Early Retirement and Youth Employment in Norway*

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Abstract

This paper studies the short term and micro level effects of the number of jobs released through early retirement on labour market outcomes for young potential entrants. Causal estimates of old-young substitution effects in local labour markets are obtained by means of instrumental variable estimation, exploiting a unique administrative data set covering the entire population of Norway and exogenous variation created by the gradual phasing in of an early retirement programme. The resulting short term substitution effects are rather sizable: For every additional early retirement pensioner there is room for *one* new labour market entrant.

JEL codes: J21, J26, H55.

Keywords: Pension reform, old-young substitution, instrumental variables, matched employer-employee register data.

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1 Introduction

The main purpose of this paper is to investigate short term and micro level effects of the vacating of jobs by elderly workers on the employment prospects of young labour market entrants. Old-young substitution on the labour market, or the idea that early retirement is needed in order to improve job opportunities for younger workers, is an argument that was sometimes used in favour of early retirement schemes, particularly in France and in the UK. This view is widely criticised amongst economists and is often referred to as the "lump of labour" fallacy, as it relies on the seemingly fallible view that the economy runs on a fixed amount of labour. Taken literally, this would imply that unemployment rates should have skyrocketed when populations increased or women entered the workforce, since the number of jobs is constant while the potential workforce increased.[1] The argument has also found little empirical support: Gruber and Wise (2010) collects 12 country studies, none of which find any evidence that retiring workers make room for the young. However, even if the "lump of labour" theory is indeed a fallacy in the long run and at the aggregate level, such substitution may still take place in the short run and at the micro level, though this will only be discernible with more detailed data than that used by Gruber and Wise (2010).

Knowledge about the magnitude and nature of the micro level substitution between older and younger workers is becoming increasingly relevant as most Western countries now are faced with greying populations. Also, knowing how much old-young substitution one may expect might prove useful with respect to the timing of reforms in countries where governments seek to restrict access to early retirement, and substantial old-young substitution could even weaken the case for higher retirement ages as such. Martins, Novo, and Portugal (2009) argue that in the presence of incentive schemes under which the wages of elderly workers exceed the value of their marginal product (Lazear (1981)), increased

[1] Similar arguments have also been used in other settings, for instance in favour of hours restrictions or "work sharing" policies, see e.g. Hunt (1999).

retirement age would be detrimental to firms' profitability. Analysing the effects of higher legal retirement age in Portugal, they find that firms hire approximately one fewer worker for each older worker that is retained due to the higher retirement age. Younger workers, and younger women in particular, are found to be most affected by the lower level of total hirings.

The on-going financial crisis has brought renewed attention to problems related to youth employment, but the question of whether unemployment is a more severe problem at early stages of the career than later on remains an open one (see e.g. Bell and Blanchflower (2011) and Nordström Skans (2011)). There is, however, a vast literature documenting the impact of unemployment on wages and on future labour market prospects[2], and the combination of high (youth) unemployment and the possibility of unemployment hysteresis implies that short term effects may have long term consequences for the economy as a whole.[3] Short term effects of the number of jobs released through early retirement are therefore highly relevant.

Using a unique administrative data set covering the full population of Norway, I construct a sample consisting of repeated cross sections of potential entrants for each of the years 1994–2004. Potential entrants are defined as individuals between 18 and 29 years of age who were not in full time employment by the end of year t. The empirical strategy in this paper is based upon estimating models of the probability of entering the labour market in year $t+1$ for young potential entrants, as a function of a proxy for the number of jobs released through early retirement in different labour market regions. The main challenge related to assessing the causal relationship between early retirement and youth employment in this setting is represented by business cycles or shocks to the economy potentially affecting employment at all ages in the same direc-

[2]See inter alia Beaudry and DiNardo (1991), Gregg (2001), Mroz and Savage (2006), Raaum and Røed (2006) and Kahn (2010).

[3]The term *hysteresis (in unemployment)* is used to describe a situation in which increases in actual unemployment are causing structural changes associated with higher equilibrium rates of unemployment, and is also sometimes used somewhat more loosely to describe any mechanism that leads to transitory shocks having persistent, but not necessarily permanent effects. See Røed (1997) for a survey of various hysteretic explanations offered in the literature.

tion. For instance, it may be that firms and workers in declining industries use early retirement as a way to cushion the effects of downsizing. This sort of endogeneity would lead to negative biases in the OLS estimator. To tackle this problem I make use of an instrumental variable strategy in which reductions in the lower age limit for early retirement provide exogenous variation in the number of jobs released through early retirement.

After having confirmed that the first stage in the instrumental variable strategy exists with a strong and highly significant relationship between the instrument and the causal variable of interest, a comparison of OLS and 2SLS point estimates confirms the suspicion that OLS estimates would be negatively biased. Moreover, the 2SLS point estimate from the baseline specification implies that for every additional early retirement pensioner there is room for *one* new labour market entrant in the studied age spans. A battery of robustness checks is performed in order to assess the support for a causal interpretation of the main results. Finally, the very same empirical framework is utilized in an assessment of the effects of the number of jobs released through early retirement on the incidence and duration of unemployment for young potential entrants. Effects on unemployment are negative, as expected, but less robust to variations in the empirical specifications than are the effects on the probability of active labour force participation, and should therefore be interpreted with caution.

The paper proceeds as follows: Section 2 describes the institutional setting and Section 3 describes the data and the construction of the sample of potential entrants. Section 4 spells out the empirical specification and discusses identification, before the main results are given in Section 5. Section 6 presents the results from a series of robustness checks, and Section 7 reports results on the incidence and duration of unemployment. Section 8 concludes.

2 Institutional setting

AFP (AvtaleFestet Pensjonsordning) is a subsidized voluntary early retirement scheme that was introduced January 1 1989 as a result of the central tariff negotiations in 1988. The scheme started out with a lower age limit at 66 years, which was reduced to 65 from January 1 1990, to 64 from October 1 1993, to 63 from October 1 1997, and finally to 62 years from March 1 1998. In the public pension scheme the retirement age is 67, and prior to 1989 there were no purely voluntary early retirement options available to workers below this age. Possible quasi-voluntary or informal exit routes were unemployment, sickness leave and disability pensions. These were claimed to be associated with social stigma, and the need for a more dignified exit from the workforce for early leavers was one of the arguments in favour of AFP.

For a worker to qualify for early retirement through AFP (hereafter referred to simply as early retirement (ER)) she had to be employed by a firm affiliated with the ER scheme when she reached the lower age limit. While all public sector firms are affiliated with the ER scheme, coverage in the private sector is limited to firms taking part in the central tariff agreements. These firms employ about half the workers in the private sector. In addition to firm affiliation there are also some rather weak requirements related to individual work histories[4] that must be met prior to retirement.

The ER scheme is fairly generous[5], as the pension level received is the same as it would be had the person continued working until the age of 67 in the job she held just prior to retirement. In addition comes a subsidy of 950 NOK/month

[4]The individual requirements are (i) at least 10 years of work experience (earnings exceeding 1 Basic Amount (BA)) since the age of 50, (ii) the average of the 10 highest yearly incomes since 1967 must exceed 2BA, (iii) current employment, and (annualized) earnings at least equal to 1BA in the year of retirement and the year before, and iv) at least 3 years of tenure in the present firm. The Basic Amount is frequently referred to as G, and is a central feature of the public pension system in Norway. It is adjusted every year, with a nominal rate of growth varying between 2 and 14% since its introduction in 1967. The average BA for 2012 is 81,153 NOK, which at the time of writing corresponds to about 10,800 EUR or 8,700 GBP.

[5]Røed and Haugen (2003) find that average replacement rates, net of taxes, for ER pension benefits, disability pension benefits and unemployment benefits are 72, 64 and 62 percent, respectively. Sickness leave is another informal exit route which gives a benefit replacement rate of 100 percent, but for a maximal duration of 12 months.

during the early retirement years, and when an ER pensioner reaches the age of 67 she will receive a public pension which is calculated as if she had been working until that age. The full costs are borne by the participating employers for pensioners below the age of 64, by means of a fund financed by fees per employee varying according to hours worked (three categories), whereas the government covers 40% of the costs for those of age 64 to 67.

Figure 1 shows the number of ER pensioners as fractions of all individuals of age 62-66 for the country as a whole over the period 1994–2004, which is the period studied in this paper. It increased from about four per cent of all individuals of age 62-66 in 1994, when the lower age limit for ER was 64, to about 25 per cent in 2004, when the lower age limit was at 62 years. There were two reductions in the lower age limit over this time period; from 64 to 63 in 1997 (October 1) and from 63 to 62 in 1998 (March 1).

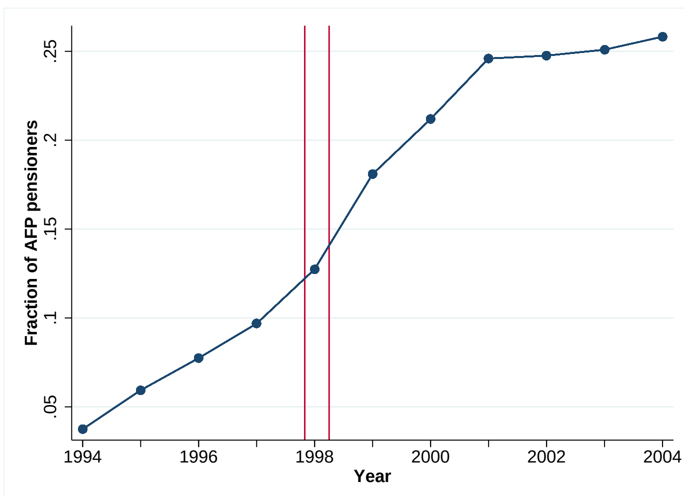

Figure 1: ER pensioners as fractions of all individuals of age 62-66, 1994-2004. Vertical lines indicate changes in the lower age limit for ER.

3 Data

The data used in this paper combines several administrative registers obtained from Statistics Norway. One is the Register of Employers and Employees, which covers the entire Norwegian working age population over the period 1992-2009 and gives both firm and individual specific information for all job spells. The data contains demographic information for all residents, identifies recipients of ER pensions, and includes individual earnings data (in terms of pension points) dating back to 1967.

When constructing the sample I restrict attention to individuals between 18 and 69 years of age in each of the years 1994–2004. Individuals are grouped into labour market regions (LMRs)[6] according to their municipality of residence in year t. The active labour force in each LMR is defined as individuals in full-time employment (working at least 35 hours/week with annualized earnings exceeding two Basic Amounts[7]) at the end of each year t. Potential entrants are individuals who were below the age of 30 and not part of the active labour force in year t; a potential entrant in year t is classified as an entrant in year $t+1$ if she was part of the active labour force of some labour market region in year $t+1$. Table 1 gives a list of variables describing the demographic composition of the active labour force in each LMR, and Figure 2 gives box plots[8] that summarize both cross-sectional variation and variation over time in selected LMR demographic characteristics. We note that the percentage of individuals between 18 and 69 years of age being part of the active labour force is typically around 40 (lower left panel), and the percentage of potential entrants around 20 (upper left panel).

Some descriptive statistics for a repeated cross-sectional sample of potential entrants over the period 1994–2004 are given in Table 2; individual characteristics in the upper panel and LMR characteristics in the lower panel. We note

[6]The labour market regions are defined on the basis of commuting patterns and documented in Bhuller (2009).

[7]See footnote 4.

[8]The lower and upper hinges of the boxes indicate the 25^{th} and 75^{th} percentiles, respectively, denoted by $x_{[25]}$ and $x_{[75]}$, and the horizontal lines cutting through the boxes indicate the median. The vertical lines below and above the boxes are called adjacent lines, and the markers on each end of the lines indicate lower and upper adjacent value, respectively. Adjacent values are calculated as described in the Stata Manual [G] Graphics: Define x_i as the ith ordered value of x, and define U as $x_{[75]} + \frac{2}{3}(x_{[75]} - x_{[25]})$ and L as $x_{[25]} - \frac{2}{3}(x_{[75]} - x_{[25]})$. The upper adjacent value is x_i such that $x_i \leq U$ and $x_{i+1} > U$, and the lower adjacent value is x_i such that $x_i \geq L$ and $x_{i+1} < L$. Observations above (below) the upper (lower) adjacent values are shown as dots the figure.

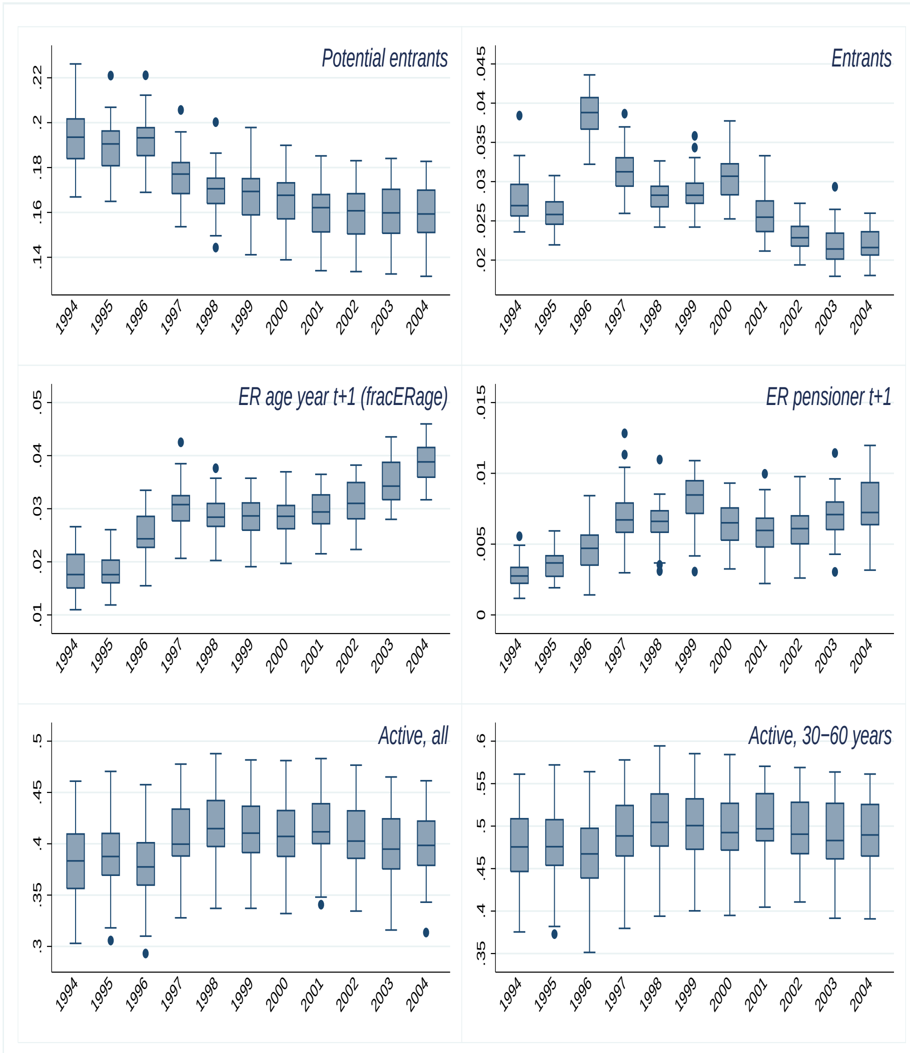

Figure 2: Box plots of the number of potential entrants, actual entrants, individuals in the active labour force in year t who reach the lower age limit for ER within year $t+1$, individuals in the active labour force in year t who receive ER pension benefits in December year $t+1$, and the number of individuals in the active labour force, as fractions of the total number of citizens in each LMR.

Table 1: Variable definitions

Variable	Definition
	The proportion of individuals in the active labour force in year t ...
fracERage	... who reach the lower age limit for ER within year $t+1$
fracpens	... who receive AFP pensions in December year $t+1$
fracmale	... who are male
fracimm	... with immigrant background
fracguest	... who are classified as guest workers
fraclow	... with low education
frachi	... with higher education

that, besides being younger and less educated, the potential entrants differ considerably from the active labour force in terms of the fraction of male and the fraction of immigrants.

4 Empirical specification and identification

A natural starting point for studying the causal relationship between the number of jobs released through ER and youth employment would be to estimate the probability of being part of the active labour force in year $t+1$ for potential entrants in each year t by means of linear regression;

$$P_{ijt} = \alpha^0 + F_j^0 + \lambda_t^0 + X_{it}^{'}\beta^0 + \delta_{OLS} fracpens_{jt} + \varepsilon_{ijt}^0 \qquad (1)$$

P_{ijt} is an indicator for participation in the active labour force in year $t+1$ for year t potential entrants, F_j are unobserved LMR level fixed effects, λ_t are year fixed effects, and X_{it} a vector of individual and (time-varying) LMR level characteristics: Dummies for age, educational attainment, immigrant status and sex; fraction of male, fraction of immigrants, fraction of guest workers, fraction with low and fraction with high education in the active labour force, and fraction of the active labour force in different industries. $fracpens_{jt}$ is the proportion of individuals in the active labour force in year t who received ER pension benefits in December year $t+1$, and is used as a proxy for the number of jobs released through ER.

Table 2: Summary statistics for potential entrants, $t \in [1994, 2004]$

Variable	**Mean**	**Std. Dev.**	**Min.**	**Max.**
Entrant	0.164	0.37	0	1
Male	0.446	0.497	0	1
Immigrant	0.096	0.294	0	1
Educational level				
Low	0.129	0.335	0	1
Intermediate	0.744	0.436	0	1
High	0.127	0.333	0	1
Age	23	3.439	18	29
	LMR statistics			
Number of ...				
... citizens	307,699	351,654	11,696	922,368
... active	141,453	167,037	3,506	430,528
... pensioners	793	972	7	3,111
Fraction of ...				
... active	0.428	0.038	0.293	0.488
... potential entrants	0.169	0.016	0.132	0.226
... entrants	0.164	0.026	0.111	0.23
fracERage	0.03	0.007	0.011	0.046
fracpens	0.006	0.002	0.001	0.013
fracmale	0.637	0.041	0.555	0.775
fracimm	0.048	0.026	0.008	0.102
fracguest	0.008	0.008	0	0.074
fraclow	0.119	0.026	0.062	0.232
fracintermed	0.619	0.068	0.482	0.754
frachi	0.261	0.081	0.111	0.412
N		5,393,241		

The OLS estimator δ_{OLS} is unbiased if there are no omitted (unobservable) variables that are correlated with both *fracpens* and the error term, i.e. if $Cov\left(fracpens_{jt}, \varepsilon^0_{ijt} | covariates\right) = 0 \, \forall \, i, j, t$. The main reason to suspect that this condition is violated is that shocks to the economy may affect employment at all ages in the same direction. For instance, it may be that firms and workers in declining industries use ER as a way to cushion the effects of downsizing. To get an idea of the likely sign of the bias in δ_{OLS}, take the downsizing mechanism as an example: If more workers tend to exit through ER and recruitment of young workers is lower in bad times than in good times, the omitted business cycle shocks will be negatively correlated with $fracpens$, while they are likely to be positively correlated with the probability of entry. Hence, δ_{OLS} will be negatively biased relative to the causal effect of $fracpens$ on P_{ijt}. Put differently, high levels of an explanatory variable would be associated with low values of the error term, biasing the coefficient downwards.[9]

To cope with the endogeneity issues associated with estimating equation (1) by means of linear regression I use *fracERage* as an instrument for *fracpens*. The first stage is given by equation (2):

$$fracpens_{jt} = \alpha^1 + F^1_j + \lambda^1_t + X'_{it}\beta^1 + \gamma fracERage_{jt} + \varepsilon^1_{ijt} \tag{2}$$

The two reductions in the lower age limit in 1997 and 1998 provide exogenous variation in the number of workers at risk of exiting through ER. Conditional on the other covariates in equation (2), much of the variation in *fracERage* is due to these reductions in the lower age limit, and *fracERage* is therefore likely to be a valid instrumental variable for *fracpens*; correlated with *fracpens*, and correlated with P_{ijt} only through *fracpens*.

Figure 3 gives a graphical depiction of the first stage regression, which reveals a clear positive relationship between *fracERage* and *fracpens*. The second stage is given by equation (3), where *fracpens* is replaced by fitted values from the

[9]This line of reasoning follows the omitted variables bias formula in Angrist and Pischke (2009, p. 60).

first stage:

$$P_{ijt} = \alpha + F_j + \lambda_t + X_{it}^{'}\beta + \delta_{2SLS}\widehat{fracpens}_{jt} + \varepsilon_{ijt} \tag{3}$$

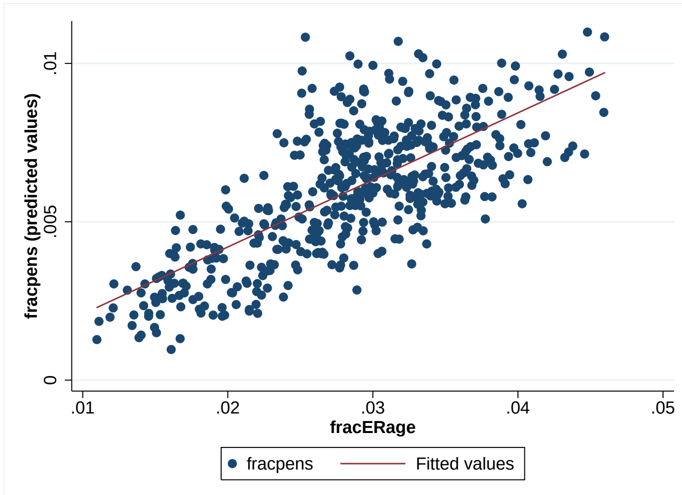

Figure 3: Scatter plot of *fracERage* against fitted values of *fracpens* from OLS on equation (2) with LMR characteristics only. The slope of the least squares regression line drawn through the points is 0.249, with a standard error of 0.022.

The validity of the instrument *fracERage* will be questionable if there are factors other than $fracpens$ that are correlated with $fracERage$ *and* with P_{ijt}. In other words, if something influences the measure of workers soon eligible for ER, i.e. the fraction of the active labour force reaching the lower age limit for ER in the next year, while also influencing the probability that young people in the same LMR will enter the active labour force. One such factor could be the demographic composition of different LMRs, which could be driven by labour market conditions. For instance, if young potential entrants tend to move away from declining LMRs to more prosperous ones, then the age cohort about to reach early retirement age would make a larger share of the active labour force in the declining LMRs. LMRs in decline would then cause higher values of *fracERage*, as younger cohorts move away, as well as lower incentives for young people to enter the active labour force. Again, high levels of the explanatory variable, this time the instrument, would be correlated with negative values of the error term in equation (3). On the other hand, one could imagine that elderly workers remaining in the active labour force is a sign of a well-functioning labour market in which also younger workers are likely to succeed. Regardless of the sign of the possible bias, one might want to include controls for the demographic composition in the regressions. One possible way of doing so is described in Section A.1 in the Appendix. With more demographic controls there is also a risk of throwing away relevant variation, however, and 2SLS point estimates from this alternative specification are generally closer to zero and less precisely estimated than those from equation (3). Since the broad picture emerging is largely the same regardless of whether additional demographic controls are included or not, I have chosen to rely mostly on the more parsimonious of the two specifications. I will comment on the relatively few significant differences across specifications when appropriate.

Examples of other factors that could somehow be correlated with both $fracERage$ and with P_{ijt} are the flows of immigrants and guest workers into different LMRs, representing shifts in the supply of certain types of workers. By comparing Figure 2 and 4 we see that the trend in $fracERage$ (the fractions of the active labour force reaching ER age in year $t+1$) is somewhat similar to that of the fraction of immigrants in the active labour force, while this is not so much the case for the fractions of guest workers. The fraction of immigrants is included as a control variable in all regressions, and robustness checks in the following section confirm that the results are not sensitive to whether or not the LMRs with the highest fractions of immigrants and guest workers are included in the sample.

Figure 5 plots the changes in the lower age limit of the early retirement scheme AFP against the nation-wide unemployment rate. We see that the ER scheme was introduced and decisions to lower the age limit to 65 and 64 were made in a period of rising unemployment. The unemployment rate had just started to decline by the time when the age limit was lowered to 64 in 1993, and was still declining when the lower age limit was set to 63 in 1997

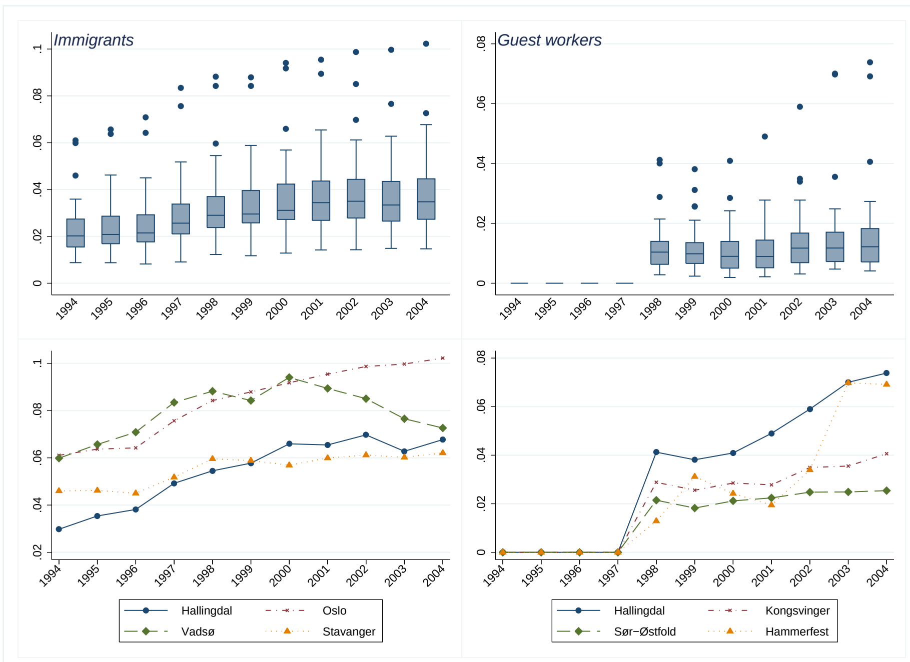

Figure 4: Box plots of the fraction of first and second generation immigrants in the active labour force (upper left panel) and the number of guest workers as fractions of all workers (upper right panel), along with the time series for the four LMRs with the highest fractions (lower panels). Guest workers are identified in the data only from 1998 and onwards.

and finally to 62 in 1998. The observation period in this paper is 1994-2004, which means that the variation in the lower age limit for ER that is used for identification of δ_{2SLS} coincides with a declining trend in the economy wide unemployment rates. This would not represent a major threat to the validity of the instrument, however, since the variation used in the second stage is what remains after conditioning on LMR fixed effects, economy wide business cycle conditions captured by year dummies, and time varying LMR characteristics contained in X_{it}. Some informal evidence in support of this claim is given in the following section: Point estimates obtained from a sample restricted to the years of reductions in the lower age limit and those just before and just after are smaller, but not significantly different from those obtained from estimations over the full observation period.

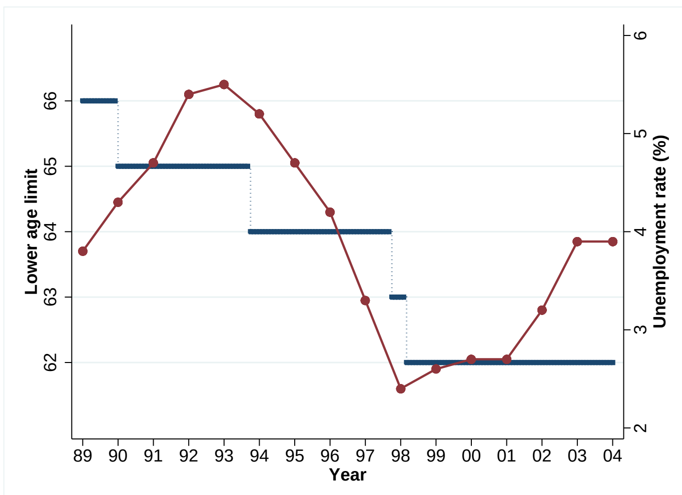

Figure 5: Lower age limits for the early retirement scheme AFP along with register based unemployment rates from the Norwegian Labour and Welfare Administration (NAV).

5 Results

Estimation results for the effects of *fracERage* on *fracpens*, i.e. equation (2), the first stage, and for the effects of *fracpens* on the probability of entry into the active labour force (P_{ijt}) are given in Table 3[10]; for the full sample of potential entrants, for men and women separately (upper panel), and separately for potential entrants divided into three groups according to educational attainment (lower panel). Starting with the first stage estimations, we note that *fracERage* has a sizable and statistically significant effect on *fracpens*: The first stage point estimate for the full sample implies 0.271 new ER pensioners for every active worker who reaches the lower age limit. Also, the first stage F-statistic against the null hypothesis that the excluded instrument is irrelevant is well above the conventional threshold of 10.

Turning to the effects of *fracpens* on P_{ijt} we first note that all OLS estimates are very small in magnitude. The importance of taking account of the endogeneity of *fracpens* becomes clear, however, when we compare OLS estimates with the 2SLS estimates: The latter being much larger in magnitude confirms our suspicion that OLS estimates would be negatively biased. Using $\hat{\delta}_{2SLS}$ from Columns 4 and 6, the mean probability of entry for male (0.21) and female (0.13) potential entrants and the sample mean of *fracpens*, we can calculate the elasticity of P_{ijt} with respect to *fracpens* at the sample mean of both variables; 0.10 (men) and 0.13 (women). Going one step further, the estimated effect of one additional ER pensioner on the mean probability of entry is $2.11 \cdot 10^{-5}$ percentage points, which implies that for every additional ER pensioner there is room for *one* new entrant.[11]

Dividing the estimation sample into sub-groups according to the potential entrants' educational attainment might give a first indication of how skills matter for the effects of jobs released through early exits. In the lower panel of Table 3, low education corresponds to compulsory education only or missing information on educational attainment; intermediate corresponds to high school (or some high school), but no tertiary education; and high corresponds to (some)

[10] More complete sets of coefficient estimates from the first and second stage regressions are provided in Table A2 and Figure A1 in the Appendix.

[11] $\Delta\bar{P} = \hat{\delta}_{2SLS} \cdot \Delta \mathit{fracpens} = 2.11 \cdot 10^{-5}$, where one new ER pensioner, keeping all else constant, implies that $\Delta \mathit{fracpens} = \frac{1}{141{,}453}$. Assuming that the number of potential entrants is constant we multiply $\Delta\bar{P}$ by the average number of potential entrants $(52,001)$ to find that for every additional ER pensioner there is room for one new entrant $(2.11 \cdot 10^{-5} \cdot 52,001 \approx 1.1)$.

university level education. The magnitude of the 2SLS point estimates indicate that effects are larger for potential entrants with higher skill levels.

Table 3: Estimation results for equations (1)–(3)

	All		**Men**		**Women**	
	OLS	IV	OLS	IV	OLS	IV
First stage						
fracERage		0.271^{***}		0.270^{***}		0.273^{***}
		(0.024)		(0.024)		(0.024)
F-statistic		126.0		123.6		127.9
Second stage						
fracpens	0.379	2.990^{**}	0.306	3.551^{**}	0.506	2.820^{**}
	(0.256)	(1.162)	(0.421)	(1.464)	(0.314)	(1.259)
N	5,393,241		2,407,898		2,985,343	
R^2	0.033		0.024		0.026	

	Level of education					
	Low		**Intermediate**		**High**	
	OLS	IV	OLS	IV	OLS	IV
First stage						
fracERage		0.262^{***}		0.270^{***}		0.292^{***}
		(0.024)		(0.024)		(0.024)
F-statistic		120.7		123.4		146.0
Second stage						
fracpens	-0.024	-0.551	0.501^{*}	3.309^{***}	-0.042	2.337
	(0.643)	(1.704)	(0.292)	(1.216)	(0.872)	(2.675)
N	695,575		4,014,499		683,167	
R^2	0.026		0.012		0.043	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable takes the value 1 if a year t potential entrant was part of the active labour force in year $t + 1$ and is zero otherwise. *fracERage* is used as IV in the 2SLS estimations.
Additional covariates are year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).
Low education corresponds to compulsory education only or missing information on educational attainment; intermediate corresponds to high school (or some high school), but no tertiary education; high corresponds to (some) university level education.

Figure 6 shows average elasticities of P_{ijt} with respect to $fracpens$ from equation (3) estimated by means of a probit model; separately for different age groups, and separately for men and women. Starting with those based on estimates from the samples with both men and women we note that the elasticities are significantly different from zero at the five percent level at age

19, 21, 24 and 28. 19 is the age at which most students finish high school and thus have to decide whether to go on to higher education or try to find a job. While most students will have had time to finish studies corresponding to a master's degree when they reach age 25 and above, 24 is not to the same extent a typical age for major schooling or work decisions. When the model is estimated separately for men and women, the elasticity is significantly different from zero only at age 28 for men, but the point estimates are positive for all ages. For women they are positive and significant at age 19, 21 and 28.[12]

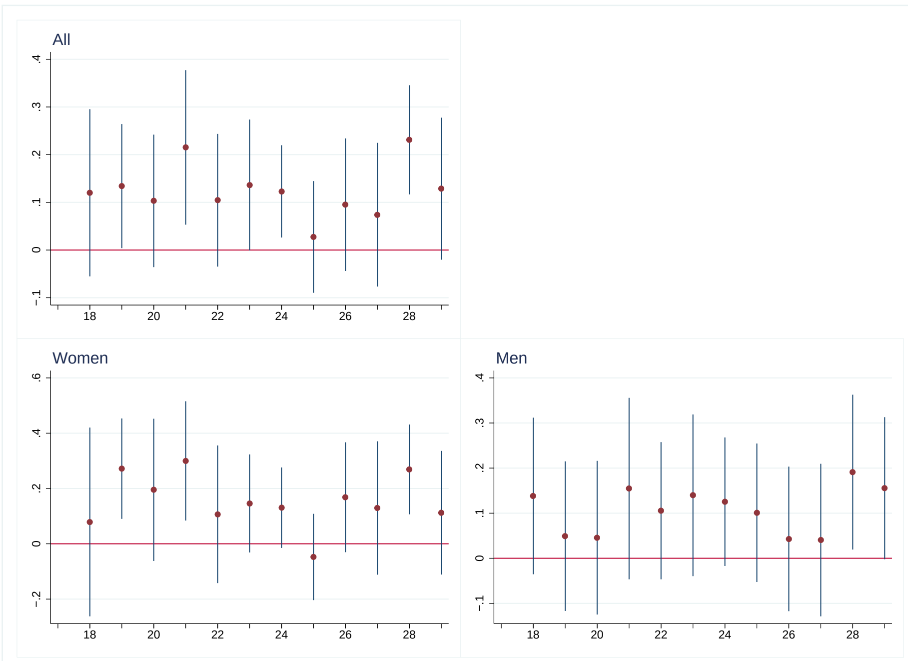

Figure 6: Average elasticities of P_{ijt} with respect to *fracpens* from equation (3) estimated as a probit model. The model is estimated separately for different age groups, and separately for men and women. Vertical spikes correspond to 95% confidence intervals.

[12]This means that out of 24 independent age and gender specific estimates, as many as 23 are positive. It is worth noting that under the binomial distribution, this outcome would have a probability of roughly one in 700,000 under the null hypothesis of no effect. Similarly, the probability of getting four out of 24 estimates significant at the 5% level is itself less than 5% under the null hypothesis.

6 Robustness

6.1 Sensitivity checks

The results of three different sensitivity checks are given in Table 4, along with results from the main specification (upper left panel). First, equations (1)-(3) are estimated on data covering the years 1996–1999 only, i.e. the years of reductions in the lower age limit and those just before and just after. This is a way of investigating whether or to what extent the results in Table 3 are driven by a declining trend in unemployment, coinciding with the reductions in the lower age limit for ER that are being used for identification of causal effects. We see from the upper right panel of Table 4 that the first stage still exists, and although the 2SLS point estimate is smaller and less precisely estimated than that from the full sample, the two point estimates are not substantially different.

To test whether inflows of immigrants and guest workers into different LMRs are likely to contaminate the results, I have estimated equations (1)-(3) on samples from which four LMRs with very high shares of immigrants and guest workers, respectively, are excluded. Results from these exercises are reported in the lower panels of Table 4, and confirm that the 2SLS point estimates are robust also to such variations in the estimation sample.

Estimation results from the specification with additional controls for the demographic composition of different LMRs are given in Table A1 in the Appendix. When comparing the 2SLS point estimates in the two tables we see that although they are smaller and less precisely estimated in Table A1 than in Table 3, the point estimates are significantly different only for male potential entrants.

As potential entrants are identified on the basis of no registered employment and/or annualised earnings below two basic amounts, and as sickness and parental leave benefits are not recorded in the data as regular earnings, some of those classified as potential entrants will actually be on leave from what would

Table 4: Estimation results for equations (1)–(3), different samples

	Full sample		**1996-1999 only**	
	OLS	IV	OLS	IV
First stage				
fracERage		0.271***		0.242***
		(0.024)		(0.051)
F-statistic		126.0		22.3
Second stage				
fracpens	0.379	2.990**	-1.532***	1.328
	(0.256)	(1.162)	(0.564)	(2.633)
N	5,393,241		1,975,479	
R^2	0.033		0.034	

	Drop four LMRs with high immigrant shares		**Drop four LMRs with high guest worker shares**	
	OLS	IV	OLS	IV
First stage				
fracERage		0.267***		0.270***
		(0.025)		(0.025)
F-statistic		108.8		114.4
Second stage				
fracpens	0.722***	3.344***	0.550**	3.543***
	(0.263)	(1.046)	(0.268)	(1.140)
N	3,520,644		5,104,542	
R^2	0.034		0.033	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 (42) LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable takes the value 1 if a year t potential entrant was part of the active labour force in year $t+1$ and is zero otherwise. *fracERage* is used as IV in the 2SLS estimations.

Additional covariates are year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).

be seen as "proper" jobs. Sickness and parental leave benefits recipients are protected against job loss while on leave, and hence one could suspect that their presence in the sample might tend to overstate the effect of $fracpens$ on the probability of entry. However, excluding some $600,000$ sickness and parental leave benefits recipients from the sample leads only to negligible changes in the 2SLS estimates (results not reported).

6.2 ER affiliated vs non-affiliated firms

Given that we have found that early exit by elderly employees does indeed improve the prospects for young potential entrants one would expect the effects to a large extent to be driven by entry into firms from which jobs have been released, that is, into ER affiliated firms. This is confirmed by the estimates reported in Table 5, where estimated effects of $fracpens$ on the probability of entry into any firm (the baseline specification) are compared with estimated effects on the probability of entry into ER affiliated firms and on the probability of entry into non-affiliated firms. 2SLS point estimates are significantly different from zero both for the probability of entry into ER firms and for the probability of entry into non-ER firms; the former is twice the size of the latter, and the sum of the two is exactly equal to the estimated effect of *fracpens* on the probability of entry into any firm (the main specification). These results confirm our expectations, and at the same time they could be taken as indications of non-negligible second order effects following the positive labour demand effect in ER firms. One example of such effects could be a positive shift in aggregate demand, if the reduced demand from ER pensioners is more than compensated by increased demand from new entrants, leading to positive shifts in the labour demand of both ER and non-ER firms. Alternatively, it could be that jobs in ER firms are given to workers in non-ER firms rather than to less experienced potential entrants, and that potential entrants in turn get some of the jobs released in non-ER firms.

Results from a model of the probability of entry into ER affiliated firms relative to entry into non-affiliated firms, that is, equation (1)-(3) estimated for actual entrants only, are given in the lower right panel of Table 5. Despite the 2SLS point estimate being highly imprecise, its sign does indicate an increase in the relative probability of entry into affiliated firms.

Table 5: Evaluating different mechanisms

	P(entry) **All**		**P(entry into ER firm)** **All**	
	OLS	IV	OLS	IV
fracpens	0.379	2.990**	-0.211	1.931**
	(0.256)	(1.162)	(0.201)	(0.898)
N	5,393,241		5,393,241	
R^2	0.033		0.018	
	P(entry into non-ER firm) **All**		**P(entry into ER firm)** **Entrants**	
	OLS	IV	OLS	IV
fracpens	0.590***	1.059*	-2.701***	0.568
	(0.181)	(0.603)	(0.868)	(3.147)
N	5,393,241		882,845	
R^2	0.016		0.020	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. *fracERage* is used as IV in the 2SLS estimations.

Additional covariates are year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).

6.3 Alternative instruments

With the very rich data we have at hand it is possible to define several potentially valid instruments for $fracpens$ that could be used in addition to or as alternatives to $fracERage$. One would be the fraction of active working in ER affiliated firms, another the fraction of active satisfying (a subset of) the individual criteria for ER[13], and yet another candidate would be the fraction of active working in ER affiliated firms *and* satisfying the individual criteria *and* reaching the lower age limit for ER in year $t+1$.

Box plots of $fracERage$ and the three other alternative instruments are given in Figure 7. The variable *frac(ERage)x(Indcrit)x(ERfrm)* in panel d is a combination of the three other instruments; its variation will be due to the interaction of the demographic composition of different LMRs, reductions in the lower age limit for ER, the composition of firms (affiliated and not affiliated with ER) and of workers (satisfying and not satisfying individual criteria). The clear advantage of this instrument over $fracERage$ is that it exploits more of the available information in the data, meaning that it should have more predictive power with respect to the endogenous variable $fracpens$. The disadvantage is that it increases the risk of biases, both due to possible errors in the procedure used to identify firms' affiliation with the ER scheme[14], and because the composition of firms and workers in different LMRs may be correlated with entry probabilities via other channels than through the fraction of pensioners. The fraction of ER affiliated firms would be positively correlated with the probability of entry if, for instance, affiliated firms are more stable, recruit more regularly and provide longer lasting jobs relative to non-affiliated firms. It could also be that the fraction of active workers satisfying the individual criteria for ER is

[13]The following individual criteria are checked for the construction of this instrument: At least 10 years after the age of 50 with earnings exceeding one BA, the average of the 10 highest earnings since 1967 corresponding to at least two BAs, and earnings in year $t-1$ corresponding to at least one BA.

[14]There is no direct information on ER affiliation at the firm level. To separate affiliated from non-affiliated firms I have identified all recipients of public and private ER pension benefits from 1992 to 2008 and tried to identify the last job prior to pension take-up for each of these individuals.

positively correlated with P_{ijt}, for instance if high fractions of workers satisfying these criteria indicate that workers are more productive and hence that the firms employing them are in some sense better, and possibly also that jobs in these regions are more stable.

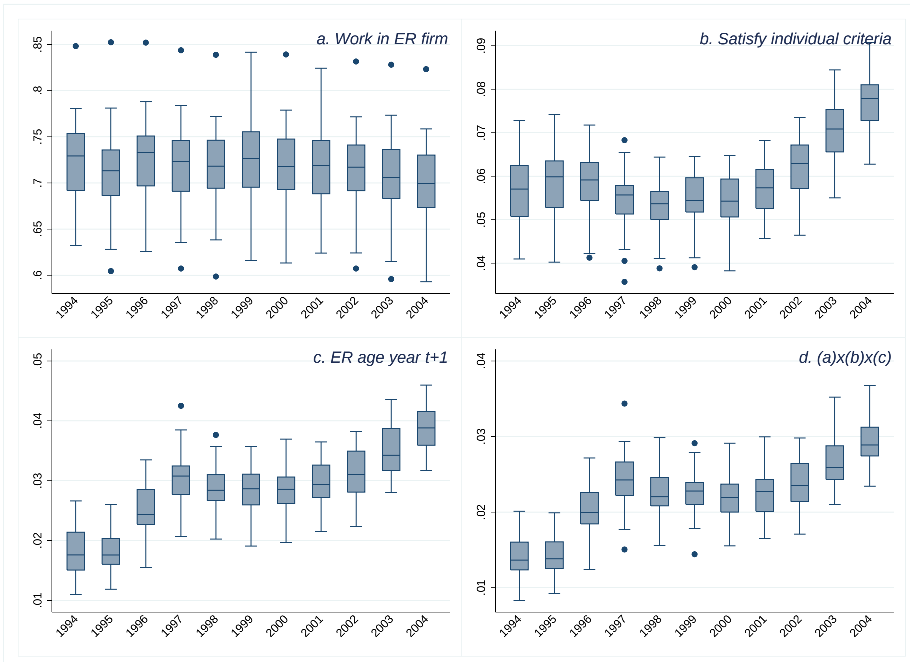

Figure 7: Box plots of alternative instruments; the fraction of active working in ER affiliated firms (panel a), the number of active satisfying individual individual ER criteria as fractions of all active (panel b), the fraction of active reaching ER age in year $t + 1$ (panel c), and the fraction of active working in ER affiliated firms *and* satisfying the individual criteria *and* reaching the lower age limit for ER in year $t + 1$ (panel d).

Figure 8 illustrates the first stage and Table 6 gives estimation results from equation (1)-(3) when *frac(ERage)x(Indcrit)x(ERfrm)* replaces $fracERage$ as instrument for $fracpens$. The first stage is indeed stronger when the refined instrument is used, and the three 2SLS point estimates are smaller in magnitude when compared to the corresponding estimates in Table 3. When $fracERage$, the fraction of active working in ER firms and the fraction of active satisfying the individual criteria are used as three separate instruments, only $fracERage$ is significantly different from zero in the first stage, and IV redundancy tests fail to reject that the two latter instruments are redundant. Moreover, the 2SLS point estimates are practically identical to those from estimations with $fracERage$ as the only instrument and are therefore not reported.

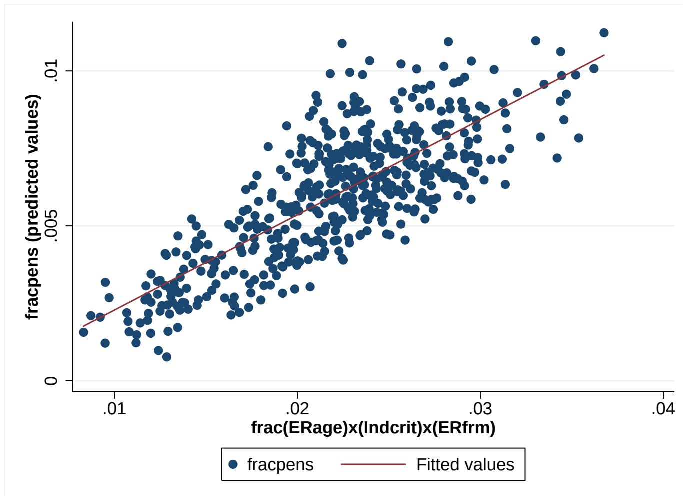

Figure 8: Scatter plot of *frac(ERage)x(Indcrit)x(ERfrm)* against fitted values of *fracpens* from OLS on equation (2) with LMR characteristics only. The slope of the least squares regression line drawn through the points is 0.326, with a standard error of 0.023.

Table 6: Estimation results for equations (1)–(3) using an alternative instrument

	All		**Men**		**Women**	
	OLS	IV	OLS	IV	OLS	IV
First stage						
(ERage)x(Indcrit)x(ERfrm)		0.351^{***}		0.348^{***}		0.353^{***}
		(0.022)		(0.022)		(0.022)
F-statistic		255.3		253.5		256.5
Second stage						
fracpens	0.379	2.237^{**}	0.306	2.695^{**}	0.506	2.066^{**}
	(0.256)	(0.896)	(0.421)	(1.204)	(0.314)	(0.952)
N	5,393,241		2,407,898		2,985,343	
R^2	0.033		0.024		0.026	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable takes the value 1 if a year t potential entrant was part of the active labour force in year $t + 1$ and is zero otherwise. *frac(ERage)x(Indcrit)x(ERfrm)* is used as IV in the 2SLS estimations.

Additional covariates are year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).

7 Effects on unemployment

Thus far the focus has been on estimating causal effects of the number of jobs released through ER on the probability of entry into the active labour force. We have established that exogenous increases in the fraction of ER pensioners do increase the probability of entry for young potential entrants, but inasmuch as the ultimate goal is to make claims about the welfare implications of ER programmes, this result is only part of the story. Knowing how much of the effect that is due to reduced educational attainment or shorter duration of education[15] on the one hand, and to a lower incidence or shorter duration of unemployment on the other hand, is another essential part of the story. This section represents a first step in that direction.

Some descriptive statistics on the incidence and duration of registered unemployment in year $t+1$ for year t potential entrants are given in Table A3 in the Appendix; estimation results from equation (1) and (3) estimated with an indicator for registered unemployment and with the number of months with registered unemployment, respectively, as dependent variables are given in Table 7. All OLS coefficients are relatively small in magnitude and all but one are non-significant. When we adjust for the endogeneity of *fracpens* by means of 2SLS estimation, however, the coefficient of interest increases considerably in magnitude and comes out with the expected negative sign. Hence, the causal effects of *fracpens* on unemployment appears to be negative. To compare these results with those in Section 5, we compute that five new ER pensioners reduce the number of year t potential entrants registered as unemployed in year $t+1$ by 1.5. The effect of one additional ER pensioner on the expected number of months in unemployment of the young potential entrants is a rather moderate $-2.79 \cdot 10^{-4}$. It should be noted, however, that when the models of the incidence

[15] Raaum and Røed (2006) find that although local labour market conditions do not appear to affect the individuals' educational attainment as such, higher local unemployment seems to postpone graduation. Possible reasons why students might choose to postpone graduation when local labour market conditions are unfavourable, they argue, are relatively low opportunity costs of educational activities under such circumstances, and relatively fierce competition for jobs requiring better exams to obtain an acceptable job.

and duration of unemployment are estimated with the additional controls for the demographic composition of LMRs, the resulting 2SLS point estimates are much smaller in magnitude, and no longer significantly different from zero.[16]

[16] I have also estimated the effects of $fracpens$ on the probability of entry, the probability of unemployment and on the number of months with registered unemployment in year $t+1$ for a sub-sample of $261,141$ potential entrants registered as unemployed in December year t. The 2SLS point estimates are generally very imprecise and results from these exercises are therefore not reported.

Table 7: Estimation results for the incidence and duration of unemployment

	All		Men		Women	
	OLS	IV	OLS	IV	OLS	IV
P(unemployment year $t+1$)						
fracpens	0.750	-4.525**	0.415	-6.079***	1.066	-3.244*
	(0.769)	(1.945)	(0.972)	(2.498)	(0.763)	(1.843)
R^2	0.026		0.029		0.017	
Months of unemployment year $t+1$						
fracpens	1.393	-22.86***	-0.101	-28.40**	2.695*	-18.46**
	(1.183)	(8.809)	(1.930)	(11.52)	(1.460)	(7.470)
R^2	0.025		0.027		0.018	
N	5,393,241		2,407,898		2,985,343	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable in the upper panel takes the value 1 if a year t potential entrant was registered as unemployed for at least one month in year $t+1$ and is zero otherwise. The dependent variable in the lower panel is the number of months of registered unemployment in year $t+1$. *fracERage* is used as IV in the 2SLS estimations.

Additional covariates are year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).

8 Conclusion

The purpose of this paper has been to investigate the short term and micro level effects of the number of jobs released through early retirement on labour market outcomes for young potential entrants. Using reductions in the lower age limit for early retirement as a source of exogenous variation in the number of jobs released, I have obtained estimates implying that for every additional early retirement pensioner there is room for *one* new entrant. Results are stronger for more educated potential entrants. We have also seen that effects on the incidence and duration of unemployment have the expected negative sign, but these appear to be less robust to alternative specifications than are those for the probability of active participation in the labour force.

References

Angrist, J. D. and J.-S. Pischke (2009). *Mostly harmless econometrics: An empiricist's companion*. Princeton University Press.

Beaudry, P. and J. DiNardo (1991). The Effect of Implicit Contracts on the Movement of Wages Over the Business Cycle: Evidence from Micro Data. *Journal of Political Economy 99*(4), pp. 665–688.

Bell, D. N. F. and D. G. Blanchflower (2011). Youth unemployment in Europe and in the United States. *Nordic Economic Policy Review 1*, pp. 11–37.

Bhuller, M. S. (2009). Inndeling av Norge i arbeidsmarkedsregioner. Notater 2009/24 (in Norwegian), Statistics Norway, Oslo.

Gregg, P. (2001). The Impact of Youth Unemployment on Adult Unemployment in the NCDS. *Economic Journal 111*(475), pp. 626–653.

Gruber, J. and D. A. Wise (Eds.) (2010). *Social Security Programs and Retirement around the World: The Relationship to Youth Employment*. University of Chicago Press.

Hunt, J. (1999). Has Work-Sharing Worked in Germany? *The Quarterly Journal of Economics 114*(1), pp. 117–148.

Kahn, L. B. (2010). The long-term labor market consequences of graduating from college in a bad economy. *Labour Economics 17*(2), pp. 303–316.

Lazear, E. P. (1981). Agency, Earnings Profiles, Productivity, and Hours Restrictions. *The American Economic Review 71*(4), 606–620.

Martins, P. S., A. A. Novo, and P. Portugal (2009). Increasing the Legal Retirement Age: The Impact on Wages, Worker Flows and Firm Performance. IZA DP No. 4187.

Mroz, T. A. and T. H. Savage (2006). The Long-Term Effects of Youth Unemployment. *The Journal of Human Resources 41*(2), pp. 259–293.

Nordström Skans, O. (2011). Comment on Bell and Blanchflower: Youth unemployment in Europe and in the United States. *Nordic Economic Policy Review 1*, pp. 39–43.

Raaum, O. and K. Røed (2006). Do Business Cycle Conditions at the Time of Labor Market Entry Affect Future Employment Prospects? *Review of Economics and Statistics 88*(2), pp. 193–210.

Røed, K. (1997). Hysteresis in Unemployment. *Journal of Economic Surveys 11*(4), pp. 389–418.

Røed, K. and F. Haugen (2003). Early Retirement and Economic Incentives: Evidence from a Quasi-natural Experiment. *Labour 17*, 203–228.

A Appendix

A.1 Alternative specification

Let the variables $\mathit{frac61}_{jt} - \mathit{frac66}_{jt}$ be defined as the proportion of individuals in the active labour force of LMR j in year t who are of age 61-66 in year t. These workers would all be of ER eligible age in year $t+1$ had the lower age limit been 62 throughout the period[17]. When these variables are included in the first stage regression, it may be written as follows:

$$\begin{aligned} \mathit{fracpens}_{jt} = \alpha^1 + F_j^1 + \lambda_t^1 + X_{it}^{'}\beta^1 + \gamma \mathit{fracERage}_{jt} \\ + \gamma_{61}^1 \mathit{frac61}_{jt} + \ldots + \gamma_{66}^1 \mathit{frac66}_{jt} + \varepsilon_{ijt}^1 \end{aligned} \tag{A1}$$

The second stage is modified accordingly:

$$\begin{aligned} P_{ijt} = \alpha + F_j + \lambda_t + X_{it}^{'}\beta + \delta_{2SLS}\widehat{\mathit{fracpens}}_{jt} \\ + \gamma_{61} \mathit{frac61}_{jt} + \ldots + \gamma_{66} \mathit{frac66}_{jt} + \varepsilon_{ijt} \end{aligned} \tag{A2}$$

Conditional on the other covariates in (A1), variation in *fracERage* is now due to the interaction between variation in the demographic composition of different LMRs and reductions in the lower age limit for ER. That is, had the lower age limit been constant at 62 throughout the period, the sum of $\mathit{frac61}_{jt} - \mathit{frac66}_{jt}$ would equal *fracERage*, and identification of the effect of *fracERage* on *fracpens* would not be possible given the presence of $\mathit{frac61}_{jt} - \mathit{frac66}_{jt}$ in the regression. The two reductions in the lower age limit in 1997 and 1998 provide variation in the number of workers at risk of exiting through ER over and above the variation in the demographic composition of LMRs.

Estimation results from equation (A1) and (A2) for the same groups of potential entrants as those in Table 3 are given in Table A1, along with the

[17] 61 year-olds born in December are not counted, since the first possible take-up of ER benefits is the month after the lower age limit is reached. 66 year-olds born in January are also not counted, since they reach the lower age limit for the public old-age pension scheme already in January year $t+1$.

corresponding OLS point estimates. When comparing the 2SLS point estimates in the two tables we see that although they are smaller and less precisely estimated in Table A1 than in Table 3, the point estimates are significantly different only for male potential entrants.

Table A1: Estimation results for equations (A1) and (A2)

	All		**Men**		**Women**	
	OLS	IV	OLS	IV	OLS	IV
First stage						
fracERage		0.292***		0.290***		0.295***
		(0.035)		(0.035)		(0.036)
F-statistic		66.5		65.6		67.1
Second stage						
fracpens	-0.410	1.902	-0.820*	0.025	-0.068	3.668**
	(0.290)	(1.236)	(0.475)	(2.060)	(0.356)	(1.617)
N	5,393,241		2,407,898		2,985,343	
R^2	0.033		0.024		0.026	

	Level of education					
	Low		**Intermediate**		**High**	
	OLS	IV	OLS	IV	OLS	IV
First stage						
fracERage		0.275***		0.289***		0.339***
		(0.033)		(0.036)		(0.037)
F-statistic		68.9		62.0		84.1
Second stage						
fracpens	0.052	-0.650	-0.265	2.633*	-0.746	4.677
	(0.718)	(2.422)	(0.330)	(1.325)	(1.018)	(3.602)
N	695,575		4,014,499		683,167	
R^2	0.026		0.012		0.043	

Standard errors in parentheses; 2SLS standard errors are adjusted for 46 LMR clusters. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable takes the value 1 if a year t potential entrant was part of the active labour force in year $t + 1$ and is zero otherwise. *fracERage* is used as IV in the 2SLS estimations.

Additional covariates are *frac61*-*frac66* and year dummies, individual characteristics (sex, immigrant status, educational attainment, age), LMR level fixed effects, and LMR level characteristics (fraction of male, fraction of immigrants, fraction of guest workers, fraction with low/high education in the active labour force, and fraction of active labour force in different industries).

Low education corresponds to compulsory education only or missing information on educational attainment; intermediate corresponds to high school (or some high school), but no tertiary education; high corresponds to (some) university level education.

A.2 Estimation results for the main specification

Table A2: Coefficient estimates, first and second stage regressions

	First stage		**Second stage**	
fracpens	-	-	2.990^{**}	(1.162)
fracERage	0.271^{***}	(0.024)	-	-
Year dummies				
1995	0.001^{***}	(0.000)	-0.007^{***}	(0.002)
1996	0.002^{*}	(0.001)	0.031	(0.017)
1997	0.003^{**}	(0.001)	0.001	(0.019)
1998	0.003^{**}	(0.001)	-0.008	(0.020)
1999	0.005^{***}	(0.001)	-0.012	(0.020)
2000	0.003^{**}	(0.001)	0.007	(0.020)
2001	0.002^{*}	(0.001)	-0.015	(0.021)
2002	0.002	(0.001)	-0.033	(0.021)
2003	0.001	(0.001)	-0.039^{*}	(0.023)
2004	0.002	(0.002)	-0.039^{*}	(0.023)
fracmale	-0.007	(0.008)	0.033	(0.143)
fracguest	0.008	(0.009)	-0.144	(0.140)
fracimm	-0.022	(0.014)	-0.018	(0.139)
fraclow	0.002	(0.010)	-0.070	(0.157)
frachi	-0.002	(0.004)	-0.004	(0.048)
immigrant	-0.000	(0.000)	-0.046^{***}	(0.002)
low education	-0.000	(0.000)	-0.065^{***}	(0.001)
male	0.000	(0.000)	0.082^{***}	(0.008)
N	5,393,241			
R^2	0.033			

Standard errors in parentheses, clustered at the LMR level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The dependent variable in the first stage is $fracpens$; the dependent variable in the second stage takes the value 1 if a year t potential entrant was part of the active labour force in year $t+1$ and is zero otherwise. *fracERage* is used as IV in the 2SLS estimation and the estimation sample includes all potential entrants. LMR level fixed effects coefficients, coefficients for age dummies and for fractions of the active labour force in different industries are omitted from the table.

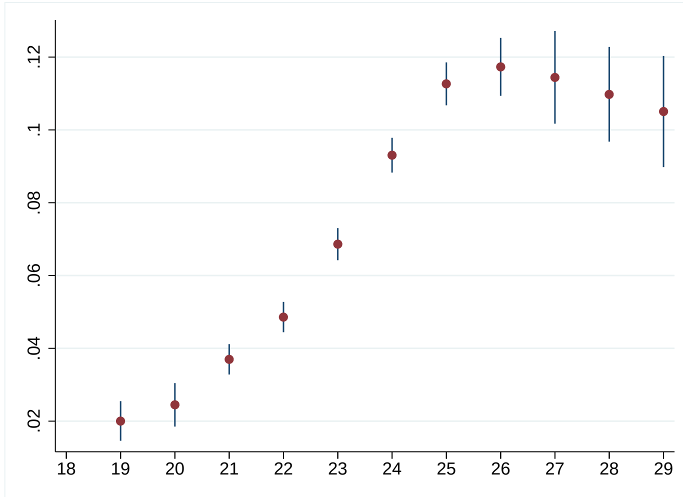

Figure A1: Coefficients for age dummies from the second stage regression for all potential entrants. Vertical spikes correspond to 95% confidence intervals, and age 18 is the reference category.

A.3 Descriptive statistics

Table A3: The incidence and duration of unemployment

	Mean	**Std. Dev**	**Min**	**Max**	**N**
Unemployed year $t+1$					
All	0.178	0.383	0	1	5,393,241
Men	0.210	0.407	0	1	2,407,898
Women	0.144	0.351	0	1	2,985,343
Months of unemployment year $t+1$					
All	0.585	1.699	0	12	5,393,241
Men	0.705	1.856	0	12	2,407,898
Women	0.488	1.554	0	12	2,985,343