women. Part-time jobs, however, increase for younger cohorts.

Figure 4.4 shows the average full-time equivalent yearly wage for those in full-time and part-time employment. Average yearly wages of men are approximately 30,000 euros at the age of 25 and about 50,000 euros at the age of 58. Female yearly average wages increase from 27,000 euros at the age of 25 to 35,000 euros at the age of 35 after which it remains relatively constant.

Decomposing the observed wages for persons in full-time and part-time employment shows that full-time wages are generally higher than the part-time wages. This applies to both men (figure 4.5) and women (figure 4.6). This observation may be explained by self-selection effects into full-time and part-time employment, e.g. persons with beneficial (observed and unobserved) characteristics tend to choose for full-time employment. The difference in full-time and part-time wages may also be well explained by the existence of a part-time wage penalty. To test the existence of selection and a part-time wage penalty, we use the model explained in the following sections.

Figure 4.3: Percentage of women in full-time employment (a) and part-time employment (b)

(a) Full-time employment (%)



(b) Part-time employment (%)

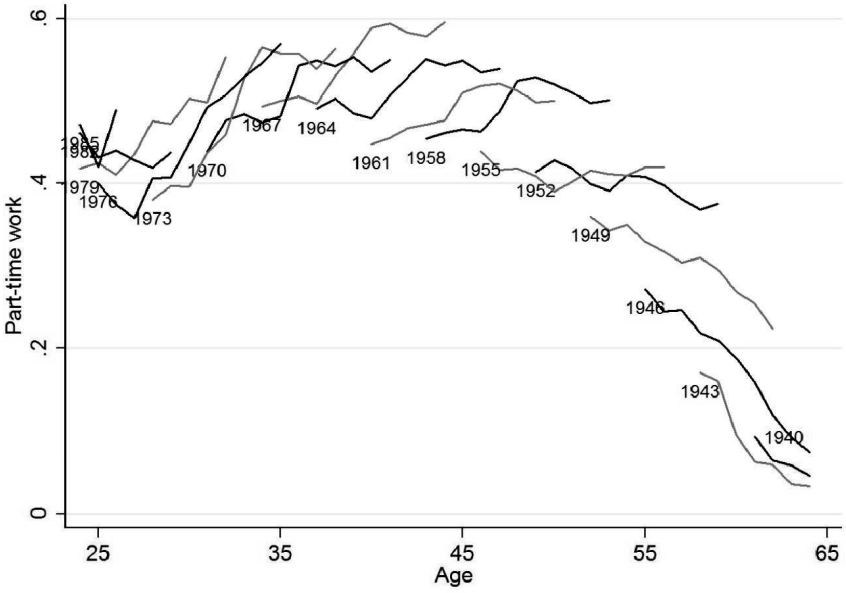
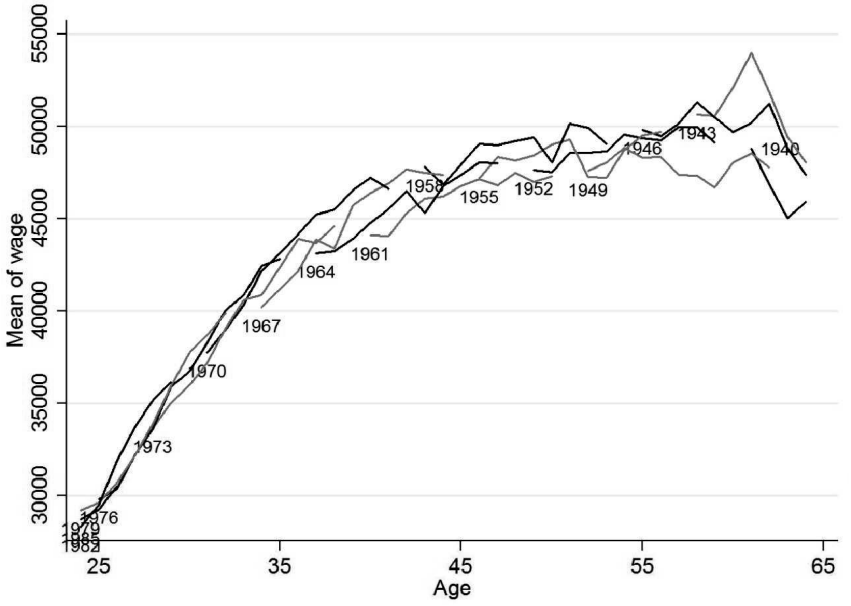


Figure 4.4: Life-cycle wages of men (a) and women (b)

(a) Wages men



(b) Wages women

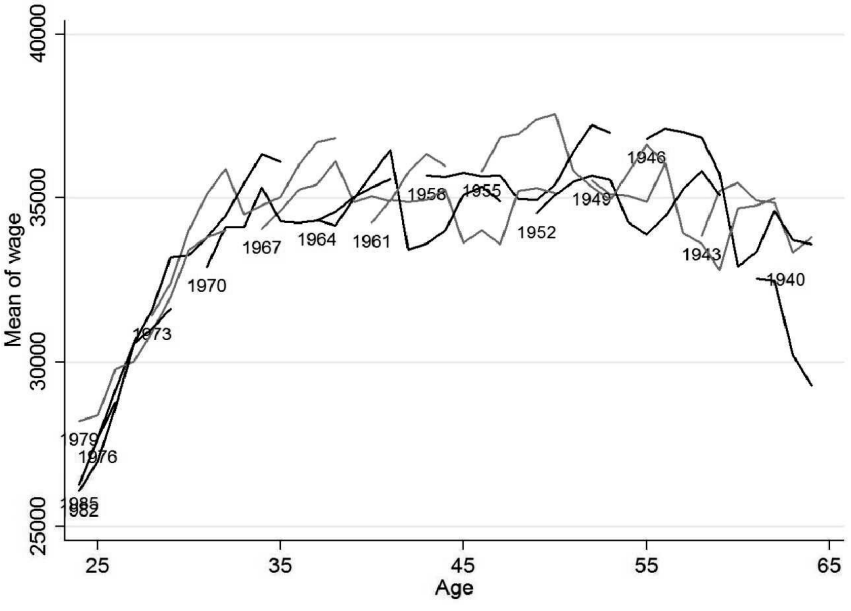
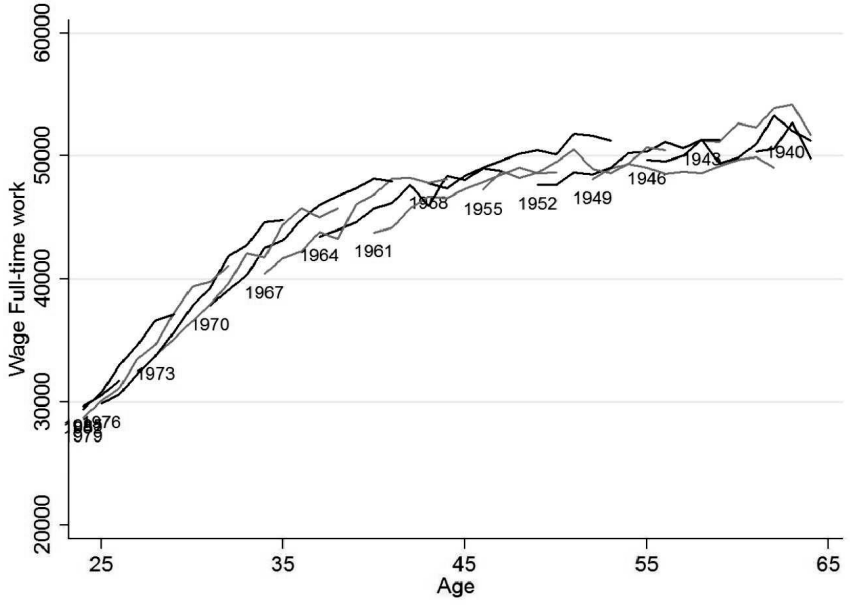


Figure 4.5: Full-time and part-time wages of men

(a) Full-time wages



(b) Part-time wages

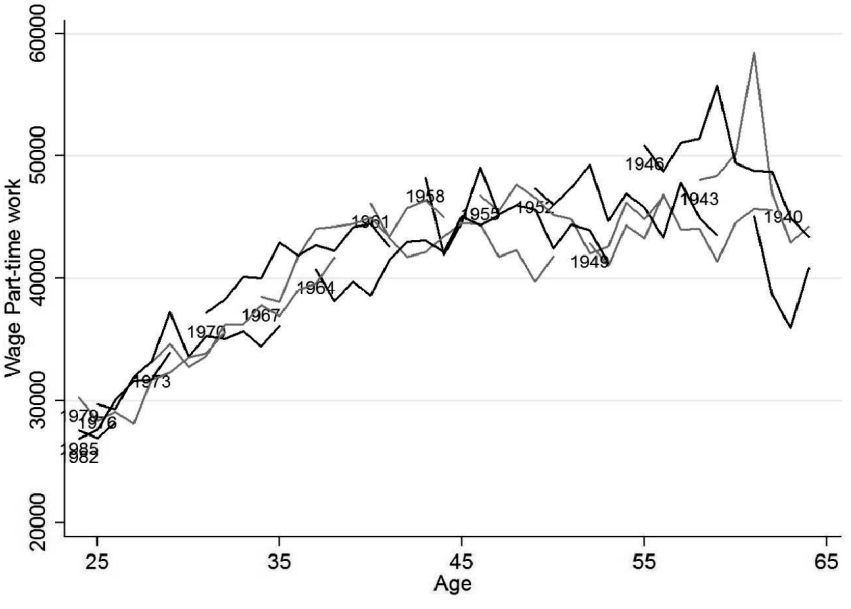
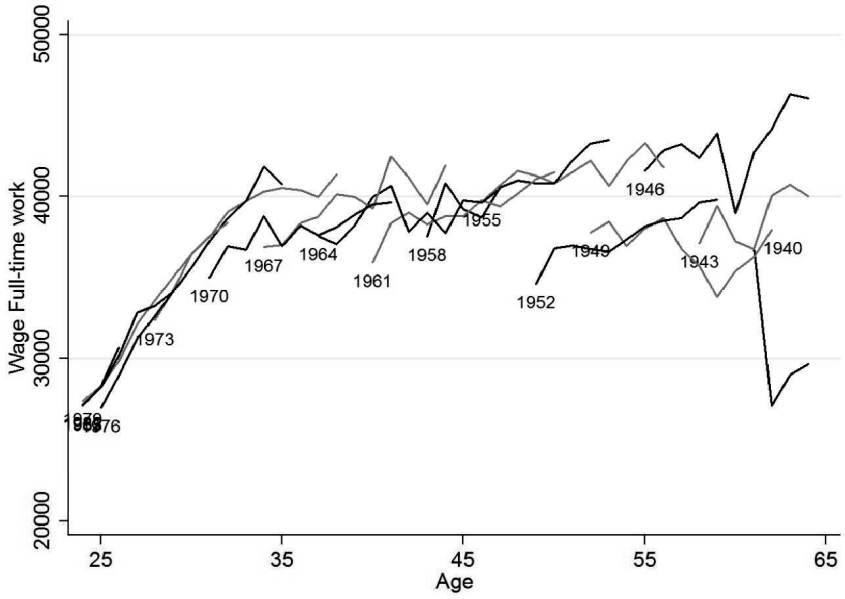
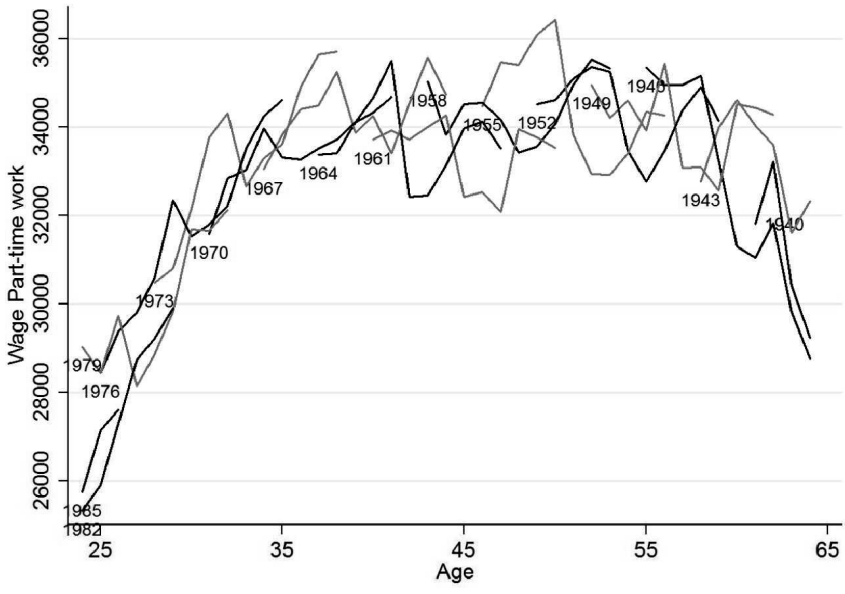


Figure 4.6: Full-time and part-time wages of women

(a) Full-time wages



(b) Part-time wages



4.3 Model

4.3.1 Panel data sample selection model

This section outlines our empirical model for analyzing wages. As discussed in section 4.2 we observe wages and the number of hours worked per year. We use a panel data sample selection model to model both wages and labor force participation at the extensive and intensive margin. The model can be written as follows:

$$y_{it}^* = x_{it}\beta + \alpha_i + u_{it} \quad i = 1, ..., N \quad t = 1, ..., T \tag{4.1}$$

$$h_{it}^* = z_{it}\gamma + \eta_i + v_{it} \tag{4.2}$$

$$y_{it} = \begin{cases} y_{it}^* & \text{if } h_{it}^* > \delta_{1t} \\ \text{unobserved} & \text{otherwise} \end{cases} \tag{4.3}$$

$$h_{it} = \begin{cases} 0 \quad \text{(no participation)} & \text{if } h_{it}^* \leq \delta_{1t} \\ 1 \quad \text{(part-time)} & \text{if } \delta_{1t} < h_{it}^* \leq \delta_{2t} \\ 2 \quad \text{(part-time)} & \text{if } \delta_{2t} < h_{it}^* \leq \delta_{3t} \\ \vdots & \\ J \quad \text{(full-time)} & \text{if } \delta_{Jt} < h_{it}^* \end{cases} \tag{4.4}$$

where y_{it} is the log full-time equivalent wage for individual i in period t. h_{it} contains J categories of labor (no labor force participation, several categories of part-time labor force participation, and full-time labor force participation). Furthermore, x_{it} and z_{it} are vectors of explanatory variables. For identification z_{it} includes variables that do not appear in x_{it} (exclusion restrictions) such as information regarding marital status, children and other household characteristics. β and γ are unknown parameter vectors to be estimated and α_i and η_i are unobserved individual specific effects, which are possibly correlated with x_{it} and z_{it}. Finally, δ_{jt} with $j = \{1, .., J\}$ are cut-off points to be estimated and u_{it} and v_{it} are unobserved disturbances, presumably not independent of each other,[9] which are assumed

[9] If u_{it} and v_{it} are independent, we do not need to worry about selection effects in the wage equation.

to follow a normal distribution with mean zero and variances $\sigma_{u,t}$ and $\sigma_{v,t}$. u_{it} and v_{it} are assumed to be uncorrelated with x_{it} and z_{it}.

To estimate the model we built upon the approaches of Rochina-Barrachina (1999) and Kalwij (2003). Kalwij (2003) proposed a new estimator for a panel data Tobit model in which the unobserved individual specific effects are allowed to correlate with the explanatory variables. The paper of Rochina-Barrachina (1999) is concerned with the estimation of a panel data sample selection model where both the selection and the regression equation contain individual effects allowed to be correlated with the observable variables.

Following Mundlak (1978) we parameterize the individual specific effect in the selection equation (4.2) as a linear function of the average explanatory variables over time plus a random individual specific effect that is assumed to be independent of the explanatory variables:[10]

$$\eta_i = \overline{z_i}\theta + c_i \tag{4.5}$$

where c_i is a random effect that is assumed to be a normally distributed random variable with mean zero and variance σ_c. Substituting (4.5) into (4.2) yields:

$$h^*_{it} = z_{it}\gamma + \overline{z_i}\theta + \mu_{it} \tag{4.6}$$

where $\mu_{it} = c_i + v_{it}$. Given the distributional assumptions it holds that $\mu_{it} \sim N(0, \sigma_{\mu,t})$, where $\sigma^2_{\mu,t} = \sigma^2_c + \sigma^2_{v,t}$. Furthermore, μ_{it} is allowed to be serially dependent (this is important, because of the term c_i).

By taking first- and higher order differences we eliminate the individual specific unobserved effects α_i without having to assume a specific parameterization of the individual unobserved effect in the wage equation (4.1). We can only observe wage differences for those observations for

[10] An application of Mundlak (1978) to panel data selection models was first used in Wooldridge (1995).

which an individual has worked at both time t and $t-m$:

$$y_{it} - y_{it-m} = \begin{cases} y^*_{it} - y^*_{it-m} & \text{if } h^*_{it-m} > \delta_{1t-m} \text{ and } h^*_{it} > \delta_{1t} \\ \text{unobserved} & \text{otherwise} \end{cases} \tag{4.7}$$

where

$$y^*_{it} - y^*_{it-m} = (x_{it} - x_{it-m})\beta + (u_{it} - u_{it-m}), m \geq 1 \tag{4.8}$$

Estimating equation (4.8) by OLS would yield inconsistent estimates of β as the conditional expectation of the error term is unlikely to be zero due to correlation between u_{it} and v_{it} (e.g. selection effects into work). Therefore, Rochina-Barrachina (1999) calculates the expectation conditional on $h^*_{it-m} > \delta_{1t-m}$ and $h^*_{it} > \delta_{1t}$.[11]

$$\begin{aligned} &E[y_{it} - y_{it-m} | x_i, z_i, h^*_{it-m} > \delta_{1t-m}, h^*_{it} > \delta_{1t}] \\ &= (x_{it} - x_{it-m})\beta + E[u_{it} - u_{it-m} | x_i, z_i, h^*_{it-m} > \delta_{1t-m}, h^*_{it} > \delta_{1t}] \\ &= (x_{it} - x_{it-m})\beta \\ &+ E[u_{it} - u_{it-m} | x_i, z_i, \mu_{it-m} > \delta_{1t-m} - z_{it-m}\gamma - \overline{z_i}\theta, \mu_{it} > \delta_{1t} - z_{it}\gamma - \overline{z_i}\theta] \end{aligned} \tag{4.9}$$

The errors $[(u_{it} - u_{it-m}), \mu_{it-m}, \mu_{it}]$ are assumed to be trivariate normally distributed conditional on x_i and z_i. Denote the correlation coefficient of μ_{it-m} and μ_{it} by ρ_{tm}. By taking the derivative of the moment generating function of the truncated multi-normal distribution with respect to t and evaluating the function in $t = 0$, Rochina-Barrachina (1999) obtains the following conditional mean of the error term $(u_{it} - u_{it-m})$:[12]

$$\begin{aligned} &E[u_{it} - u_{it-m} | \mu_{it-m} > \delta_{1t-m} - z_{it-m}\gamma - \overline{z_i}\theta, \mu_{it} > \delta_{1t} - z_{it}\gamma - \overline{z_i}\theta] \\ &= \pi_{1tm}\lambda_{1itm}(M_{it-m}, M_{it}, \rho_{tm}) + \pi_{2tm}\lambda_{2itm}(M_{it-m}, M_{it}, \rho_{tm}) \end{aligned} \tag{4.10}$$

[11] The method of Rochina-Barrachina (1999) is a specific case of our general model presented in equations (4.1)-(4.4) in which only information on work versus no work is used. Equation (4.4) contains two categories: no participation and participation.

[12] This result is based on calculating the first moment of the truncated multivariate normal distribution as in Tallis (1961).

where

$$M_{it-m} = (-\delta_{1t-m} + z_{it-m}\gamma + \overline{z_i}\theta)/\sigma_{\mu,t-m} \tag{4.11}$$

$$M_{it} = (-\delta_{1t} + z_{it}\gamma + \overline{z_i}\theta)/\sigma_{\mu,t} \tag{4.12}$$

and

$$\lambda_{1tm}(M_{it-m}, M_{it}, \rho_{tm}) = \frac{\phi(M_{it-m})\Phi\left((M_{it} - \rho_{tm}M_{it-m})/\sqrt{1-\rho_{tm}^2}\right)}{\Phi^2(M_{it-m}, M_{it}; \rho_{tm})} \tag{4.13}$$

$$\lambda_{2tm}(M_{it-m}, M_{it}, \rho_{tm}) = \frac{\phi(M_{it})\Phi\left((M_{it-m} - \rho_{tm}M_{it})/\sqrt{1-\rho_{tm}^2}\right)}{\Phi^2(M_{it-m}, M_{it}; \rho_{tm})} \tag{4.14}$$

Applying OLS on the sample of first- and higher order differences will yield consistent estimates of β if the selection correction terms (4.10) are added to (4.1). If added to the regression equation, the new error term $\xi_{it} \equiv (u_{it} - u_{it-m}) - (\pi_{1tm}\lambda_{1itm} + \pi_{2tm}\lambda_{2itm})$ has a conditional expectation of zero by construction.

Panel data sample selection model with part-time employment 4.3.2

The proposed method by Rochina-Barrachina (1999) takes into account the binary selection of work versus no work. We argue that more information regarding the correlation between u_{it} and v_{it} can be added to the model by additionally taking into account labor supply at the intensive margin.

By using an ordered selection equation instead of a binary selection equation, we are able to take into account the extra information available from observing part-time and full-time work. Thus, instead of only correcting for systematic differences between those who work and those who do not work, we also take into account unobserved differences between those who work part-time and full-time.

We extend equation (4.9) by taking into account the lower- and upper thresholds of working hours categories, which yields

$$\begin{aligned}
&E[y_{it} - y_{it-m} | x_i, z_i, \delta_{j,t} < h^*_{it} \leq \delta_{j+1,t}, \delta_{j,t-m} < h^*_{it-m} \leq \delta_{j+1,t-m}] \\
&= (x_{it} - x_{it-m})\beta \\
&+ E[u_{it} - u_{it-m} | x_i, z_i, \delta_{j,t} < h^*_{it} \leq \delta_{j+1,t}, \delta_{j,t-m} < h^*_{it-m} \leq \delta_{j+1,t-m}] \\
&= (x_{it} - x_{it-m})\beta \\
&+ E[u_{it} - u_{it-m} | x_i, z_i, G_{it-m} \leq \mu_{it-m} < H_{it-m}, G_{it} \leq \mu_{it} < H_{it}] \quad (4.15)
\end{aligned}$$

where

$$H_{it-m} = -\delta_{j,t-m} + z_{it-m}\gamma + \overline{z_i}\theta \quad (4.16)$$

$$G_{it-m} = -\delta_{j+1,t-m} + z_{it-m}\gamma + \overline{z_i}\theta \quad (4.17)$$

$$H_{it} = -\delta_{j,t} + z_{it}\gamma + \overline{z_i}\theta \quad (4.18)$$

$$G_{it} = -\delta_{j+1,t} + z_{it}\gamma + \overline{z_i}\theta \quad (4.19)$$

and where j is the working hours category of individual i at time t. For persons who do not work at time t, $\delta_{0,t} = -\infty$. For these people, $H_{it} = \infty$. Similarly, for persons engaged in full-time work at time t, $\delta_{J+1,t} = \infty$ such that $G_{it} = -\infty$.

As in the framework of Rochina-Barrachina (1999), the errors $[(u_{it} - u_{it-m}), \mu_{it-m}, \mu_{it}]$ are assumed to be trivariate normally distributed conditional on x_i and z_i. Denote the correlation coefficient of μ_{it-m} and μ_{it} by ρ_{tm}. We can write out the conditional mean in (4.15) by:

$$\begin{aligned}
&\mathrm{E}(u_{it} - u_{it-m} | x_i, z_i, G_{it-m} \leq \mu_{it-m} < H_{it-m}, G_{it} \leq \mu_{it} < H_{it}) = \quad (4.20) \\
&\pi_{1tm}\lambda_{1itm}(\rho_{tm}, b_{it}, a_{it-m}, b_{it-m}) \\
&+ \pi_{2tm}\lambda_{2itm}(\rho_{tm}, a_{it}, a_{it-m}, b_{it-m}) \\
&+ \pi_{3tm}\lambda_{3itm}(\rho_{tm}, a_{it}, b_{it}, b_{it-m}) \\
&+ \pi_{4tm}\lambda_{4itm}(\rho_{tm}, a_{it}, b_{it}, a_{it-m})
\end{aligned}$$

where

$$a_{it-m} = \frac{G_{it-m}}{\sigma_{\mu,t-m}} \tag{4.21}$$

$$b_{it-m} = \frac{H_{it-m}}{\sigma_{\mu,t-m}} \tag{4.22}$$

$$a_{it} = \frac{G_{it}}{\sigma_{\mu,t}} \tag{4.23}$$

$$b_{it} = \frac{H_{it}}{\sigma_{\mu,t}} \tag{4.24}$$

with $\sigma_{\mu,t}$ and with $\sigma_{\mu,t-m}$ being the variances of the error term of the selection equation for time t and $t-m$ respectively and where

$$\lambda_{1itm}(\rho_{tm}, b_{i,t}, a_{it-m}, b_{it-m}) = \frac{\phi(b_{it}) \left[\Phi\left((b_{it-m} - \rho_{tm} b_{it}) / \sqrt{1-\rho_{tm}^2} \right) - \Phi\left((a_{it-m} - \rho_{tm} b_{it}) / \sqrt{1-\rho_{tm}^2} \right) \right]}{\Phi^2(b_{it-m}, b_{it}; \rho_{tm}) - \Phi^2(a_{it-m}, a_{it}; \rho_{tm})} \tag{4.25}$$

$$\lambda_{2itm}(\rho_{tm}, a_{it}, a_{it-m}, b_{it-m}) = \frac{\phi(a_{it}) \left[\Phi\left((b_{it-m} - \rho_{tm} a_{it}) / \sqrt{1-\rho_{tm}^2} \right) - \Phi\left((a_{it-m} - \rho_{tm} a_{it}) / \sqrt{1-\rho_{tm}^2} \right) \right]}{\Phi^2(b_{it-m}, b_{it}; \rho_{tm}) - \Phi^2(a_{it-m}, a_{it}; \rho_{tm})} \tag{4.26}$$

$$\lambda_{3itm}(\rho_{tm}, a_{it}, b_{it}, b_{it-m}) = \frac{\phi(b_{it-m}) \left[\Phi\left((b_{it} - \rho_{tm} b_{it-m}) / \sqrt{1-\rho_{tm}^2} \right) - \Phi\left((a_{it} - \rho_{tm} b_{it-m}) / \sqrt{1-\rho_{tm}^2} \right) \right]}{\Phi^2(b_{it-m}, b_{it}; \rho_{tm}) - \Phi^2(a_{it-m}, a_{it}; \rho_{tm})} \tag{4.27}$$

$$\lambda_{4itm}(\rho_{tm}, a_{it}, b_{it}, a_{it-m}) = \frac{\phi(a_{it-m}) \left[\Phi\left((b_{it} - \rho_{tm} a_{it-m}) / \sqrt{1-\rho_{tm}^2} \right) - \Phi\left((a_{it} - \rho_{tm} a_{it-m}) / \sqrt{1-\rho_{tm}^2} \right) \right]}{\Phi^2(b_{it-m}, b_{it}; \rho_{tm}) - \Phi^2(a_{it-m}, a_{it}; \rho_{tm})} \tag{4.28}$$

For the derivation of this result by calculating the first moment of the doubly truncated multivariate distribution, we refer to Appendix 4.A. As

$\xi_{itm} \equiv (u_{it} - u_{it-m}) - (\pi_{1tm}\lambda_{1itm} + \pi_{2tm}\lambda_{2itm} + \pi_{3tm}\lambda_{3itm} + \pi_{4tm}\lambda_{4itm})$ has a conditional expectation of zero by construction, taking into account both the lower- and upper thresholds of working hours categories results in four selection correction terms; two more than the binary selection approach of Rochina-Barrachina (1999).[13]

4.3.3 Experience and unemployment

Labor market experience has a positive return on the wage rate (see for example Dustmann and Meghir 2005). On the other hand, unemployment has a negative effect on post-unemployment wages (see for example Schmieder et al. 2013). In our proposed model in section 4.3.2, we are able to take into account information regarding experience by investigating wage differences between t and $t-m$ ($m = \{1, 2, 3, ..., 10\}$). When people experience years of unemployment between t and $t-m$ we take this explicitly into account in the model by including a variable indicating the number of years without labor income between time t and $t-m$. This provides information about how wage growth is influenced by years of unemployment. Our large data set allows us to investigate the effect of unemployment on wage growth for men and women at different ages and during different stages of the business cycle. Since the effect of the number of years unemployed on wage growth may be nonlinear, we include the number of years unemployed as a linear spline with knots at 0, 1 and 3 years of unemployment. This linear spline takes into account that the effect of unemployment on wage growth may be different in the first year compared to the second and third year and four or more years.

4.3.4 Estimation

To estimate the model we use a two-step estimation procedure. In the first step we deal with the selection equation. We estimate the following bivariate ordered probit model for each $s = \{t, t-m\}$.

[13] Technically, this is a consequence of the difference in analyzing the first moment of a singly and doubly truncated multivariate normal distribution.

$$h^*_{it-m} = z_{it-m}\gamma_{t-m} + \overline{z_i}\theta_{t-m} + \mu_{it-m} \tag{4.29}$$

$$h^*_{it} = z_{it}\gamma_t + \overline{z_i}\theta_t + \mu_{it} \tag{4.30}$$

$$h_{is} = \begin{cases} 0 \quad \text{(no participation)} & \text{if } h^*_{is} \leq \delta_{1s} \\ 1 \quad \text{(part-time)} & \text{if } \delta_{1s} < h^*_{is} \leq \delta_{2s} \\ 2 \quad \text{(part-time)} & \text{if } \delta_{2s} < h^*_{is} \leq \delta_{3s} \\ \vdots & \\ J \quad \text{(full-time)} & \text{if } \delta_{Js} < h^*_{is} \end{cases} \tag{4.31}$$

The bivariate ordered probit model takes into account the correlation between μ_{it} and μ_{it-m}. This is necessary because we assume that this error-term has a time-constant individual component (c_i in $\mu_{it} = c_i + v_{it}$, see section 4.3).

In the second step we construct the correction terms λ_{1itm}, λ_{2itm}, λ_{3itm} and λ_{4itm} by using the estimates $\hat{a}_{it}$, $\hat{a}_{it-m}$, $\hat{b}_{it}$, $\hat{b}_{it-m}$, $\hat{\sigma}_{\mu,t}$, $\hat{\sigma}_{\mu,t-m}$ and $\hat{\rho}_{tm}$. $\hat{\lambda}_{1itm}$, $\hat{\lambda}_{2itm}$, $\hat{\lambda}_{3itm}$ and $\hat{\lambda}_{4itm}$ are used as additional regressors in the wage equation to obtain consistent estimates of β by OLS on the sample of wages observed in t and $t-m$.[14,15] $\binom{M}{2} \times 4$ selection terms are added to the wage equation.[16] We estimate

$$y^*_{it} - y^*_{it-m} = (x_{it} - x_{it-m})\beta + \sum_{c=1}^{4} \pi_{ctm}\lambda_{citm} + (u_{it} - u_{it-m}),$$
$$m \geq 1 \tag{4.32}$$

[14] We use bootstrapped standard errors for inference in the two-stage approach (Wooldridge 2002).

[15] Note that our estimation approach is slightly different from the approach taken in Rochina-Barrachina (1999). Rochina-Barrachina (1999) estimates separate OLS regressions for each s and uses a minimum distance estimator on the separate OLS regressions to obtain the regression results. We, on the other hand, estimate one OLS regression on first- and higher order differences. Both approaches assume that the effects are the same for each s.

[16] $M = 10$. The bivariate model consists of pairs of 2. We obtained 4 selection correction terms per combination.

4.4 Estimation results

4.4.1 Selection equation

We model the first-stage bivariate ordered probit models with four ordered categories of labor force participation: 1) no participation, 2) participation lower than or equal to 50% of the full-time working hours, 3) more than 50% but less than 100% of the full-time working hours, and 4) working full-time.[17]

We allow for a semi-parametric specification of age effects by using age-dummies as explanatory variables in vector z. Following Ermisch and Wright (1993) and Paci et al. (1995), we use information on marital status (dummies for married, divorced and widowed) and children (the number of children and age of the youngest child) as exclusion restrictions in z_{it}. Furthermore, we use a dummy variable that indicates whether an individual has a partner aged 62 or older. As an additional control variable we include a dummy for first-generation immigrants. $\overline{z_i}$ includes the individual's time-averages of the marital status dummies, the variables providing information on children and the dummy whether there is a partner aged 62 or older present in the household.

The bivariate ordered selection model is estimated for each combination of t and $t-m$.[18] The separate estimations capture period and cohort differences in labor force participation.

4.4.2 Wage equation

The main equation (4.1) contains a flexible semi-parametric specification of age- and period effects (following Kalwij and Alessie 2007). However, age, period, and cohort effects (captured in the individual effect) cannot be identified empirically because the calendar year is equal to the year of birth plus age thereby spanning up the vector space. To identify age, period, and cohort effects we follow the identification restriction proposed by

[17] We do a sensitivity check with more part-time employment categories.

[18] In our case with data from 2001-2011, this implies separate estimations for $(2002, 2001)$, $(2003, 2002)$, $(2003, 2001)$, ..., $(2011, 2010)$, ..., $(2011, 2001)$.

Deaton and Paxson (1993). This means that we assume that all remaining period effects add up to zero and are orthogonal to a linear time trend.

The estimated period effects (not reported here) are generally significant but their effects on wages are rather small compared to the age effects, e.g. most of the wage growth is a result of age and cohort effects.

Selection over the life-cycle

Figure 4.7 shows the estimated age coefficients of the wage regressions for men and women. The figure indicates the differences in the estimated age coefficients for 1) a model without selection correction (solid line), 2) a model with binary selection correction as in Rochina-Barrachina (1999) (dotted line) and 3) a model with ordered selection correction as proposed in this paper (dashed line).

First focusing on the model without selection (solid line), the results can be interpreted as follows. A 64 year old male has a 60% higher wage than a 24 year old male. A 64 year old female has an approximately 50% higher wage than a 24 year old female. Wages slightly decrease after the age of 58. The wage at age 64 is significantly lower than the wage at age 58.[19] The wage at age 64 is comparable to the wage at age 46.[20] Among women, we observe a lower wage growth from the age of 47 as the slope of the wage curve decreases.

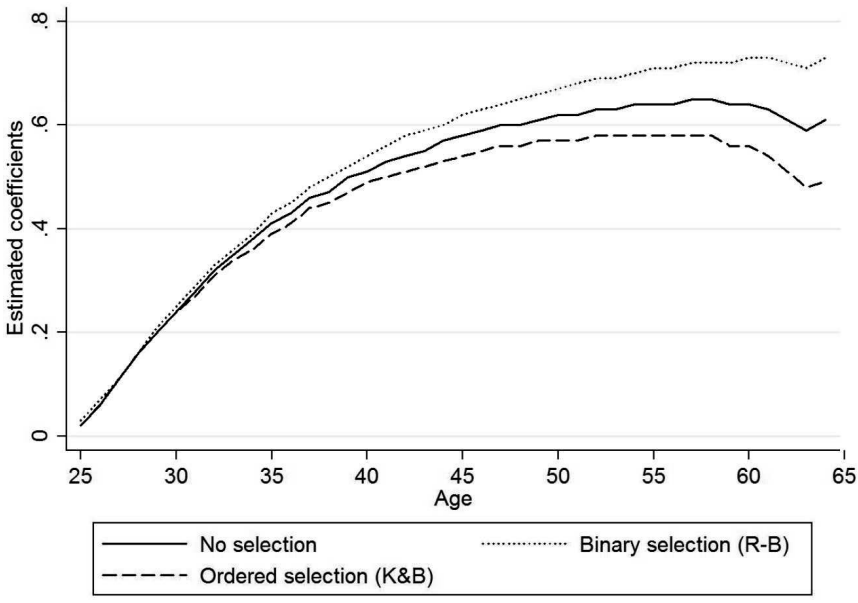
To test for selection, we follow Rochina-Barrachina (1999) who argues that a valid test of no selection is a *Wald-test* of the joint significance of the selection terms. In the binary selection model this means a Wald-test on $\binom{M}{2} \times 2$ coefficients. In the ordered selection model this means a Wald-test on $\binom{M}{2} \times 4$ coefficients.

For men, the selection correction terms of the binary selection correction are jointly significant.[21] Estimated age coefficients are higher than in the model without selection correction. This suggests that correcting the wages for persons whose wages are not observed gives a higher age effect on wages than the model without selection correction. This result suggest

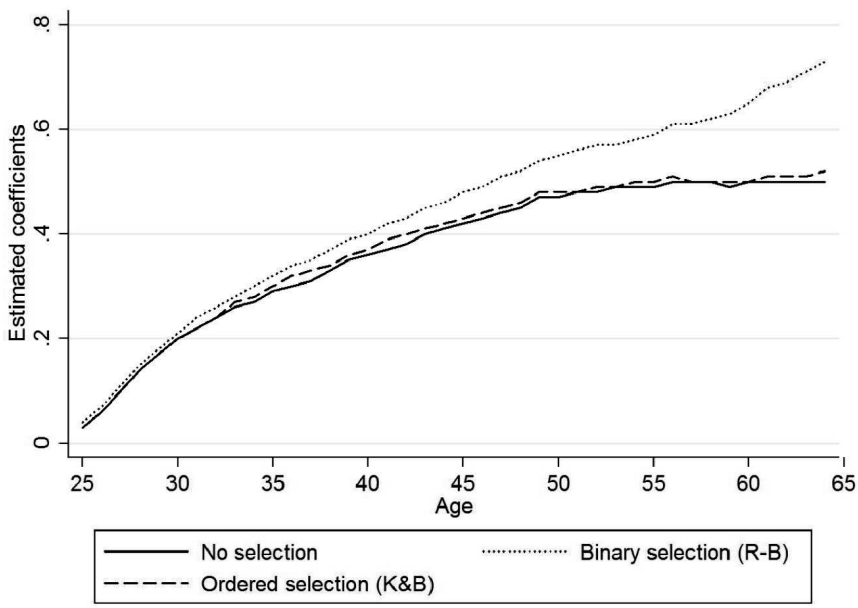
[19] $H_0 : \beta_{58} = \beta_{64}$. H_0 rejected, p-value= 0.00.
[20] $H_0 : \beta_{46} = \beta_{64}$. H_0 can not be rejected, p-value= 0.20.
[21] P-value= 0.00.

Figure 4.7: Binary versus ordered selection correction regressions and regressions without selection correction of men (a) and women (b)

(a) Estimated coefficients men



(b) Estimated coefficients women



the existence of negative selection into work over the life-cycle among men, e.g. men with worse observed and unobserved characteristics tend to work.

Testing the joint significance of the selection correction terms of the ordered selection correction are jointly significant.[22] Again, we find an inverse U-shape of wages over age that is even more pronounced than in the model with binary selection correction and the model without selection correction. The results show that, wages drop by a substantial 9%-points from the peak at age 55 to the wage at age 64. Correcting the wages for persons whose wages are not observed by the ordered selection procedure gives a lower age effect on wages than the model without selection correction. This result suggests positive selection into work over the life-cycle among men. Positive selection seems especially pronounced towards the end of the career. Such positive selection into work would have remained unnoticed in a model that corrects for selection by using a binary indicator. In stead, the model with the binary selection procedure suggests that there is negative selection into work among men. These different results indicate that correcting for selection into work and working hours simultaneously may lead to different conclusions than correcting solely for selection into work.

For women, the selection correction terms of the binary selection correction are jointly significant.[23] The estimation results show that women's wages tend to increase over the life-cycle. Furthermore, we find that correcting for selection with the binary selection indicator suggests the existence of negative selection into work.

Testing the joint significance of the selection correction terms for the ordered selection model indicates that selection is present.[24] Whereas the model with the binary selection indicator suggests the existence of negative selection among women, the model with the ordered selection rule suggests that this negative selection is much smaller. Especially among older women.

[22] P-value= 0.00.
[23] P-value= 0.00.
[24] P-value= 0.00.

Table 4.2: Effect of career breaks (years) on wage

	Men		**Women**	
	Coeff.	S.D.	Coeff.	S.D.
Career break $= 1$	-0.11***	0.01	-0.07***	0.01
$1 <$ Career break ≤ 3	-0.05***	0.01	-0.05***	0.01
Career break > 3	0.01	0.02	0.02	0.02

Based on a binary selection indicator we would conclude that negative selection over the life-cycle is present among men and women. However, adding information from working hours decisions makes us conclude that positive selection exists over the life-cycle among men while the negative selection for women is much smaller than suggested by the model with binary selection terms.

Career breaks

The effects of a career break, defined as a year in which one does not receive labor income, on the life-cycle wage is estimated by a linear spline for 1, 2-3 and 4+ years of a career break. A semi-parametric linear spline is used because of possible non-linear effects, e.g. the effect a the first year may be different from the effect of 4+ years.

The estimated coefficients of the linear spline function of career breaks are shown in table 4.2 and can be interpreted as follows. Males who suffer from a career break of at most one year have a 11% lower wage than men without a career break. Men who suffer from a career break of at most 2-years (not necessarily subsequently) face an additional 5% lower wage. An additional third year lowers the wage with another 5%. An additional year after 3 years does not significantly reduce the wage anymore.

Women with a one-year career break face a 7% lower wage than women without a career break. An additional second and third year reduce the wage by 5%. The effects of a career break are not significantly different from zero thereafter.

4.4.3 Education-specific selection in full-time and part-time employment

In this section, we estimate education-specific life-cycle wage profiles. To analyze selection into full-time and part-time wages, we estimate separate wage equations for full-time (β_{FT}) and part-time work (β_{PT}).

Taking into account education may be relevant as wage growth may differ between educational levels (Connolly and Gottschalk 2006). To analyze whether selection into full-time and part-time wages differs between educational levels, we estimate separate wage equations for different educational levels. We use the international ISCED3 standard to define low education (ISCED3= 1 or ISCED3= 2) and high education (ISCED3= 3).

In the selection equation, we use a simplified version of the specification of the bivariate ordered probit model due to the loss of observations when using educational information as explained in section 4.2. Basically, we no longer use a semi-parametric specification of the age effects but assume the effects of age on working hours to be quadratic. Furthermore, we assume that the age effects on hours decisions may differ between low- and high education. The same exclusion restrictions are used as in the earlier specification of the bivariate ordered probit model.

The results (figure 4.8) suggest that there is positive selection in part-time and full-time employment among men. For low-educated men we find positive selection in both part-time[25] and full-time employment.[26] However, we do not find any wage growth over the life-cycle in the part-time wage profile for low-educated men as the estimated age-coefficients are not significantly different from zero. Among high-educated men, we only find significant positive selection into full-time employment.[27] Selection into part-time employment is not significant among high-educated men.[28]

Analyzing the selection effects into part-time and full-time employment among low-educated women shows that there is positive selection into

[25] P-value= 0.00.
[26] P-value= 0.02.
[27] P-value= 0.03.
[28] P-value= 0.11.

both part-time work[29] and full-time work[30] (figure 4.9). When correcting for selection, all age effects on wages become insignificant among low-educated women working part-time, e.g. there is no significant wage growth over the life-cycle. Corrections for selection into part-time and full-time employment among high-educated women shows positive selection in both part-time[31] and full-time work[32] (figure 4.8).

4.4.4 Education-specific part-time wage penalties

Noticeable observations in figures 4.8 and 4.9 are 1) wage growth is steeper over the life-cycle for full-time employment and 2) wage growth is steeper for high-educated persons. This is true for both men and women.

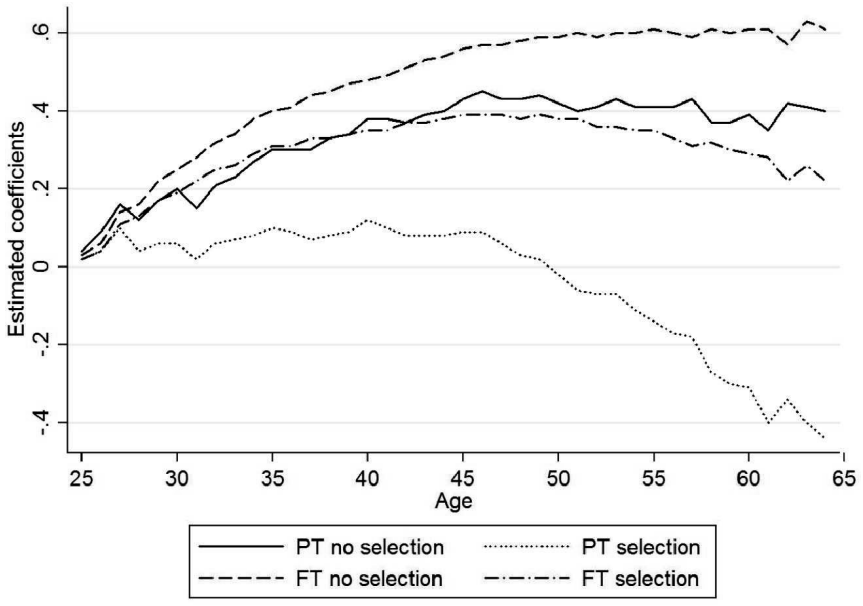
The seminal work of Mincer (1974) focused on the relationship between education and wage. Since Mincer (1974) many more advanced techniques to identify the causal link between education and wages have emerged as summarized by Card (1999) and virtually all studies find positive returns to education. More recently Connolly and Gottschalk (2006) found that also wage growth differs among educational levels which explains our second observation in figures 4.8 and 4.9.

Our first observation suggests the existence of a part-time wage penalty. Most studies find an existing part-time wage penalty although in some papers the penalty almost disappears when controlling for differences in personal- and job characteristics (for example, Manning and Petrongolo 2008). Others studies still find a part-time wage penalty after controlling for such observables with substantial cross-country variation (for example, Gornick and Jacobs 2002). A part-time wage *premium* is also found empirically (for example, Pissarides et al. 2005). However, aforementioned studies do not control for differences in unobserved characteristics such as ability, tastes and preferences for full-time work. If there are unobserved differences between persons choosing for part-time and full-time employment, such as suggested by Hakim (1997), results of aforemen-

[29] P-value= 0.02.
[30] P-value= 0.03.
[31] P-value= 0.00.
[32] P-value= 0.00.

Figure 4.8: Part-time and full-time wage regressions for (a) low-educated and (b) high-educated men

(a) Estimated coefficients low-educated



(b) Estimated coefficients high-educated

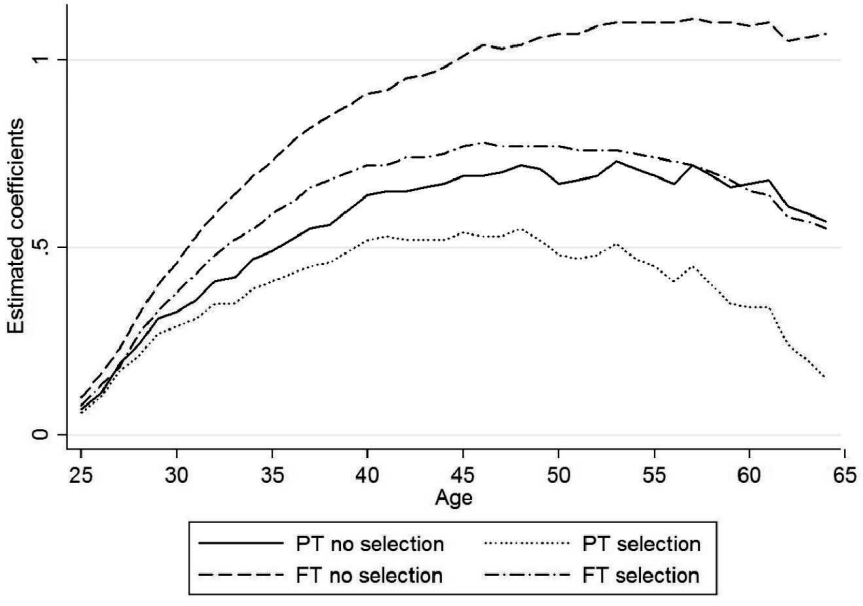
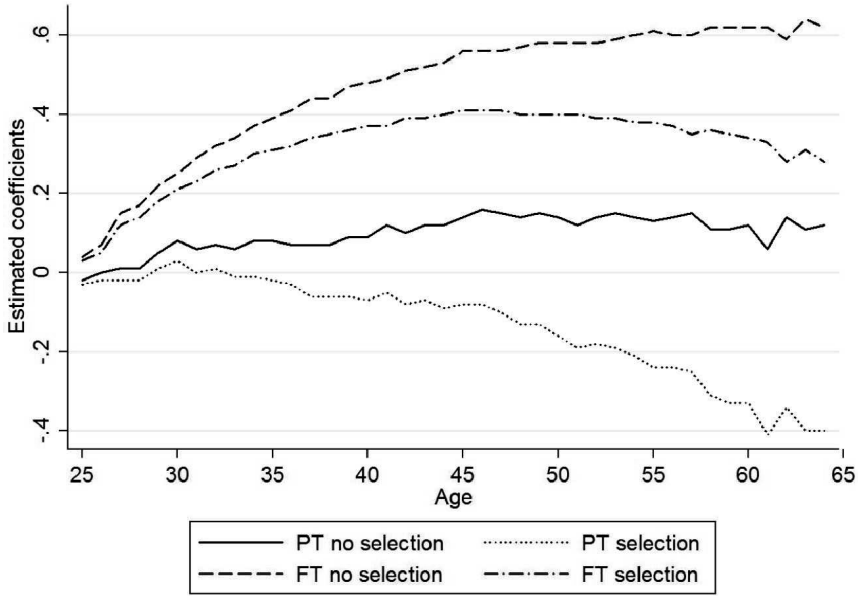
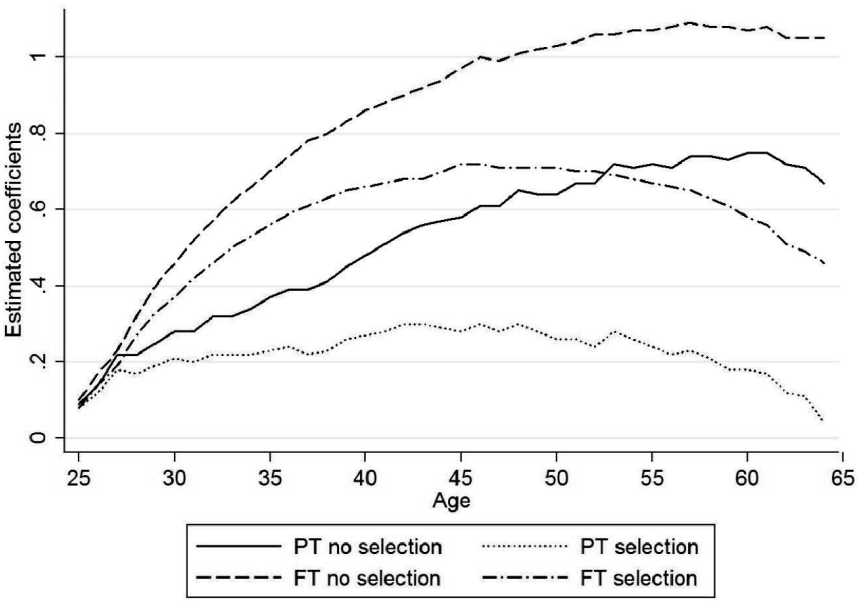


Figure 4.9: Part-time and full-time wage regressions for (a) low-educated and (b) high-educated women

(a) Estimated coefficients low-educated



(b) Estimated coefficients high-educated



tioned studies are likely to be biased as most studies on the part-time wage penalty are based on cross-sectional data and do not control for selection.[33]

Compared to these earlier studies, we take into account selection effects into part-time and full-time work. This is important as Aaronson and French (2004) and Casanova (2013) also find substantial part-time wage penalties for men (25% and 34% of a full-time wage respectively) approaching the state-pension eligible age for who part-time employment often functions as a partial retirement route. Dustmann and Schmidt (2000) do take into account selection into full-time and part-time employment when calculating wage differentials. However, their main interest lies in the wage differential between native- and migrant women and estimate these wage differentials for women in full-time and part-time employment.

Using the results in figures 4.8 and 4.9 we can test the existence of a part-time wage penalty by testing whether all differences in the age coefficients of part-time and full-time age dummies are jointly equal to zero.[34] Among high-educated men, the difference between estimated age coefficients of the part-time and full-time model is not jointly significantly different from zero meaning that we do not find a significant part-time wage penalty over the life-cycle among high-educated men.[35] We do find joint significance of the difference between estimated age coefficients for low-educated men which suggests the existence of a part-time wage penalty.[36] The part-time wage penalty is present among both low-educated[37] and high-educated[38] women.

To get an idea about the magnitude of the part-time wage penalty, we use simulation to calculate the part-time wage penalty for the mean using the following procedure: 1) we derive the unobserved heterogeneity α_i for each person, 2) we calculate the average of α_i and every variable in vector x_{it} assuming $u_{it} = 0$, 3) we set $\lambda_{1itm} = 0$, $\lambda_{2itm} = 0$, $\lambda_{3itm} = 0$ and $\lambda_{4itm} = 0$ since λ_{1itm}, λ_{2itm}, λ_{3itm} and λ_{4itm} are included in the regression to obtain

[33] Ermisch and Wright (1993) do correct for selection with cross-sectional data.
[34] $H_0 : \beta_{PT,25} - \beta_{FT,25} = ... = \beta_{PT,64} - \beta_{FT,64} = 0$.
[35] P-value= 0.45.
[36] P-value= 0.03.
[37] P-value= 0.00.
[38] P-value= 0.00.

correct estimates of β. The selection correction terms are, however, of no relevance in this simulation excercise. 4) We predict the full-time and part-time wage for the average person and 5) we calculate the differences of full-time and part-time wages for the mean and the associated variance using bootstrap.[39] The interpretation of this method is that we calculate the differences for the average person (in terms of observed and unobserved characteristics) who worked either full-time or part-time. Results are shown in table 4.3.

Table 4.3 shows the simulated part-time and full-time wages for the mean as well as the absolute difference and the relative difference (the part-time wage penalty).[40] We find that the part-time wage penalties of -16% and 14% are not statistically different from zero[41] for low-educated and high-educated men respectively (see table 4.3). Correcting for selection, career breaks and education, the size of the part-time wage penalty is negligible over the life-cycle among men. For low-educated and high-educated women we find a significant part-time wage penalty of 30% and 34% respectively (see table 4.3).

These results, ofcourse, depend on the age that is used in the simulation exercise. We use the average age that is observed in our data. Using a lower (higher) age is likely to give a smaller (larger) part-time wage penalty as the differences between full-time and part-time wages increases over the life-cycle because of cumulative effects in experience.[42] Manning and Robinson (2004), Hirsch (2005) and Russo and Hassink (2008) find that the part-time wage penalty is small or absent at the start of a career but develops over the life-cycle. This can be explained by the lower experience among part-time workers as well as by the observed lower incidence of promotions among part-time workers compared to full-time workers (Russo and Hassink 2008). These two effects tend to accumulate over

[39] Using 1500 replications.

[40] Please note that observed- and unobserved characteristics are averaged within gender and not over gender. This implies that part-time and full-time wages can be compared within gender but not between gender. Only the relative wage penalty can be compared between gender.

[41] The variance of this penalty is relatively large and therefore the difference in wage is insignificantly different from zero.

[42] In this simulation, we assume people to work either part-time of full-time during their whole working life.

Table 4.3: Part-time wage penalty for the mean for educational levels

	Men				Women			
	Low-educated		*High-educated*		*Low-educated*		*High-educated*	
	Mean	Variance	Mean	Variance	Mean	Variance	Mean	Variance
$\hat{\bar{y}}_{PT}$	47,225	12,110	48,970	7,310	29,215	4,027	36,717	3,286
$\hat{\bar{y}}_{FT}$	40,696	1,420	56,690	1,489	41,626	1,621	57,774	1,487
Absolute difference	-6,529	12,286	7,719	7,467	12,411	4,430	21,056	3,564
Relative difference	-16%		14%		30%		34%	

the life-cycle which explains the increasing gap between full-time and part-time wages (Russo and Hassink 2008). The fact that we only find a part-time wage penalty among women may be explained by compensation for the ability to combine work with care (Boeri and Van Ours 2008).

Sensitivity analyses 4.4.5

To determine the robustness of our results we perform two sensitivity analyses. Firstly, we analyze the consequences of increasing the number of part-time employment categories in the selection equation. Secondly, we discuss the possible endogeneity of the careeer breaks and the effect on the conclusions.

Increasing part-time employment categories

The paper argues that adding additional information regarding the intensive margin of participation is important in estimating wage profiles as it gives more information regarding unobserved characteristics that remain unnoted in selection correction models that only take into account the extensive margin. To prove this, we also compare our baseline results ($J = 4$) with an extended model with more part-time employment categories ($J = 7$).

We increase the number of ordered categories in the selection equation to $J = 7$: 1) no participation, 2) full-time factor between 0% and 20%, 3) full-time factor between 20% and 40%, 4) full-time factor between 40%

and 60%, 5) full-time factor between 60% and 80%, 6) full-time factor between 80% and 100%, and 7) working full-time (100%). The percentage of men observed in these categories is 26%, 0.4%, 1%, 2%, 2%, 10% and 59% respectively. The percentage of women observed in these categories is 42%, 1%, 5%, 12%, 12%, 12% and 16% respectively.

Figure 4.10 shows the estimation results when we take into account 7 working hours categories in stead of 4 in part-time and full-time wage equations.

The estimation results are highly comparable for the full-time wage profile. We observe that using $J = 7$ in stead of $J = 4$ in the first-stage causes the part-time wage profile to increase for men and decrease for women. So, the number of hours categories taken into account does matter for the second-stage wage profiles. For future research, we would like to increase J as long as there are a sufficient amount of observations in each j.

Career breaks

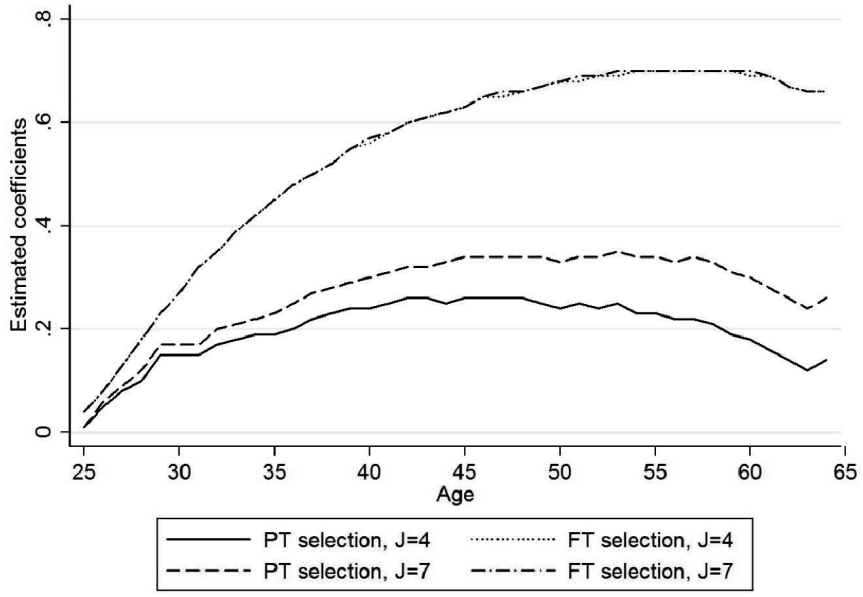
Since perons with a low-wage potential may also be subject to a career break more often (because of unemplyment, for example), the coefficients of the effects of career breaks on wages should be interpreted with caution. To check that the possible endogeity of the career breaks does not affect the main conclusions of the paper, we estimated a model excluding career break variables. Comparing the estimated wage profiles without career breaks to the model including career break variables indicates that the estimated age coefficients are highly similar. Therefore, we conclude that the inclusion of career break variables does not affect the main conclusions of the paper despite the possible endogenity of the variables.

4.5 Conclusion

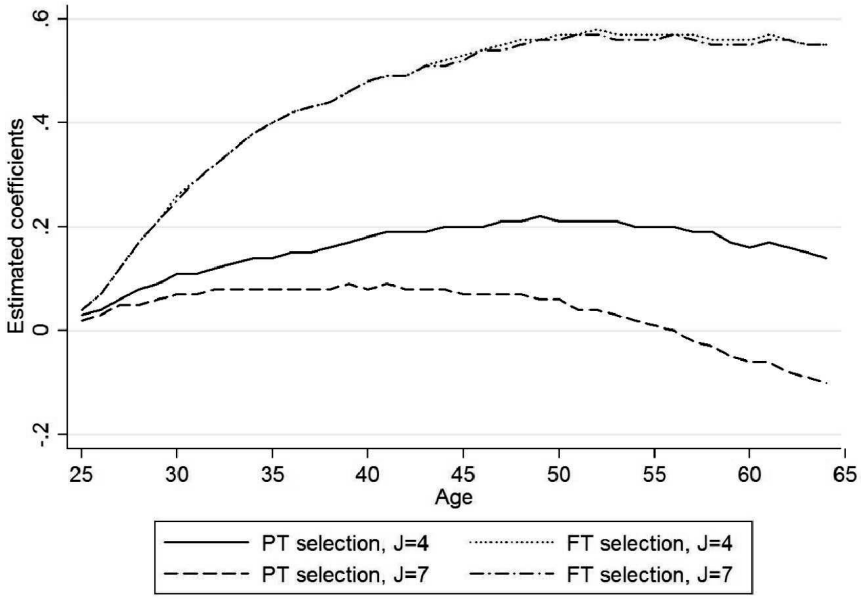
To gain insight in consumption and savings behavior over the life-cycle and to assess the adequacy of retirement savings, it is important to model life-cycle earnings as labor income usually is the primary source of income

Figure 4.10: Ordered selection correction regressions with J=7 for men (a) and women (b)

(a) Estimated coefficients men



(b) Estimated coefficients women



(Scholz et al. 2006). Also, earnings are often directly related to the accumulation of (occupational) pension rights over the life-cycle. Conclusions regarding the adequacy of retirement savings depend on a correct specification of the wage equation. However, most life-cycle models neglect the selection into wages while selection into work is likely to be nonrandom (Casanova 2013).

This paper proposes a new estimator to estimate life-cycle wage profiles using a panel data sample selection model that takes into account information about part-time and full-time work. Our proposed new estimator is an extension of the method proposed by Rochina-Barrachina (1999). Rochina-Barrachina (1999) proposes a binary selection equation to correct for selection into work. We propose an ordered selection equation to correct for selection into work and the number of hours of work simultaneously. By taking into account the number of hours that people work, extra information is available about unobserved characteristics in the wage equation. This is especially relevant for the analysis of wages over the life-cycle as women who work full-time or have a large part-time job during the upbringing of young children may have different unobserved characteristics compared to women in small part-time jobs. Also, men who retire partially may be a selective group with different observed and unobserved characteristics than men who do not retire gradually. The estimator proposed in this paper is applied to estimate life-cycle wage profiles and to analyze selection into part-time and full-time employment as well as the part-time wage penalty over the life-cycle conditional on possible career breaks.

Using the binary selection correction proposed by Rochina-Barrachina (1999) we find negative selection into work over the life-cycle among men and women. However, adding information regarding hours decisions by using the ordered selection correction proposed in this paper we find positive selection into work over the life-cycle among men and less substantial negative selection among women. This difference indicates that it is important to take into account both participation and hours decisions to account for unobserved heterogeneity in wages. This is strengthened by our analysis that increases the number of hours categories. The positive selection suggests that persons with more affluent observed and unobserved

characteristics tend to work over the life-cycle whereas persons with less beneficial observed and unobserved characteristics are less likely to be employed. Career breaks have a substantial negative effect on life-cycle wages with an effect of 11% (men) and 7% (women) of the first year which increases up to 21% (men) and 17% (women) from the third year.

Education-specific life-cycle wage profiles for low- and high-educated persons show that both selection effects and part-time wage penalties may differ between these groups. Among men, we generally find positive selection. The part-time wage penalty over the life-cycle is not significantly different from zero for low- and high-educated men. Positive selection into part-time and full-time employment is found among both low-educated and high-educated women. Estimating the life-cycle wage profiles separately for low- and high-educated women substantially gives an average part-time wage penalty of 30% and 34% for low- and high-educated women respectively.

The paper shows the existence of selection into work over the life-cycle for both men and women. This has consequences for applications in which estimating life-cycle earnings processes are crucial. The extra information regarding unobserved individual heterogeneity that the proposed estimator incorporates in estimating life-cycle wages makes it well-applicable to models that depend on life-cycle earnings processes such as life-cycle models of consumption and savings (Scholz et al. 2006), earnings inequality (Cappellari 2004) and microsimulation models of future pension accumulation (Borella 2004).

4.A Derivation of the selection terms

To derive the selection correction terms based on the ordered selection equation, we need to calculate the first moment of the doubly truncated trivariate normal distribution. To calculate this first moment, we follow the approach of Manjunath and Wilhelm (2012) using the moment generating function (*m.g.f.*) of a doubly truncated multivariate normal distribution. The m.g.f. of a doubly truncated trivariate normal distribution[43] that is truncated in **a** and **b** yields (see equation 5 in Manjunath and Wilhelm 2012)

$$m(\boldsymbol{t}) = e^{\frac{1}{2}\boldsymbol{t'\Sigma t}} \int_{\mathbf{a^*}}^{\mathbf{b^*}} \phi_{\alpha\boldsymbol{\Sigma}}(\boldsymbol{x}) d\boldsymbol{x} \tag{4.33}$$

where $\boldsymbol{x}$ is a three-dimensional normal density $\boldsymbol{x'} = \begin{bmatrix} x_1 & x_2 & x_3 \end{bmatrix}$ with location parameter $\boldsymbol{\mu} = 0$ and covariance matrix $\boldsymbol{\Sigma}$. $\phi_{\alpha\boldsymbol{\Sigma}}(\boldsymbol{x})$ is the trivariate normal distribution defined as[44]

$$\phi_{\alpha\boldsymbol{\Sigma}}(\boldsymbol{x}) = \frac{1}{\alpha(2\pi)^{3/2}|\Sigma|^{1/2}} exp\left(-\frac{1}{2}\boldsymbol{x'\Sigma}^{-1}\boldsymbol{x}\right) d\boldsymbol{x} \tag{4.34}$$

with α being the fraction of the multivariate normal distribution after truncation, $\mathbf{a^{*\prime}} = \begin{bmatrix} a_1^* & a_2^* & a_3^* \end{bmatrix}$ and $\mathbf{b^{*\prime}} = \begin{bmatrix} b_1^* & b_2^* & b_3^* \end{bmatrix}$, such

$$a_1^* = a_1 - \boldsymbol{\Sigma t} \tag{4.35}$$
$$a_2^* = a_2 - \boldsymbol{\Sigma t} \tag{4.36}$$
$$a_3^* = a_3 - \boldsymbol{\Sigma t} \tag{4.37}$$
$$b_1^* = b_1 - \boldsymbol{\Sigma t} \tag{4.38}$$
$$b_2^* = b_2 - \boldsymbol{\Sigma t} \tag{4.39}$$

[43] $[(u_{it} - u_{it-m}), \mu_{it-m}, \mu_{it}]$ are assumed to be trivariate normally distributed conditional on x_i and z_i.

[44] See Muthen (1990) for a derivation of the doubly truncated bivariate normal distribution and Tallis (1961) for a derivation of the singly truncated multivariate normal distribution.

$$b_3^* = b_3 - \boldsymbol{\Sigma t} \tag{4.40}$$

with

$$\boldsymbol{t'} = \begin{bmatrix} t_1 & t_2 & t_3 \end{bmatrix} \tag{4.41}$$

We are interested in $\mathrm{E}(x_1 | a_2 \leq x_2 \leq b_2, a_3 \leq x_3 \leq b_3)$ with $a_1 = -\infty$ and $b_1 = \infty$. Therefore, we need to take the partial derivative of the m.g.f. (equation 4.33) with respect to t_1. Using the chain rule for calculating derivatives gives

$$\frac{\partial m(\boldsymbol{t})}{\partial t_1} = e^{\frac{1}{2}\boldsymbol{t'\Sigma t}} \frac{\partial \Phi_{\alpha\boldsymbol{\Sigma}}}{\partial t_1} + \Phi_{\alpha\boldsymbol{\Sigma}} \frac{\partial e^{\frac{1}{2}\boldsymbol{t'\Sigma t}}}{\partial t_1} \tag{4.42}$$

Inserting the trivariate normal distribution and applying *Leibniz's rule* for differentiation under the integral sign we get

$$\begin{aligned}
\frac{\partial \phi_{\alpha\boldsymbol{\Sigma}}}{\partial t_1} &= \frac{\partial}{\partial t_1} \int_{a_1^*}^{b_1^*} \int_{a_2^*}^{b_2^*} \int_{a_3^*}^{b_3^*} \phi_{\alpha\boldsymbol{\Sigma}}(x_1, x_2, x_3) dx_3 dx_2 dx_1 = \\
&- \int_{a_2^*}^{b_2^*} \int_{a_3^*}^{b_3^*} \phi_{\alpha\boldsymbol{\Sigma}}(b_1^*, x_2, x_3) dx_3 dx_2 + \int_{a_2^*}^{b_2^*} \int_{a_3^*}^{b_3^*} \phi_{\alpha\boldsymbol{\Sigma}}(a_1^*, x_2, x_3) dx_3 dx_2 \\
&- \sigma_{12} \int_{a_1^*}^{b_1^*} \int_{a_3^*}^{b_3^*} \phi_{\alpha\boldsymbol{\Sigma}}(x_1, b_2^*, x_3) dx_3 dx_1 + \sigma_{12} \int_{a_1^*}^{b_1^*} \int_{a_3^*}^{b_3^*} \phi_{\alpha\boldsymbol{\Sigma}}(x_1, a_2^*, x_3) dx_3 dx_1 \\
&- \sigma_{13} \int_{a_1^*}^{b_1^*} \int_{a_2^*}^{b_2^*} \phi_{\alpha\boldsymbol{\Sigma}}(x_1, x_2, b_3^*) dx_2 dx_1 + \sigma_{13} \int_{a_1^*}^{b_1^*} \int_{a_2^*}^{b_2^*} \phi_{\alpha\boldsymbol{\Sigma}}(x_1, x_2, a_3^*) dx_2 dx_1
\end{aligned} \tag{4.43}$$

and

$$\frac{\partial e^{\frac{1}{2}\boldsymbol{t}'\boldsymbol{\Sigma}\boldsymbol{t}}}{\partial t_1} = e^{\frac{1}{2}\boldsymbol{t}'\boldsymbol{\Sigma}\boldsymbol{t}} \sum_{k=1}^{3} \sigma_{1k} t_k \tag{4.44}$$

Evaluating the derivative $\frac{\partial m(\boldsymbol{t})}{\partial t_1}$ at $\boldsymbol{t} = 0$ in order to compute the first moment ($\mathrm{E}(X_1)$) gives $\frac{\partial e^{\frac{1}{2}\boldsymbol{t}'\boldsymbol{\Sigma}\boldsymbol{t}}}{\partial t_1} = 0$, $a_1^* = a_1$, $a_2^* = a_2$, $a_3^* = a_3$, $b_1^* = b_1$, $b_2^* = b_2$ and $b_3^* = b_3$ such that

$$\begin{aligned}
\alpha \mathrm{E}(X_1) = & \left. \frac{\partial}{\partial t_1} \right|_{t_1=0} \int_{a_1}^{b_1} \int_{a_2}^{b_2} \int_{a_3}^{b_3} \phi_{\boldsymbol{\Sigma}}(x_1, x_2, x_3) dx_3 dx_2 dx_1 = \\
& - \int_{a_2}^{b_2} \int_{a_3}^{b_3} \phi_{\boldsymbol{\Sigma}}(b_1, x_2, x_3) dx_3 dx_2 + \int_{a_2}^{b_2} \int_{a_3}^{b_3} \phi_{\boldsymbol{\Sigma}}(a_1, x_2, x_3) dx_3 dx_2 \\
& - \sigma_{12} \int_{a_1}^{b_1} \int_{a_3}^{b_3} \phi_{\boldsymbol{\Sigma}}(x_1, b_2, x_3) dx_3 dx_1 + \sigma_{12} \int_{a_1}^{b_1} \int_{a_3}^{b_3} \phi_{\boldsymbol{\Sigma}}(x_1, a_2, x_3) dx_3 dx_1 \\
& - \sigma_{13} \int_{a_1}^{b_1} \int_{a_2}^{b_2} \phi_{\boldsymbol{\Sigma}}(x_1, x_2, b_3) dx_2 dx_1 + \sigma_{13} \int_{a_1}^{b_1} \int_{a_2}^{b_2} \phi_{\boldsymbol{\Sigma}}(x_1, x_2, a_3) dx_2 dx_1
\end{aligned} \tag{4.45}$$

Since $a_1 = -\infty$ and $b_1 = \infty$, the terms $-\int_{a_2}^{b_2} \int_{a_3}^{b_3} \phi_{\alpha\boldsymbol{\Sigma}}(b_1, x_2, x_3) dx_3 dx_2$ and $\int_{a_2}^{b_2} \int_{a_3}^{b_3} \phi_{\alpha\boldsymbol{\Sigma}}(a_1, x_2, x_3) dx_3 dx_2$ are zero.

$$\begin{aligned}
\alpha \mathrm{E}(X_1) = & -\sigma_{12} \phi(b_2) \int_{a_1}^{b_1} \int_{a_3}^{b_3} \phi \left(\frac{x_1 - \rho_{12} b_2}{\sqrt{1-\rho_{12}^2}}, \frac{x_3 - \rho_{23} b_2}{\sqrt{1-\rho_{23}^2}}, \rho_{13} \right) dx_3 dx_1 \\
& + \sigma_{12} \phi(a_2) \int_{a_1}^{b_1} \int_{a_3}^{b_3} \phi \left(\frac{x_1 - \rho_{12} a_2}{\sqrt{1-\rho_{12}^2}}, \frac{x_3 - \rho_{23} a_2}{\sqrt{1-\rho_{23}^2}}, \rho_{13} \right) dx_3 dx_1 \\
& - \sigma_{13} \phi(b_3) \int_{a_1}^{b_1} \int_{a_2}^{b_2} \phi \left(\frac{x_1 - \rho_{13} b_3}{\sqrt{1-\rho_{13}^2}}, \frac{x_2 - \rho_{23} b_3}{\sqrt{1-\rho_{23}^2}}, \rho_{12} \right) dx_2 dx_1
\end{aligned}$$

$$+\sigma_{13}\phi(a_3)\int_{a_1}^{b_1}\int_{a_2}^{b_2}\phi\left(\frac{x_1-\rho_{13}a_3}{\sqrt{1-\rho_{13}^2}},\frac{x_2-\rho_{23}a_3}{\sqrt{1-\rho_{23}^2}},\rho_{12}\right)dx_2dx_1 \tag{4.46}$$

Here, ρ_{12}, ρ_{13} and ρ_{23} are the correlation coefficients between. We can rewrite equation (4.46) (see Manjunath and Wilhelm 2012) such that

$$\begin{aligned}\alpha \mathrm{E}(X_1) = &-\sigma_{12}\phi(b_2)\left[\Phi\left(\frac{b_3-\rho_{23}b_2}{\sqrt{1-\rho_{23}^2}}\right)-\Phi\left(\frac{a_3-\rho_{23}b_2}{\sqrt{1-\rho_{23}^2}}\right)\right]\\ &+\sigma_{12}\phi(a_2)\left[\Phi\left(\frac{b_3-\rho_{23}a_2}{\sqrt{1-\rho_{23}^2}}\right)-\Phi\left(\frac{a_3-\rho_{23}a_2}{\sqrt{1-\rho_{23}^2}}\right)\right]\\ &-\sigma_{13}\phi(b_3)\left[\Phi\left(\frac{b_2-\rho_{23}b_3}{\sqrt{1-\rho_{23}^2}}\right)-\Phi\left(\frac{a_2-\rho_{23}b_3}{\sqrt{1-\rho_{23}^2}}\right)\right]\\ &+\sigma_{13}\phi(a_3)\left[\Phi\left(\frac{b_2-\rho_{23}a_3}{\sqrt{1-\rho_{23}^2}}\right)-\Phi\left(\frac{a_2-\rho_{23}a_3}{\sqrt{1-\rho_{23}^2}}\right)\right]\end{aligned} \tag{4.47}$$

such that the first moment of X_1 in the doubly truncated trivariate normal distribution becomes

$$\mathrm{E}(X_1) = \frac{\alpha \mathrm{E}(X_1)}{\alpha} = \frac{\alpha \mathrm{E}(X_1)}{\Phi_2(b_2,b_3,\rho_{23})-\Phi_2(a_2,a_3,\rho_{23})} \tag{4.48}$$

with $\Phi_2(.)$ being the bivariate normal distribution. $\Phi_2(b_2,b_3,\rho_{23})-\Phi_2(a_2,a_3,\rho_{23})$ is the fraction of the trivariate normal distribution after truncation, e.g. a normalization of the terms in equation (4.47). The four terms in equation (4.47) are the four selection correction terms where $-\sigma_{12}$, σ_{12}, $-\sigma_{13}$ and σ_{13} are the coefficients to be estimated in the wage equation ($\pi_1 tm$, $\pi_2 tm$, $\pi_3 tm$ and $\pi_4 tm$ in equation (4.20)). a_2, a_3, b_2, b_3 and ρ_{23} are to be estimated in the first-stage selection equation (subscript 2 denoted as t and 3 denoted as $t-m$ in equation (4.15)). $a_2 = H_{it}/\sigma_t$, $b_2 = G_{it}/\sigma_t$, $a_3 = H_{it-m}/\sigma_{t-m}$, $b_3 = G_{it-m}/\sigma_{t-m}$, $\rho_{23} = \rho_{tm}$ and H_{it}, H_{it-m}, G_{it} and G_{it-m} defined as in equations (4.16) to (4.19). With x_1 being the error term of the wage

equation and x_2 and x_3 being the error terms of the selection equations we get

$$\begin{aligned} &\mathrm{E}(x_1|a_2 \leq x_2 \leq b_2, a_3 \leq x_3 \leq b_3) = \\ &\mathrm{E}(u_{it} - u_{it-m}|G_{it-m} \leq \mu_{it-m} \leq H_{it-m}, G_{it} \leq \mu_{it} \leq H_{it}) \end{aligned} \tag{4.49}$$

4.B First-stage regression results

Since we estimate the first-stage bivariate ordered probit model for every combination of t and $t-m$ for $t = \{2002, ..., 2011\}$ and $m = \{1, .., 10\}$, we end up having 55 different models to construct the selection correction terms λ_{1itm}, λ_{2itm}, λ_{3itm}, λ_{4itm}. We report the estimation results for the combination 2002 and 2001 and the combination 2011 and 2010 in tables 4.4-4.5 for men and women respectively. Apart from the sign and significance, the reported coefficients have no direct interpretation.

Age-effects are with respect to the baseline of age 25. We estimate the selection equations for persons born no later than 1980. As a consequence, the baseline of age-effects shifts from $t = 2006$. Coefficients should be interpreted with respect to the estimated parameters δ_{1t}, δ_{2t}, δ_{3t}, δ_{1t-m}, δ_{2t-m} and δ_{3t-m} that indicate the thresholds between the $J = 4$ labor supply categories for time t and t respectively. ρ_{tm} indicates the correlation between the error terms at time t and $t-m$ in the selection equation.

The estimation results (see tables 4.4-4.5) show that the likelihood of participation, and especially full-time work, decreases with age. This is true for both men and women, although the decline over age is relatively smaller for men than for women. Especially in the earlier years. Also, the first-stage regressions suggest that life-cycle participation decisions changed over time. The decrease in the probability to participate over the life-cycle is much larger in 2002 than in 2011. This suggests that labor force participation over the life-cycle increased over time. For women, this can also be concluded from the part-time factor in table 4.1. For men, the differences in labor force participation seem to be concentrated at the end of the career.

We find that immigrant men are significantly and substantially less likely to work (full-time). Being married is positively related to the labor force attachment among men, but only in the later years. The number of children decreases the labor force participation while the effect of having a partner of age 62 or older is usually not significant. Being married or divorced increases labor force participation among men. For women, we find a significant and substantial negative association between the labor force participation, children and having a partner of age 62 or older. Furthermore, a woman is less likely to work (full-time) if she is an immigrant, married or widowed.

A final interesting result from the first-stage equations are the estimates of ρ_{tm}. We find that $\hat{\rho}_{tm}$ decreases for higher m (e.g. the correlation between the error terms of the selection equation decreases if the period between the choices is longer). $\hat{\rho}_{tm}$ is rather constant over time in both the bivariate ordered probit model and the bivariate probit model, but $\hat{\rho}_{tm}$ is generally higher in the bivariate probit models than in the bivariate ordered probit models. However, estimating the wage model with bivariate ordered probit selection correction while using $\hat{\rho}_{tm}$ from the bivariate probit model gives highly similar results as the estimates using $\hat{\rho}_{tm}$ from the bivariate ordered probit model as presented in figure 4.7 (dashed line). Hence, the difference in the estimates of the wage profiles of the model with binary selection correction (dotted line in figure 4.7) and the ordered selection correction (dashed line figure 4.7) is not a consequence of the difference in the correlation of the error terms estimated by the two approaches.

Table 4.4: Estimation results first-stage selection equation, men

	$t = 2002$		$t - m = 2001$		$t = 2011$		$t - m = 2010$	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Age 25	ref.		ref.					
Age 26	0.00	0.03	-0.09	0.07				
Age 27	-0.12*	0.07	-0.06	0.06				
Age 28	-0.11	0.07	-0.07	0.06				
Age 29	-0.10	0.07	0.13*	0.07				
Age 30	0.05	0.07	0.02	0.06			ref.	
Age 31	-0.06	0.07	-0.04	0.06	ref.		0.11	0.08
Age 32	-0.08	0.06	-0.10*	0.06	0.11	0.07	-0.16**	0.07
Age 33	-0.17***	0.06	-0.04	0.06	-0.15***	0.07	-0.05	0.07
Age 34	-0.09	0.07	-0.07	0.06	-0.05	0.07	-0.03	0.07
Age 35	-0.16**	0.07	-0.12*	0.06	0.00	0.07	-0.22***	0.07
Age 36	-0.16**	0.07	-0.15**	0.06	-0.21***	0.07	-0.12*	0.07
Age 37	-0.24***	0.07	-0.10*	0.06	-0.12*	0.07	-0.16**	0.07
Age 38	-0.19***	0.06	-0.09	0.06	-0.17**	0.07	-0.11	0.07
Age 39	-0.16**	0.06	-0.07	0.07	-0.08	0.07	-0.19***	0.07
Age 40	-0.16**	0.07	-0.17***	0.06	-0.20***	0.07	-0.27***	0.07
Age 41	-0.22***	0.07	-0.20***	0.06	-0.22***	0.07	-0.26***	0.07
Age 42	-0.28***	0.07	-0.16**	0.06	-0.22***	0.07	-0.27***	0.07
Age 43	-0.23***	0.07	-0.27***	0.07	-0.27***	0.07	-0.22***	0.07
Age 44	-0.27***	0.07	-0.24***	0.07	-0.24***	0.07	-0.31***	0.07
Age 45	-0.28***	0.07	-0.19***	0.07	-0.31***	0.07	-0.31***	0.07
Age 46	-0.24***	0.07	-0.24***	0.07	-0.31***	0.07	-0.23***	0.07
Age 47	-0.31***	0.07	-0.23***	0.07	-0.20***	0.07	-0.15***	0.07
Age 48	-0.29***	0.07	-0.28***	0.07	-0.16**	0.07	-0.29***	0.07
Age 49	-0.39***	0.07	-0.32***	0.07	-0.32***	0.07	-0.31***	0.07
Age 50	-0.38***	0.07	-0.33***	0.07	-0.32***	0.07	-0.28***	0.07
Age 51	-0.39***	0.07	-0.29***	0.07	-0.29***	0.07	-0.29***	0.07
Age 52	-0.34***	0.07	-0.22***	0.07	-0.32***	0.07	-0.38***	0.07
Age 53	-0.36***	0.07	-0.38***	0.07	-0.33***	0.07	-0.25***	0.08
Age 54	-0.47***	0.07	-0.48***	0.07	-0.27***	0.08	-0.35***	0.07
Age 55	-0.55***	0.07	-0.46***	0.06	-0.40***	0.07	-0.38***	0.07
Age 56	-0.59***	0.07	-0.67***	0.07	-0.39***	0.08	-0.51***	0.07
Age 57	-0.76***	0.08	-0.77***	0.07	-0.52***	0.07	-0.45***	0.07
Age 58	-0.89***	0.07	-0.93***	0.07	-0.51***	0.07	-0.51***	0.07
Age 59	-1.05***	0.07	-1.21***	0.08	-0.59***	0.07	-0.62***	0.08
Age 60	-1.37***	0.08	-1.48***	0.08	-0.67***	0.08	-0.72***	0.08
Age 61	-1.78***	0.08	-1.83***	0.08	-0.82***	0.08	-0.96***	0.08
Age 62	-2.11***	0.09	-2.13***	0.09	-1.09***	0.08	-1.43***	0.08
Age 63	-2.35***	0.10	-2.45***	0.10	-1.59***	0.08	-1.69***	0.08
Age 64	-2.53***	0.10			-1.83***	0.08		
Number of children	-0.02*	0.01	-0.02**	0.01	-0.02***	0.01	-0.02**	0.01
Single	ref.		ref.		ref.		ref.	
Married	-0.02	0.03	0.01	0.03	0.19***	0.04	0.19***	0.04
Divorced	-0.09*	0.05	-0.08*	0.05	0.17***	0.05	0.18***	0.06
Widowed	0.12	0.17	-0.04	0.15	0.10	0.14	0.07	0.14
Immigrant	-0.49***	0.03	-0.49***	0.03	-0.46***	0.03	-0.47***	0.03
Partner 62+	0.11*	0.06	0.02	0.07	-0.02	0.04	0.04	0.05
F-test $\overline{z_i}$	220.86***				64.67***			
δ_{1t}	-0.90***	0.04			-0.84***	0.05		
δ_{2t}	-0.86***	0.04			-0.79***	0.05		
δ_{3t}	-0.60***	0.04			-0.37***	0.05		
δ_{1t-m}	-0.84***	0.04			-0.84***	0.05		
δ_{2t-m}	-0.79***	0.04			-0.78***	0.05		
δ_{3t-m}	-0.54***	0.04			-0.41***	0.05		
ρ_{tm}	0.97***	0.02			0.97***	0.02		
Obs.	24,129				20,886			
Chi^2	3,247				2,088			

Table 4.5: Estimation results first-stage selection equation, women

	$t = 2002$		$t - m = 2001$		$t = 2011$		$t - m = 2010$	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Age 25	ref.		ref.					
Age 26	-0.02	0.05	0.12*	0.06				
Age 27	0.04	0.07	0.04	0.06				
Age 28	-0.04	0.07	-0.03	0.06				
Age 29	-0.13*	0.07	-0.07	0.06				
Age 30	-0.13*	0.06	-0.12**	0.06			1.19***	0.08
Age 31	-0.18***	0.06	-0.19***	0.06	0.10	0.06	1.14***	0.08
Age 32	-0.31***	0.06	-0.18***	0.06	0.08	0.06	1.04***	0.08
Age 33	-0.30***	0.06	-0.30***	0.06	-0.03	0.06	1.04***	0.08
Age 34	-0.38***	0.06	-0.31***	0.06	ref.		1.05***	0.08
Age 35	-0.37***	0.06	-0.35***	0.06	0.02	0.06	1.00***	0.08
Age 36	-0.44***	0.06	-0.39***	0.06	-0.06	0.06	0.99***	0.08
Age 37	-0.46***	0.06	-0.39***	0.06	-0.05	0.06	0.94***	0.08
Age 38	-0.44***	0.06	-0.33***	0.06	-0.11*	0.06	0.84***	0.08
Age 39	-0.45***	0.06	-0.37***	0.06	-0.17***	0.06	0.97***	0.08
Age 40	-0.44***	0.07	-0.45***	0.06	-0.04	0.06	0.86***	0.08
Age 41	-0.50***	0.07	-0.37***	0.06	-0.18***	0.06	0.88***	0.07
Age 42	-0.42***	0.07	-0.33***	0.06	-0.14***	0.06	0.87***	0.08
Age 43	-0.39***	0.07	-0.38***	0.06	-0.16***	0.06	0.93***	0.08
Age 44	-0.42***	0.07	-0.52***	0.06	-0.08	0.06	0.84***	0.08
Age 45	-0.58***	0.07	-0.51***	0.06	-0.16***	0.06	0.86***	0.08
Age 46	-0.56***	0.07	-0.54***	0.07	-0.20***	0.06	0.91***	0.08
Age 47	-0.61***	0.07	-0.65***	0.06	-0.13**	0.06	0.86***	0.07
Age 48	-0.66***	0.07	-0.68***	0.07	-0.13**	0.06	0.88***	0.08
Age 49	-0.73***	0.07	-0.74***	0.07	-0.17***	0.07	0.77***	0.08
Age 50	-0.80***	0.07	-0.75***	0.07	-0.26***	0.06	0.75***	0.08
Age 51	-0.81***	0.07	-0.92***	0.07	-0.27***	0.06	0.82***	0.08
Age 52	-0.94***	0.07	-0.99***	0.07	-0.22***	0.07	0.75***	0.08
Age 53	-1.01***	0.07	-1.04***	0.07	-0.28***	0.07	0.67***	0.08
Age 54	-1.12***	0.07	-1.24***	0.07	-0.40***	0.07	0.60***	0.08
Age 55	-1.26***	0.07	-1.24***	0.07	-0.42***	0.07	0.58***	0.08
Age 56	-1.30***	0.08	-1.29***	0.08	-0.49***	0.07	0.49***	0.08
Age 57	-1.34***	0.08	-1.45***	0.08	-0.55***	0.07	0.39***	0.08
Age 58	-1.55***	0.09	-1.57***	0.09	-0.63***	0.07	0.24***	0.08
Age 59	-1.65***	0.09	-1.78***	0.09	-0.80***	0.07	0.31***	0.08
Age 60	-1.87***	0.10	-1.84***	0.10	-0.71***	0.08	ref.	
Age 61	-2.07***	0.12	-2.10***	0.10	-1.03***	0.08	-0.16***	0.09
Age 62	-2.34***	0.11	-2.65***	0.14	-1.16***	0.08	-0.38***	0.09
Age 63	-2.69***	0.14	-2.87***	0.18	-1.47***	0.08	-0.67***	0.09
Age 64	-3.14***	0.22			-1.84***	0.09		
Number of children	-0.08***	0.01	-0.10***	0.01	-0.08***	0.01	-0.09***	0.01
Single	ref.		ref.		ref.		ref.	
Married	-0.23***	0.04	-0.21***	0.04	0.01	0.04	0.01	0.05
Divorced	-0.09	0.06	-0.12**	0.05	0.09	0.06	0.14**	0.06
Widowed	-0.22**	0.10	-0.18**	0.09	0.00	0.08	0.07	0.10
Immigrant	-0.28***	0.03	-0.26***	0.03	-0.34***	0.03	-0.34***	0.05
Partner 62+	0.09**	0.04	0.01	0.04	-0.04	0.02	0.05	0.03
F-test $\overline{z_i}$	358.02***				215.18***			
δ_{1t}	-1.40***	0.05			-1.09***	0.05		
δ_{2t}	-1.05***	0.05			-0.72***	0.05		
δ_{3t}	-0.20***	0.05			0.38***	0.05		
δ_{1t-m}	-1.32***	0.04			-0.09	0.07		
δ_{2t-m}	-0.97***	0.04			0.28***	0.07		
δ_{3t-m}	-0.15***	0.04			1.35***	0.07		
ρ_{tm}	0.95***	0.02			0.96***	0.03		
Obs.	21,547				20,515			
Chi^2	4,819				2,686			

5 Early retirement across Europe. Does non-standard employment increase participation of older workers?

Abstract

In many European countries, the labor market participation of older workers is considerably lower than the labor market participation of prime-age workers. This study analyzes the variation in labor market withdrawal of older workers across 13 European countries over the period 1995-2008. We seek to contribute to existing macro-econometric studies by taking non-standard employment into account, by relating the empirical model more explicitly to microeconomic theory on retirement decisions and by using a two-step IV-GMM estimator to deal with endogeneity issues. The analysis leads to the conclusion that part-time employment is negatively related to labor market withdrawal of older men. This relationship is less strong among women. Additionally, we find that part-time employment at older ages does not decrease the average actual hours worked. Furthermore, the results show a positive relationship between unemployment among older workers and early retirement similar to previous studies.

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5.1 Introduction

The aging of the population is an important challenge for most of the Western welfare systems in the near future. While an increasing life expectancy will lead to higher expenditures on pensions and healthcare, there will be fewer younger people who pay taxes and contributions to finance the welfare systems. In this respect, Barr (2006) argues that *"The problem is not that people are living longer, but that they retire too early."* Pestieau (2003) argues that the financial sustainability of pension systems is substantially affected by the low participation rates of older workers. Although the participation rates of older workers have been rising for both men and women in many European countries, the participation rates are still low compared to those of prime-age workers. Furthermore, the developments in the participation rates of older workers vary considerably across countries. In Italy, the participation rate of persons aged 55-64 years increased from 29.5 to 35.5 percent between 1995 and 2008, in Austria from 30.8 to 41.9 percent and in the Netherlands from 30 to 54.7 percent. This study aims to analyze the variation in early labor market withdrawal across countries and over time.

One prominent explanation for the low participation rates of older people is that once older people are unemployed or receive disability benefits,[1] relatively few of them start working again before they reach the statutory retirement age. As a result, such social insurance programs function in practice quite often as an arrangement to smoothen the transition from work to retirement, next to formal retirement programs. Gruber and Wise (2004) and Blondal and Scarpetta (1998) indicate this importance of social insurance programs as early retirement mechanisms based on coherent country-specific micro-econometric analyses. In the last decades of the twentieth century, almost all European countries had strong disincentives to work at older ages because of such social insurance programs (Gruber and Wise 1998). Since the 1990s, many governments have started to reform welfare state institutions to reduce the disincentives to work as well as encouraging employers to maintain older workers (Casey et al. 2003, Sonnet

[1] Which may be agreed upon by employer and employee.

et al. 2014). Hence, these reforms may have contributed to the increase of participation rates of older workers across Europe.

Another class of factors that may explain the developments in participation rates of older people is the rise of non-standard forms of employment. Chen et al. (2013) observe rapid increases of non-standard employment in many OECD countries over the past decades, especially among people aged 50-64 years. The largest increase has been observed for part-time employment. As older workers generally tend to have a relatively strong preference for leisure (Kantarci and Van Soest 2008), they often take up part-time work before full retirement (Morris and Mallier 2003). Hence, part-time work provides the opportunity to retire gradually by providing a bridge between full-time employment and retirement (Cahill et al. 2006, Gustman and Steinmeier 1984, Kim and DeVaney 2005, Quinn and Kozy 1996, Ruhm 1990, 2006). Reday-Mulvey and Delsen (1996) indicate the importance of such 'bridge jobs' across OECD countries. Similarly, older people may opt for self-employment because it provides flexibility in working hours such that self-employment may be used as gradual retirement mechanism similar to part-time employment (Bruce et al. 2000, Gu 2009). Self-employment rates are found to be relatively high among the 50+ population (Zissimopoulos and Karoly 2009) and should therefore be taken into account in analyzing retirement behavior. Morris and Mallier (2003) show that the high and increasing importance of such non-standard employment opportunities in European countries can be related to the countries' patterns of labor force participation at older ages. They argue that the increases in non-standard employment are both related to voluntary decisions to decrease working hours prior to retirement and in anticipation to declining opportunities in both full-time employment (Dorn and Sousa-Poza 2010) and early retirement possibilities (Casey et al. 2003). However, whether non-standard forms of employment can explain the cross-national variation in labor market participation of older workers has not been analyzed thus far.

The amount of comparative macro-econometric work on labor participation among older workers is rather limited. Existing studies (Blake and Sangnier 2011, Blondal and Scarpetta 1998, Duval 2003, Faggio and Nickell 2007, Johnson 2000) found that generous social insurance- and

early retirement programs enhance early labor market withdrawal. Furthermore, they found that also high unemployment rates lead to lower participation rates among older workers. With respect to these empirical analyses at the macro level, we seek to make three contributions. First, we extend the analysis by taking part-time employment and self-employment into account. Second, we strengthen the theoretical underpinning of the macroeconomic empirical analysis of mutually exclusive labor market states and retirement by relating it more explicitly to microeconomic theory on retirement decisions. Third, for the analysis of panel data for 13 European countries, we use a two-step GMM estimator to deal with endogeneity issues.

The remainder of the paper is structured as follows. In section 5.2.2, we describe the theoretical framework to explain retirement decisions and we discuss the model to be estimated in section 5.2.3. Section 5.3 describes our data and variables and presents the developments in early retirement and non-standard employment across countries and over time. Section 5.4 presents our estimation results of the pooled time-series cross-country regression analyses. In section 5.5, we discuss the robustness of our results based on a variety of sensitivity analyses. Section 5.6 concludes.

5.2 Model

5.2.1 Literature review

A natural starting point for explaining individual retirement decisions is by analyzing the difference in the expected present value of retiring immediately and postponing retirement.[2] The expected present value is the utility stream, which is an arbitrary function of consumption and leisure possibilities, coming from one of the decisions, corrected for time-

[2]Such models are often solved by way of an Option Value Model (Stock and Wise 1990) or Dynamic Programming (Rust 1989). The main difference between the two approaches is the way of modeling uncertainty (Belloni 2008). However, since the actual theoretical modeling of uncertainty is not important in our empirical application, the option value- and the dynamic programming model reduce to one theoretical framework that can be used to study retirement decisions.

preferences that attach different weights to consumption in current and future periods. If the difference in the expected present value of postponing retirement is higher than the expected present value of retiring, the person delays retirement up to the point where the expected present value of retirement is higher than the value one would receive from postponing retirement. In this way, retirement decisions can be attributed to, for example, labor income, income from social insurances, or social security wealth which affect the opportunity costs of retirement (leisure). Evaluating the opportunity costs of several labor market decisions allows the model not only to include the decision when to retire, but how to withdraw from the labor market as well. A person will choose a period of using social insurance before retirement if the expected present value of doing so exceeds the expected present value of the other possibilities.

Applying this microecnomic theory on retirement behavior has been primarily focused on country-specific analyses, e.g. modeling microeconomic responses to country-specific retirement incentives in national institutions.[3] This approach, however, makes it hard to explain the differences in retirement patterns observed across countries. Piekkola and Deschryvere (2010) are the first to analyze retirement decisions with three countries based on the option value model using micro data from ECHP. The cross-country dimension, focusing on Belgium, Germany and Finland, remains limited however.

Cross-country macroeconomic reduced form approaches by Blondal and Scarpetta (1998), Johnson (2000), Duval (2003) and Blake and Sangnier (2011) are consistent with microeconomic theory although the empirical model is not directly derived from it. Nevertheless, these studies are the first to put the effects of early retirement incentives on early retirement in an international perspective using macro data. Blondal and Scarpetta (1998), Johnson (2000), Duval (2003) and Blake and Sangnier (2011) include

[3]Structural estimation applications can be found in, among others, Berkovec and Stern (1991), Rust and Phelan (1997), Heyma (2004), Van der Klaauw and Wolpin (2008) and Gustman and Steinmeier (2005). Examples of reduced form approaches can be found in Brugiavini and Peracchi (2003) and Mastrogiacomo et al. (2004) (using binary nonlinear models), Kerkhofs et al. (1999), Euwals et al. (2010) (using duration models). A multinomial logit approach is taken by, for example, De Vos et al. (2012), Zucchelli et al. (2012), Emmanoulidi and Kyriazidou (2012) and Been and Knoef (2013). The latter three also take non-standard employment into account in the retirement decisions.

approximations of disincentives to work at older ages stemming from early retirement programs including unemployment- and disability-related schemes. Although such a macroeconomic approach is unable to cope with individual and household characteristics and their influence on retirement decisions, it is able to analyze the effects of retirement incentives embedded in national institutions and to explain the observed differences in early retirement across countries.

To underpin our cross-country analysis of early retirement patterns, we extend the microeconomic framework with non-standard employment as labor market states. Two important types of cross-country heterogeneity in non-standard employment are part-time employment and self-employment (Chen et al. 2013), especially for persons approaching retirement (Morris and Mallier 2003).

5.2.2 Theoretical framework to model retirement decisions

Following microeconomic theory, we assume that persons maximize their current and future utility given their time-preferences (ρ)[4] and preferences for consumption (c) and leisure (l) and constraints that coincide with labor market choices. Since we are interested in retirement decisions in a context with regular employment, social insurance and non-standard employment possibilities, we assume six possible labor market states (j) for older persons: full-time employment (FT), part-time employment (PT), self-employment (SE), disability insurance (DI), unemployment insurance (UI) and early retirement (R). Disability insurance is often assumed to be a substitute to unemployment (Autor and Duggan 2003, Hassink et al. 1997, Koning and Van Vuuren 2010) and, more specifically, as a way to induce early retirement (Euwals et al. 2012, Riphahn 1997). Therefore, disability insurance is an important early retirement possibility in many countries (Gruber and Wise 1998).

[4] A higher value of ρ indicates a higher weight for current utility streams than for future utility streams such that the individual is relatively time-impatient.

Persons compare and reevaluate utility (u) streams coming from these six labor market states (j) in each period (t)[5] and so maximize their intertemporal utility given their characteristics that determine utility directly (ϑ) and the current institutional setting (s):[6]

$$U_t = \sum_{\tau=t}^{T} (1+\rho)^{t-\tau} u_\tau(c_{i\tau}, l_{i\tau}, j_{i\tau}; s_{i\tau}, \vartheta_{i\tau}) \tag{5.1}$$

Not only is the labor market state j associated with its own combination of consumption and leisure possibilities, it also enters the utility function directly such that some states can provide more satisfaction than other states.[7] Furthermore, the institutional setting determines to what extent states can be chosen (e.g. eligibility criteria) and how attractive these states are in terms of consumption possibilities (e.g. replacement rates, duration, tax differences) and leisure possibilities (e.g. extra requirements, hours flexibility).

Instead of solving and estimating a structural model for the retirement decision (Heyma 2004), we approximate equation (5.1) with a linear value function V (Mastrogiacomo et al. 2004). Here, V can be viewed as direct utility received from consumption, leisure, personal characteristics and the institutional setting captured by the different retirement possibilities[8]

$$V_{ij}(t) = \vartheta_{it}\beta_j + Z_{it}\theta_j + \epsilon_{it} \tag{5.2}$$

[5]Until the period of full retirement (T) which is assumed to be an absorbing state. Empirically, it is observed that full retirement is not always an absorbing state: older people may re-enter the labor market after full withdrawal (Maestas 2010). This assumption has, however, no consequences for our empirical reduced form model that is estimated using macro data.

[6]The institutional setting (s) implicitly includes the generosity of social insurance schemes as early retirement routes.

[7]In this way, active states may explicitly provide more utility than inactive states because it may make older workers feel 'useful'. A meta-analysis by Pinquart (2002) suggests positive associations between being employed and having a higher purpose in life at older ages.

[8]It is common in the literature to include an individual specific parameter as well in equation (5.2). We omitted this individual specific parameter from the equation since we do not intend to estimate equation (5.2) at the micro level using individual specific information.

such that (early) retirement is preferred if

$$V_{iR}(t) > V_{ik}(t) \tag{5.3}$$

for

$$k = \{FT, PT, SE, DI, UI\} \tag{5.4}$$

Here, ϑ is the vector of all observed characteristics that describe preferences for consumption and leisure and can therefore be seen as 'taste shifters'. Z includes labor market state-specific variables that indicate the incentives associated with each labor market state given by the institutional framework (s). The error component, ϵ, is assumed to follow a Type I extreme value distribution meaning that equation (5.2) is estimated by a multinomial logit model.

The mutually exclusive labor market states j and the baseline are modeled explicitly by normalizing the coefficients of the baseline to zero for identification.

$$P(j_t|\vartheta_{it}, Z_{it}) = \frac{\exp(\vartheta_{it}\beta_j + Z_{it}\theta_j + \epsilon_{it})}{\sum_{q=1}^{J} \exp(\vartheta_{it}\beta_q + Z_i\theta_q + \epsilon_{it})} \tag{5.5}$$

An increase in the probability to use a particular retirement route is always relative to the baseline $j = R$. Instead of estimating equation (5.5), we aim to estimate a macroeconomic analogy to the microeconomic reduced form approximation of microeconomic retirement behavior.

5.2.3 Empirical model

At the macro level, we do not have information on individual choices regarding labor market states j. Only the aggregate of individual choices is observed at the macro level. Therefore, we use information on early retirement, full-time employment, part-time employment, self-employment,

disability and unemployment rates in the macroeconomic analogy to equation (5.5). Since these rates are no binary indicators of j, as in the multinomial logit case at the micro level, we are unable to use an explicit multinomial logit procedure. Instead, we estimate equation (5.6) that explicitly allows for analyzing the mutual exclusiveness between early retirement and the aforementioned labor market states.

$$R_{it} = \beta_0 + \vartheta_{ct}\beta_1 + DI_{ct}\beta_2 + UI_{ct}\beta_3 + SE_{ct}\beta_4 + PT_{ct}\beta_5 + \gamma_t + \delta_c + \epsilon_{ct} \quad (5.6)$$

Here, R, DI, UI, PT and SE are the early retirement, disability, unemployment, part-time employment and self-employment rates respectively.[9] c now represents a country instead of an individual i as in equation (5.5). Country-fixed effects and time-effects are captured by γ and δ respectively. These fixed effects capture the unobserved heterogeneity in, for example, the cross-country differences in social acceptance of early retirement. The error term, ϵ, follows an i.i.d. normal distribution. Early retirement is the dependent variable since we are particularly interested in the substitution effects of labor market states with regard to early retirement and not so much in the substitution effects among labor market states in general.

Clearly, R, DI, UI, PT and SE are jointly, and so endogenously, determined in the model. Applying a simple OLS to equation (5.6) would yield biased and inconsistent estimates of the coefficients due to the endogeneity in equation (5.6). Therefore, we estimate equation (5.6) by a two-step procedure as in equation (5.7) where ω indicates the error term of the first-stage regression and ϵ indicates the error-term of the second step regression:

$$\begin{aligned}
R_{ct} &= \beta_0 + \vartheta_{ct}\beta_1 + DI_{ct}\beta_2 + UI_{ct}\beta_3 + SE_{ct}\beta_4 + PT_{ct}\beta_5 + \gamma_t + \delta_c + \epsilon_{ct} \\
DI_{ct} &= \pi_{0,DI} + \vartheta_{ct}\pi_{1,DI} + Z_{ct}\pi_{2,DI} + \omega_{ct,DI} \\
UI_{ct} &= \pi_{0,UI} + \vartheta_{ct}\pi_{1,UI} + Z_{ct}\pi_{2,UI} + \omega_{ct,UI} \\
SE_{ct} &= \pi_{0,SE} + \vartheta_{ct}\pi_{1,SE} + Z_{ct}\pi_{2,SE} + \omega_{ct,SE}
\end{aligned}$$

[9] Note that the full-time employment rate is captured by β_0.

$$PT_{ct} = \pi_{0,PT} + \vartheta_{ct}\pi_{1,PT} + Z_{ct}\pi_{2,PT} + \omega_{ct,PT} \tag{5.7}$$

Here, different characteristics, included in ϑ, and the institutional setting related to the labor market state, included in Z, determine the relative attractiveness of the labor market states. ϑ contains country-specific effects in the empirical model. Z includes labor market institutions affecting DI, UI, SE and PT but not directly affecting R (see section 5.2.5 for an extended discussion of the instruments). Compared to the microeconomic framework in equation (5.2), the relative attractiveness of the labor market states is now expressed as the aggregate of individual choices at the macro level by taking rates.

As suggested by the theoretical framework and captured by Z in equation (5.2), a different institutional setting may give a different relative attractiveness of labor market states and therefore influence retirement decisions.

For example, higher UI benefits, captured in Z, increase the attractiveness of UI resulting in a higher aggregate rate of UI. At the same time, the higher UI benefits decrease resulting in a lower aggregate rate of FT. Since the institutional setting is generally considered to be exogenous, the theoretical framework provides us with a set of instruments that can be used in the system of equations (equation 5.7) as long as the instruments used in equation (5.7) are not influenced by the aggregate retirement decisions.

5.2.4 Finding the appropriate estimator

Equation (5.7) suggests a two-step procedure with instruments for DI, UI, SE and PT. However, the second-stage error terms and labor market state specific first-stage error terms are possibly cross-correlated which would be ignored by a simple two-stage procedure. Ignoring the covariance structure of the error terms results in an inefficient 2SLS estimator. To increase efficiency, a 3SLS procedure could be applied (Zellner and Theil 1962). 3SLS accounts for both endogeneity (like in 2SLS estimation) and the covariance structure of the error terms across equations in the system

(like in SUR estimation). 3SLS estimations are widely applied in economics and they can be found in various contexts.[10]

Nevertheless, IV-GMM estimation as a 2SLS-estimator may be preferred to 3SLS in cases of heteroskedasticity since heteroskedasticity leads to an inconsistent 3SLS estimator resulting in inconsistent standard errors (Wooldridge 2002). In the case of IV-GMM, the error terms are i.i.d., but a possible correlation structure among the error terms is not taken into account. Wooldridge (2002) argues that there is generally not much reason to choose 3SLS over a GMM estimator if the assumption of homoskedasticity does not hold. Using a two-step IV-GMM estimator is also preferred to the regular 2SLS estimator as it relaxes the i.i.d. assumptions of the error terms resulting in efficiency gains in a situation of arbitrary heteroskedasticity.

As usual in IV regressions, we need validity and relevance of the instruments to justify our instruments. In a situation with possibly weak instruments, a Fuller-*k* estimator can be applied which is suggested to be more robust to possibly weak instruments than the IV-GMM estimator as the Fuller-*k* estimator is median unbiased (Stock et al. 2002). Even in a situation with weak instruments, LIML estimators such as the Fuller-*k* are almost unbiased (Blomqvist and Dahlberg 1999). As a consequence, LIML estimators have a relatively large variance and can be regarded as a conservative estimator when dealing with many weak instruments and small sample sizes (Blomqvist and Dahlberg 1999).

Instruments 5.2.5

Blondal and Scarpetta (1998), Johnson (2000) and Duval (2003) use several labor market institutions such as unemployment benefit replacement rates, employment protection legislation (EPL) and labor taxes as regressors in explaining early retirement and argue that these labor market states are exogenous in an OLS framework. However, as these studies note, the effect of labor market institutions such as unemployment benefits and EPL on early retirement run through their effect on unemployment

[10] A small selection of 3SLS applications in economics: Buck and Hakim (1982), Audretsch and Feldman (1996), Burnside (1996), Barro (2000), Glewwe et al. (2001), Burton et al. (2002), Butkiewicz and Yanikkaya (2005) and Brown and Alexander (2005).

rates.[11] Therefore, we apply a two-step procedure using labor market institutions as instruments for the aggregated levels of DI, UI, PT and SE. For identification, the instruments should 1) be correlated with DI, UI, PT and SE (instrument relevance) and 2) be uncorrelated with the error term, ϵ, in equation (5.7) (instrument validity). If these conditions are fulfilled, the instruments are associated with changes in DI, UI, PT and SE but do not lead to changes in R aside from their indirect effect through DI, UI, PT and SE. As we will explain in detail below, labor market institutions associated with DI, UI, PT and SE are theoretically proven to be relevant as well as valid instruments.[12]

For the selection of instruments we follow the literature on the effects of labor market institutions - such as unemployment benefits, employment protection legislation (EPL) and active labor market policies (ALMP) - on unemployment rates (Bassanini and Duval 2009, Belot and Van Ours 2004, Blanchard and Wolfers 2000, Elmeskov et al. 1998, Nickell 1998, Nickell et al. 2005, Scarpetta 1996). This literature shows that more generous benefits make unemployment more attractive and so increase unemployment rates. The effects of EPL are somewhat more ambiguous. Nickell (1998) argues that the effects of EPL on the unemployment rate are likely to be small because EPL mainly tends to decrease flows into- and out of unemployment. Nevertheless, decreasing inflows and outflows from unemployment may decrease short-term unemployment and increase long-term unemployment (Nickell 1998). An excellent overview of the theory and empirics of EPL by Addison and Teixeira (2003) shows that estimations of the effect of EPL on unemployment rates can be both positive and negative. However, Addison and Teixeira (2003) also indicate that most of the effects that are estimated to be significant show that EPL and unemployment rates are positively related. As an additional determinant of unemployment rates, the effects of ALMP spending are generally found to be small in economic terms (Card et al. 2010, Kluve 2010). Nevertheless, ALMP spending is important to take into account as some countries may have high initial UI benefit levels but also many reintegration measures.

[11] This, of course, does not apply to the included variables that indicate the disincentives of postponing retirement such as the implicit tax rate on continued work.

[12] Empirical tests also prove that instruments are both relevant and valid as we will show in section 5.5.

Replacement rates of UI benefits, EPL and ALMP spending are proven to be relevant instruments according to the aforementioned literature. There is no *ex ante* reason to belief that there is a direct relationship between replacement rates of UI benefits and ALMP spending and early retirement (validity). However, this does not hold for EPL. Dorn and Sousa-Poza (2010) find that strict EPL leads to higher shares of involuntary early retirement.

A similar approach as Scarpetta (1996), Nickell (1998), Nickell et al. (2005), Elmeskov et al. (1998), Blanchard and Wolfers (2000), Belot and Van Ours (2004) and Bassanini and Duval (2009) - focusing on cross-country self-employment instead of unemployment - has been employed by Torrini (2005), Robson (2003) and Parker and Robson (2004). Torrini (2005) focuses on explaining variation in self-employment rates by cross-country and time differences in tax incentives. Robson (2003) focuses on cross-country and time differences in EPL on self-employment (rates) and Parker and Robson (2004) focus on cross-country and time differences in self-employment due to (dis)incentives of income taxes and unemployment benefit replacement rates. All three papers suggest that labor market institutions, as used in the literature focusing on unemployment, are relevant in explaining self-employment as well. The results suggest that high taxes on labor induce self-employment (Parker and Robson 2004, Robson 2003, Torrini 2005) while high unemployment benefit replacement rates (Parker and Robson 2004, Robson 2003) and EPL (Robson 2003) reduce self-employment. Therefore, the labor market variables (replacement rates of UI benefits particularly) used to instrument the unemployment rate are also expected to be relevant instruments for the self-employment rate. As a tax incentive, we include the implicit tax on self-employment as an additional instrument for self-employment. We do not use the implicit tax on labor as this variable may also be relevant for the marginal decision to work in regular employment at older ages.

It is harder to find valid instruments for the part-time employment rates as most determinants (e.g. possible relevant instruments) of part-time employment (Dorn and Sousa-Poza 2010), such as the fertility rate, child benefits and female labor supply, possibly also have a direct effect on labor market withdrawal at older ages (e.g. non-valid). However,

Buddelmeyer et al. (2008) suggest that EPL is an important determinant of part-time employment from a macro perspective. Based on their findings and earlier work, Buddelmeyer et al. (2008) suggest that strict EPL increases the advantages of hiring part-time employment in firms. This effect can either be direct (e.g. EPL strictly limits the use of part-time employment) or indirect (e.g. to ease the firms' burden of highly rigid employment legislation associated with full-time employment). Either way, EPL seems to be relevant for unemployment, self-employment and part-time employment, but the validity of this instrument is questionable (Dorn and Sousa-Poza 2010). Therefore, we do not use the level of EPL as an instrument but the difference between EPL for regular and temporary work instead.[13] The difference between EPL for regular and temporary jobs may be relevant for the decision to work part-time as this difference may lead to spill-overs between full-time and part-time work given that part-time jobs are generally more often of a temporary nature (OECD 2002) and because part-time employment is complementary to temporary employment (Buddelmeyer et al. 2008). Additionally, EPL_{diff} may be valid as it is likely that only the level of EPL is important for involuntary early retirement.

Fairly little research has been devoted to the cross-country variation in disability rates. OECD (2009a) shows that there is large heterogeneity in disability rates among countries and that such heterogeneity can be explained by cross-country differences in compensation, such as accessibility and generosity of disability benefits (both positively related to disability rates), and (re)integration policies such as vocational rehabilitation programs, subsidized employment and other activation policies that provide incentives to work (all negatively related to disability rates). Furthermore, OECD (2009a) argues that UI benefit replacement rates may also be of importance in modeling disability rates since it is a crude measure of alternative benefit options. OECD (2009a) finds that higher UI benefit replacement rates are negatively related to disability rates. Following aforementioned study, we instrument disability rates with UI replacement rates and ALMP spending. Furthermore, we use the information regarding

[13] $EPL_{diff} = EPL_{regular} - EPL_{temporary}$

the spending on disability which is publicly available at the macro level.[14] Additionally, we include the replacement rates of social assistance benefits as an additional instrumental variable in the sensitivity analyses as these benefits may be a true alternative benefit option. Social assistance benefits may also be important to include in the analysis as exhausted UI benefits are usually followed-up by social assistance benefits. This is especially relevant for older persons bridging the gap between employment and retirement with UI. Additionally, we are able to use the compensation and integration indices as used in OECD (2009a).

A final note on the instruments is that we use the same full set of instruments for each endogenous variable to account for the alternative options.

Data and definitions 5.3

Data 5.3.1

To identify early retirement, we use a variable that describes labor market withdrawal of persons aged 55-64 compared to below 55 following Duval (2003).[15] More specifically, the dependent variable is constructed as

$$R = \frac{AR_{25-54} - AR_{55-64}}{AR_{25-54}} \cdot 100 \tag{5.8}$$

where AR indicates the activity rate. By relating the activity rate of persons aged 55-64 to the activity rate of persons aged 25-54 we take into account cross-country differences in participation of prime-age persons who are not confronted with retirement decisions yet. By using this relative measure, we also correct for within-country cohort differences in participation which are especially relevant among women (Ebbinghaus 2006). Blondal and Scarpetta (1998), Johnson (2000), Duval (2003) and

[14] We use the transformed natural logarithm of per capita disability spending.

[15] We also use different age spans to disaggregate early retirement approximations (Ebbinghaus 2006). Results are largely robust to this baseline specification of the dependent variable.

Blake and Sangnier (2011) explicitly focus on males only. We, however, analyze both men and women in separate analyses. *R* can be interpreted as the percentage decrease in activity rates of the population aged 55-64 compared to 25-54, e.g. a macroeconomic measure of labor supply at the extensive margin.

For the dependent variable, the study relies on activity rates, defined as the sum of the employed and unemployed workers as a percentage of the population.[16] We use activity rates for both men and women and for different age groups, which are taken from the Labour Force Survey, provided by Eurostat (2014).

For the independent variables, we use a number of labor market indicators for persons aged 55-64. The unemployment rate is measured as the number of unemployed employees as a percentage of the labor force aged 55-64 and is publicly available as a macro indicator at Eurostat. Disability rates, self-employment rates and part-time employment rates are calculated ourselves using the Labour Force Survey micro data (Eurostat 2014) as the publicly available macro data are insufficient for our purposes.[17]

In general, constructing internationally comparable disability rates from survey data is hard as Banks et al. (2004) show. There is no internationally comparable data on disability rates among older persons. The OECD provides cross-country rates of disability for people aged 20-64, but the cross-country comparability is questionable (OECD 2010a). To identify disability, we use persons aged 55-64 with no job during the reference week who do not search for a job because they indicate to suffer from sickness or disability.[18] The disability rate is then constructed by relating

[16] According to Eurostat, the activity rate represents the number of persons in the labor force (employed and unemployed persons) as a percentage of the total population of the same age. Using this statistic, it is assumed that persons in unemployment have not retired yet. If one assumes the older unemployed to have retired instead, the employment rate would be a more appropriate indicator to construct the dependent variable (Ebbinghaus 2006). We follow Blondal and Scarpetta (1998), Johnson (2000), Duval (2003) and Blake and Sangnier (2011) by using activity rates. A sensitivity analysis (table 5.5) where employment rates are used instead of activity rates shows that the results are robust.

[17] Do-files to construct these measures are available upon request.

[18] In LFS: $WSTATOR = 5$, $SEEKWORK = 3$ and $SEEKREAS = 2$.

the total of identified disabled persons to the total of persons aged 55-64 for whom these variables are non-missing.[19]

We identify part-time employment as persons aged 55-64 working as an employee with a part-time job.[20] This excludes self-employed working part-time as we only want to include part-time paid-employment in part-time employment. Part-time employment is standardized across countries with a measure that identifies a person to be in part-time employment if a person's usual working hours are less than the working hours considered to be 'normal'. Normal working hours is considered to be a full-time working week at the national, regional, industrial or unit level. Hence, the part-time employment indicator is a relative indicator that accounts for local standards with respect to working hours. The part-time employment rate is measured as part-time employment as a percentage of the total of paid-employment. This implies a percentage of total paid-employment for whom the identifying variables are non-missing.

Self-employment is identified as persons aged 55-64 working as a self-employed in industries other than agriculture.[21] Issues may arise regarding the cross-country measurement of self-employment. The self-employment definition of Eurostat is standardized to a broad definition of all workers who are not in (paid) employment. This includes sole or joint owners of unincorporated enterprises (unless their main activity is in paid-employment), unpaid family workers, outworkers and people in self-sufficiency (Eurostat 2014).[22] However, issues regarding the cross-country comparability of self-employment arise with regard to the question whether to treat incorporated enterprises as self-employment or not (OECD 2005). Given our set of countries, this comparability issue

[19]However, Eurostat acknowledges that breaks in the data are present for $SEEKREAS = 2$ for some countries due to survey improvement. If this is observed for several years in a country and the variable $MAINSTAT$ is available for that country, we base the disability rates on the $MAINSTAT$ variable instead of basing the rate on the variables $WSTATOR$, $SEEKWORK$ and $SEEKREAS$. In the cases that $MAINSTAT$ is not available, we delete the disability rates that are subject to the break from the data set (this occurs in Ireland (2005), France (2003-2008), Portugal (1995-1997) and UK (2005)). Hence, the disability rate variable should be interpreted with caution, although Eurostat confirms that there is no other option to construct disability rates.

[20]In LFS: $WSTATOR = 1, 2$, $STAPRO = 3$ and $FTPT = 2$.

[21]In LFS: $WSTATOR = 1, 2$, $STAPRO = 1, 2$ and $NA11s = 1, 2$.

[22]This standardized definition of self-employment is also used by the OECD for cross-country comparisons (OECD 2005).

is small and unlikely to affect the results since only Norway considers the incorporated self-employed to be in paid-employment. We exclude self-employment in agriculture because self-employment is traditionally high in the agricultural sector (Parker 2004, Torrini 2005) while agricultural self-employment is unlikely to function as a bridge between full-time work and full labor market withdrawal.

The self-employment rate is measured as a percentage of total employment between the age of 55 and 64, which is the convention in the literature (Chen et al. 2013, Robson 2003). This implies a percentage of the total of paid- and self-employment for whom the identifying variables are non-missing.

5.3.2 Descriptive statistics

Table 5.1 shows the activity rates of men and women for 15 European countries between 1995 and 2008. For persons aged 25-54, activity rates are generally higher than the activity rates of persons aged 55-64. Furthermore, the table also indicates that the activity rates of persons aged 55-64 have increased over time. Table 5.2 presents the developments in early retirement among men and women (based on the information from table 5.1). Among men, labor market withdrawal is the highest in France and Belgium. In 2008 for example, the activity rate among French men between 55 and 64 years old was 54.9 percent lower than the activity rate among the men between 25 and 54. For women, labor market withdrawal was the highest in Belgium. Across the board, labor market withdrawal between 55 and 64 is more prevalent among women then among men, except for Finland in 2008. Between 1995 and 2008, early retirement declined across all countries. Only among Danish men and Greek women, early retirement increased substantially. The decline was the sharpest in the Netherlands, both among men and women.

The disability, unemployment, self-employment and part-time employment rates for the population between 55 and 64 are presented in

figure 5.1.[23] The disability rate is especially high in Nordic countries and the Netherlands (about 15-20 percent on average), while disability rates are only decreasing in Finland and the Netherlands. The other Nordic countries seem to have high and persistent disability rates among 55 to 64 year olds. On the other hand, the unemployment rates are, on average, rather low in these countries (except for Finland) compared to the relatively high unemployment rates in Germany (about 12 percent), Spain (about 9 percent). Moreover, unemployment rates also vary within countries over time.

Self-employment rates are rather stable over time in most countries although Greece, Norway, Sweden and the UK show a slightly increasing trend in self-employment among the 55-64 population over time. Only Belgium and Italy show marked decreases in self-employment. Interestingly, self-employment rates are relatively high in the Southern European countries and in Ireland while, simultaneously, these countries show relatively low rates of disability and unemployment. This may suggest that self-employment is a reaction to a lack of alternatives at old-age.[24]

[23] Disability-, self-employment- and part-time employment rates for Germany are only available from 2002 as the LFS micro data is unavailable before 2002 due to an improvement in the survey method.

[24] This is also more or less suggested by Been and Knoef (2013).

Table 5.1: Activity rates of workers aged 25-54 and 55-64

	Men						Women					
	Activity rate 25-54			Activity rate 55-64			Activity rate 25-54			Activity rate 55-64		
	1995	2008	Change	1995	2008	Change	1995	2008	Change	1995	2008	Change
Austria	93.6	93.0	-0.6	44.0	52.8	8.8	73.1	81.5	8.4	18.7	31.6	12.9
Belgium	92.3	92.3	0.0	34.9	44.4	10.5	67.7	79.0	11.3	13.5	27.9	14.4
Denmark	92.1	93.3	1.1	70.3	66.9	-3.4	83.0	86.4	3.4	40.2	53.0	12.8
Finland	90.8	91.2	0.4	44.6	60.6	16.0	84.4	85.7	1.3	41.4	58.8	17.4
France	95.2	94.4	0.8	36.5	42.6	6.1	77.5	83.4	5.9	27.5	37.6	10.1
Germany	93.1	93.5	0.4	54.4	67.2	13.8	73.3	80.5	7.2	31.5	50.5	19.0
Greece	94.6	94.4	-0.2	61.8	60.9	-0.9	55.2	69.4	14.2	24.7	28.6	3.9
Ireland	90.9	91.3	0.4	64.5	68.6	4.1	54.6	71.8	17.2	20.2	42.2	22.0
Italy	90.4	91.0	0.6	46.4	47.0	0.6	53.4	65.2	11.8	14.1	24.7	10.6
Netherlands	92.4	94.5	2.1	41.1	65.9	24.8	66.0	82.5	16.5	19.0	43.5	24.5
Norway	91.2	91.3	0.1	72.3	75.0	2.7	83.4	85.5	2.1	59.4	64.7	5.3
Portugal	93.5	93.2	-0.3	63.1	63.0	-0.1	73.4	82.9	9.5	33.9	46.6	12.7
Spain	91.7	92.6	0.9	54.0	65.1	11.1	52.6	74.7	22.1	18.5	34.2	15.7
Sweden	90.4	93.1	2.7	68.6	76.5	7.9	86.8	87.6	0.8	61.9	69.0	7.1
UK	92.7	91.7	-1.0	62.4	69.9	7.5	74.1	78.7	4.6	40.7	50.2	9.5
Mean	92.3	92.7	-0.6	54.6	61.8	7.2	70.6	79.7	9.1	31.0	44.2	13.2

Source: Own calculations based on Eurostat (2014). The first available year for Norway is 2000.

Part-time employment has been traditionally relatively low in most Southern European countries, although the part-time employment rates have been rising in virtually all countries as shown in the figures. Also in countries with relatively high initial rates of part-time employment, such as Germany and the Netherlands, part-time employment increased between 1995 and 2008. These developments suggest that the decline in early retirement as observed in table 5.2 might be related to the increase in part-time employment.

Below, we employ regression analyses to examine such relationships in more detail. Unfortunately, not all instruments are available for all countries. Implicit tax rates on self-employment and social assistance benefits replacement rates are unavailable for Germany and Greece respectively.

Results from reduced form retirement model 5.4

As mentioned in section 5.2.4, 3SLS regression may be preferred to IV-GMM regression if homoskedasticity of the error terms and a covariance structure among the error terms is present. A Breusch-Pagan LM Test, an LR Test and a Wald Test indicate that heteroskedasticity is present in the model.[25] Therefore, using a two-step IV-GMM estimator is preferred.

Validity of the instruments is tested using the Hansen's J-statistic.[26] We perform robustness checks of the results with instruments that are certain to be exogenous, such as the lags of the endogenous regressors.

Relevance of the instruments is tested using an F-test in the first-stage regressions[27] as well as by the Anderson-Rubin statistic that is robust to

[25] P-value = 0.000 for all three tests implemented with the Stata program written by Shehata (2011). This applies to separate analyses of both men and women.

[26] P-values of the Hansen's J-statistic are reported in the regression tables and show that the instruments are valid in all baseline regressions.

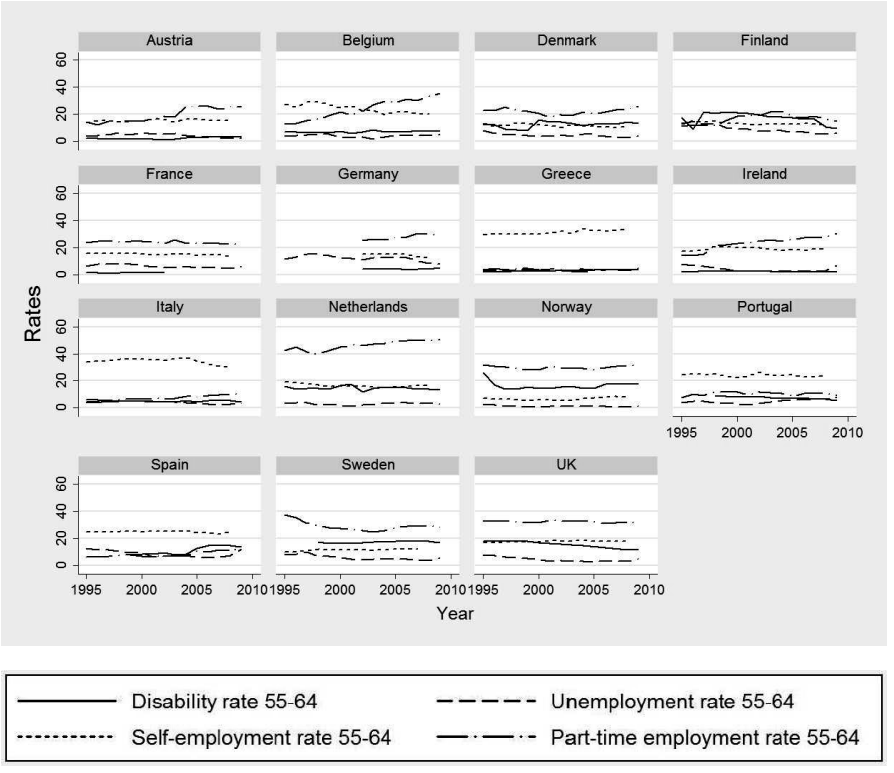
[27] The F-tests of our first-stage regressions in the baseline model (Model 2 in table 5.3 and table 5.4) reject the null-hypothesis of non-relevance of the instruments at 1%, 5% and 1% level for part-time employment, self-employment and unemployment respectively. The F-tests indicate that the set of instruments is relevant for these endogenous variables in the baseline specification. Regarding disability, we cannot reject the null-hypothesis. However, if we add interactions between two institutional variables, as in Model 4 and 5, the null-hypothesis can be rejected. T-tests indicate that the relevance of a single instrument may differ between the instrumented variables DI, UI, SE and PT, but they are included as we tend to keep the set of instruments the same in all first-stage regressions.

Table 5.2: Labor market withdrawal of workers aged 55-64 relative to workers aged 25-54

	Men			Women		
	1995	2008	Change	1995	2008	Change
Austria	53.0	43.2	-9.8	74.4	61.2	-13.2
Belgium	62.2	51.9	-10.3	80.1	64.7	-15.4
Denmark	23.7	29.3	5.6	51.6	38.7	-12.9
Finland	50.9	33.6	-17.3	51.0	31.6	-19.4
France	61.7	54.9	-6.8	64.5	54.8	-9.8
Germany	41.6	28.1	-13.4	57.0	37.3	-19.8
Greece	34.7	35.5	0.8	55.3	58.8	3.5
Ireland	29.0	24.9	-4.2	63.0	41.2	-21.8
Italy	48.7	48.4	-0.3	73.6	62.1	-11.5
Netherlands	55.5	30.3	-25.3	71.2	47.3	-23.9
Norway	20.7	18.6	-2.1	28.8	23.7	-4.1
Portugal	32.5	32.4	-0.1	53.8	43.8	-10.0
Spain	41.1	29.7	-11.4	64.8	54.2	-10.6
Sweden	24.1	17.8	-6.3	28.7	21.2	-7.5
UK	32.7	23.7	-9.0	45.1	35.8	-9.3
Mean	40.8	34.2	-6.6	57.5	45.1	-12.4

Source: Own calculations based on Eurostat (2014). The first available year for Norway is 2000.

Figure 5.1: Disability-, unemployment-, self-employment- and part-time employment rates of persons aged 55-64 over time.



Source: Own calculations based on Eurostat (2014).

weak instruments.[28] To test the robustness of the results to possibly weak instruments, we apply a Fuller-*k* estimator. The results (not reported here) show that the main conclusions are robust to the Fuller-*k* estimator.[29]

In all estimations, we use HAC-corrected standard errors, correcting for both heteroskedasticity and autocorrelation which are generally present in macroeconomic retirement indicators (Ebbinghaus 2006).[30] Based on the results of the Wooldridge (2002) test for autocorrelation in panel data, we reject the null-hypothesis of no first-order autocorrelation (p-value = 0.000).

Table 5.3 and table 5.4 show the estimation results for men and women respectively. A regular fixed-effects regression is used in Model 1. This model indicates that the disability-, unemployment- and self-employment rate increase early retirement while part-time employment decreases early retirement. However, as these variables are endogenous, estimation results are biased and we should use instrumental variables to infer correct relationships.

Model 2 uses the IV-GMM approach proposed in this paper and indicates that a higher unemployment rate among persons aged 55-64 increases the percentage change in activity rates from 25-54 to 55-64, e.g. more people enter early retirement.[31] This applies to both men and women. More specifically, we find that a 1 percentage point increase in the unemployment rate increases early retirement by about 1.6 and 2.5 percentage points among men and women respectively.[32] We do not find significant effects of

[28]The AR statistic shows that the instruments are relevant in the baseline model (Model 2 in table 5.3 and table 5.4). The null-hypothesis of relevance of the instruments can not be rejected. However, this does not mean that there is no weak instruments problem.

[29]The unknown parameter of the Fuller-*k* estimator is set equal to 4 following Hahn et al. (2004).

[30]Estimations are corrected for autocorrelation by using a Bartlett kernel function with a bandwidth of 2. A common choice for the bandwidth is $T^{1/3}$ with T the time-dimension (Baum et al. 2007). In our case, T has a minimum of 7 and a maximum of 14.

[31]Estimated coefficients of the first-stage regressions that are significantly different from zero are in line with the expected signs.

[32]This is slightly higher than the 0.6-0.9 percentage points found in the preferred model by Blondal and Scarpetta (1998) and lower than the 1.2-9 percentage points found in the preferred model by Duval (2003). Please note that their dependent variable also differs. Duval (2003) uses a similar approach as we do, but employs smaller age-groups in activity rates. Blondal and Scarpetta (1998) only take into account the participation rate of 55-64 year olds.

the disability rate, which may be related to the difficulties of constructing the variable as mentioned in section 5.3. Model 3 therefore uses a different series of disability rates.[33] Results are nevertheless robust.

Self-employment is not found to have a significant effect on activity at older ages. Regarding the part-time employment rate, we find that a 1 percentage point increase in the part-time employment rate decreases early retirement by 1.7 percentage points among men. Among women, we do not find such a relationship between part-time employment and early retirement in Model 2. Models 3, 4, 5, 7, 10 and 11 do find this relationship between part-time employment and early retirement among women. The effect is generally smaller among women than among men which may be explained by the different purposes part-time employment has for men and women over the life-cycle. Part-time employment as gradual retirement mechanism is likely to be more important among men than among women (Peracchi and Welch 1994).

Please note that estimated coefficients larger than 1.0 do not necessarily reflect an extra effect upon direct labor supply effects as not all variables use the same denominator. Activity rates, unemployment rates and disability rates use the total labor force as denominator, self-employment rates the total of employment and part-time employment rates the total of people in paid-employment. Next to that, the dependent variable is a relative indicator indicating the percentage change in activity of 55-64 year olds compared to 25-54 year olds.

Subsequently, we examine the robustness of the results with respect to different specifications. In Model 4 and 5, we add interactions between the instruments as additional instruments since institutions can interact in their impact on macroeconomic outcomes (Belot and Van Ours 2004). Results are largely robust except that part-time employment now seems to decrease early retirement among women similar to men. Model 6 indicates that this effect can be primarily attributed to voluntary part-time employment measured as the share of total part-time employment that is part-time employed because of other reasons than that the person could

[33]These disability rates are solely based on the variable *WSTATOR*, *SEEKWORK* and *SEEKREAS* in the LFS micro data without imputing results from the variable *MAINSTAT* and deleting possibly false rates due to breaks in the data.

not find a full-time job.[34] So, we define involuntary part-time employment as a consequence of labor demand. A 1 percentage point higher share of voluntary part-time employment decreases early retirement by about 2 and 1.3 percentage points among men and women respectively. These results suggest that people participate longer in the labor force when they can voluntarily choose for a reduction of their working hours. Without the option of part-time employment, more persons would exit the labor market rather than working full-time. Involuntary part-time employment does not lead to less early retirement, because these persons prefer a full-time job per definition. Hence, it is likely that these persons in involuntary part-time employment do not use part-time employment as partial retirement mechanism and prefer continue working (full-time) instead of retiring early.

Model 7 includes the state pension eligible age as an additional exogenous variable. This additional explanatory variable may correct for the precariousness of sustained labor force participation at older ages due to the graying population and the reforms undertaken to improve the sustainability of the pension system. The pension eligible age does not seem to affect early retirement and earlier results regarding unemployment and part-time employment are robust. The insignificance of the state pension eligible age may be explained by the fact that this variable does not vary much within countries over time (see also Blake and Sangnier 2011). Most of the variation in the variable comes from cross-country variation. If Model 2 is performed on a subsample of countries that have a state pension eligible age that is 65 or above,[35] results of Model 2 are robust for men (not reported here). Significant effects are lost when this is done for women as a consequence of a large drop in observations (not reported here), e.g. many countries have a state pension eligible age for women that is below 65.[36] Blake and Sangnier (2011) did find significant effects of the state pension eligible age on activity rates among older persons.

[34] Whether a person works part-time because the persons could not find a full-time job is based on the variable *FTPTREAS* in the LFS micro data.

[35] Austria, Belgium, Denmark, Finland, Germany, Greece (from 1999), Ireland, Italy (from 2002), Netherlands, Norway, Portugal, Spain, Sweden, and the UK have a state pension eligible age of 65 or above.

[36] Only Denmark, Finland, Germany, Ireland, Netherlands, Norway, Portugal (from 1999), Spain and Sweden have a female state pension eligible age of 65 or more.

However, they found the effect to be negative among persons aged 55-59 and to be positive among persons aged 60-64. In our case of looking at the aggregate of persons aged 55-64, these effects may cancel out.

Model 8 and 9 include an indicator for the implicit tax rate on continued work at older ages as explanatory variables of early retirement. Model 8 includes the implicit tax rate for the 55+ population and model 9 for the 60+ population. Both indicators are included as approximations of incentives to work at older ages. The implicit tax rate on continued work is an approximation of the importance of official early retirement schemes (Blake and Sangnier 2011). The implicit tax rate on continued work gives an indication of the marginal benefits of continued working, e.g. the marginal costs of retiring early. Although this concept is different from replacement rates for early retirement, Duval (2003) argues that cross-country differences in implicit tax rates on continued work are good indicators of cross-country differences in the level of generosity of early retirement. Unfortunately, data on implicit tax rates on continued work are highly unbalanced. Therefore, we use an alternative indicator as suggested by Duval (2003).[37]

We construct an indicator of the implicit tax on continued work that is equal to 0 if the implicit tax rate is in the first quartile[38] of the cross-country and -time distribution of implicit tax rates. Similarly, the indicator equals 1 if the tax rate is in the second or third quartile and 2 if the tax rate is in the fourth quartile. Unlike Johnson (2000) (0.6-1.7 percentage points), Duval (2003) (0.6-1.7 percentage points) and Blake and Sangnier (2011) (no effect among 55-59 year old persons and 0-0.7 percentage points among persons aged 60-64) we do not find significant effects of the implicit tax rate on continued work. A possible explanation for not finding significant effects of the implicit tax rate on continued work is that the marginal costs of retiring may already be implicitly defined by the alternative labor market options relative to full-time employment. Nevertheless, we still observe

[37]Duval (2003) makes a difference between countries with a low, medium and high level of implicit tax rates on continued work and constructs an indicator that is equal to 0, 1 or 2 respectively.

[38]The first quartile of the 55+ (60+) indicator stops at an implicit tax rate of 21.40 (20.3). The fourth quartile begins at an implicit tax rate of 58.2 (74.2). The distribution has a mean and median of 40.9 (47.4) and 44.1 (35.58) respectively. The lowest value in the distribution is 2.80 (-0.15). The highest value is 101.4 (105.1).

that a higher part-time employment rate decreases early retirement among men. For women, the results of Model 8 and 9 are similar to Model 2. A more practical explanation for not finding an effect of the implicit tax on continued work is the loss of heterogeneity because of constructing the aforementioned indicator. If we would use the original data from Duval, our regressions would end up having only 39 observations. However, specifying a simple fixed-effects model with HAC-corrected standard errors and the implicit tax rate on continued work as the only explanatory variable does show that a higher implicit tax rate on continued work increases early retirement significantly (based on 89 observations).

Model 10 shows the results of the IV-GMM estimation with a different set of instruments. Whereas the previous estimations are based on a set of instruments that include labor market institutions only, Model 10 presents estimation results with a set of instruments that consists of both the labor market institutions and the first lag of the endogenous variables. Also this model indicates that unemployment increases early retirement, part-time employment decreases early retirement and self-employment does not decrease early retirement (no effect among men and a positive effect among women). To test the robustness of our results to a set with highly exogenous instruments, we perform a regression with only the first and second lags of the endogenous regressors of instruments. The estimated effects (not reported here) are slightly smaller than in the baseline regression, but the estimated effects are highly robust.

Finally, Model 11 presents the estimation results for a specification in which the compensation and reintegration policies of disability insurance are taken into account in the first-stage as these indicators seem to partially explain cross-country differences in disability rates (OECD 2009a). Based on the underlying data of figures 4.7 and 4.8 presented in OECD (2009a), we constructed time-varying indices of compensation and integration policies regarding disability insurance.[39] The main conclusions do not alter by including these instruments that are highly likely to be valid.[40]

[39]For an explanation of these indices we refer to Annex 4.A2 in OECD (2009a).

[40]The C-statistic confirms that we can assume these indices as well as the DI spending to be exogenous with p-value=0.224 and p-value=0.274 respectively (orthogonality is tested subsequently).

Table 5.3: Regression results of labor market withdrawal, men 55-64[a]

	Model 1[b]	Model 2	Model 3[c]	Model 4[d]	Model 5[e]	Model 6	Model 7	Model 8	Model 9	Model 10[f]	Model 11[g]
Disability rate 55-64	0.38**	-1.48	-0.08	-0.37	0.16	-2.15	-0.96	-1.54	-2.04	0.23	-0.34
	(0.15)	(2.32)	(1.17)	(0.70)	(0.59)	(3.18)	(1.63)	(2.70)	(2.81)	(0.41)	(0.63)
Unemployment rate 55-64	0.89***	1.61*	2.71**	1.59***	1.51***	1.32	1.33**	1.41	2.00**	1.41***	1.63***
	(0.20)	(0.87)	(1.20)	(0.48)	(0.48)	(1.06)	(0.59)	(1.00)	(0.98)	(0.34)	(0.50)
Self-employment rate 55-64	0.55**	-1.92	-2.19	0.18	-0.53	-2.44	-0.76	-2.34	-2.00	-0.45	0.22
	(0.22)	(1.81)	(1.91)	(0.99)	(1.04)	(2.26)	(1.21)	(2.19)	(2.15)	(0.57)	(1.32)
Part-time employment rate 55-64	-0.44***	-1.70**	-1.92**	-0.91***	-0.79**		-1.25**	-1.69**	-1.68**	-0.82***	-0.92**
	(0.12)	(0.80)	(0.88)	(0.34)	(0.35)		(0.51)	(0.83)	(0.85)	(0.21)	(0.39)
Voluntary part-time employment rate 55-64						-1.96*					
						(1.06)					
State pension eligible age							0.75				
							(3.46)				
Implicit tax rate on continued work 55+ (indicator)								2.00			
								(3.06)			
Implicit tax rate on continued work 60+ (indicator)									4.89		
									(3.89)		
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations (N x T)	177	144	156	144	144	143	144	144	144	132	122
Hansen's J statistic (p-value reported)[h]	n.a.	0.83	0.85	0.25	0.26	0.83	0.55	0.79	0.68	0.48	0.29

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions include the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Model 1 is a regular FE model assuming exogeneity of the regressors.

[c] Model 3 uses DI rates solely based on the micro variables WSTATOR, SEEKWORK and SEEKREAS.

[d] Model 4 additionally includes the interaction between UI benefits replacement rate and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[e] Model 5 additionally includes the interaction between the implicit tax rate on self-employment and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[f] Model 10 additionally includes the first lag of the endogenous variables as instruments.

[g] Model 11 additionally includes the DI compensation and integration indices of OECD (2009a) as instruments. Indices are unavailable for France and Italy.

[h] Hansen's J statistic H0: valid instruments.

Table 5.4: Regression results of labor market withdrawal, women 55-64[a]

	Model 1[b]	Model 2	Model 3[c]	Model 4[d]	Model 5[e]	Model 6	Model 7	Model 8	Model 9	Model 10[f]	Model 11[g]
Disability rate 55-64	0.41**	0.76	-0.62	-0.35	0.21	-0.31	3.66	0.91	0.80	0.23	-0.60
	(0.19)	(1.71)	(0.76)	(0.94)	(0.77)	(1.57)	(2.91)	(2.40)	(2.19)	(0.43)	(0.80)
Unemployment rate 55-64	1.43***	2.49***	3.31***	2.43***	2.35***	2.35***	2.13**	2.50**	2.61***	1.55***	2.61***
	(0.24)	(0.89)	(0.77)	(0.54)	(0.54)	(0.69)	(1.50)	(1.19)	(0.98)	0.42	(0.62)
Self-employment rate 55-64	0.74***	0.77	0.30	0.83	1.47	0.62	3.79	0.68	0.70	1.11*	1.84
	(0.22)	(1.44)	(1.19)	(0.90)	(1.00)	(1.17)	(3.71)	(1.76)	(1.68)	(0.62)	(1.45)
Part-time employment rate 55-64	-0.71***	-0.93	-1.54**	-1.15***	-0.81***		-0.44**	-0.92	-0.84	-0.78***	-1.03**
	(0.14)	(0.65)	(0.70)	(0.34)	(0.30)		(0.96)	(0.73)	(0.73)	(0.25)	(0.44)
Voluntary part-time employment rate 55-64						-1.30**					
						(0.61)					
State pension eligible age							2.60				
							(2.39)				
Implicit tax rate on continued work 55+ (indicator)								0.20			
								(4.07)			
Implicit tax rate on continued work 60+ (indicator)									1.64		
									(2.90)		
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations (N x T)	177	144	156	144	144	143	144	144	144	132	122
Hansen's J statistic (p-value reported)[h]	n.a.	0.31	0.22	0.37	0.28	0.26	0.80	0.36	0.30	0.11	0.13

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions include the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Model 1 is a regular FE model assuming exogeneity of the regressors.

[c] Model 3 uses DI rates solely based on the micro variables WSTATOR, SEEKWORK and SEEKREAS.

[d] Model 4 also includes the interaction between UI benefits replacement rate and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[e] Model 5 also includes the interaction between the implicit tax rate on self-employment and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[f] Model 10 additionally includes the first lag of the endogenous variables as instruments.

[g] Model 11 additionally includes the DI compensation and integration indices of OECD (2009a) as instruments. Indices are unavailable for France and Italy.

[h] Hansen's J statistic H0: valid instruments.

So far, we have assumed that the endogenous independent variables can be explained by structural labor market institutions. However, unemployment can also be explained by cyclical effects (Nickell et al. 2005). To account for such cyclical effects, we use labor demand shocks[41] as an additional instrumental variable as well. The results (not reported here) are highly robust to the inclusion of this cyclical instrument.

Sensitivity of the dependent variable 5.5

Section 5.4 shows that early retirement is increased by higher rates of unemployment and decreased by higher rates of part-time employment. We do not find effects for the disability rate. Self-employment rates are largely insignificant implying that the self-employment rate does not decrease early retirement. The current section provides evidence on the robustness of the results to the use of variations on the measure of labor market withdrawal.

Employment rates 5.5.1

Firstly, we compare the results of Model 2 using activity rates (similar to Blake and Sangnier 2011, Blondal and Scarpetta 1998, Duval 2003, Johnson 2000) to a model that identifies retirement from employment rates (as suggested by Ebbinghaus 2006). Mentioned earlier in the paper, using activity rates (including employment and unemployment) assumes that unemployment is not considered to be early retirement while using employment rates assumes that unemployment is similar to non-participation

[41] Following Nickell et al. (2005) labor demand shocks are modeled by including the estimated residual ($\hat{\varepsilon}_\tau$) of the following equation as a variable in the first-stage regressions:

$$ER_\tau = \theta_0 + LCG_\tau\theta_1 + RGDPG_\tau\theta_2 + \varepsilon_\tau \tag{5.9}$$

Where ER is the employment rate for people aged 15-64, LCG is the growth of labor costs and $RGDPG$ is the growth of real GDP. The equation is estimated separately for each of our 13 countries using OLS with Newey-West standard errors to correct for autocorrelation. The results of our baseline specification are robust for different assumed lags of autocorrelation in the labor demand shock equation.

Table 5.5: Retirement indicator based on employment rates[a]

	Men	Women
Disability rate 55-64	-0.94	1.69
	(1.88)	(2.54)
Unemployment rate 55-64	2.24***	2.90***
	(0.77)	(0.87)
Self-employment rate 55-64	-1.18	-0.99
	(1.83)	(2.39)
Part-time employment rate 55-64	-1.51**	-1.45**
	(0.59)	(0.70)
Fixed effects	Yes	Yes
Observations (N x T)	149	149
Hansen's J statistic (p-value reported)[b]	0.56	0.41

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions includes the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Hansen's J statistic H0: valid instruments.

at older ages. Table 5.5 shows that the results are robust for the use of either employment- or inactivity rates in the proposed retirement indicator, although the effect of part-time employment is more pronounced among women now.

5.5.2 Age-windows

Secondly, we analyze the robustness of the estimation results by using retirement indicators that are disaggregated to smaller age-spans. Table 5.6 and table 5.7 show the results of the disaggregated retirement indicators for men and women respectively. For both men and women, the estimation

results of Model 2 in table 5.3 and table 5.4, which is Model 1 in table 5.6 and table 5.7, show similar patterns: disability rates are insignificant, unemployment is positive and significant, self-employment is insignificant and part-time employment is negative and significant in most cases.

For men, the magnitude of the coefficients indicates that the effect of unemployment on early retirement is the largest for the 60-64 year olds (Model 3 and Model 6). Regarding part-time employment, we observe that the effects on early retirement are the largest for the total group of men aged 55-64. Nevertheless, patterns vary only marginally between Model 1-6.

Among women, we find similar results: unemployment rate effects seem to be larger for the group of women aged 60-64. Comparing participation of women aged 55-64 and 55-59 to the group of 25-54 did not result in finding significant effects of part-time employment. However, comparing women aged 55-64, 55-59 and 60-64 to a smaller basis (e.g. 50-54, 50-54 and 55-59 respectively) indicates that part-time employment does decrease early retirement significantly.

Hours decisions 5.5.3

The estimation results so far suggest that part-time employment reduces early retirement and this seems to be mainly driven by voluntary part-time employment. This would suggest that inducing part-time employment by partial retirement programs increases labor force participation at older ages. However, such partial retirement programs are only efficient if they induce people to work part-time who would otherwise have fully retired. Such programs are inefficient if it induces people to work part-time who would have worked full-time in absence of partial retirement possibilities.

Results in micro studies are ambiguous with respect to the efficiency of part-time retirement options. Support for a positive effect of part-time retirement on hours worked is found by Wadensjo (2006) (using Swedish data), Kapteyn et al. (2007) (using Dutch data) and Kantarci and Van Soest (2008) (using Dutch data). Allen et al. (2001) (using US data) and Ilmakunnas and Ilmakunnas (2006) (using Finnish data) find negative effects on hours.

Table 5.6: Sensitivity of the results to decreases in the age-window of the dependent variable, men[a]

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
	55-64 to 25-54	55-59 to 25-54	60-64 to 25-54	55-64 to 50-54	55-59 to 50-54	60-64 to 55-59
Disability rate 55-64	-1.48	-0.66	-0.57	-2.78	-0.98	-1.18
	(2.32)	(2.19)	(2.16)	(3.24)	(2.01)	(1.86)
Unemployment rate 55-64	1.61*	1.48**	3.19***	1.72	1.84**	2.55**
	(0.87)	(0.74)	(0.97)	(1.30)	(0.85)	(1.00)
Self-employment rate 55-64	-1.92	-1.45	-2.31	-3.22	-1.58	-2.36
	(1.81)	(1.84)	(2.01)	(2.65)	(2.08)	(1.79)
Part-time employment rate 55-64	-1.70**	-1.57**	-1.72*	-1.99*	-1.48**	-1.52**
	(0.80)	(0.73)	(0.64)	(1.11)	(0.67)	(0.61)
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations (N x T)	144	144	149	144	149	149
Hansen's J statistic (p-value reported)[b]	0.83	0.69	0.59	0.86	0.41	0.55

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions include the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Hansen's J statistic H0: valid instruments.

Table 5.7: Sensitivity of the results to decreases in the age-window of the dependent variable, women[a]

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
	55-64 to 25-54	55-59 to 25-54	60-64 to 25-54	55-64 to 50-54	55-59 to 50-54	60-64 to 55-59
Disability rate 55-64	0.76	1.33	1.53	-1.14	0.77	0.12
	(1.71)	(2.42)	(2.28)	(1.78)	(2.57)	(1.71)
Unemployment rate 55-64	2.49***	1.91	3.89***	2.18***	2.04**	2.77**
	(0.88)	(1.23)	(0.91)	(0.68)	(0.81)	(1.07)
Self-employment rate 55-64	0.77	1.51	-1.97	-0.98	-1.53	-2.32
	(1.44)	(1.89)	(2.25)	(1.46)	(2.26)	(2.02)
Part-time employment rate 55-64	-0.93	-1.12	-1.25**	-1.32**	-1.42**	-1.13**
	(0.65)	(0.84)	(0.63)	(0.54)	(0.64)	(0.52)
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations (N x T)	144	144	149	144	149	149
Hansen's J statistic (p-value reported)[b]	0.31	0.33	0.65	0.36	0.30	0.48

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions include the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Hansen's J statistic H0: valid instruments.

To address this issue, we construct a macro variable indicating the average actual hours worked by persons aged 55-64 (including people that do not work). This macro variable measures the number of hours actually worked during the reference week in the main job.[42] Table 5.8 presents the cross-country trends in hours worked among persons aged 55-64 and shows that the average number of hours worked among the total of men is substantially higher than among women in most countries. On the other hand, the growth in the actual hours worked has increased more substantially among women. Greece, Ireland, Portugal and Norway even have a decreasing trend in working hours among men and an increasing trend among women.

The largest increases in average hours worked for both men and women can be found in Finland and the Netherlands with increases of more than five hours on average. As depicted in figure 5.1, these countries also showed relatively strong increases in part-time employment in the period 1995-2008. To test the effect of part-time employment on labor supply decisions at the intensive margin, we use the same analytical framework as in the case of labor supply effects at the extensive margin. Instead of using the retirement indicator based on activity rates, we now use the average actual hours worked as a dependent variable in the regression analysis.

The estimation results indicate that an increase in the unemployment rate of one percentage point decreases the average hours worked by about 0.5 hours among men and 0.6 hours among women. More interestingly, the estimation results show that a higher part-time employment rate not only increases the labor supply of older workers at the extensive margin but that part-time employment does not decrease the average number of hours worked either. A one percentage point higher part-time employment rate increases the average actual number of hours worked by about 0.4 hours among men. The effect is smaller, or even absent, among women, but the results indicate that part-time employment possibilities do not have a negative effect on the total labor supply at older ages.

[42] This variable is based on the *HWACTUAL* variable included in the LFS micro data.

Table 5.8: Average number of hours worked, 55-64

	Men			Women		
	1995	2008	Change	1995	2008	Change
Austria	16.0	19.0	3.0	6.0	8.2	2.2
Belgium	12.7	14.6	1.9	3.9	6.6	2.7
Denmark	21.6	22.0	0.4	9.4	14.2	4.8
Finland	9.9	16.8	6.9	9.6	15.2	5.6
France	13.1	13.9	0.8	7.6	9.7	2.1
Germany	18.9	22.6	3.6	8.5	11.7	3.2
Greece	25.7	24.5	-1.2	8.9	9.7	0.8
Ireland	26.0	25.5	-0.5	5.7	10.6	4.9
Italy	17.0	16.7	-0.3	4.6	7.0	2.4
Netherlands	14.1	19.7	5.6	3.4	9.0	5.6
Norway	26.0	23.2	-2.8	14.6	15.8	1.2
Portugal	24.4	21.5	-2.9	12.0	13.4	1.4
Spain	18.4	21.6	3.2	6.0	8.7	2.7
Sweden	20.2	23.7	3.5	15.1	18.1	3.0
UK	20.7	23.5	2.8	9.0	11.9	2.9
Mean	19.0	20.6	1.6	8.3	11.3	3.0

Source: Own calculations based on Eurostat (2014). The first available year for Germany is 2002.

Table 5.9: Labor supply effects at the intensive margin, 55-64[a]

	Men			Women		
	Model 1	Model 2[b]	Model 3[c]	Model 1	Model 2[b]	Model 3[c]
Disability rate 55-64	0.44	0.36	-0.37	-0.57	-0.02	-0.15
	(0.57)	(0.41)	(0.37)	(0.69)	(0.24)	(0.26)
Unemployment rate 55-64	-0.50*	-0.47**	-0.58***	-0.70***	-0.59***	-0.63***
	(0.29)	(0.21)	(0.19)	(0.24)	(0.14)	(0.14)
Self-employment rate 55-64	0.67	0.35	0.75	0.34	-0.08	0.08
	(0.64)	(0.37)	(0.62)	(0.67)	(0.21)	(0.40)
Part-time employment rate 55-64	0.47**	0.30**	0.38*	0.30	0.22***	0.26**
	(0.20)	(0.12)	(0.20)	(0.20)	(0.08)	(0.13)
Fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations (N x T)	149	149	149	149	149	149
Hansen's J statistic (p-value reported)[d]	0.87	0.24	0.51	0.38	0.45	0.40

[a] HAC standard errors using Bartlett kernel bandwith=2 to correct for autocorrelation in parentheses. * Significant at the .10 level; ** at the .05 level; *** at the .01 level using t-statistics. Countries included: Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and the United Kingdom. Germany, Greece and Luxembourg have one or more empty variables. All first-stage regressions include the instruments: UI benefits replacement rate, EPL_{diff}, Implicit tax rate on self-employment, ALMP spending, DI spending, and SA benefits replacement rate.

[b] Model 2 also includes the interaction between UI benefits replacement rate and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[c] Model 3 also includes the interaction between the implicit tax rate on self-employment and EPL_{diff} and the interaction between the implicit tax rate on self-employment and EPL_{diff}.

[d] Hansen's J statistic H0: valid instruments.

Summary and discussion 5.6

In many European countries, the labor market participation of older workers is considerably lower than the labor market participation of prime-age workers. However, this gap between older and prime-age workers declined considerably between 1995 and 2008. One of the factors that may have contributed to the decline in early labor market withdrawal is the rise of non-standard forms of employment among older workers. Since non-standard employment provides downwards flexibility in working hours, older workers with a relatively strong preference for leisure could use these forms of employment as a bridge between full-time employment and retirement. Moreover, non-standard employment may be chosen because of less generous social insurance programs that were often used as early retirement routes in the past. The existing empirical literature on the relationship between non-standard employment and labor market withdrawal consists of micro-level studies. To analyze the variation in early labor market withdrawal across 13 European countries, this study extends the models employed in macro-level retirement studies with non-standard employment in addition to social insurance schemes. We regress early labor market withdrawal on part-time employment, self-employment, unemployment and disability rates, instrumented by institutional variables.

Our analysis finds that for men, part-time employment is a substitute for full early labor market withdrawal. The results suggest that this is mainly because of the possibility to reduce working hours as we find that specifically voluntary part-time employment induces labor force participation at older ages. Among women, this effect of part-time employment on early retirement is smaller and somewhat more ambiguous. This might be due to the fact that part-time work plays a different role in the careers of men than in the careers of women. During their prime age, men work relatively more full-time than women, as women tend to combine the flexibility of part-time employment with raising children. Hence, men use part-time employment as a step in a gradual transition from full-time employment to retirement, whilst women do not work more in part-time at the end of their career than before (Peracchi and Welch 1994).

With regard to self-employment, our results provide no evidence for substitution effects between self-employment among older workers and early retirement. Our finding that part-time employment may function as a substitute to early retirement, whilst self-employment does not, is in line with the findings of other studies on non-standard employment. Results from recent studies at the micro-level (Been and Knoef 2013, Emmanoulidi and Kyriazidou 2012) indicate that older workers primarily choose for self-employment as a way to end unemployment and much less as a way to reduce working hours in paid employment. In contrast, part-time employment among older workers usually follows after full-time employment. In terms of microeconomic theory, our findings suggest that the combination of leisure and consumption while working in part-time gives on average more utility to older workers than retiring early, whereas the combination of leisure and consumption while being self-employed does not. This difference may be owing to several factors. For instance, starting a new business requires a certain investment in terms of working hours, whilst the income is often uncertain.

Furthermore, we find complementary effects between unemployment rates and early retirement of both men and women. Unemployment among older workers contributes to early labor market withdrawal. In contrast, our results provide no evidence for complementary effects between disability rates and early retirement. However, this result should be taken with caution, because the data on the beneficiaries of disability benefits might be troublesome due to cross-country incomparability as well as breaks in the LFS data in some countries.

As a wider implication, our results suggest that facilitating part-time work might contribute to higher labor market participation among older workers at the extensive margin. However, facilitating part-time work could also induce a reduction in working hours among persons who would otherwise have remained working in full-time employment. Our analysis suggests that increases in part-time employment did not have negative effects on the labor supply at the intensive margin across countries. For men, the results even suggest clear positive effects. This indicates that part-time work schemes may actually increase the labor supply at both the extensive and the intensive margin at older ages.

Descriptive statistics 5.A

Table 5.10: Dependent and independent variables (raw data)

Variable	Obs.	Mean	S.D.	Source
Dependent variables				
Activity rate 55-64 (males)	205	56.89	11.62	Eurostat (2014)
Activity rate 25-54 (males)	205	92.34	1.42	Eurostat (2014)
Activity rate 55-64 (females)	205	36.47	15.17	Eurostat (2014)
Activity rate 25-54 (females)	205	75.01	8.91	Eurostat (2014)
Average hours worked 55-64 (males)	203	19.76	4.31	Eurostat (2014)
Average hours worked 55-64 (females)	203	9.70	3.84	Eurostat (2014)
Endogenous independent variables				
Disability rate 55-65	183	9.53	6.17	Eurostat (2014)[a]
Unemployment rate 55-64	210	5.39	3.12	Eurostat (2014)
Self-employment rate 55-64	202	18.90	7.65	Eurostat (2014)[a]
Part-time employment rate 55-64	202	21.15	11.39	Eurostat (2014)[a]
Voluntary part-time employment 55-64	201	20.47	11.16	Eurostat (2014)[a]
Instrumental variables				
EPL_{diff}	210	0.12	1.24	OECD (2013a)
ALMP expenditure (% GDP)	210	0.90	0.48	OECD (2012b)
DI expenditure (p.c.)	210	880.02	485.32	OECD (2012b)
UI replacement rate	208	60.48	13.00	Van Vliet et al. (2012)
SA replacement rate	195	46.30	8.59	Wang and Van Vliet (2014)
Implicit tax rate on self-employment	174	15.76	6.20	Eurostat (2013)
DI compensation index	180	26.36	4.08	OECD (2009a)
DI integration index	180	18.69	5.49	OECD (2009a)
Additional control variables				
Statutory retirement age (males)	210	64.82	1.58	OECD (2012a)
Statutory retirement age (females)	210	63.23	2.89	OECD (2012a)
Implicit tax continued work (55+)	76	43.14	20.75	OECD (2013b)
Implicit tax continued work (60+)	76	52.49	30.44	OECD (2013b)

[a] Own caluclation.

Table 5.11: Pairwise correlation coefficients of endogenous variables and instruments[a]

	DI rate 55-64	UI rate 55-64	SE rate 55-64	PT rate 55-64
EPL_{diff}	0.19	0.10	-0.41	0.44
	(0.01)	(0.13)	(0.00)	(0.00)
ALMP expenditure	0.33	0.23	-0.50	0.44
	(0.00)	(0.00)	(0.00)	(0.00)
DI expenditure	0.70	-0.13	-0.77	0.63
	(0.00)	(0.05)	(0.00)	(0.00)
UI replacement rate	0.28	0.15	-0.43	0.25
	(0.00)	(0.03)	(0.00)	(0.00)
SA replacement rate	0.06	-0.36	0.03	0.00
	(0.42)	(0.00)	(0.69)	(0.95)
Implicit tax rate on self-employment	0.37	0.00	-0.35	0.08
	(0.00)	(0.98)	(0.00)	(0.28)
DI compensation index	0.52	0.04	-0.49	-0.08
	(0.00)	(0.60)	(0.00)	(0.30)
DI integration index	0.43	0.00	-0.78	0.46
	(0.00)	(0.95)	(0.00)	(0.00)

[a] P-value in parentheses.

6 The necessity of self-employment towards retirement: Evidence from labor market dynamics and search requirements for unemployment benefits

Abstract

This paper investigates whether individuals at the end of working life choose self-employment out of necessity and to what degree job search requirements for unemployment benefits induce people to become self-employed. For this purpose we model labor market transitions at older ages using a dynamic multinomial logit model with unobserved heterogeneity. The results indicate that at the end of the career individuals with a weak labor market position have a relatively high probability to become self-employed, e.g. to end or avoid a period of unemployment or inactivity (necessity driven self-employment). Contrasting some earlier work, the results do not suggest that self-employment is used as a gradual retirement route for employees. A difference-in-differences analysis shows that job

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search requirements for unemployed older workers increased the outflow from unemployment and decreased the inflow into unemployment, but did not increase self-employment out of necessity or opportunity.

6.1 Introduction

In virtually all OECD countries, labor force participation rates of the 50+ population decreased in the period from the 1960s to the mid-1990s (OECD 2011c). This was partially due to generous unemployment insurance, disability insurance and early retirement schemes (Gruber and Wise 1998).[1] Since the mid-1990s aging has raised concerns about the sustainability of the welfare state and social insurance reforms have been undertaken to increase the labor force participation of the 50+ population. As a result, the share of people being active in both paid-employment[2] and self-employment increased.

This paper focuses on self-employment at older ages and the introduction of job search requirements for unemployed older workers. Interestingly, self-employment is found to be relatively high among the 50+ working population, compared to other age groups (Hurd 1996, Karoly and Zissimopoulos 2004, Zissimopoulos and Karoly 2007). Taking into account self-employment is therefore important for understanding pathways to retirement (Zissimopoulos and Karoly 2009).

This paper's contribution to the literature is twofold. First, this study contributes to the literature on the importance of necessity and opportunity driven self-employment. In the literature, two main hypotheses have risen to explain self-employment at older ages. First, self-employment may be chosen out of necessity, to end or to avoid unemployment.[3] The 50+ population particularly faces difficulties finding a new job once unemployed

[1] Country-specific analyses of the effects of such schemes on early retirement can be found in Bould (1980), Hogarth (1988), Ruhm (1995), Riphahn (1997), Kerkhofs et al. (1999), Hernoes et al. (2000), Roed and Haugen (2003), Friedberg and Webb (2005), Van Vuren and Van Vuuren (2007), Euwals et al. (2010), Euwals et al. (2012), De Vos et al. (2012).

[2] Defined as being an employee.

[3] E.g. Taylor (1999), Reize (2000), Earle and Sakova (2000), Kuhn and Schuetze (2001), Kellard et al. (2002), Rissman (2003) and Glocker and Steiner (2007).

(Chan and Stevens 2001, Maestas and Li 2006). Second, self-employment may be chosen as an opportunity to reduce working hours and enhance gradual retirement.[4] To investigate the nature of self-employment at older ages we test 1) whether transitions from unemployment to self-employment are important and increase with age,[5] 2) whether high unemployment rates push workers from paid-employment to self-employment,[6] and 3) whether the introduction of job search requirements for unemployed older workers increases self-employment. For the last test we use a Dutch UI reform which introduced job search requirements for unemployed persons between the age of 57.5 and 63 as from January 2004. Before this reform unemployed older workers did not have to search for a job in order to receive unemployment benefits. The reform implied an exogenous and unanticipated shock in the attractiveness of unemployment as a pathway to retirement. Whereas Lammers et al. (2013) and Hullegie and Van Ours (2013) investigate the effect of this reform on the outflow from welfare and substitution effect with regard to disability and early retirement,[7] we focus on the effect of mandatory search requirements on the entry of self-employment. Self-employment may increase when unemployment becomes less attractive as an exit route to retirement. As far as we know, there are no other studies that investigated the effect of job search requirements on substitution between unemployment and self-employment as an exit route to retirement.[8]

[4]This is suggested by Fuchs (1982), Hurd (1996), Bruce et al. (2000), Morris and Mallier (2003), Zissimopoulos and Karoly (2007), Giandrea et al. (2008), and Gu (2009).

[5]Parker and Rougier (2007) find that transitions from unemployment to self-employment are relatively important and argue that this indicates necessity-driven self-employment at older ages.

[6]Several studies find that high unemployment rates increase self-employment propensities, e.g. Benedict and Hakobyan (2008), Kim and Cho (2009), and Congregado et al. (2012). This latter effect is known as the *recession push hypothesis*. This hypothesis is, however, not confirmed in all papers (Moore and Mueller 2002 and Tapia 2008). Among others, Carrasco (1999) finds that self-employment becomes more attractive when the economic situation improves (the *prosperity pull hypothesis*).

[7]Lammers et al. (2013) and Hullegie and Van Ours (2013) both find that the 2004 UI reform significantly increased exits from unemployment to paid-employment. Lammers et al. (2013) also find substitution effects between unemployment insurance and disability insurance.

[8]For an overview of the literature regarding the effects of job search requirements in unemployment, see Fredriksson and Holmlund (2006).

Our second contribution concerns the effect of job search requirements on the inflow to unemployment. We expect that the introduction of search requirements for unemployed older workers lowers the inflow into unemployment, since job search requirements make unemployment a less attractive exit route to retirement. Other studies that investigate the inflow into unemployment are focused on entrance requirements to unemployment insurance (e.g. Christofides and McKenna 1996, Green and Riddell 1997) and on the level and/or duration of benefits (e.g. Andersen and Meyer 1997, Lalive et al. 2006, Tuit and Van Ours 2010, Winter-Ebmer 2003). Lalive et al. and Tuit et al., for example, focus on unemployed older workers and show that benefit duration affects the inflow to unemployment insurance. The bulk of the literature on search requirements is focused on the effects of exiting unemployment instead of the inflow to unemployment (Fredriksson and Holmlund 2006).

This paper analyzes labor market transitions using a dynamic multinomial logit model.[9] This model allows us to study the pathways through which people enter self-employment, to study the effect of the unemployment rate on transitions to self-employment, and to study the effect of the introduction of job search requirements on labor market transitions using a difference-in-differences approach. We correct for unobserved heterogeneity by allowing for correlated random effects (Wooldridge 2010) and we take into account the initial conditions problem by using the method of Wooldridge (2005). Estimating a dynamic multinomial logit model avoids a possible sample selection bias, which may occur when considering binomial estimates for a transition. To estimate the model, the paper takes advantage of the long panel dimension of the Dutch Income Panel data (1989-2009). The Dutch Income Panel is a large administrative dataset and since we are not able to estimate the model for all observations at once, we use two subsamples of the data (such that all observations are used) and apply minimum distance.

Our main finding is that at the end of the career unemployed individuals have a relatively high probability to enter self-employment (necessity driven) and this effect is found to be significantly increasing

[9]This model has also been used by Cappellari et al. (2010), Constant and Zimmerman (2004), Caliendo and Uhlendorff (2008) and Martinez-Granado (2002).

with age. For men in paid-employment the results show significant evidence for the recession push hypothesis. For inactive men and women in paid-employment, on the other hand, we find that a low unemployment rate increases the probability to enter self-employment. At lower ages, self-employment entry is most likely from inactivity. In the highest age-category, self-employment entry from unemployment and inactivity are not significantly different. Introducing job search requirements for the unemployed at the end of their working life increased exits from unemployment. This reform, however, did not increase self-employment out of necessity (we find no significant increase in flows from unemployment to self-employment due to the reform). Finally, job search requirements have decreased the inflow to unemployment.

The structure of the paper is as follows. The next section describes the Dutch unemployment insurance system. Section 6.3 presents the model, and section 6.4 describes the data. Section 6.5 reports the estimation results, after which section 6.6 provides some discussion and section 6.7 concludes the paper.

Unemployment insurance towards retirement 6.2

As this paper focuses on self-employment and unemployment as exit routes to retirement, this section provides an overview of the Dutch UI benefit system. In the 1990s unemployment was an attractive exit route for older workers because of generous arrangements and easy eligibility rules. As from the age of 57.5 people had the possibility to use UI benefits up to the mandatory retirement age without having to search for a job. Unemployment was, therefore, used frequently as an exit route to retirement. The number of UI beneficiaries expanded and, in light of the aging population, reforms have been undertaken.[10]

This paper investigates the effect of a UI reform introduced on January 1st 2004, which implied that unemployed persons older than 57.5 years

[10]For an international comparison of unemployment as an early retirement route, see Gruber and Wise (1998).

were no longer exempted from the requirement to search actively for a job. Search requirements involve that persons in unemployment 1) have a mandatory intake meeting at the unemployment office, where individual criteria are made regarding the expected activities undertaken during unemployment that are *ex post* testable,[11] 2) have the obligation to accept suitable job-offers, where suitable job offers are defined by the educational level and the time spent in unemployment, 3) have to make a sufficient number of applications,[12] where sufficiency is individually determined and related to the labor market, the number of available vacancies and personal health, 4) have to participate in educational programs and job search assistance when they are assumed to not to be able to find work within six months, and 5) have regular report meetings every 4-6 weeks in addition to the mandatory intake meeting and the follow-up to explain the further procedures.

The baseline from which individual arrangements are made is the requirement of applying for a job once a week on average. An automatic exception is made for individuals starting their own business. Furthermore, exceptions are made for persons participating in care or volunteering for at least 20 hours per week for a period of at most six months, individuals taking part in an educational program, people of age 64, or persons older than 62 years and 2 months who already received UI benefits for at least a year in 2004. The first two exceptions are made because they may increase the probability to find a job. The latter two exceptions are made because of a transitory regime. The strictness of job monitoring in the Netherlands is high[13] and due to the risk of substantial financial sanctions we can reasonably assume people to be complying with the search requirements (Verveen et al. 2005). The reform also implied that, after some time, people have to accept all job offers irrespective of their educational level.

[11]The employability of an individual is determined by objective characteristics such as profession, education, age and experience as well as the subjective impression of the caseworker during the interview.

[12]The following options are considered to be an application: letter, e-mail, phone call or nuncupative contact with a company, registering at an agency, having a job interview and doing an assessment.

[13]From an international perspective, Venn (2012) ranks the Netherlands among countries with a high strictness of job search monitoring. The OECD indicator suggests that monitoring job search is stricter in the Netherlands than in countries such as the US, Canada and Scandinavian countries.

Fulfilling above mentioned requirements, together with eligibility requirements that people have worked at least 26-out-of-36 weeks, gives persons the right to receive UI benefits. Until October 2006 the maximum UI benefits duration for receiving 70% of previous earnings was age-dependent and amounted to a maximum of 42, 48 and 60 months for persons aged 50-54, 55-59 and 60-64 respectively. Until August 2003 persons aged 57.5+ could, in principal, even extend the benefit period up to the age of 65 by using extended UI benefits. These extended UI benefits amounted 70% of minimum wage. From August 2003, extended UI benefits were abolished simultaneously with the introduction of the so called IOAW-benefits[14] targeted at unemployed 50+ individuals. The only difference between the extended UI benefits and the IOAW for older unemployed is that receiving the latter depends on the income of the spouse while extended benefits were unconditional on the income of the spouse. Single households are therefore indifferent between receiving extended UI benefits or IOAW benefits.

In October 2006, both benefits and the duration of benefits were moderated for all UI recipients and the maximum UI benefit duration was made conditional on the employment history, with a maximum of 38 months. However, after 38 months of UI benefits, unemployed elderly can obtain social benefits from *IOAW* and the *IOW*[15] (implemented in August 2003 and December 2009, respectively) to complement household income up to subsistence level without asset-based means testing (and for the IOW also unconditional on the income of a partner). Furthermore, self-employed elderly individuals with a low income who have to stop their business can receive benefits to complement their income up to subsistence level, without the strict asset-based means testing from social assistance benefits.[16]

[14] *Wet inkomensvoorziening oudere en gedeeltelijk arbeidsongeschikte werkloze werknemers.*

[15] *Inkomensvoorziening oudere werklozen.*

[16] This program is called the IOAZ (*Wet Inkomensvoorziening oudere en gedeeltelijk arbeidsongeschikte gewezen zelfstandigen.*)

6.3 Model

6.3.1 Exit routes to retirement

This section describes the model we use to investigate labor market transitions among the 50+ population. The exit route to retirement can be seen as the outcome of a maximization process, in which individuals reevaluate their optimal labor market status each period, given their preferences and the constraints that coincide with each labor market state. Individuals compare utility streams associated with different exit routes and choose the alternative with the highest utility stream. More specifically, we define the inter-temporal utility of individual i as follows:

$$U_{it} = \sum_{\tau=t}^{T} (1+\rho)^{t-\tau} u_\tau(c_{i\tau}, l_{i\tau}, j_{i\tau}; s_{i\tau}, v_{i\tau}) \tag{6.1}$$

where $c_{i\tau}$ and $l_{i\tau}$ denote consumption and leisure of individual i in time period τ implicitly defined by labor market state j. ρ is the discount factor and T the time horizon of the individual. In our model we distinguish between four mutually exclusive labor market states: paid-employment ($j = 1$), self-employment ($j = 2$), unemployment insurance ($j = 3$), and inactivity ($j = 4$).[17] Each labor market status is associated with it's own consumption and leisure possibilities, but labor market status itself may also influence the utility function directly. E.g., conditional on leisure and consumption, some people receive a higher utility from self-employment than from paid-employment, due to characteristics of self-employment such as the independence and flexibility that self-employment provides.

Social insurance rules $s_{i\tau}$ that hold for individual i in period τ influence the exit route to retirement. An increase of job search requirements, for example, decreases the amount of leisure and so the value of unemployment as a retirement route. Furthermore, transitions from self-employment or inactivity to unemployment are not possible because only persons in paid-employment are eligible for UI benefits. Finally, observed and unobserved characteristics $v_{i\tau}$ influence the utility function indirectly through

[17] Inactivity includes individuals in disability, welfare, early retirement, and individuals without personal income.

preferences. For example, age, the number of children in the household, and education may influence the utility perceived from consumption and leisure.

Equation (6.1) provides a guideline for the empirical specification of the model. It shows that individuals choose the exit route that maximizes their utility over consumption, leisure, and labor market status. Furthermore, individual characteristics and social insurance rules affect current and future labor market statuses. For the empirical implementation of the problem, like Blau (1998) and Mastrogiacomo et al. (2004), we approximate the value function U_{it} for individual i who chooses labor market status j at time t with a linear function:

$$V_{ij}(t) = X_{it}\beta_j + Z_{it-1} \otimes [1 \quad AGE'_{it} \quad YEAR'_{it}]\gamma_j + Z_{it-1}UR_t\theta_j + D_{ijt} + \mu_{ij} + \epsilon_{ijt}, \quad (6.2)$$

where X_{it} is a vector of observed personal and household characteristics that influence preferences as shown in (6.1). Z_{it-1} is a vector of dummy variables indicating lagged labor market status. AGE_{it} and $YEAR_{it}$ are vectors of dummy variables indicating age and year categories. These are interacted with Z_{it-1} to allow for mobility differences across age and periods. UR_t is the unemployment rate in period t, which we interact with Z_{it-1} to take into account that the unemployment rate may affect individuals with various previous employment states differently. The treatment variables function D contains variables and interactions that we use to identify the effect of the job search requirements introduced in 2004 and will be explained in section 6.3.2.

Finally, the terms μ_{ij} describe individual specific unobserved heterogeneity and ϵ_{ijt} are i.i.d. error terms, which we assume to be independent of the explanatory variables and to follow a Type I extreme value distribution. Hence, the probability for individual i to have labor market status j at time $t > 0$ can be written as

$$P(j_t|X_{it}, Z_{it-1}, AGE_{it}, YEAR_{i,t}, UR_t, D_{ijt}, \mu_{i1}, ..., \mu_{iJ}) =$$

$$\frac{\exp(X_{it}\beta_j + Z_{it-1} \otimes [1 \quad AGE'_{it} \quad YEAR'_{it}]\gamma_j + Z_{it-1}UR_t\theta_j + D_{ijt} + \mu_{ij})}{\sum_{k=1}^{J} \exp(X_{it}\beta_k + Z_{it-1} \otimes [1 \quad AGE'_{it} \quad YEAR'_{it}]\gamma_k + Z_{it-1}UR_t\theta_k + D_{ikt} + \mu_{ik})}, \tag{6.3}$$

where J denotes the number of mutually exclusive labor market states distinguished in the model. To identify the model, $\beta_1, \gamma_1, \theta_1$ and μ_{i1} are normalized to zero (paid-employment is the reference category). The unobserved heterogeneity or random effects $\mu_i = (\mu_{i2}, \mu_{i3}, \mu_{i4})'$ are assumed to follow a multivariate normal distribution with mean zero and variance Σ_μ.

Introducing unobserved heterogeneity has the advantage that the irrelevance of independent alternatives (IIA) property of the multinomial logit model is avoided. Furthermore, allowing for unobserved heterogeneity within choice possibilities will give true, instead of spurious, state dependence in the model. The initial labor market status Z_{i0} is not fixed or exogenous and, as in most papers, we do not have the entire history of the process generating individual's employment dynamics available. Therefore, the initial conditions problem arises, which is discussed in Heckman (1981). To deal with this problem Heckman (1981) proposed to estimate a static multinomial logit model for the initial state with different slope parameters and without lagged labor market status, simultaneously with the dynamic model. Several studies investigating transitions between multiple states have used this method, e.g. Gong et al. (2000), Uhlendorff (2006) and Cappellari et al. (2010). In this paper we will use an alternative approach, proposed by Wooldridge (2005), to take into account the initial conditions problem. In the method of Wooldridge (2005), individual specific heterogeneity terms are modeled conditional on the initial condition, the initial value of the lagged dependent variable, and the individual mean of time-varying covariates

$$\mu_{ij} = \alpha_{0j} + Z_{i0}\alpha_{1j} + X_i\alpha_{2j} + a_{ij} \quad j = 2, 3, 4 \tag{6.4}$$

where Z_{i0} is the vector of initial conditions and X_i the vector of the individual mean of time-varying covariates. The remaining stochastic element, a_{ij}, is assumed to follow a multivariate normal distribution with

mean zero and variance Σ_a. In other words,

$$\begin{pmatrix} a_{i2} \\ a_{i3} \\ a_{i4} \end{pmatrix} = L \begin{pmatrix} \eta_{i2} \\ \eta_{i3} \\ \eta_{i4} \end{pmatrix} \quad \text{with} \begin{pmatrix} \eta_{i2} \\ \eta_{i3} \\ \eta_{i4} \end{pmatrix} \sim N \left(\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \right), \quad (6.5)$$

where L is the Cholesky matrix of Σ_a which has to be estimated (the unique lower triangular matrix such that $LL' = \Sigma_a$). In this way, we allow for unobserved heterogeneity within and between choice possibilities.

Applying the Wooldridge correction for initial conditions in the way explained above, automatically results in a Correlated Random Effects model (Mundlak 1978). Applying this Correlated Random Effects regression has the advantage of allowing for correlation between observed- and unobserved heterogeneity similar to a fixed effects model, even in an unbalanced panel (Wooldridge 2010).

Akay (2011) studied the performance of the Wooldridge method, compared to the Heckman method. He found that the method proposed by Wooldridge works well for moderately long panels (5-8 periods) and that all methods perform equally well for panels of long duration (longer than 15-20 periods)[18]. For short panels, Rabe-Hesketh and Skrondal (2013) find that the bias practically disappears when the initial-period explanatory variables are included as additional regressors. Examples of other studies that used the Wooldridge approach are Devicienti and Poggi (2011), Michaud and Tatsiramos (2011), Haan and Wrohlich (2011), Buddelmeyer et al. (2010) and Christelis and Sanz-de Galdeano (2011).

Identifying the effects of job search requirements 6.3.2

The 2004 UI reform, described in section 6.2, provides an exogenous source of variability in the data. These search requirements decrease leisure in the unemployment state. This implies that the UI reform makes the value of unemployment relatively lower compared to paid-employment, self-employment and inactivity. As from 2004 individuals of age 57.5 and older are no longer exempted from job search requirements. To infer causal

[18] In this paper we have a long panel of 21 periods available.

effects of job search requirements, we apply a difference-in-differences framework. In this framework, we compare the inflow to and the outflow from unemployment before and after the reform for the 57.5+ population (for whom job search requirements were no longer exempted), relative to those younger than 57.5 (for whom nothing changed). We assume that in absence of the reform there would not be a discontinuous change in labor market transitions for 57.5+ individuals relative to those younger than 57.5 after the reform.[19]

Formally, the difference-in-differences framework is implemented in equation (6.2) using the treatment variable function D which is given by

$$D_{ijt} = [PE_{it-1} \quad UI_{it-1}] \otimes [G_{it} \quad P_{it} \quad G_{it} \cdot P_{it}]\delta_j \tag{6.6}$$

where G_{it} is a dummy variable indicating the treatment group, which is equal to one if a person is between the ages of 58 and 63 (at December 31th) and zero otherwise.[20] Only, due to a transitional regime, persons older than 62 years and 2 months who were already unemployed for a minimum of one year at the time the reform was implemented were not affected by the reform and are classified as belonging to the control group. P_{it} indicates the treatment period (2004-2009), and $G_{it} \cdot P_{it}$ is one for those persons that are treated. Finally, by interacting the treatment variables with indicators for paid-employment (PE) and unemployment (UI) in the previous period, we investigate the effects of the reform on the outflow from unemployment and on the inflow from paid-employment to unemployment.

Lammers et al. (2013), who exploit the same policy reform, notice that anticipation of the policy change can result in selective inflow into unemployment around the time the policy was initiated, but found no evidence of this. Probably, since none of the individuals flowing into UI in 2003 were exempted from the new rules, speeding up the firing procedure

[19] Placebo tests will follow to verify this common trends assumption.

[20] Since we have yearly data we cannot identify effects that start during a year. The smallest bias is introduced when we define individuals to belong to the treatment group as from the year in which they become 58. Taking the year in which people become 57 increases the bias, since all individuals born after June do not reach the age of 57.5 during that year. Furthermore, also those born from January to June have a smaller bias when the treatment group starts as from the year in which individuals become 58.

could not prevent them from the new search requirements after the age of 57.5. Therefore, we can reasonably assume that the introduction of the reform was unanticipated. Another type of anticipation effect may well arose before the reform. If before 2004 unemployed individuals who were close to 57.5 were already reducing their search capacity in anticipation of the removal of the search requirement after the age of 57.5, the labor market transitions of those younger than 57.5 are also affected by the reform. Hullegie and Van Ours (2013) find that individuals already reduced their search intensity about two months prior to the age of 57.5 in the period before 2004, meaning that persons anticipated the abolishment of search requirements at older ages. If indeed the treated group would be all individuals as from the age of 57 and 4 months (57.5 minus 2 months), we would change our definition of the treatment group. We would indicate persons born in January or February to be treated as from the year in which they become 57 (instead of 58), so to reduce the bias resulting from the yearly observations. A robustness check (not reported here) in which the treatment group also consists of persons of age 57 who were born in January or February shows that the results hardly change.

The 2004 UI reform did not change the UI benefit level and -duration, but only introduced mandatory job search requirements that increased the number oblifations to receive unemployment benefits. To make sure that we only measure the effects of the introduction of job search requirements on the first of January 2004 and not the abolition of extended benefits in August 2003, we exploit the fact that the reform of August 2003 did not affect singles (as mentioned in section 6.2) in the robustness checks.

Estimation 6.3.3

We estimate the model's parameters using maximum likelihood. The likelihood contribution of an individual i with observed labor market states $j_1, ..., j_M$ is

$$L_i(j_1, ..., j_M | X, Z, AGE, YEAR, UR, D, a_i; \alpha, \beta, \gamma, \theta, \delta) =$$

$$\prod_{t=1}^{M_i}\prod_{j=1}^{J}\left(\frac{\exp(X_{it}\beta_j + Z_{it-1}\otimes[1 \quad AGE'_{it} \quad YEAR'_{it}]\gamma_j + Z_{it-1}UR_t\theta_j + D_{ijt} + Z_{i0}\alpha_{1j} + X_i\alpha_{2j} + a_{ij})}{\sum_{k=1}^{J}\exp(X_{it}\beta_k + Z_{it-1}\otimes[1 \quad AGE'_{it} \quad YEAR'_{it}]\gamma_k + Z_{it-1}UR_t\theta_k + D_{ikt} + Z_{i0}\alpha_{1k} + X_i\alpha_{2k} + a_{ik})}\right)^{I(j=j_t)} \tag{6.7}$$

where M_i is the last observation for individual i. We do not observe the individual specific effects a_i $(= (a_{i2}, a_{i3}, a_{i4}))$. This term has to be integrated out, such that the likelihood contribution becomes

$$L_i(j_1,...,j_M|X,Z,AGE,YEAR,UR,D,a_i;\alpha,\beta,\gamma,\theta,\delta) = \int_{-\infty}^{\infty} L_i(j_1,...,j_M|X,Z,AGE,YEAR,UR,D,a_i;\alpha,\beta,\gamma,\theta,\delta)da_i \tag{6.8}$$

We evaluate the integral using Maximum Simulated Likelihood (for details, see Gourieroux and Monfort 1993, Hajivassiliou and Ruud 1994). We apply Halton draws instead of random draws, as they are found to give more precise estimation results (Bhat 2001, Train 2000).

Due to our large dataset (164,620 men and 161,487 women) we are unable to estimate our dynamic multinomial logit model with unobserved heterogeneity for all observations at once. Hence, we draw a random sample of individuals. To increase the efficiency of the estimated coefficients we estimate the model on two subsamples of the data, such that all observations are used, and apply minimum distance (Chamberlain 1984), where we restrict the estimates of the two subsamples to be the same. This method is applicable to all kind of situations in which (complicated) models have to be estimated with large data sets.

6.4 Data

6.4.1 Data and definitions

Data are from the Dutch Income Panel Study 1989-2009 (IPO, Inkomens Panel Onderzoek, CBS 2009b), gathered by Statistics Netherlands. IPO is an administrative dataset that contains a representative sample of the Dutch population. About 95,000 individuals are selected, based on their

national security number, and followed over time. Detailed information is available, most particularly from the tax office, on income, wealth, gender, age, marital status, children, ethnicity, homeownership and labor market status.

A major advantage of having administrative data is the number of observations and the high level of representativeness. It is a well-known fact that the rich and the poor are often underrepresented in surveys, but also that self-employed individuals are often underrepresented. Another advantage of IPO is that we have a long time span available (21 years) and that we have no endogenous panel attrition, since panel attrition only occurs as a result of emigration or death.

In this paper we select men and women between the ages of 50 and 63.[21] To define labor market status we use an individuals's main source of income during a year of observation. We make one exception for self-employment, namely, we also indicate someone to be self-employed when the person has a negative profit (a loss) while income from wealth (rents and dividends) is larger than any other component that year. This, for example, allows us to take into account start-ups.[22]

The analysis also uses additional published data of Statistics Netherlands about the macroeconomic unemployment rate and the consumer price index (CPI). The unemployment rate decreased from 6.9% in 1989 to 2.6% in 2009, with peaks in 1994 (7.5%) and 2004 (4.5%).

Descriptive analysis 6.4.2

Table 6.1 describes individual- and household characteristics. We distinguish individuals in the treatment and the control group, in the treatment and control period. Men and women are analyzed separately, because their retirement routes may be quite different. Within control and treatment groups we do not find large differences over time in personal and household characteristics. Only, the share of men and women with a

[21] Individuals of age 64 are excluded from the UI reform that we investigate. Becoming unemployed at the age of 64 implies being exempted from search requirements.

[22] Income from self-employment denotes income from profit, freelancing or from being a director/major shareholder.

partner decreased about 10%-points between the control and the treatment period for the control group.

Labor market statuses, on the other hand, changed substantially between the pre- and post-reform period. Paid-employment increased at the expense of inactivity, especially among women in the treatment group. This can be explained by cohort effects, as found by Euwals et al. (2011). About 10% of the people are self-employed and only about 2-5% of these people receive a substantial amount of labor income in addition to the profit from their business (at least half of their profit). Furthermore, only 10 to 15% of the unemployed received a substantial amount of labor income (at least half of their unemployment benefits). This reassures us that that we do not have to worry about only using the main income source to define labor market status.

Income from wealth offers some information about relative wealth differences between individuals. Since labor market status influences wealth (e.g. wealth may decline in a period of unemployment), we use initial wealth in our analysis. We find that young cohorts receive a higher income from financial wealth than old cohorts and that homeownership has increased among younger cohorts. On the other hand, also mortgages have increased (probably largely due to tax incentives and eased loan restrictions). Also, younger cohorts tend to receive a slightly higher share of their income from wealth from risky assets such as stocks and bonds. Especially in the treatment period.

Transition matrices in tables 6.2 and 6.3 present labor market transitions. The diagonals of table 6.2 show that year to year transitions out of paid-employment, self-employment and inactivity diminished between the control and treatment period. In contrast, yearly transitions out of unemployment increased between the control and treatment period (10% in the treatment group and 17% in the control group). People who leave unemployment move into paid-employment, self-employment and inactivity. In the treatment group transitions from unemployment to self-employment increased from 0.49% to 1.25%. This may be due to the introduction of job search requirements, however, also in the control group we find an increase (from 1.88% to 3.97%). Transitions from unemployment to paid-employment increased from 1.80% to 4.69% in the treatment group

Table 6.1: Descriptive statistics[a]

	1989-2003 (control period)				2004-2009 (treatment period)			
	Age 50-57 (Control group)		Age 58-63 (Treatment group)		Age 50-57 (Control group)		Age 58-63 (Treatment group)	
	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.
Men								
Personal and household characteristics								
Age	53.34	2.28	60.41	1.71	53.48	2.30	60.38	1.69
Birth year	1943	4.83	1936	4.70	1953	2.88	1946	2.28
Immigrant	0.08	0.27	0.07	0.25	0.10	0.30	0.08	0.27
Partner	0.87	0.34	0.96	0.19	0.77	0.42	0.93	0.25
Children	0.17	0.38	0.05	0.22	0.23	0.42	0.05	0.22
Number of children[b]	1.53	0.87	1.57	0.89	1.55	0.76	1.51	0.79
Age youngest child[b]	12.48	4.51	10.73	5.58	12.45	4.28	11.44	5.32
Labor market status								
Paid-employment (PE)	0.65	0.48	0.29	0.45	0.70	0.46	0.42	0.49
Self-employment (SE)	0.12	0.32	0.09	0.28	0.13	0.33	0.10	0.30
Unemployment (UI)	0.02	0.14	0.06	0.23	0.02	0.15	0.04	0.19
Inactive (IA)	0.21	0.41	0.57	0.50	0.15	0.36	0.44	0.50
Partial paid-employment								
SE and PE[c]	0.02	0.14	0.02	0.15	0.03	0.16	0.02	0.14
UI and PE[d]	0.11	0.32	0.04	0.19	0.13	0.34	0.07	0.25
Financial variables (expressed in 2010 euro's using the CPI)								
Income financial wealth (t=0)[e]	636.83	12341.01	562.77	4711.88	1034.79	14542.46	720.33	14974.99
Homeowner (t=0)	0.57	0.50	0.48	0.50	0.67	0.47	0.63	0.48
Income housing wealth (t=0)[f]	-457.32	4678.91	341.83	3711.48	-2037.59	5190.78	-770.87	5237.12
Mortgage (t=0)[g]	66.14	133.67	36.01	82.34	134.01	244.75	73.02	120.75
Risky assets (t=0)[h]	1.45	61.82	1.45	64.16	3.50	111.22	1.59	42.30
Observations	69,916		39,928		31,951		22,825	
Women								
Personal and household characteristics								
Age	53.35	2.28	60.43	1.71	53.46	2.31	60.38	1.69
Birth year	1943	4.82	1936	4.72	1953	2.93	1946	2.28
Immigrant	0.07	0.26	0.07	0.25	0.10	0.30	0.08	0.26
Partner	0.93	0.25	0.99	0.11	0.82	0.38	0.97	0.16
Children	0.09	0.29	0.02	0.14	0.13	0.34	0.02	0.13
Number of children[b]	1.39	0.79	1.61	0.94	1.35	0.64	1.57	0.81
Age youngest child[b]	13.35	4.26	8.74	6.28	13.63	3.77	9.03	6.49
Labor market status								
Paid-employment (PE)	0.33	0.47	0.12	0.32	0.53	0.50	0.23	0.42
Self-employment (SE)	0.07	0.26	0.03	0.18	0.09	0.29	0.07	0.26
Unemployment (UI)	0.02	0.13	0.02	0.13	0.02	0.13	0.02	0.14
Inactive (IA)	0.58	0.49	0.83	0.37	0.36	0.48	0.67	0.47
Partial paid-employment								
SE and PE[c]	0.02	0.15	0.01	0.12	0.05	0.21	0.03	0.17
UI and PE[d]	0.11	0.31	0.04	0.19	0.16	0.36	0.05	0.23
Financial variables (expressed in 2010 euro's using the CPI)								
Income financial wealth (t=0)[e]	971.19	20363.13	833.57	4525.06	1477.91	25362.00	1270.65	28053.39
Homeowner (t=0)	0.54	0.50	0.45	0.50	0.64	0.48	0.59	0.49
Income housing wealth (t=0)[f]	-165.83	4536.95	489.45	3415.68	-1536.87	5348.39	-394.90	5086.26
Mortgage (t=0)[g]	55.09	119.75	29.30	64.23	118.09	392.71	61.26	135.57
Risky assets (t=0)[h]	0.00	0.00	0.00	0.00	3.20	102.92	0.00	0.00
Observations	67,716		40,551		31,095		22,116	

[a] Only 5% of the men and 3% of the women aged between 58 and 63 years between 2004-2009 are in a transitory arrangement (e.g. persons aged 62+ who received UI benefits for at least a year in 2004).

[b] Conditional on having at least one child.

[c] Partial SE shows the percentage of individuals whose main source of income is profit from business, but who also receive a substantial amount of labor income (at least half of profit from business).

[d] Partial UI shows the percentage of individuals whose main source of income are unemployment benefits, but who also receive a substantial amount of labor income (at least half of the unemployment benefits).

[e] Income from financial wealth is the sum of interest and dividends, minus interest payments for debts other than mortgage debt at the household level.

[f] Income from housing wealth is the imputed rent minus the interest payments from mortgages at the household level.

[g] Mortgage shows the mortgage interest payments divided by the rental value of the house at the household level (this information gives some idea about the loan to value).

[h] Risky assets shows the percentage of income from total wealth that is generated by stocks and bonds at the household level.

and from 15.99% to 26.84% in the control group. Among the individuals active in the labor market, self-employment is higher in the treatment than the control group. This may be due to necessity reasons (it is generally more difficult for older men to find a job), but also preferences may play a role (gradual retirement through self-employment). Transitions from paid-employment to self-employment do not change very much but we do observe a decline in the share of employed people moving to unemployment, especially in the treatment group, who were confronted with the search requirements of the 2004 UI reform. For treated men we find that transitions from paid-employment to unemployment declined from 2.49% to 1.41%, compared to a decline only from 1.29% to 1.18% in the control group.

Similar patterns emerge for women. The major difference compared to men is that relatively more women are inactive. Transitions in tables 6.2 and 6.3 are not conditional on observed and unobserved characteristics. Therefore, information on state dependence may be spurious. In the following section we take into account background characteristics and unobserved heterogeneity.

Table 6.2: Average year-to-year transitions, men

Year 1989-2003 (control period)	Age 50-57 (control group) Year t					Age 58-63 (treatment group)				
Year $t-1$	PE	SE	UI	IA	Total	PE	SE	UI	IA	Total
PE	94.60	1.10	1.29	3.01	100.00	73.57	0.98	2.49	22.96	100.00
SE	5.86	86.32	0.13	7.70	100.00	3.67	82.15	0.04	14.15	100.00
UI	15.99	1.88	64.35	17.78	100.00	1.80	0.49	80.61	17.09	100.00
IA	4.54	3.72	0.79	90.95	100.00	1.73	1.23	0.36	96.68	100.00
Total	64.30	12.22	2.23	21.25	100.00	24.44	8.55	5.81	61.21	100.00
Year 2004-2009 (treatment period)	Year t									
Year $t-1$	PE	SE	UI	IA	Total	PE	SE	UI	IA	Total
PE	96.16	0.93	1.18	1.73	100.00	81.17	0.89	1.41	16.53	100.00
SE	3.38	92.27	0.03	4.31	100.00	3.12	85.61	0.29	10.98	100.00
UI	26.84	3.97	53.50	15.69	100.00	4.69	1.25	72.93	21.13	100.00
IA	3.96	3.12	1.05	91.87	100.00	1.32	1.58	0.56	96.53	100.00
Total	69.80	12.92	2.22	15.06	100.00	38.02	10.31	3.80	47.87	100.00

Table 6.3: Average year-to-year transitions, women

Year 1989-2003 (control period)	Age 50-57 (control group) Year t					Age 58-63 (treatment group)				
Year $t-1$	PE	SE	UI	IA	Total	PE	SE	UI	IA	Total
PE	90.46	1.52	1.40	6.63	100.00	72.63	1.95	2.18	23.24	100.00
SE	6.89	83.00	0.15	9.96	100.00	4.96	71.45	0.09	23.49	100.00
UI	14.89	1.85	64.78	18.48	100.00	2.56	0.92	79.67	16.85	100.00
IA	2.50	1.28	0.17	96.05	100.00	0.85	0.60	0.06	98.48	100.00
Total	32.62	7.31	1.66	58.40	100.00	9.93	3.35	1.73	85.00	100.00
Year 2004-2009 (treatment period)	Year t									
Year $t-1$	PE	SE	UI	IA	Total	PE	SE	UI	IA	Total
PE	94.81	1.17	1.06	2.96	100.00	81.56	1.45	1.06	15.93	100.00
SE	4.93	89.82	0.09	5.16	100.00	3.45	84.73	0.09	11.73	100.00
UI	21.04	4.10	55.74	19.13	100.00	6.27	1.42	72.93	19.37	100.00
IA	3.07	1.77	0.54	94.62	100.00	0.80	1.07	0.15	97.98	100.00
Total	53.03	9.92	1.69	35.37	100.00	21.32	7.43	2.02	69.22	100.00

Results 6.5

Estimation results 6.5.1

Tables 6.4 and 6.5 show the estimation results of our baseline model for men and women, respectively.[23] The results provide evidence of self-employment out of necessity among older workers. First, after controlling for individual- and household characteristics as well as unobserved heterogeneity, the results show that between the ages of 54 and 63 unemployed individuals are significantly more likely to enter self-employment than paid-employed individuals and this increases with age (necessity hypothesis I at the end of the table). This is in line with Zissimopoulos and Karoly (2009) who show that propensity of self-employment entry from unemployment and disability relative to paid-employment increases with age among older workers. Second, γ_4 and γ_8 in the self-employment equation do not indicate that transitions from paid-employment to self-employment increase with age, such as the opportunity hypothesis of self-employment as a bridge to retirement would suggest. In fact, the probability of flowing from paid-employment to self-employment even decreases with age among men.

Tables 6.4 and 6.5 show that inactive men of age 50-57 between 1999 and 2009 and inactive women of age 50-53 between 1999-2009 are more likely to become self-employed than their unemployed counterparts (necessity hypothesis II). For women this only holds for the age group 50-53 between 1999 and 2003 (table 6.5). Table 6.6 shows that inactive men who enter self-employment were often depending on income from disability, wealth or the income of a spouse in the previous period while women were often relying on the income of a partner. Furthermore, individuals flowing from disability, early retirement, or social assistance to self-employment had a relatively low income, compared to all people in the same labor market status. This may indicate the necessity of self-employment. Only men for whom income from wealth is the main income source are becoming self-employed more often when they have a relatively large income,

[23] In our estimation procedure we use 50 Halton draws. The baseline results are robust for 100 and 200 Halton draws.

suggesting that not all flows from inactivity to self-employment are driven by necessity.

With regard to the macroeconomic unemployment rate, the results for men show that a higher unemployment rate not only leads to more transitions from paid-employment to unemployment, but also to relatively more transitions from paid-employment to self-employment. This suggests that self-employment is not only chosen to end a spell of unemployment but also as a way of avoiding unemployment, consistent with the *recession push hypothesis* found in Benedict and Hakobyan (2008), Kim and Cho (2009), and Congregado et al. (2012). For women, on the other hand, we find that a higher unemployment rate reduces the probability of flowing from paid-employment to self-employment which is consistent with the *prosperity pull hypothesis* found by Carrasco (1999). The difference between men and women can be explained by the fact that men are more often the main income earner of a household. A higher unemployment rate does not lead to significantly more or less transitions from unemployment or inactivity to self-employment.[24] As expected, people in unemployment are significantly more likely to stay in unemployment when the unemployment rate is high.

In line with Lammers et al. (2013) and Hullegie and Van Ours (2013) the results show that job search requirements for unemployed individuals between the ages of 58 and 63 have increased transitions out of unemployment (δ_2 in the unemployment equation of tables 6.4 and 6.5). Our results show that the introduction of search requirements did not increase transitions from paid employment or unemployment to self-employment, relative to paid employment. Apparently, individuals that are confronted with search requirements are (at least partly) able to find a job. For women we find a significantly negative treatment effect for transitions between unemployment and inactivity. This means that as a result of the treatment, the growth in transitions between unemployment and paid employment is significantly higher than for transitions between unemployment and inactivity. Finally, necessity-hypothesis III in tables 6.4 and 6.5 shows that after the treatment individuals entering self-employment between the ages

[24]The sum of θ_1 and θ_3 and the sum of θ_1 and θ_4 are not significantly different from zero in the self-employment equation.

of 58 and 63 are still significantly more often coming from unemployment than from paid-employment.

In addition to previous research, our approach does not only allow us to investigate the effect of job search requirements on the outflow from unemployment, but also to investigate the effect on the inflow to unemployment. δ_1 in the unemployment equation of tables 6.4 and 6.5 show that the introduction of job search requirements significantly reduced transitions from paid-employment to unemployment. For women we find a significantly weak positive effect of the treatment on transitions from paid employment to inactivity, suggesting substitution effect between unemployment and inactivity as retirement routes.

The lower parts of tables 6.4 and 6.5 show the variances and covariances of the random effects. We allow for flexible correlated random effects that take into account, for example, unobserved differences in education and ambition. When we would not take into account these effect, we would find a higher state dependence (spurious versus true state dependence). The estimates show that the random effect for self-employment plays a significant role and is more important than the idiosyncratic error term (which has a variance of $\pi^2/6$, by normalization). This means that, compared to paid-employment, time invariant unobserved characteristics play a substantial role in the choice for self-employment. The random effect for unemployment is only significant for women and the random effect for inactivity is significant for both men and women. These random effects are less important than the idiosyncratic error term. The covariances of the random effects for self-employment and unemployment are significantly positive, meaning that unobserved characteristics that are related with a high probability of self-employment are also related with a high probability of unemployment. The covariance of the random effect for self-employment and inactivity is positive for men and negative for women. This difference between genders may be explained by the fact that for women inactivity often means having no personal income (relying on the income of a spouse), whereas for men inactivity often means early retirement or disability. Finally, for women we find a significantly positive covariance between unemployment and inactivity. This is reasonable as both states imply non-participation. The significance of the covariances

show us that it is important to model self-employment, unemployment and inactivity simultaneously.

In table 6.7 we extend the baseline model with financial variables and health status in the initial state. We use the initial state since, for example, wealth may decline when people become unemployed or inactive or when people start their own business (endogeneity). Also, liquidity constraints may be important for transitions to self-employment. Panel A shows that homeownership and financial wealth are associated with a higher probability of entering self-employment for men. For women, only homeownership is associated with a higher probability to enter self-employment. It is interesting to see that mortgages are negatively associated with inactivity. The financial variables are endogenous, e.g. risk loving individuals may hold more risky assets and may be more likely to be self-employed. The treatment effects, however, hardly change with the inclusion of financial variables.

Health, measured by receiving disability benefits in the first period of observation, is negatively associated with self-employment and positively associated with unemployment and inactivity, compared to paid-employment (panel B in table 6.7). This is in line with Parker and Rougier (2007), who show that a poor health status decreases the probability of self-employment entry relative to retirement entry among older persons. Results of Zissimopoulos and Karoly (2007), however, indicate that limiting health conditions increase the probability of self-employment entry from paid-employment among older persons.

6.5.2 Robustness checks

This section presents three types of robustness checks, 1) two placebo tests to verify the common trends assumption, 2) robustness checks with regard to the time span of the sample around the treatment, and 3) a robustness check that ensures us to measure the effects of the introduction of job search requirements and not the abolition of extended benefits.

In the first placebo test we estimate the treatment effects for people of age 56-57, just prior to the group that actually received the treatment. In the second placebo test we estimate the treatment effects for the period

Table 6.4: Estimation results baseline model[a] (men)

		Self-employment		Unemployment		Inactivity	
Effects relative to paid-employment		Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Mobility							
γ_1	SE_{t-1}	3.40***	0.40			2.68***	0.36
γ_2	UI_{t-1}	2.66**	1.06	2.51***	0.43	2.60***	0.47
γ_3	IA_{t-1}	2.77***	0.45			4.93***	0.24
Age groups							
γ_4	PE_{t-1}·Age 54-57	-0.18*	0.10	0.07	0.08	-0.38***	0.06
γ_5	SE_{t-1}·Age 54-57	-0.07	0.11			-0.19	0.12
γ_6	UI_{t-1}·Age 54-57	0.58	0.37	1.70***	0.15	0.21	0.17
γ_7	IA_{t-1}·Age 54-57	0.49***	0.14			0.01	0.08
γ_8	PE_{t-1}·Age 58-63	-0.40**	0.16	0.81***	0.11	0.55***	0.07
γ_9	SE_{t-1}·Age 58-63	-0.19	0.14			-0.24*	0.13
γ_{10}	UI_{t-1}·Age 58-63	1.44***	0.45	3.94***	0.21	0.92***	0.22
γ_{11}	IA_{t-1}·Age 58-63	0.17	0.16			0.17*	0.09
Time periods							
γ_{12}	PE_{t-1}·Year 94-98	-0.28**	0.13	-0.63***	0.09	0.49***	0.06
γ_{13}	SE_{t-1}·Year 94-98	0.85***	0.14			1.66***	0.15
γ_{14}	UI_{t-1}·Year 94-98	0.31	0.59	-0.07	0.22	0.91***	0.25
γ_{15}	IA_{t-1}·Year 94-98	0.52***	0.16			1.40***	0.09
γ_{16}	PE_{t-1}·Year 99-03	-0.22	0.19	-1.05***	0.12	1.20***	0.09
γ_{17}	SE_{t-1}·Year 99-03	0.72***	0.20			2.05***	0.21
γ_{18}	UI_{t-1}·Year 99-03	-1.07	0.80	0.10	0.29	1.56***	0.32
γ_{19}	IA_{t-1}·Year 99-03	0.51**	0.24			1.73***	0.14
γ_{20}	PE_{t-1}·Year 04-09	-0.22	0.22	-0.90***	0.15	2.26***	0.12
γ_{21}	SE_{t-1}·Year 04-09	1.36***	0.22			3.42***	0.22
γ_{22}	UI_{t-1}·Year 04-09	0.06	0.71	-0.24	0.29	2.36***	0.31
γ_{23}	IA_{t-1}·Year 04-09	0.91***	0.25			3.27***	0.15
Unemployment rate (UR)							
θ_1	UR	0.09**	0.04	0.10***	0.03	-0.10***	0.02
θ_2	$SE_{t-1} \cdot UR$	-0.07	0.06			-0.02	0.06
θ_3	$UI_{t-1} \cdot UR$	-0.25*	0.15	0.18***	0.06	0.01	0.07
θ_4	$IA_{t-1} \cdot UR$	-0.17**	0.07			-0.11***	0.04
Treatment							
δ_1	PE_{t-1}· treatment	0.09	0.20	-0.50***	0.14	0.08	0.08
δ_2	UI_{t-1}· treatment	-0.62	0.57	-0.81***	0.25	-0.17	0.28
Personal and household characteristics							
β_1	Birth year	0.00	0.02	0.02	0.02	-0.14***	0.01
β_2	Immigrant	-0.45***	0.12	0.28***	0.07	-0.08	0.05
β_3	Partner	-0.06	0.09	0.30***	0.09	0.20***	0.05
β_4	Number of children	0.05	0.06	0.03	0.05	0.14***	0.03
β_5	Age youngest child	0.00	0.01	-0.03***	0.01	-0.02***	0.00
β_0	Constant	-14.37	43.38	-46.08	37.30	262.87***	23.18
σ^2_{se}		4.05***	0.24				
$\sigma_{se,ui}$		0.34**	0.14				
$\sigma_{se,ia}$		-0.79***	0.07				
σ^2_{ui}		0.03	0.03				
$\sigma_{ui,ia}$		-0.04	0.03				
σ^2_{ia}		0.37***	0.05				

	Age 50-53		Age 54-57		Age 58-63	
Necessity-hypothesis I: year 89-93	1.90[b]***	0.72	2.67[c]***	0.73	3.74***	0.73
Necessity-hypothesis I: year 94-98	2.49[d]***	0.56	3.25***	0.57	4.33***	0.58
Necessity-hypothesis I: year 99-03	1.05**	0.48	1.81***	0.50	2.89***	0.50
Necessity-hypothesis I: year 04-09	2.18***	0.33	2.95***	0.34	4.02***	0.46
Necessity-hypothesis II: year 89-93	-0.37[e]	0.73	-0.27[f]	0.74	0.90	0.74
Necessity-hypothesis II: year 94-98	-0.59[g]	0.58	-0.49	0.59	0.69	0.59
Necessity-hypothesis II: year 99-03	-1.95***	0.49	-1.86***	0.50	-0.68	0.50
Necessity-hypothesis II: year 04-09	-1.22***	0.34	-1.12***	0.35	0.06	0.45
Necessity-hypothesis III: year 04-09	-	-	-	-	3.32***	0.41

[a] * Significant at the 0.10 level; ** at the 0.05 level; *** at the 0.01 level. The log-likelihood of the estimations on the subsample and the complement are -27,215.74 and -22,063.05 respectively providing an $LR\ \chi^2$ of 1,094.63 and 892.83. Initial conditions corrections are included in the estimation. Necessity hypothesis I tests whether unemployed individuals have a higher probability to enter SE than paid-employed individuals. Necessity hypothesis II tests whether unemployed individuals have a higher probability to enter SE than inactive individuals. Hypothesis III is the same as hypothesis I, but with the treatment. In the hypotheses we assume an unemployment rate of 3%.

[b] $H_0 : \gamma_2 + 3 \times \theta_3 = 0$

[c] $H_0 : \gamma_2 + (\gamma_6 - \gamma_4) + 3 \times \theta_3 = 0$

[d] $H_0 : \gamma_2 + (\gamma_{14} - \gamma_{12}) + 3 \times \theta_3 = 0$

[e] $H_0 : (\gamma_2 - \gamma_3) + 3 \times (\theta_3 - \theta_4) = 0$

[f] $H_0 : (\gamma_2 - \gamma_3) + (\gamma_6 - \gamma_7) + 3 \times (\theta_3 - \theta_4) = 0$

[g] $H_0 : (\gamma_2 - \gamma_3) + (\gamma_{14} - \gamma_{15}) + 3 \times (\theta_3 - \theta_4) = 0$

[h] $H_0 : \gamma_2 + (\gamma_{10} - \gamma_8) + (\gamma_{22} - \gamma_{20}) + 3 \times \theta_3 + (\delta_2 - \delta_1) = 0$

Table 6.5: Estimation results baseline model[a] (women)

		Self-employment		Unemployment		Inactivity	
Effects relative to paid-employment		Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Mobility							
γ_1	SE_{t-1}	3.47***	0.44			2.69***	0.42
γ_2	UI_{t-1}	1.20	1.23	1.44***	0.50	0.63	0.57
γ_3	IA_{t-1}	1.01**	0.40			3.81***	0.22
Age groups							
γ_4	PE_{t-1}·Age 54-57	-0.35***	0.11	0.06	0.10	-0.32***	0.06
γ_5	SE_{t-1}·Age 54-57	-0.21*	0.12			-0.22	0.14
γ_6	UI_{t-1}·Age 54-57	1.08***	0.40	1.13***	0.17	0.45**	0.19
γ_7	IA_{t-1}·Age 54-57	0.23**	0.11			-0.07	0.07
γ_8	PE_{t-1}·Age 58-63	-0.18	0.17	0.89***	0.16	0.42***	0.08
γ_9	SE_{t-1}·Age 58-63	-0.09	0.16			0.13	0.16
γ_{10}	UI_{t-1}·Age 58-63	1.53***	0.55	3.30***	0.26	1.37***	0.27
γ_{11}	IA_{t-1}·Age 58-63	0.46***	0.15			0.35***	0.09
Time periods							
γ_{12}	PE_{t-1}·Year 94-98	0.65***	0.15	-0.74***	0.11	0.53***	0.08
γ_{13}	SE_{t-1}·Year 94-98	1.25***	0.19			0.99***	0.20
γ_{14}	UI_{t-1}·Year 94-98	0.86	0.68	-0.09	0.25	1.87***	0.30
γ_{15}	IA_{t-1}·Year 94-98	0.87***	0.14			1.05***	0.09
γ_{16}	PE_{t-1}·Year 99-03	-0.06	0.21	-1.41***	0.14	0.88***	0.11
γ_{17}	SE_{t-1}·Year 99-03	0.62**	0.24			0.94***	0.27
γ_{18}	UI_{t-1}·Year 99-03	0.23	0.87	0.14	0.33	2.06***	0.40
γ_{19}	IA_{t-1}·Year 99-03	0.70***	0.21			1.42***	0.12
γ_{20}	PE_{t-1}·Year 04-09	0.34	0.23	-1.17***	0.17	1.43***	0.13
γ_{21}	SE_{t-1}·Year 04-09	1.63***	0.26			1.88***	0.27
γ_{22}	UI_{t-1}·Year 04-09	1.58*	0.83	0.54	0.35	3.53***	0.39
γ_{23}	IA_{t-1}·Year 04-09	1.49***	0.23			2.52***	0.14
Unemployment rate (UR)							
θ_1	UR	-0.09*	0.04	0.06	0.04	-0.12***	0.02
θ_2	$SE_{t-1} \cdot UR$	-0.07	0.07			-0.17**	0.07
θ_3	$UI_{t-1} \cdot UR$	-0.14	0.18	0.35***	0.08	0.08	0.09
θ_4	$IA_{t-1} \cdot UR$	0.06	0.06			0.07**	0.03
Treatment							
δ_1	PE_{t-1}· Treatment	-0.02	0.20	-0.80***	0.19	0.18*	0.09
δ_2	UI_{t-1}· Treatment	-1.02	0.69	-1.19***	0.32	-0.91***	0.35
Personal and household characteristics							
β_1	Birth year	0.02	0.02	0.01	0.03	-0.14***	0.01
β_2	Immigrant	-0.32***	0.12	0.32***	0.10	0.04	0.06
β_3	Partner	0.00	0.10	0.40***	0.14	0.34***	0.07
β_4	Number of children	0.05	0.08	0.04	0.14	0.05	0.06
β_5	Age youngest child	0.00	0.01	-0.04**	0.02	-0.01*	0.01
β_0	Constant	-43.30	44.32	-26.69	49.71	267.45***	28.05
σ^2_{se}		3.10***	0.19				
$\sigma_{se,ui}$		0.27*	0.14				
$\sigma_{se,ia}$		0.62***	0.09				
σ^2_{ui}		0.55***	0.13				
$\sigma_{ui,ia}$		0.15**	0.07				
σ^2_{ia}		1.50***	0.08				

	Age 50-53		Age 54-57		Age 58-63	
Necessity-hypothesis I: year 89-93	0.77[b]	0.82	2.20[c] ***	0.81	2.49***	0.89
Necessity-hypothesis I: year 94-98	0.99[d]	0.64	2.41***	0.64	2.70***	0.73
Necessity-hypothesis I: year 99-03	1.07**	0.45	2.49***	0.43	2.78***	0.50
Necessity-hypothesis I: year 04-09	2.02***	0.40	3.45***	0.35	3.73***	0.55
Necessity-hypothesis II: year 89-93	-0.41[e]	0.82	0.44[f]	0.81	0.66	0.89
Necessity-hypothesis II: year 94-98	-0.42[g]	0.65	0.43	0.64	0.65	0.73
Necessity-hypothesis II: year 99-03	-0.88*	0.45	-0.03	0.43	0.19	0.50
Necessity-hypothesis II: year 04-09	-0.32	0.40	0.53	0.35	0.76	0.54
Necessity-hypothesis III: year 04-09	-	-	-	-	2.74***	0.47

[a] * Significant at the 0.10 level; ** at the 0.05 level; *** at the 0.01 level. The log-likelihood of the estimations on the subsample and the complement are -23,215.31 and -20,049.67 respectively providing an LR χ^2 of 1,008.45 and 756.42. Initial conditions corrections are included in the estimation. Necessity hypothesis I tests whether unemployed individuals have a higher probability to enter SE than paid-employed individuals. Necessity hypothesis II tests whether unemployed individuals have a higher probability to enter SE than inactive individuals. Hypothesis III is the same as hypothesis I, but with the treatment. In the hypotheses we assume an unemployment rate of 3%.

[b] $H_0 : \gamma_2 + 3 \times \theta_3 = 0$

[c] $H_0 : \gamma_2 + (\gamma_6 - \gamma_4) + 3 \times \theta_3 = 0$

[d] $H_0 : \gamma_2 + (\gamma_{14} - \gamma_{12}) + 3 \times \theta_3 = 0$

[e] $H_0 : (\gamma_2 - \gamma_3) + 3 \times (\theta_3 - \theta_4) = 0$

[f] $H_0 : (\gamma_2 - \gamma_3) + (\gamma_6 - \gamma_7) + 3 \times (\theta_3 - \theta_4) = 0$

[g] $H_0 : (\gamma_2 - \gamma_3) + (\gamma_{14} - \gamma_{15}) + 3 \times (\theta_3 - \theta_4) = 0$

[h] $H_0 : \gamma_2 + (\gamma_{10} - \gamma_8) + (\gamma_{22} - \gamma_{20}) + 3 \times \theta_3 + (\delta_2 - \delta_1) = 0$

Table 6.6: Main income source and income level for those moving from inactivity to self-employment

	Age 50-53			Age 54-57			Age 58-63		
	%[a]	Median income[b]		%	Median income		%	Median income	
Men		SE_t	All_t		SE_t	All_t		SE_t	All_t
$Disability_{t-1}$	18	18,645	22,006	27	19,616	22,452	26	17,863	23,082
$Early\ retirement_{t-1}$	7	17,788	30,424	8	26,620	35,015	16	29,240	35,453
$Social\ assistance_{t-1}$	17	6,423	13,670	5	3,199	13,462	5	8,1401	13,016
$Wealth_{t-1}$	32	47,982	38,164	39	54,470	40,581	43	44,307	27,687
$Income\ spouse_{t-1}$	26	0	0	21	0	0	11	0	0
Women									
$Disability_{t-1}$	3	14,108	13,776	6	13,029	13,723	6	8,743	13,735
$Early\ retirement_{t-1}$	5	14,085	21,426	3	11,846	21,340	16	10,144	19,672
$Social\ assistance_{t-1}$	5	16,279	15,076	5	18,587	15,002	2	11,735	14,897
$Wealth_{t-1}$	8	37,246	17,020	10	7,476	17,363	10	11,516	20,197
$Income\ spouse_{t-1}$	79	0	0	76	0	0	66	0	0

[a] % refers to the percentage of inactive persons in $t-1$ who enter self-employment from a certain category.
[b] The table shows median personal total income in period $t-1$ for those individuals moving from a certain inactivity category to self-employment and for all individuals in that inactivity category in $t-1$.

2002-2003, which is the period just before the period in which the reform was actually introduced. The results in panel A of table 6.8 are reassuring in that we do not find significant effects from the fake treatments on the inflow and outflow from unemployment.

The robustness check in panel B of table 6.8 shows that also after reducing the time window to the period 1999-2009, search requirements still increase the outflow from unemployment for men and women. However, the inflow to unemployment is no longer significantly affected by the reform. Table 6.8 only shows the coefficients of the treatment effects. Conclusions with regard to mobility and the macroeconomic unemployment rate do not change.

Using yearly data makes it hard to disentangle the effects of the job search requirements introduced in January 2004 and the abolition of extended benefits in August 2003. To ensure that our treatment effect measures the effect of the introduction of search requirements we exploit the fact that the abolition of the extended UI benefits did not change the generosity of the UI system for single persons as mentioned in section 6.2. This robustness check is also exploited by Lammers et al. (2013). We ensure measuring treatment effects of job search requirements by adding interaction terms with singles.
In this way we can test whether the treatment effects for single persons are significantly different from the treatment effects estimated in the baseline

Table 6.7: Estimation results extended models[a]

		Self-employment		Unemployment		Inactivity	
Effects relative to paid-employment		Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Panel A. Financial variables							
Men[b]							
β_6	$\text{Homeowner}_{t=0}$	0.23***	0.07	-0.30***	0.06	-0.08**	0.03
β_7	$\text{Mortgage}_{t=0}/10^6$	0.21	0.21	-0.03	0.27	-0.35**	0.18
β_8	$\text{Financial wealth}_{t=0}/10^5$	0.70***	0.25	-0.40	0.66	-0.02	0.16
β_9	$\text{Net housing wealth}_{t=0}/10^5$	-0.14	0.48	0.09	0.84	0.38	0.36
β_{10}	$\text{Risky assets}_{t=0}/10^3$	0.43	0.47	0.06	0.37	-0.39*	0.20
δ_1	$\text{PE}_{t-1}\cdot$ Treatment	0.13	0.21	-0.53***	0.14	0.08	0.08
δ_2	$\text{UI}_{t-1}\cdot$Treatment	-0.52	0.59	-0.76***	0.26	-0.12	0.28
Women[c]							
β_6	$\text{Homeowner}_{t=0}$	0.42***	0.07	-0.29**	0.07	0.00	0.05
β_7	$\text{Mortgage}_{t=0}/10^6$	0.18	0.30	-0.07	0.29	-0.56**	0.23
β_8	$\text{Financial wealth}_{t=0}/10^5$	0.26	0.25	-0.30	0.72	-0.01	0.20
β_9	$\text{Net housing wealth}_{t=0}/10^4$	0.08	0.08	-0.03	0.10	0.03	0.06
β_{10}	$\text{Risky assets}_{t=0}/10^2$	0.18	0.28	0.14	0.28	0.15	0.19
δ_1	$\text{PE}_{t-1}\cdot$ Treatment	0.05	0.21	-0.86***	0.20	0.16*	0.10
δ_2	$\text{UI}_{t-1}\cdot$Treatment	-1.08	0.70	-1.27***	0.33	-1.03***	0.37
Panel B. Health							
Men[d]							
β_6	$\text{Health}_{t=0}$	-0.42***	0.09	0.71***	0.07	0.75***	0.04
δ_1	$\text{PE}_{t-1}\cdot$ Treatment	0.12	0.20	-0.50***	0.14	0.10	0.08
δ_2	$\text{UI}_{t-1}\cdot$Treatment	-0.62	0.57	-0.65***	0.25	-0.02	0.28
Women[e]							
β_6	$\text{Health}_{t=0}$	-0.54***	0.12	1.22***	0.09	0.70***	0.06
δ_1	$\text{PE}_{t-1}\cdot$ Treatment	-0.02	0.20	-0.88***	0.19	0.17*	0.09
δ_2	$\text{UI}_{t-1}\cdot$Treatment	-1.05	0.69	-1.01***	0.32	-0.78**	0.35

[a] * Significant at the 0.10 level; ** at the 0.05 level; *** at the 0.01 level All regressions include the variables from the baseline regression.

[b] Financial variables are jointly significant with $\chi^2(15) = 95.87$ and $p-value = 0.000$. The log-likelihood of the estimations on the subsample and the complement are -24,793.58 and -22,030.97 respectively providing an $LR\ \chi^2$ of 980.40 and 901.99.

[c] Financial variables are jointly significant with $\chi^2(15) = 99.08$ and $p-value = 0.000$. The log-likelihood of the estimations on the subsample and the complement are -19,339.93 and -20,029.29 respectively providing an $LR\ \chi^2$ of 889.92 and 767.43.

[d] $\text{Health}_{t=0}$ equals 1 if a person received disability benefits in the initial period observed and 0 otherwise. The log-likelihood of the estimations on the subsample and the complement are -27,076.24 and -21,939.01 respectively providing an $LR\ \chi^2$ of 1,056.45 and 866.57.

[e] $\text{Health}_{t=0}$ equals 1 if a person received disability benefits in the initial period observed and 0 otherwise. The log-likelihood of the estimations on the subsample and the complement are -23,100.08 and -19,991.02 respectively providing an $LR\ \chi^2$ of 993.69 and 743.59.

regression. If the treatment effects are significantly different this is likely to be a consequence of partially measuring the effects of the abolition of the extended UI benefits among non-singles.

Panel C in table 6.8 indicates that δ_1 and δ_2 are highly comparable to δ_1 and δ_2 from the baseline regression of males in table 6.4. ϑ_1 and ϑ_2 in panel C are not significantly different from δ_1 and δ_2 which implies that the treatment effects of singles are not different from the treatment effects of non-singles. Stated differently, it is likely that δ_1 and δ_2 only capture the effects of the introduced search requirements in 2004 among men.

Among women, the estimation results indicate that δ_1 and δ_2 are highly comparable to δ_1 and δ_2 from the baseline regression of in table 6.5, except that δ_2 in the self-employment equation is now significantly negative at the 0.10 level whereas this coefficient was only close to the 0.10 significance level in the baseline regression, e.g. the search requirements decreases the flow from unemployment to self-employment relative to flows from unemployment to paid-employment. ϑ_1 and ϑ_2 in panel C are not significantly different from zero among women, except the coefficient ϑ_1 in the self-employment equation at the 0.05 level. This coefficient shows that the treatment decreases the probability of flowing from paid-employment to self-employment relative to staying in paid-employment among single women.[25] However, this does not affect our necessity-hypotheses. Instead, the 'pure' effect of search requirements suggests that more women remained in paid-employment relative to flowing from paid- to self-employment. So, inducing extra obligtions in unemployment did not make self-employment more attractive as a way to reduce active hours spent. This interpretation is consistent with the results of the baseline regression in table 6.5.

Finally, conclusions do not change when we test the robustness of the results with regard to different model specifications, e.g. sensitivity analyses of the age and time categories as well as the categories in the multinomial dependent variable (not reported here).

Since the data set only contains yearly information, we do not observe within-year transitions. For example, if someone's main source of income in year $t-1$ was unemployment, but he also received a substantial amount

[25] $H_0 : \delta_1 + \vartheta_1 = 0$ is rejected at the 0.10 level for the self-employment category.

Table 6.8: Robustness checks[a]

	Self-employment		Unemployment		Inactivity	
Effects relative to paid-employment	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Panel A. Placebo tests						
Men: placebo age 56-57						
δ_1 $PE_{t-1}\cdot$ Treatment	0.20	0.27	0.29	0.19	-0.08	0.13
δ_2 $UI_{t-1}\cdot$Treatment	-0.39	0.96	0.28	0.33	0.17	0.38
Women: placebo age 56-57						
δ_1 $PE_{t-1}\cdot$ Treatment	-0.06	0.16	0.01	0.14	0.07	0.07
δ_2 $UI_{t-1}\cdot$Treatment	-0.12	0.52	-0.04	0.25	0.13	0.29
Men: placebo year 2000-2003						
δ_1 $PE_{t-1}\cdot$ Treatment	-0.23	0.31	0.18	0.21	0.21**	0.09
δ_2 $UI_{t-1}\cdot$Treatment	0.36	1.40	0.32	0.44	0.20	0.47
Women: placebo year 2000-2003						
δ_1 $PE_{t-1}\cdot$ Treatment	-0.11	0.24	0.34	0.21	0.18**	0.09
δ_2 $UI_{t-1}\cdot$Treatment	-12.91	556.15	0.34	0.34	-0.04	0.40
Panel B. Smaller time window						
Men						
δ_1 $PE_{t-1}\cdot$ Treatment	-0.02	0.32	0.05	0.22	0.10	0.11
δ_2 $UI_{t-1}\cdot$ Treatment	-0.44	1.12	-0.70*	0.41	-0.43	0.43
Women						
δ_1 $PE_{t-1}\cdot$ Treatment	0.06	0.24	0.06	0.21	0.36***	0.10
δ_2 $UI_{t-1}\cdot$Treatment	-0.92	0.75	-0.79**	0.35	-0.47	0.39
Panel C. Single						
Men						
δ_1 $PE_{t-1}\cdot$ Treatment	0.07	0.21	-0.51***	0.14	0.09	0.08
ϑ_1 $PE_{t-1}\cdot$ Treatment $\cdot$ Single	0.16	0.35	0.06	-0.09	0.09	0.11
δ_2 $UI_{t-1}\cdot$ Treatment	-0.42	0.61	-0.67***	0.27	-0.03	0.30
ϑ_2 $UI_{t-1}\cdot$ Treatment $\cdot$ Single	-0.80	0.93	-0.52	0.35	-0.52	0.40
Women						
δ_1 $PE_{t-1}\cdot$ Treatment	0.11	0.21	-0.87***	0.21	0.21**	0.09
ϑ_1 $PE_{t-1}\cdot$ Treatment $\cdot$ Single	-0.81**	0.37	0.27	0.26	-0.13	0.11
δ_2 $UI_{t-1}\cdot$Treatment	-1.37*	0.71	-1.43***	0.34	-1.05***	0.39
ϑ_2 $UI_{t-1}\cdot$Treatment $\cdot$ Single	-11.47	310.90	0.63	0.49	0.69	0.56

[a] * Significant at the 0.10 level; ** at the 0.05 level; *** at the 0.01 level. Results of the different robustness checks are estimated separately. All regressions include the variables from the baseline regression including the initial conditions correction and correlated random effects parameters. *Single* is a binary variable with a value of one for single individuals and zero otherwise.

of labor income, than we indicate this person as unemployed in year $t-1$. Table 6.1 already showed that partial unemployment and partial self employment are not very important. As a last robustness check we added variables to the model indicating partial unemployment and partial self-employment. Including these variables in the baseline specification does not affect the conclusions (not reported here).

6.5.3 Simulation

To facilitate the interpretation of the estimation results in the baseline model outlined above, we use the baseline estimates to simulate transition probabilities for a reference individual with specific values assigned to the covariates. Here, we take as a reference a native male and female with a partner, without children in the same household, and of age 60 in the year 2006.[26] For the initial labor market status we take the average of the sample and the random effects are set to zero. First we present the simulation results without the treatment effect, after that we show how the transition rates would change when the treatment is taken into account. Standard errors are based on a parametric bootstrap over the asymptotic distribution of our estimates.

When we compare the simulation results in table 6.9 with transition rates in the right bottom of tables 6.2 and 6.3 we find that state dependence is far less important when observed and unobserved heterogeneity are taken into account, especially for the self-employed. This is in line with the relatively high variance of the random effect for self-employment found in tables 6.4 and 6.5. Although the probabilities to enter self-employment are low, this probability is higher for individuals in unemployment than for individuals in paid employment or inactivity.

The last two rows of table 6.9 present the treatment effects. Job search requirements between the ages of 58 and 63 reduced the probability to stay in unemployment for men significantly with 15% (12%-points) and for women insignificantly with 19% (7%-points). These individuals now move to paid employment and inactivity. Because of the reform the

[26] The unemployment rate in 2006 was 3.6%

probability for men to move from unemployment to paid employment increased significantly with 93% (from 2.14% to 4.14%) and the probability to move from unemployment to inactivity increased significantly with 63% (from 15.61% to 25.45%). For women the probability to move from unemployment to paid employment increased significantly with 165% (from 1.19% to 1.96%) and the probability to move from unemployment to inactivity increased significantly with 8% (from 61.59% to 66.40%).[27] In fact, it seems that the mandatory search requirements increased the probability of finding a paid job at older ages while decreasing the probability of using unemployment as an early retirement route. Similar effects have been found by Lammers et al. (2013) who focus on substitution effects between unemployment and disability in specific. We find that most treated individuals moving from unemployment to inactivity enter early retirement (almost 60% for both men and women). About 27% of the treated men and 17% of the women enter disability, and the remaining 13% (men) and 23% (women) enter social assistance or become dependent on income from wealth or a partner. Self-employment (out of necessity) did not increase because of the reform.

Our analysis also allows us to study the effect of job search requirements on the inflow to unemployment. Job search requirements reduced the probability to enter unemployment significantly with about 40% for men (from 2.77% to 1.65%) and about 59% for women (from 1.71% to 0.70%). Mandatory job search requirements, however, did not induce more people to stay in paid-employment. Substitution effects towards other exit routes are mainly observed, suggesting that these options are still more attractive than using self-employment as an opportunity to reduce working hours. The still relatively generous early retirement and social insurance schemes may also explain the minor importance of for example bridge-jobs, which are often found among elderly in the US.

[27] The relative increase of paid employment is higher than the relative increase of inactivity, as was already suggested by the significantly negative coefficient δ_2 in the inactivity equation of table 6.5.

Table 6.9: Simulation results[a b]

	Men				Women			
	Year t							
Year $t-1$	PE	SE	UI	IA	PE	SE	UI	IA
PE	73.13	0.19	2.77	23.90	48.86	1.09	1.71	48.34
	(1.41)	(0.03)	(0.34)	(1.44)	(2.37)	(0.19)	(0.32)	(2.43)
SE	12.67	4.58	0.52	82.23	7.94	17.34	0.36	74.36
	(1.34)	(0.57)	(0.07)	(1.57)	(1.01)	(1.84)	(0.06)	(2.14)
UI	2.14	0.29	81.95	15.61	1.19	1.12	36.10	61.59
	(0.75)	(0.12)	(1.97)	(1.88)	(0.32)	(0.59)	(5.55)	(5.61)
IA	1.75	0.22	0.07	97.95	0.63	0.28	0.03	99.06
	(0.12)	(0.03)	(0.00)	(0.13)	(0.05)	(0.03)	(0.00)	(0.06)
Treatment effects								
PE	-0.61	0.01	-1.12	1.72	-3.86	-0.11	-1.01	4.97
	(2.01)	(0.04)	(0.35)	(1.42)	(2.18)	(0.21)	(0.29)	(2.24)
UI	2.00	0.00	-11.85	9.84	1.96	-0.08	-6.70	4.81
	(0.75)	(0.16)	(2.83)	(2.73)	(0.71)	(0.71)	(5.11)	(5.34)

[a] This table presents a simulated transition matrix for a reference individual, which is a native male or female with a partner, without children in the same household, and of age 60 in the year 2006.

[b] Standard errors in parentheses (1500 bootstrap replications).

Discussion 6.6

A few points remain for discussion. An explanation why unemployed individuals have a higher probability to enter self-employment than paid-employed individuals may be that part-time employment is widely available in the Netherlands and is an effective way to reduce working hours for those in paid-employment.[28]

Another explanation for necessity reasons outweighing opportunity reasons may be that moving from paid-employment to self-employment can have a negative effect on occupational pension accumulation. Since occupational pensions are generally not accumulated during unemployment, pension accumulation considerations are far less important for transition from unemployment to self-employment. Zissimopoulos and Karoly (2007) find that having access to pension coverage in paid-employment reduces the probability to enter self-employment. Moore and Mueller (2002), on the other hand, find no effects of pensions in paid-employment on self-employment entry.

[28] Emmanoulidi and Kyriazidou (2012) indeed find that in Britain part-time paid employment is more often used as an exit from paid employment than self-employment.

A final point of discussion is the absence of education and health shocks in the analysis. The unobserved heterogeneity term corrects for unobserved differences in education levels, but is unable to correct for health shocks. Zucchelli et al. (2012) show that ill-health and health shocks do not increase the probability of using self-employment as retirement mechanism, however. Instead, health seems to be an important determinant for retiring early. Therefore, including health indicators in the analysis will likely be relevant for transitions to and from inactivity, but probably does not affect our conclusions about the nature of choosing self-employment as an exit route to retirement. All the more because in the Netherlands those who are in bad health are selected into disability insurance, which is financially more attractive than unemployment insurance or early retirement schemes (De Vos et al. 2012) and probably usually also more attractive than starting an own business.

For future research it would be interesting to investigate how income develops when people make a transition from paid-employment or unemployment to self-employment or inactivity. Substantial tax advantages of self-employment (beyond the scope of this paper) are also relevant in this context.

6.7 Conclusion

This paper examines whether individuals at the end of working life choose self-employment out of necessity and to what degree the introduction of search requirements for unemployment benefits induce people to become self-employed. For this purpose we model transitions between labor market states for people at older ages using a dynamic multinomial logit model with unobserved heterogeneity.

Our empirical specification allows us to measure the role of necessity-driven factors by analyzing the labor market position of people that enter self-employment and, from a macroeconomic perspective, how the unemployment rate affects inflow into self-employment. The effects of search requirements are examined using a Dutch UI reform in 2004, that introduced search requirements for people older than 57.5 years.

The main empirical findings can be summarized as follows. After correcting for observed and unobserved heterogeneity, unemployed and inactive individuals have a higher probability to enter self-employment at the end of working life than those in paid-employment. Furthermore, mobility from paid-employment to self-employment is relatively low and does not increase with age (as would be the case when self-employment would be chosen out of opportunity to reduce working hours at the end of working life). This indicates that at older ages necessity reasons are important to become self-employed. Moreover, the unemployment rate has a positive effect on transitions from paid-employment to self-employment among men. This is in line with the recession push hypothesis, which suggests that men in paid-employment become self-employed at older ages in order to avoid a period of unemployment. For women, on the other hand, we find a negative effect of the unemployment rate on transitions from paid-employment to self-employment, which is consistent with the prosperity pull hypothesis (e.g. they are more likely to start self-employment when the unemployment rate is low). For inactive men and women the prosperity pull hypothesis also holds. At lower ages, self-employment entry is most likely from inactivity. In the highest age-category, self-employment entry from unemployment and inactivity are not significantly different, suggesting that transitions from unemployment to self-employment become increasingly important over age.

The introduction of job search requirements at the end of working life have stimulated people to exit unemployment and discouraged people to enter unemployment. The reform, however, did not increase necessity or opportunity driven self-employment. Individuals that are confronted with search requirements are partly able to find a job, but there are also large substitution effects between unemployment and inactivity (mostly early retirement) which suggests that these options are still more attractive than using self-employment as a retirement mechanism.

Taken together, our findings suggest that at the end of working life individuals with a relatively weak labor market position are more likely to switch to self-employment. The results do not suggest that self-employment is used as a gradual retirement route. Job search requirements in UI increase the outflow from unemployment and decrease the inflow to

unemployment, but do not increase self-employment out of necessity or opportunity.

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Nederlandse samenvatting

Pensioenen, Pensionering, en de Financiële Positie van Ouderen

Dit proefschrift is een bundeling van vijf studies met betrekking tot *pensioenen, pensionering en de financiële positie van ouderen*, met als doel bij te dragen aan onze kennis van daadwerkelijke pensioenopbouw, pensioneringsgedrag en de financiële positie gedurende pensionering.

Langetermijnontwikkelingen, zoals het verouderen van de bevolking, de toename in de gemiddelde levensverwachting en de gevolgen van de recente financiële crisis, hebben bijgedragen aan de discussie rondom de houdbaarheid van het pensioenstelsel. De veroudering van de bevolking zorgt voor een toename van het aantal gepensioneerden ten opzichte van het aantal niet-gepensioneerden en daarmee voor een toenemende belasting- en premiedruk op de beroepsbevolking. Deze druk wordt versterkt door de lage geboortecijfers. De toename van de gemiddelde levensverwachting zorgt er voor dat gepensioneerden gemiddeld genomen langer pensioenuitkeringen ontvangen. De recente financiële crisis heeft een grote negatieve invloed gehad op de kapitaalreserves van pensioenfondsen, met kortingen op pensioenuitkeringen tot gevolg. De huidige pensioenstelsels lijken onvoldoende voorbereid te zijn voor deze langetermijnontwikkelingen en korte termijn volatiliteit.

Veel OECD-landen hebben hervormingen in de pensioensystemen voorgesteld en geïmplementeerd om zodoende de pensioensystemen houdbaar te maken voor eerdergenoemde ontwikkelingen. Veel van deze hervormingen hebben betrekking op het verhogen van de effectieve en officiële pensioenleeftijd, het verlagen van de pensioenuitkeringen en het verhogen

van de pensioenbijdragen. De verhoging van de pensioenleeftijd heeft als doel de periode waarin men pensioenuitkeringen ontvangt te verkleinen en de periode waarin men bijdraagt aan de pensioenopbouw te vergroten. De meest in het oog springende hervormingen ter verhoging van de effectieve pensioenleeftijd betreffen het minder aantrekkelijk of beschikbaar maken van vervroegde uittreedroutes.

Deze hervormingen impliceren dat publieke- en private pensioenen minder royaal zijn, met als gevolg dat er veel discussie is ontstaan omtrent de financiële positie van huidige en toekomstige gepensioneerden. Dit proefschrift heeft tot doel om de effecten van dergelijke ontwikkelingen op pensioneringsgedrag en de toerekendheid van (toekomstige) pensioen in kaart te brengen en beter te begrijpen.

Pensioensystemen zijn onderhevig aan padafhankelijkheid en langetermijncontracten. Pensioenhervormingen, zoals de verhoging van de pensioenleeftijd, zijn daarom veelal politiek controversieel en moeilijk te implementeren. Als gevolg daarvan hebben veel landen pensioenen geprivatiseerd, wat zo veel wil zeggen als een verschuiving van het belang van publieke pensioenen naar private pensioenen en van 'defined benefit' pensioenregelingen naar 'defined contribution' pensioenregelingen. Deze verschuiving reduceert de druk van een vergrijzende samenleving op de overheidsfinanciën, maar kan de inkomensongelijkheid onder ouderen doen toenemen.

Hoofdstuk 2 analyseert in welke mate een verschuiving van publieke- naar private pensioenen een grotere inkomensongelijkheid en hogere armoede onder ouderen tot gevolg heeft. *A priori* is de verwachting dat een verschuiving van publieke- naar private pensioenen een grotere inkomensongelijkheid en armoede onder ouderen tot gevolg heeft, omdat private socialezekerheidsarrangementen over het algemeen minder herverdelend zijn dan publieke socialezekerheidsarrangementen. Een analyse van macroeconomische data van 15 Europese landen voor de periode 1995-2007 bevestigt deze hypothese niet. Er is geen bewijs dat verschuivingen van publieke- naar private pensioen gepaard gaan met een grotere inkomensongelijkheid of armoede onder ouderen. Dit geldt voor een breed scala aan methodologische benaderingen. De benadering in hoofdstuk 2 geeft

enig inzicht in de gevolgen van verschuivingen binnen pensioensystemen op het financiële welzijn van ouderen.

Om een beeld te krijgen van het financiële welzijn van ouderen ontwikkelt *Hoofdstuk 3* een integrale benadering om het totale beschikbare inkomen te bezien tijdens pensionering, voor zowel huidige als voor toekomstige gepensioneerden. Dit hoofdstuk beargumenteert dat deze integrale benadering van belang is in de analyse van de toereikendheid van pensioenen, omdat verschillende componenten mogelijkerwijs als substituten kunnen fungeren. De benadering onderscheidt privévermogen naast het publieke- en private pensioen. Privévermogen kan bestaan uit spaargeld, effecten, een eigen bedrijf en de netto waarde van het eigen huis (woningwaarde minus hypotheekschuld). Het niet meenemen van het privévermogen zou leiden tot een onderschatting van het inkomen dat beschikbaar is tijdens pensionering. Het privévermogen kan namelijk geconsumeerd gedurende het pensioen.

Daarnaast laat het hoofdstuk zien dat het noodzakelijk is om de analyse van de toereikendheid van pensioen te baseren op microeconomische data. Op deze manier illustreert de analyse de heterogeniteit in de daadwerkelijke opbouw van pensioenen, gegeven de huidige regels omtrent pensioenopbouw. De benadering in hoofdstuk 3 laat echter (ook) zien dat zo'n 31% van alle huishoudens een bruto vervangingsratio heeft dat minder is dan de 70%, een norm die in de internationale literatuur veelal als voldoende toereikend wordt verondersteld. Een doorsnee huishouden heeft een bruto vervanginsratio van 83% en een gemiddeld totaal beschikbaar pensioeninkomen van 33.000 euro (gestandaardiseerd). Het publieke- en private pensioen draagt voor zo'n 75% bij aan het totaal beschikbare inkomen gedurende het pensioen. Het huis draagt zo'n 10% bij, waarbij aangenomen wordt dat het huis niet verkocht en opgegeten wordt. Op basis van de analyse kan er gesteld worden dat er kwestbare groepen zijn, waarvan het pensioen onvoldoende toereikend lijkt te zijn op basis van bruto vervangingsratios en absolute niveaus van het beschikbare pensioen gedefinieerd. Dit zijn zelfstandig ondernemers, alleenstaande vrouwen, eerste-generatie immigranten en huishoudens waarin men een

bijstandsuitkering, werkloosheidsuitkering of arbeidongeschiktheidsuitkering ontvangt.

Hoofdstuk 3 laat zien dat zo'n 35% van het beschikbare inkomen tijdens pensionering afkomstig is van het werkgerelateerde pensioen. Dit werkgerelateerde pensioen wordt opgebouwd over de levenscyclus en is afhankelijk van het loon gedurende de levenscyclus. De relatie tussen loon en pensioenopbouw is in andere landen, zoals Noorwegen, Frankrijk, het VK en de VS ook aanwezig in het publieke pensioen. Daarmee bestaat er een directe relatie tussen het loon dat men verdient gedurende de gehele carrière en het pensioen dat beschikbaar is ten tijde van pensionering. Inzicht in het loonprofiel over de levenscyclus is daarom belangrijk voor de analyse van het pensioeninkomen. Loonprofielen spelen daarnaast een centrale rol in levenscyclusmodellen van consumptie- en spaargedrag. De conclusies van deze modellen hangen af van een correcte specificatie van het loonprofiel. Echter, de meeste levenscyclusmodellen corrigeren niet voor de mogelijke selectie-effecten in lonen terwijl lonen naar alle waarschijnlijkheid niet aselect geobserveerd worden. Lonen worden namelijk alleen geobserveerd bij personen die werkzaam zijn in loondienst. Personen die werken hebben mogelijk een ander potentieel loon dan personen die niet werken. Niet corrigeren voor deze selectie in lonen kan de schattingen van loonprofielen beïnvloeden. *Hoofdstuk 4* ontwikkelt een nieuwe methode om te corrigeren voor dergelijke selectie-effecten met paneldata. Deze nieuwe econometrische schatter corrigeert simultaan voor participatiebeslissingen op de extensieve marge (keuze werk versus niet-werk) en de intensieve marge (uren beslissing) waar andere schatters slechts corrigeren voor participatiebeslissingen op de extensieve marge. Het toevoegen van informatie omtrent part-time en full-time beslissingen geeft extra informatie omtrent de niet-geobserveerde achtergrondkarakteristieken die het loon kunnen beïnvloeden (bijvoorbeeld preferenties voor werken, bekwaamheid en inspanning).

Het toepassen van de nieuwe schatter op Nederlandse data laat zien dat het uitmaakt voor de resultaten of je deze extra informatie over participatie op de intensieve marge meeneemt. Een verdere decompositie van het loonprofiel naar part-time en full-time lonen en opleidingsniveau's toont

aan dat de grootte van de selectie-effecten kan verschillen tussen part-time en full-time lonen en opleidingsniveau's, maar dat selectie in zowel part-time als full-time werk over het algemeen positief is. Het bestaan van dergelijke selectie-effecten pleit voor het corrigeren voor selectie in lonen in modellen die gebruik maken van het schatten van loonprofielen, zoals het levenscyclusmodel.

Op basis van de empirische analyse in hoofdstuk 4 zijn er tevens substantiële negatieve effecten van een carrièregat op het loon. Wanneer mannen in totaal minimaal drie jaar niet gewerkt hebben, hebben zij een loon dat gemiddeld genomen 21% lager ligt dan het loon van mannen zonder carrièregat. Dit effect is 17% voor vrouwen. Daarnaast is er voor vrouwen een substantiele substantiële discontering op part-time lonen: part-time lonen zijn in verhouding gemiddeld 30% lager dan full-time lonen gecorrigeerd voor geobserveerde en niet-geobserveerde kenmerken.

Waar hoofdstuk 4 zich concentreert op de selectie-effecten van part-time werk, heeft *hoofdstuk 5* betrekking op de mate waarin niet-standaard werk, zoals part-time werk, leidt tot minder vervroegde uittreding op de arbeidsmarkt. Vervroegde uittreding is een belangrijke verklaring voor de afname van het aantal werkzame jaren in verschillende Europese landen. Een hogere participatiegraad onder ouderen zou de financiële houdbaarheid van pensioensystemen echter vergroten. Daarnaast leidt een hogere participatiegraad onder ouderen ook tot een toename in de toereikendheid van het pensioen voor huishoudens in pensionering. Allereerst betekent een langere carrière dat men cumulatief meer pensioen opbouwt. Ten tweede zorgt de mogelijkheid om volledige pensionering uit te stellen voor een inkomensstroom naast het gebruikelijke pensioeninkomen.

De belangrijkste oorzaken voor de lage participatie onder ouderen zijn de negatieve prikkels tot werken door formele (voorheen VUT) en informele uittreedroutes (voorheen WW en WAO/WIA) voor vervroegde uittreding. Hervormingen in deze uittreedroutes hebben de negatieve prikkels tot werken ten dele weggenomen met als gevolg dat de participatiegraad onder ouderen gestegen is ten opzichte van de jaren '90. Desondanks is de participatiegraad onder ouderen nog steeds relatief laag in vergelijking met jongere leeftijdscategorieën. Tegelijkertijd is er een

flinke toename in niet-standaard vormen van werk onder ouderen, zoals zelfstandig ondernemerschap en part-time werk. Beide geven flexibiliteit in het aantal uren dat men werkt en kunnen daarom gebruikt worden als brug van full-time werk naar volledige pensionering. Hoofdstuk 5 laat zien dat het relatief grote en groeiende belang van deze vormen van niet-standaard werk leidt tot minder vervroegde uittreding op de arbeidsmarkt in 13 Europese landen. Meer specifiek vindt hoofdstuk 5 dat een toename van het percentage van werkende 55-64 jarigen dat part-time werkt van 1% leidt tot een reductie van vervroegde uittreding van 1,7%. Dit lijkt voornamelijk een effect van een graduele tranisitie van full-time werk naar volledige pensionering, aangezien het effect grotendeels verklaard wordt aan de hand van de vrijwillige keuze voor part-time werk. Het effect van part-time werk op de vervroegde uittreding onder vrouwen is kleiner of zelfs afwezig afhankelijk van de gehanteerde specificatie. In de analyse worden geen significante effecten van zelfstandig ondernemerschap op vervroegde uittreding gevonden. De resultaten zijn uiterst robuust zowel op de extensieve marge (participatie beslissing) als de intensieve marge (uren beslissing).

De bevinding dat zelfstandig ondernemerschap niet leidt tot minder vervroegde uittreding is in lijn met de resultaten van *hoofdstuk 6*. De keuze voor niet-standaard vormen van werk hoeft niet per definitie een vrijwillige keuze te zijn. Het kan ook een reactie zijn op het gebrek aan andere mogelijkheden in (full-time) werk en uittreedroutes via werkloosheid en arbeidsongeschiktheid. De internationale literatuur laat zien dat zelfstandig ondernemerschap relatief belangrijk is bij 50-plussers. Eén van de verklaringen hiervoor is dat ouderen een relatief sterke preferentie hebben voor vrije tijd en dat zelfstandig ondernemerschap de flexibiliteit biedt om het aantal werkzame uren te verminderen. Hoofdstuk 6 vindt echter geen bewijs voor deze veelgebruikte verklaring. Middels drie empirisch toetsbare hypothesen omtrent de arbeidsmarktdynamiek van ouderen toegepast op Nederlandse microeconomische data, blijkt uit de resultaten in hoofdstuk 6 dat er bewijs is voor zelfstandig ondernemerschap dat is geboren uit noodzaak onder ouderen. Ouderen hebben moeite om na werkloosheid een betaalde baan te vinden. De drie hypothesen omtrent de

transities tussen werkloosheid en zelfstandig ondernemerschap suggereren dat ouderen met een relatief zwakke arbeidsmarktpositie veelal kiezen voor een periode als zelfstandig ondernemer. In de analyse van het hoofdstuk wordt er geen bewijs gevonden voor de veelgebruikte verklaring dat zelfstandig ondernemerschap onder ouderen een brug zou vormen tussen full-time werk en volledige pensionering. Een verklaring hiervoor is dat part-time werk deze rol veelal vervult in Nederland; dit in tegenstelling tot de VS, waar part-time werk minder gebruikelijk is.

De resultaten in hoofdstuk 5 en 6 tonen aan dat part-time werk mogelijkerwijs gebruikt wordt als graduele uittreedroute in Nederland. Als part-time werk leidt tot een hogere arbeidsparticipatie onder ouderen (zoals gevonden in hoofdstuk 5) kan dit bijdragen aan de financiële houdbaarheid van het pensioenstelsel en de toereikendheid van pensioenen (zoals geanalyseerd in hoofdstuk 2 en 3). Desalniettemin geeft de analyse in hoofdstuk 4 tegelijkertijd aan dat personen met relatief goede achtergrondkarakteristieken zichzelf in part-time werk selecteren. Personen met minder goede karakteristieken, die logischerwijs vaker een ontoereikend pensioen krijgen, hebben minder mogelijkheden en/of kiezen er minder vaak voor om de voordelen van part-time werk te benutten.

Curriculum Vitae

Jim Been (10/24/1987, Amsterdam) is a researcher in Economics in the project *"Pension savings and consumption needs of current and future retirees"* at the Department of Economics of the Leiden Law School, Leiden University. His research mainly focuses on applying econometrics to aging issues. Several grants have been received from Instituut Gak and Netspar to perform the research. Jim's papers were presented at several international top conferences and seminars, such as the *EEA*, *ESEM*, *IAAE*, *EALE*, *ESPE*, *IIPF* and *IZA*. Policy-relevant work was also presented at the *OECD*, *CPB Netherlands Bureau for Economic Policy Analysis*, *Scientific Council for Government Policy* and *Netherlands Ministry of Social Affairs*. Some of the work has been published in scientific and policy-related journals.

In 2014, Jim was a visting researcher at the *RAND Center for the Study of Aging* at RAND Corp., Santa Monica, California. He has been affiliated as a Netspar Junior Research Fellow, Fellow of the UBS International Center of Economics in Society and Ambassador of the Leiden Convention Bureau of the City of Leiden since 2014.

Jim wrote his Ph.D. thesis in Economics at the Department of Economics, Leiden University. Next to writing his thesis, he assisted evidence-based policymaking by advising the Dutch Socio-Economic Council (*SER*) and the *CPB* Netherlands Bureau for Economics Policy Analysis using econometric models he developed. In addition he participated as a researcher in the *OECD* project of *Retirement Savings Adequacy*.

During the period as a Ph.D. student, Jim received formal education at the *Tinbergen Institute* Graduate School in Economics, Econometrics and Finance with a specialization in Labor Economics and Applied Microe-

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Long-term trends, such as the aging of the population, the increased life-expectancy, and the consequences of the recent financial crisis, have raised concerns about the sustainability of pension systems. Consequently, many OECD countries have proposed and implemented reforms to alleviate the pension system from the pressure of demographic aging and to create sustainable pension systems for the future. Many of the reforms implemented are related to increasing both the statutory and effective retirement age, making pension benefits less generous and increasing contributions. As a consequence, the proposed and implemented reforms have raised a lot of discussion about the financial position of current and future retirees. This thesis collects five studies regarding Pensions, Retirement, and the Financial Position of the Elderly and aims to understand the effects of aging on people's retirement behavior and the adequacy of their (future) pensions. The thesis focuses on the role of self-employment and part-time employment in retirement behavior and the financial resources available at retirement.

This is a volume in the series of the Meijers Research Institute and Graduate School of the Leiden Law School of Leiden University. This study is part of the Law School's research program on 'Reform of Social Legislation'.

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