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Labour supply effects of early retirement provision

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Abstract:
The main objective of this paper is to estimate labour supply effects of an early retirement programme in Norway. Detailed administrative data are employed in order to characterize full paths towards retirement and account for substitution from other exit routes, such as unemployment and disability insurance. By exploiting a reduction in the lower age limit for early retirement as a source of exogenous variation in individual eligibility I obtain robust difference-in-differences and triple differences estimates indicating that more than two out of three pensioners would still be working at the age of 63 had the age limit been 64 rather than 62. Hence, although successful in creating a more dignified exit route for early leavers, the programme also generated substantial costs in terms of inducing others to retire earlier.

Keywords: Induced retirement, Pension reform, Matched employer-employee register data, Difference-in-differences.

JEL classification: H55, I38, J26.

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Sammendrag

I denne artikkelen estimeres sysselsettingseffekter av førtidspensjonsordningen AFP, slik ordningen var fram til 2011. Hovedformålet med artikkelen er å besvare følgende spørsmål: hvor mange førtidspensjonister ville ha vært i jobb ved ulike aldre dersom AFP ikke hadde vært en mulig vei ut av arbeidsmarkedet? Administrative registerdata danner grunnlaget for den empiriske analysen. I den første delen av artikkelen sammenliknes eldre arbeidstakere i bedrifter med og uten AFP-tilknytning. Her finner jeg at omtrent 40 prosent av alle AFP-pensjonister ville ha stått i jobb til de nådde "normal" pensjonsalder (67) hvis førtidspensjon med AFP ikke hadde vært et alternativ. Med andre ord ville over halvparten av alle førtidspensjonister ha trukket seg ut av arbeidsmarkedet uansett, de fleste via uførepensjon. Videre sammenlikner jeg to grupper arbeidstakere i AFP-bedrifter som sto overfor ulik aldersgrense for AFP. Her finner jeg at så mange som to av tre AFP-pensjonister ville ha jobbet som 63-åringer dersom aldersgrensen var 64 i stedet for 62. Førtidspensjonsordningen AFP har altså hatt en sterk negativ effekt på sysselsettingen blant eldre, samtidig som den har sørget for en mer verdig avgang for arbeidstakere som ellers ville ha vært nødt til å basere seg på ordninger som uføretrygd eller ledighetstrygd.

1 Introduction

The objective of this paper is to investigate the net induced retirement effects of an early retirement programme in Norway (AFP) in order to answer the following question: How many new exits are induced by early retirement programmes? Over the last decades, the fiscal sustainability of public pension systems in most industrialised countries have been put under heavy pressure due to a combination of increased longevity and lower age at retirement, the latter probably driven to a large extent by early retirement schemes. Although the trend towards early retirement now appears to have come to an end, life expectancy is still increasing, and more so than average age at retirement. Encouraging longer working lives is therefore one of the main pillars of recent policy advice from the OECD (see e.g. OECD (2011)), and reforms aimed at increasing the age at retirement are looming in most Western countries. With this need for reforms, the need for solid knowledge about the labour supply responses to changes in early retirement schemes is intensified.

Many countries have programmes or institutions allowing different groups of workers to permanently retire prior to the normal retirement age. Public pension systems often operate with both an early retirement age and a normal retirement age, and unemployment and disability insurance programmes sometimes provide preferential treatment for elderly workers. While such programmes are effective in insuring workers against productivity shocks occurring at late stages of the career, the fact that shocks to individual productivity are unverifiable to the insurer might create moral hazard problems efficiency losses. The success of early retirement programmes therefore hinges on the magnitude of induced retirement effects, or the extent to which workers are induced to retire early even if they are not hit by negative productivity shocks.

When evaluating the effects of programmes connected to the labour market, the main challenge lies in the non-observabililty of counterfactual outcomes, or the simple fact that we do not know how workers would have behaved had the programme not been in operation. In this particular setting, the aim is to estimate *net* induced retirement effects of an early retirement programme, taking account of any substitution from other informal exit routes, such as disability or unemployment, that are known to play roles similar to that of early retirement programmes in many countries (see e.g. Gruber and Wise (1999)). To obtain estimates with causal interpretations I will be exploiting two potentially exogenous sources of variation in individual eligibility for early retirement in order to define treatment and comparison group workers, and take the observed outcomes of comparison group workers as an approximation to the unobserved counterfactual outcomes of treatment group workers.

The early retirement programme AFP (hereafter simply referred to as ER) was phased in during the nineties in terms of step-wise reductions of the lower age limit, from 66 to 62. For a worker to be qualified for retirement prior to the normal retirement age of 67 she must meet a set of individual eligibility criteria and be employed by an ER affiliated firm. The fact that not all firms are affiliated with the ER programme represents a first source of potentially exogenous variation in retirement opportunities: Using samples of workers facing the same lower age limit for ER I obtain average treatment effects of ER affiliation on labour market outcomes at different ages. These will have a causal interpretation provided that firm level affiliation is as good as random from each worker's point of view, conditional on a set of observable characteristics. The results from this part indicate that close to 40% of all ER pensioners would still be working at the age of 66.5 had early retirement not been an option; a rather substantial labour supply response, given the age for which the treatment effect is estimated. On the other hand, a majority of ER pensioners would have left the labour market at this age regardless of the ER programme, typically relying on disability insurance benefits. However, there are indications that the identifying assumption for this first strategy might not be satisfied, despite the rich set of firm and worker characteristics contained in the data.

A second set of identification strategies exploits a difference in the lower age limit between two different birth cohorts as a source of exogenous variation in individual eligibility for ER. One approach restricts attention to workers in affiliated firms and relies on observed labour market outcomes *at each age* being influenced by year of birth only through the lower age limit. Another approach is one in which the identifying assumption instead is related to *changes over time* in labour market outcomes, and where the observed outcomes of workers in non-affiliated firms are being used to take account for possible contamination due to different labour market conditions for the two cohorts. This gives robust triple difference average treatment effects indicating that as many as two out of three ER pensioners would still be working at the age of 63 had the lower age limit been 64 rather than 62. At this age there is only relatively modest substitution from other exit routes.

The most important general insights from this papers may be summarised in two points. First, the sizable labour supply effects of the ER programme clearly suggest that economic incentives have a central role in determining the timing of retirement. On the other hand, the non-negligible benefit substitution effects indicate that mechanisms other than economic incentives are also playing important roles in pushing or pulling elderly workers out of the labour market. The fact that the lion's share of the substitution is from disability pension benefits suggests that the most important shocks against which elderly workers need insurance are health related.

The paper proceeds as follows: Section 2 reviews some related literature before a brief history of the ER programme and its main features is given in Section 3. Section 4 describes the data and the constructed samples of elderly workers. Section 5 spells out the main econometric framework and employs this to investigate the induced retirement effects of the early retirement programme and substitution effects from other exit routes and into early retirement, for two cohorts of workers faced with different lower age limits. In Section 6 the focus is on estimating the effects of reducing the lower age limit from 64 to 62 on labour market outcomes at age 63. Section 7 concludes.

2 Related literature

The evidence on labour supply effects of ER provision in the existing literature is rather mixed. Baker and Benjamin (1999) find that ER reforms in Canada led to a marked increase in pension receipt, but had little effect on labour supply. The findings suggest that the new recipients were loosely attached to the labour market and would therefore not have worked even without having access to ER pension benefits. On the other hand, both Staubli and Zweimüller (2012) and Manoli and Weber (2012) document positive and non-negligible labour supply effects of increased early retirement age in Austria. Staubli and Zweimüller report that between 30 and 40 percent of all workers who retire later do prolong their employment, but there is also substantial substitution to unemployment insurance benefits. Neither of the two studies find evidence of significant substitution to disability pensions.

The early retirement programme AFP has existed for more than three decades and its effects have been assessed in several studies. Two studies in which the quasi-natural experiment property of the programme plays an important role are Røed and Haugen (2003) and Bratberg, Holmås, and Thøgersen (2004). Røed and Haugen find that two out of three pensioners would have stayed employed had early retirement not been an option, and that the programme does not substitute for disability pensions or long-term unemployment. In contrast, Bratberg et al. find that at least half the ER pensioners would have stayed in the labour force without the scheme, but also that substitution from other exit routes is quite substantial. This latter finding stands in clear contrast to the absence of such substitution effects reported by Røed and Haugen. The two analyses are based on the same types of data covering overlapping time periods, but a major difference is that Røed and Haugen include both public and private sector workers in their sample, whereas Bratberg et al. restrict attention to the private sector. The fact that workers in the two sectors are widely different in many relevant aspects, including personal characteristics, retirement options, and observed retirement behaviour, may well explain large parts of the discrepancies in terms of estimated substitution effects.[1]

Other related studies include Karlström, Palme, and Svensson (2008), Mastrobuoni (2009), Kyyrä (2010) and Inderbitzin, Staubli, and Zweimüller (2012). Karlström et al. employ a difference-in-differences strategy to study the effects of an abolition of special eligibility rules for a Swedish disability insurance programme affecting workers in the age bracket 60-64. Prior to the reform in 1997 the Swedish programme was similar to AFP in several aspects, with medical requirements being part of the eligibility rules as a notable exception. The empirical results give no support for the intended increases in employment, but there are indications of increased utilisation of unemployment and sickness insurance benefits. Kyyrä (2010) focuses mostly on disability pensions, but effects of an early retirement programme are identified by exploiting different lower age limits for private and public sector workers. If the observed differences between workers in the two sectors are indicative of systematic differences also in terms of unobservable characteristics, any causal interpretation of results based on such an identification strategy might be questionable.

Mastrobuoni (2009) studies the effects of benefit cuts driven by increased normal retirement age on retirement behaviour in the US, and finds that the mean retirement age of the affected cohorts increased by half as much as the increase in retirement age. Finally, Inderbitzin et al. (2012) study how extended unemployment insurance benefits for older workers in Austria affect the incidence of ER and other labour market exit routes. They find that the reform led to a dramatic increase in the incidence of ER, and also that the relatively young responded by sequential take-up of unemployment and disability insurance benefits, while older workers had a tendency to substitute from disability to unemployment insurance benefits.

This contribution adds to the existing literature in the following ways: First, a clear distinction between different exit routes is possible due to detailed register data covering the full population of Norway. Second, the combination of these data and an exogenous policy change enables identification of policy relevant causal effects of early retirement programmes that are not readily available in other settings. The paper extends the literature on this particular ER programme by covering a relatively long time period (1993–2003) and characterising the full paths towards retirement for several groups of elderly workers; public sector workers and private sector workers in affiliated and non-affiliated

[1] Another contribution is Hernæs, Sollie, and Strøm (2000); they find that although financial incentives do influence retirement behaviour, a rather generous bonus is needed in order to make eligible individuals postpone retirement by one year.

firms. Importantly, the analyses are performed separately for public and private sector workers.

3 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. Another argument was that a new early retirement scheme was needed to make room for younger people in the labour market.[2]

The ER scheme is fairly generous[3], 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 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.

While the public pension scheme provides universal coverage, the ER scheme covers the public sector and about half the employees in the private sector, namely those employed in firms taking part in the central tariff agreements. In addition to firm affiliation there are also some rather weak requirements related to individual work histories[4].

[2] 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. The argument has also found little empirical support, see e.g. Gruber and Wise (2010).

[3] 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.

[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

4 Data, sample and descriptive statistics

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 and disability pensions, sick leave benefits and unemployment benefits, and includes individual earnings data (in terms of pension points) dating back to 1967.

The main sample consists of two cohorts, or sub-samples, faced with different age limits: (i. Age limit 64) individuals born between January 1 and May 31 1933 (turned 60 in 1993 (base year)), and (ii. Age limit 62) individuals born between January 1 and May 31 1937 (turned 60 in 1997 (base year)). Hence, individuals in cohort (i) turned 64 just before the age limit was lowered to 63, while individuals in cohort (ii) turned 62 shortly after the age limit was set to 62. Figure 1 illustrates the sampling; how it relates to the reductions in the lower age limit for ER, and how different age limits have applied to different groups of workers over time.

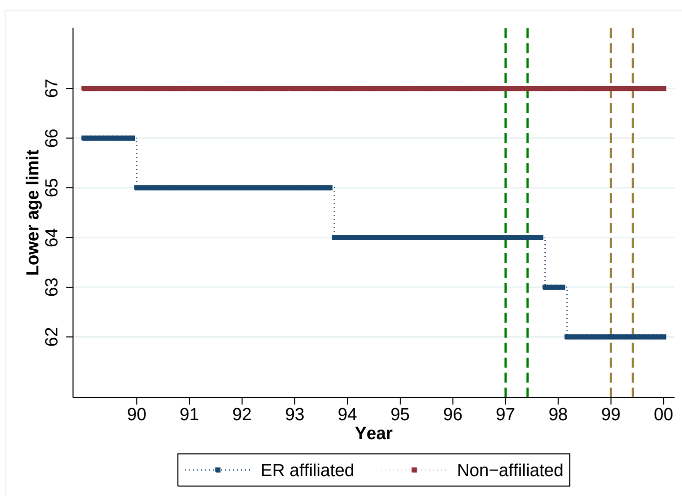

Figure 1: Lower age limits for the for workers in affiliated and non-affiliated firms. Green (brown) vertical lines indicate when workers in the "old" ("young") cohort turned 64 (62).

Private and public sector workers who met the following two ER eligibility

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.

criteria the year they turned 60 (at the latest) are included in the sample: (a) At least 10 years of work experience (i.e. earnings exceeding 1BA[5]) since the age of 50, and (b) the average of the 10 highest yearly incomes since 1967 above 2BA. We also require tenure in the base year firm at least from the age of 59, so that all individuals would have the opportunity to take out ER pensions the month after they turned 62 had the lower age limit been 62, which it was for one of the two sub-samples.[6] The final requirement is that individuals lived through their 67^{th} birthday. Based on information from the month after they turned 61, 61.5, 62, 62.5, ..., 66, 66.5, all individuals in the sample are classified into one of seven labour force states: Work, part-time work and ER benefits[7], ER, disability, long term sick leave, unemployment, and 'other'. Details regarding the classification of individuals into labour force states are provided in the Appendix.

Descriptive statistics for each of the two sub-samples are given in Table 1 and 2, along with observed labour market outcomes at the age of 66.5. Workers are divided into three groups according to the ER affiliation of the firm for which they were working at the age of 60; ER-affiliated private sector firms, non-affiliated private sector firms, and public sector firms.[8] The fraction of males, the fraction of full-time workers and the average number of days on sick leave (during 1992/1996) is higher among workers in ER affiliated firms than among the non-affiliated, and lowest in the public sector, while the incidence of unemployment (in 1992/1996) is clearly lowest for workers in the public sector, followed by those in affiliated private firms. There is also a marked difference in the distribution across industries: The fraction of workers in Manufacturing is higher and the fraction of workers in the Retail, Finance and Service industries lower among affiliated firms than among non-affiliated firms, while public sector firms are concentrated in the Service and Transport industries. Affiliated firms and public sector firms are also considerably bigger in terms of number of employees than the non-affiliated private sector firms.

As for the observed outcomes at age 66.5, relatively few of the workers in ER affiliated firms were still working, but we also note that the fractions of workers on disability and on sick leave are lower for affiliated than for non-affiliated firms. About eight percent of the workers in non-affiliated firms are classified into the states 'ER' and 'part-time work and ER benefits'. This may be explained either by (i) people changing jobs after the age of 60, as only three years of tenure are

[5]See footnote 4.

[6]The conditioning on individual eligibility criteria is imposed in order to avoid selection biases arising from the fact that workers staisfying these criteria are likely to be more closely attached to the labour market than those who do not. Given the requirement of employment at the age of 60, however, only 6% of the workers fail to satisfy one or more of the criteria.

[7]Combinations of part-time work and receipt of ER pension benefits (phased retirement) were allowed from October 1 1997.

[8]Details regarding the procedure used to identify each firm's affiliation with the ER scheme are provided in the Appendix.

required for ER benefits take-up, or (ii) mis-classifications; “wrong” main job at the age of 60 or firms being classified as non-affiliated while in reality they were affiliated.[9] We will return to this issue in the discussion of estimation results below.

The two tables also reveal some notable differences between the two birth cohorts: Unemployment is more common for the older than for the younger cohort, both when measured during the year prior to the base year and when measured at the age of 66.5. Workers in the younger cohort are more educated, they are more likely to be working in the Retail industry and less likely to be working in Manufacturing, and they are more likely to be living in Eastern Norway than are workers in the older cohort. Finally, the relative frequencies of ER pensioners is markedly higher for the younger cohort, at the same time as combining work and ER benefits appears to have become more common over time (this comes as no big surprise, however, since such combinations were only possible as of October 1997).

Figure 2 and 3 show how relative frequencies of individuals in the different labour force states change as individuals grow older, for the sample with age limit 64 and 62, respectively. More than 80 percent are still working at the age of 61, but the fraction still working declines steadily for all groups. For workers in ER-affiliated private and public sector firms, the fraction still working drops when they reach the lower age limit, at the same time as there is a marked increase in the fractions on ER pensions. This phenomenon is observed in many countries and commonly referred to as “spikes” in benefit receipt and retirement propensities at the age at which benefit eligibility begins[10]. Otherwise the trends appear to be fairly similar across the different groups of workers.

5 The effects of having access to an early retirement programme

5.1 Econometric model

The point of departure for the econometric analyses is an investigation of the induced retirement effects of ER for two cohorts of workers faced with different lower age limits. Each worker’s treatment status is in this section determined by her employer’s affiliation with the ER programme. Let K_1 and K_0 denote

[9]Labour market mobility for elderly workers is generally found to be rather low, meaning that the first possible explanation is not likely to be the most important of the two. For a sample of workers who became eligible in 1997, Røed and Haugen (2003) found job changes 3-4 years prior to ER eligibility to be extremely rare.

[10]See inter alia Rust and Phelan (1997), Baker and Benjamin (1999), Gruber and Wise (2004), Behaghel and Blau (2012) and Manoli and Weber (2012).

Table 1: Descriptive statistics, sample with age limit 64

	Private sector				**Public**	
	ER-affiliated		**Non-affiliated**			
Pension points 1993	4.4	(1.48)	4.0	(1.78)	3.9	(1.50)
Experience[a]	25.5	(3.25)	24.4	(4.16)	23.9	(4.42)
Days on sick leave[b]	12.0	(34.3)	10.8	(32.9)	9.0	(27.4)
Unemployed 1992[c]	4.3	(75)	4.0	(60)	0.4	(11)
Male	73.0	(1263)	58.9	(876)	44.7	(1371)
Married	81.2	(1407)	80.5	(1198)	78.7	(2416)
Hours/week ≥ 30	87.1	(1508)	75.4	(1122)	69.2	(2124)
Education						
Low	68.6	(1188)	63.4	(943)	53.4	(1638)
Intermediate	26.3	(455)	28.5	(424)	26.0	(798)
High	5.1	(89)	8.1	(121)	20.6	(632)
Firm size (no. of employees)						
< 10	4.4	(77)	52.8	(786)	1.0	(30)
10-49	16.6	(288)	27.9	(415)	4.4	(35)
50-99	10.5	(181)	7.5	(111)	5.7	(175)
100-499	31.0	(537)	7.0	(104)	30.9	(949)
≥ 500	37.5	(649)	4.8	(72)	58.0	(1779)
Industry						
(1) Agriculture/fisheries	0.6	(11)	3.5	(52)	0.1	(3)
(2) Petrol	3.3	(58)	0.5	(7)	0.0	(0)
(3) Manufacturing	57.0	(987)	16.1	(240)	1.0	(29)
(4) Elecricity/gas/water	0.2	(4)	1.3	(19)	2.9	(90)
(5) Construction	5.2	(91)	6.0	(89)	4.1	(27)
(6) Retail	15.9	(276)	35.2	(523)	1.0	(29)
(7) Transport	6.0	(104)	8.6	(128)	11.3	(347)
(8) Finance	7.6	(131)	13.8	(205)	0.8	(25)
(9) Service	4.0	(70)	15.1	(25)	78.8	(2418)
Regions						
(1) East	37.4	(647)	40.3	(600)	35.8	(1099)
(2) Center	12.6	(219)	12.4	(184)	15.2	(465)
(3) South/West	44.2	(766)	40.9	(608)	38.8	(1191)
(4) North	5.8	(100)	6.4	(96)	10.2	(313)
Outcomes at age $66\frac{1}{2}$						
Work	11.3	(196)	34.7	(517)	16.7	(512)
ER	46.0	(797)	8.4	(125)	33.7	(1034)
Work and ER	3.4	(59)	0.8	(12)	4.5	(137)
Disability	20.0	(347)	28.4	(422)	24.0	(737)
Sick leave	2.3	(40)	4.6	(68)	1.9	(58)
Unemployment	8.1	(140)	8.1	(120)	0.4	(11)
Other	8.8	(153)	15.0	(224)	18.9	(579)
N	1732		1488		3068	

The table reports means (percentages) for continuous (discrete) variables. Standard deviations (frequencies) in parentheses.

[a] Number of years with income $> 1BA$ (after 1967).

[b] Number of days during 1992.

[c] $= 1$ if unemployed at some point during 1992.

Table 2: Descriptive statistics, sample with age limit 62

	Private sector				**Public**	
	ER-affiliated		**Non-affiliated**			
Pension points 1997	4.6	(1.40)	4.3	(1.76)	4.0	(1.46)
Experience[a]	27.9	(4.16)	26.7	(5.05)	25.9	(5.23)
Days on sick leave[b]	11.9	(32.6)	10.8	(32.9)	10.0	(30.8)
Unemployed 1996[c]	2.7	(60)	2.1	(28)	0.5	(17)
Male	70.2	(1563)	60.5	(801)	41.4	(1317)
Married	78.6	(1748)	80.1	(1060)	76.1	(2421)
Hours/week ≥ 30	88.4	(1967)	76.3	(1009)	72.1	(2294)
Education						
Low	63.2	(1407)	59.4	(786)	49.0	(1561)
Intermediate	30.4	(676)	30.8	(407)	23.8	(758)
High	6.4	(142)	9.8	(130)	27.1	(864)
Firm size (no. of employees)						
< 10	5.3	(117)	59.2	(783)	0.8	(27)
10-49	17.1	(381)	26.3	(348)	4.0	(127)
50-99	11.7	(260)	4.4	(58)	4.6	(147)
100-499	30.7	(683)	6.5	(86)	28.9	(919)
≥ 500	35.2	(784)	3.6	(48)	61.7	(1963)
Industry						
(1) Agriculture/fisheries	1.1	(25)	2.0	(27)	0.4	(14)
(2) Petrol	3.6	(81)	0.8	(10)	0.0	(0)
(3) Manufacturing	48.6	(1082)	11.3	(150)	0.7	(21)
(4) Elecricity/gas/water	0.5	(12)	1.0	(13)	2.5	(78)
(5) Construction	5.4	(121)	6.1	(81)	3.3	(106)
(6) Retail	18.8	(418)	37.5	(496)	1.2	(39)
(7) Transport	7.1	(158)	7.9	(104)	6.6	(211)
(8) Finance	7.4	(165)	7.1	(94)	0.5	(16)
(9) Service	7.3	(163)	26.3	(348)	84.8	(2698)
Regions						
(1) East	49.5	(1102)	48.3	(639)	43.4	(381)
(2) Center	9.7	(215)	10.4	(137)	13.4	(425)
(3) South/West	35.3	(785)	36.0	(476)	32.8	(1045)
(4) North	5.5	(123)	5.4	(71)	10.4	(332)
Outcomes at age $66\frac{1}{2}$						
Work	8.9	(199)	33.6	(444)	14.4	(459)
ER	60.7	(1351)	6.6	(88)	42.7	(1358)
Work and ER	5.2	(115)	1.6	(21)	10.2	(326)
Disability	16.5	(366)	31.0	(410)	22.1	(704)
Sick leave	1.3	(29)	5.9	(78)	2.1	(68)
Unemployment	0.8	(17)	4.6	(61)	0.2	(7)
Other	6.6	(148)	16.7	(221)	8.2	(261)
N	2225		1323		3183	

The table reports means (percentages) for continuous (discrete) variables. Standard deviations (frequencies) in parentheses.

[a] Number of years with income $> 1BA$ (after 1967).

[b] Number of days during 1996.

[c] $= 1$ if unemployed at some point during 1996.

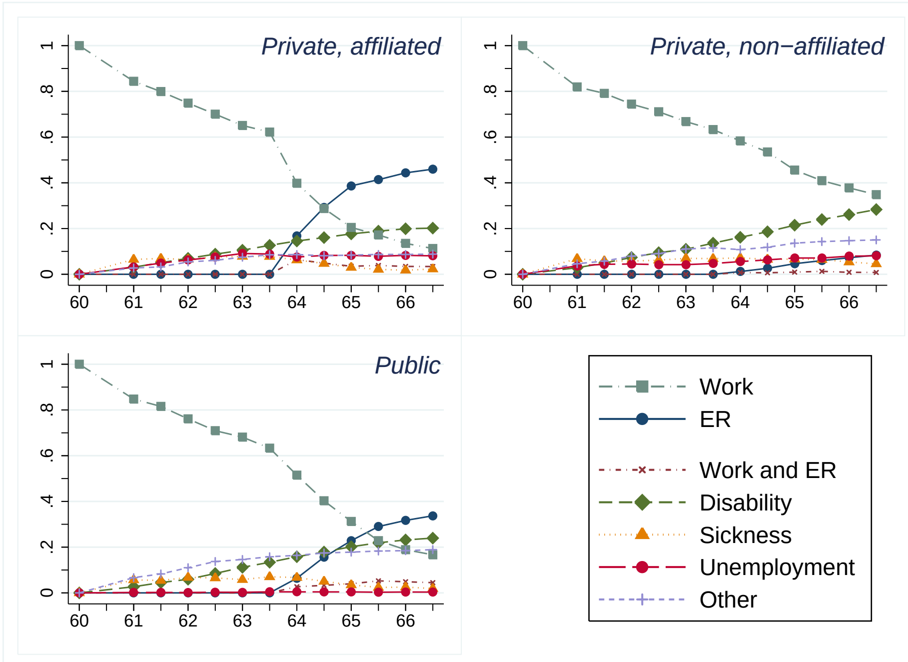

Figure 2: Observed relative frequencies at different ages, age limit 64.

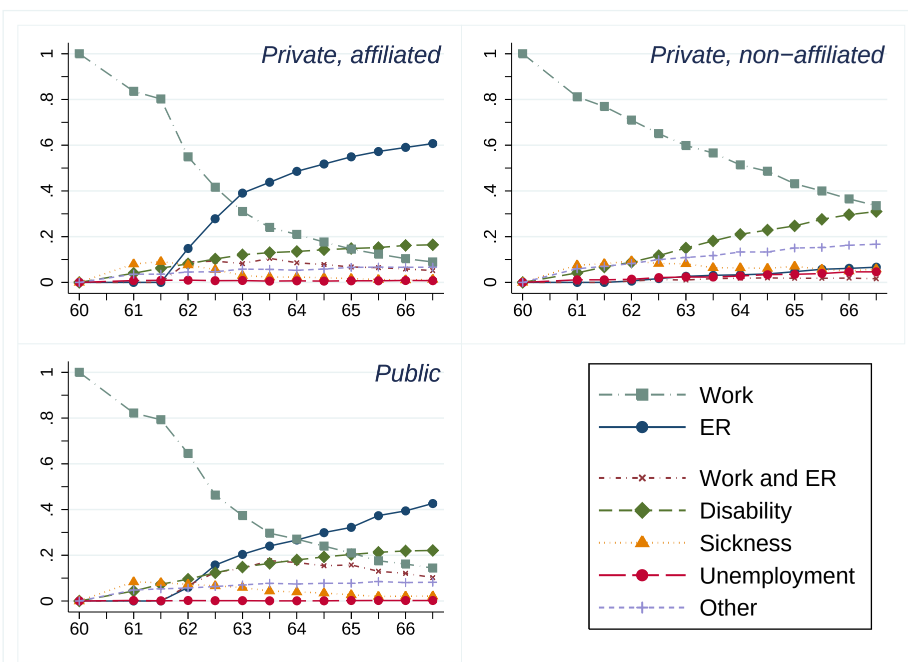

Figure 3: Observed relative frequencies at different ages, age limit 62.

the sets of potential outcomes for individuals with and without an ER entitlement: $K_1 = \{$*work, work and ER, ER, disabled, sick, unemployed, other*$\}$ for an individual with ER entitlement, and $K_0 = \{$*work, disabled, sick, unemployed, other*$\}$ for an individual without ER entitlement. Define two sets of dummies Y^k_{1it}, Y^k_{0it} such that $Y^k_{1it} = 1$ if an entitled individual i's outcome is $k \in K_1$ at age t and 0 otherwise, and Y^k_{0it} is defined accordingly for non-entitled individuals. $E(Y^k_{sit}) = \Pr\left(Y^k_{sit} = 1\right) \equiv P^k_{sit}, s = 0, 1$ is then the relative frequency of individuals in state $k \in K^s$ with eligibility status s. Let D_i be an indicator taking the value 1 if individual i is entitled to ER (works in an ER affiliated firm) and 0 otherwise, and assume that the selection into affiliated firms may be described by a utility function V:

$$V_i = \mu_v(Z_i) + u_{vi}, \qquad D_i = \mathbb{1}\{V_i > 0\}, \tag{1}$$

where Z_i and u_{vi} are observed and unobserved factors determining choices, and $\mathbb{1}\{\cdot\}$ is the indicator function. Write potential outcomes in terms of observed (X_i) and unobserved outcome-specific variables:

$$Y^k_{1it} = X_i'\beta^j + \gamma^k_t + u^k_{1it} \tag{2}$$

$$Y^k_{0it} = X_i'\beta^k + u^k_{0it} \tag{3}$$

γ^k_t is the treatment effect, that is, the effect of ER affiliation on the propensity to be in state k at age t, assumed to be constant across individuals. We normalize the unobserved variables u_{vi}, u^k_{1it} and u^k_{0it} to have zero means and follow the common practice of imposing additive separability between observed and unobserved variables. We also assume that all observed variables are exogenous. If

$$Y^k_{1it}, Y^k_{0it} \perp\!\!\!\perp D_i \quad \forall t, \tag{4}$$

that is, if potential outcomes at all ages are independent of the individual's treatment status, then the OLS estimator $\hat{\hat{\gamma}}^k_t = \bar{Y}^k_{1t} - \bar{Y}^k_{0t}$ from the regression

$$Y^k_{it} = \alpha^k + \gamma^k_t D_i + \varepsilon^k_{it} \tag{5}$$

is an unbiased estimator of the average treatment effect, ATE, i.e. the average effect of ER-affiliation on the propensity to choose outcome j at age t. We will refer to $\hat{\hat{\gamma}}^k_t$ as the *ATE under no selection*.

Assumption (4) holds whenever D is as good as randomly assigned, but is likely to be violated in our case, as workers are not distributed randomly across firms. When (4) fails to hold true one has to consider alternative identification strategies. One is known as the conditional independence assumption;

$$(Y_{1it}^k, Y_{0it}^k) \perp\!\!\!\perp D_i | W_i \quad \forall t, \qquad (CIA)$$

where $W = (X, Z)$ is the conditioning set. We will from here on assume that the sets X and Z are the same, and refer to the set X as the full set of observable covariates. Under the CIA, the OLS estimator $\hat{\gamma}_t^k$ from the regression

$$Y_{it}^k = X_i'\beta^k + \gamma_t^k D_i + \varepsilon_{it}^k \tag{6}$$

is an unbiased and consistent estimator of the ATE. $\hat{\gamma}_t^k$ will be referred to as the *ATE under selection on observables.*

In this setting, assuming conditional independence is equivalent to assuming that conditional on a set of observed covariates, ER affiliation is as good as randomly assigned from the point of view of the workers. This assumption is untestable, but Røed and Haugen (2003) point out two assumptions that are both necessary for the CIA to hold. First, one needs to assume that workers do not self-select into affiliated firms as part of an early retirement strategy. Using the same sources of data as in this paper, but different samples of workers, Røed and Haugen find no indications of this type of strategic behavior among elderly workers. Secondly, the CIA rules out any systematic unobserved differences between workers in affiliated and non-affiliated firms, that is, $Cov(u_{1t}^k, u_v) = 0$ and $Cov(u_{0t}^k, u_v) = 0 \quad \forall k, t$. Non-affiliated firms are firms that do not take part in the central tariff agreements, and are typically smaller and operate in different industries than affiliated firms (see Table 1 and 2 above). Given that the firms are different, one may also suspect that the jobs and the workers occupying the jobs are different across affiliated and non-affiliated firms, even in terms of unobservables. Røed and Haugen compare the retirement behavior of non-eligible workers in affiliated and non-affiliated firms, and find that workers in affiliated firms do have a higher voluntary retirement propensity and a higher disability propensity than those in non-affiliated firms, even when they are not eligible for ER. They claim nevertheless that these differences are fully accounted for by observed explanatory variables, and argue that having variables reflecting firm size and industry as part of the conditioning set is of fundamental importance.

5.2 Results

ATEs from OLS on equation (5) are depicted in Figure 4 for private sector workers with age limit 64 and in Figure 5 for private sector workers with age limit 62, for the outcomes 'work', 'disability', 'sickness leave', 'unemployment', and 'other'. Starting with the sample with age limit 64, there are no significant differences in the propensity to be in the state 'work' between the affiliated and the non-affiliated until the lower age limit is reached. From then on the difference is around -20 percentage points at each age. The pattern is similar for the state 'sickness leave', but with more moderate differences; they are significantly negative from age 64.5, and fairly stable at around -3 percentage points. Workers in affiliated firms were more likely than those in non-affiliated firms to receive unemployment benefits, from age 62 until age 64.5. For the states 'disability' and 'other' the differences are increasingly negative from the age at which the lower age limit is reached. Hence, judging from the pre-treatment trends in Figure 4, non-affiliated workers should provide a good approximation to the counterfactual behaviour of affiliated workers for all outcomes except 'unemployment' and 'other'. We note that individuals working in large manufacturing firms at the age of 60 make a large share of the unemployed, and large manufacturing firms are also over-represented among the ER-affiliated firms.

Turning to the sample with age limit 62, we observe that affiliated workers start off with a somewhat higher propensity to be in the state 'work' before the difference drops about the same way as for those with age limit 64 once they reach the lower age limit. The pattern is once again similar for the state 'sickness leave', with differences around -4 percentage points. For 'disability' and 'other' we observe the same falling tendency as for the sample with age limit 64, and we note that affiliated workers are significantly less likely to be in the state 'unemployment' from the age of 62.5. For this cohort, non-affiliated workers should make a good comparison group for affiliated workers for all outcomes except the outcome 'other'.

Table 3 and 4 show observed outcomes at age 63 and 66.5 for the two cohorts, and give two sets of estimated differences: ATEs under no selection and ATEs under selection on observables. The covariates in X are years of experience, days on sick leave during 1992/1996, a dummy for receipt of unemployment benefits in 1992/1996, and dummies for full time work, educational attainment, gender, industry, geographical location and firm size (1993/1997). Starting with Panel I of Table 3, the ATEs under no selection are significantly different from zero only for the outcomes 'unemployment' and 'other' (this was also clear from Figure 4), but there are, however, substantial differences between the ATEs under no selection and the ATEs under selection on observables, especially for the outcomes 'work' and 'unemployment'.

Turning to the outcomes at age 66.5 (Panel II), we first note that the magnitude of the estimated treatment effect for the outcome 'work' is reduced by more than 40 percent when the full set of controls is added, and the two confidence intervals are non-overlapping. For the outcome 'other', the treatment effect

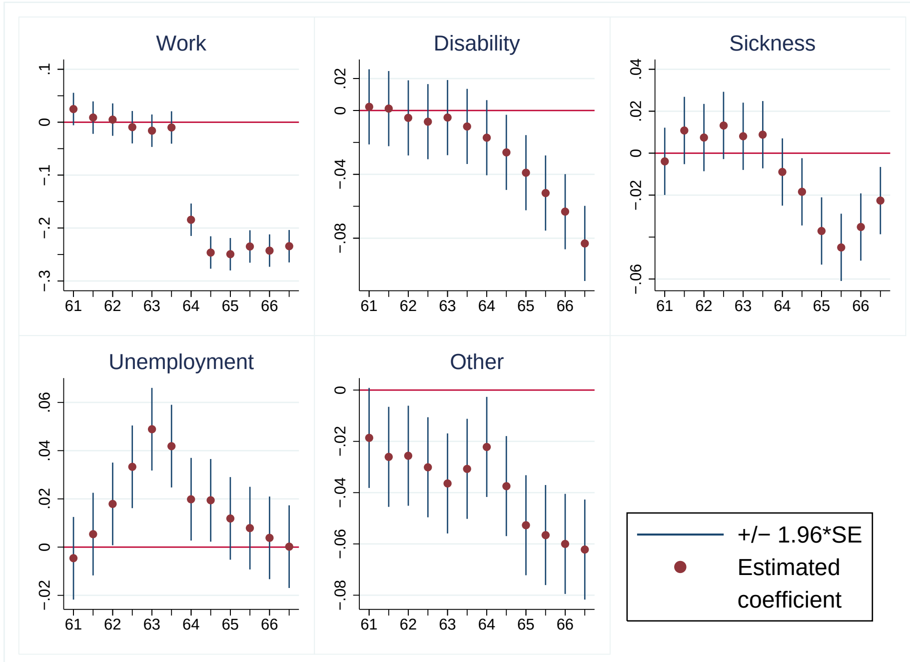

Figure 4: ATEs under no selection (from OLS on equation (5)), for private sector workers at different ages. Age limit 64.

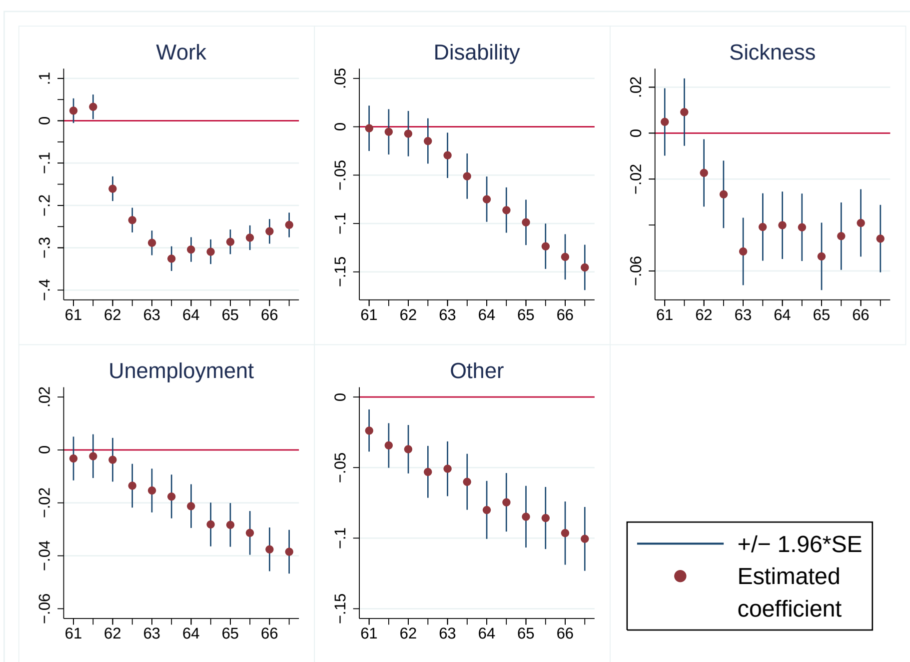

Figure 5: ATEs under no selection (from OLS on equation (5)), for private sector workers at different ages. Age limit 62.

more than doubles, and the total substitution effect[11] increases by 65 percent, from -16.8 to -27.8 percentage points. While the ATE under no selection indicates that more than half (23.4/46) the ER pensioners would still be working at the age of 66.5 if ER was not an option, one would say less than one third (13.9/46) based on the ATE under selection on observables.

The marked differences between the two sets of average treatment effects may be taken as a call for caution when it comes to making causal judgments based on comparisons between workers in affiliated and non-affiliated firms, especially since there are notable differences at age 63, one year before the cohort reaches the lower age limit for ER. As can be seen in Table 1 and 2, the two groups are clearly different in terms of observable characteristics, and hence there may also be systematic differences in terms of unobservables that are not fully accounted for by the set of control variables. The covariates that alter the average treatment effects the most are the dummies for industry and firm size. Firm size is positively related to ER affiliation (Table 1 and 2), to the probability of having an occupational pension (cf. Hernæs et al. (2011)), and possibly also to the probability of having firm provided early retirement options other than AFP. It may also be that workers in larger firms are different in terms of unobserved personal traits compared to those in small or medium sized firms.

The differences between two sets of average treatment effects are much less pronounced for the cohort with age limit 62, and all confidence intervals but two are overlapping (Table 4). The total substitution effects of Panel II are about twice those of Panel I, and this difference is largely due to an inflow of non-affiliated workers into the states 'disability' and 'other'. At the age of 66.5, the substitution from the four states 'disability', 'sick leave', 'unemployment' and 'other' now adds up to -34 percentage points, and the ATE under selection on observables indicates that about one third (19.5/60.7) of the ER pensioners would be working in the absence of the ER programme. The relative magnitudes of the effect of ER affiliation on the propensity to be working at the age of 66.5 are thus about the same for the two sub-samples, and so is the total substitution from other exit routes.

We have seen that about eight percent of the workers classified as non-affiliated at the age of 60 were receiving ER pensions at the age of 66.5. This mis-classification will cause the estimated average treatment effects to be biased downwards in magnitude (see e.g. Hausman (2001)) and is sometimes referred to as "contamination bias" (see Bratsberg, Fevang, and Røed (2010)). Assuming

[11] What is from here and onwards referred to as the 'total substitution effect' is simply the sum of the estimated average treatment effects of AFP on the outcomes Disability, Sickness leave, Unemployment and Other.

that all workers classified as affiliated were indeed affiliated and that the take-up rate of ER pension benefits is the same regardless of how the workers are classified in this paper, we compute that 18.6 percent (9.2/0.494) of the non-affiliated workers in the older sub-sample were actually eligible for ER at the age of 66.5, and so were 12 percent (8.2/0.659) of the non-affiliated workers in the younger sub-sample. We may adjust for contamination bias by dividing the estimated average treatment effects by the estimated fraction of non-affiliated workers: For the birth cohort with age limit 64 the adjusted effect of ER on the propensity to be working at the age of 66.5 is calculated as $-13.9/(1-0.186) = -17.1$. For the cohort with age limit 62 the adjusted effect is $-19.5/(1-0.124) = -22.3$. From these adjusted estimates it follows that nearly 40 percent of all ER pensioners would still be working at the age of 66.5 had early retirement not been an option.

Table 3: ATEs of ER affiliation for private sector workers with age limit 64

	Observed (%)		ATEs under no selection		ATEs under selection on observables[a]	
	ER	No ER				
I. Age 63						
Work	65.1	66.7	-1.6	[-4.9, 1.7]	7.2	[2.8, 11.6]
Disability	10.5	10.9	-0.4	[-2.6, 1.7]	-0.8	[-3.6, 1.9]
Sick leave	7.8	7.0	0.8	[-1.0, 2.6]	1.6	[-1.0, 4.1]
Unemployment	9.1	4.2	4.9	[3.1, 6.6]	-1.3	[-3.7, 1.1]
Other	7.4	11.1	-3.6	[-5.6, -1.6]	-6.7	[-9.4, -3.9]
ER	0.0	0.0				
Work and ER	0.0	0.0				
II. Age $66\frac{1}{2}$						
Work	11.3	34.7	-23.4	[-26.2, -20.7]	-13.9	[-17.7, -10.1]
Disability	20.0	28.4	-8.3	[-11.3, -5.4]	-9.2	[-13.1, -5.4]
Sick leave	2.3	4.6	-2.3	[-3.5, -1.0]	-0.9	[-2.6, 0.9]
Unemployment	8.1	8.1	0.0	[-1.9, 1.9]	-4.4	[-7.0, -1.8]
Other	8.8	15.0	-6.2	[-8.4, -4.0]	-13.3	[-16.2, -10.3]
ER	46.0	8.4	37.6	[34.8, 40.5]	38.8	[34.9, 42.8]
Work and ER	3.4	0.8	2.6	[1.6, 3.6]	2.9	[1.4, 4.3]
Observations	1732	1488				

95% confidence intervals in brackets. The "ATEs under no selection" are $\hat{\hat{\gamma}}_t^k$ from equation (5), while "ATEs under selection on observables" refers to $\hat{\gamma}_t^k$ from equation (6), for $t = \{63, 66.5\}$.

[a] Included controls are years of experience, days on sick leave during 1992, a dummy for receipt of unemployment benefits in 1992, and dummies for full time work, educational attainment, gender, industry, geographical location and firm size (1993).

Table 4: ATEs of ER affiliation for private sector workers with age limit 62

	Observed (%)		ATEs under no selection		ATEs under selection on observables[a]	
	ER	No ER				
I. Age 63						
Work	31.0	59.9	-28.8	[-32.1, -25.6]	-19.7	[-24.2, -15.2]
Disability	12.1	15.0	-2.9	[-5.2, -0.6]	-3.8	[-6.8, -0.8]
Sick leave	3.0	8.2	-5.1	[-6.6, -3.7]	-4.0	[-6.1, -1.9]
Unemployment	0.8	2.3	-1.5	[-2.3,-0.7]	-0.3	[-1.4, 0.8]
Other	5.8	10.9	-5.1	[-6.9, -3.3]	-9.9	[-12.4, -7.4]
ER	39.1	2.6	36.4	[33.7, 39.1]	28.5	[24.7, 32.3]
Work and ER	8.2	1.1	7.2	[5.6, 8.7]	9.3	[7.1, 11.4]
II. Age $66\frac{1}{2}$						
Work	8.9	33.6	-24.6	[-27.1, -22.1]	-19.5	[-23.0, -16.0]
Disability	16.4	31.0	-14.5	[-17.3, -11.8]	-13.3	[-17.0, -9.6]
Sick leave	1.3	5.9	-4.6	[-5.7, -3.4]	-4.2	[-5.8, -2.5]
Unemployment	0.8	4.6	-3.8	[-4.8,-2.9]	-2.6	[-4.0,-1.2]
Other	6.6	16.7	-10.0	[-12.1, -8.0]	-14.4	[-17.1, -11.6]
ER	60.7	6.6	54.1	[51.2, 56.9]	48.9	[44.9, 52.8]
Work and ER	5.2	1.6	3.6	[2.3, 4.9]	5.0	[3.2, 6.9]
Observations	2225	1323				

95% confidence intervals in brackets. The "ATEs under no selection" are $\hat{\tilde{\gamma}}_t^k$ from equation (5), while "ATEs under selection on observables" refers to $\hat{\gamma}_t^k$ from equation (6), for $t = \{63, 66.5\}$.

[a] Included controls are years of experience, days on sick leave during 1996, a dummy for receipt of unemployment benefits in 1996, and dummies for full time work, educational attainment, gender, industry, geographical location and firm size (1997).

6 The impacts of different lower age limits

6.1 Partial ATEs of early retirement eligibility

In this section we will see how different lower age limits affect labour market outcomes for ER affiliated elderly workers. First, we estimate equation (5) and (6) with the treatment indicator D_i replaced by δ_i, which takes the value 1 if individual i has age limit 62, on a sample consisting of affiliated workers from the two sub-samples described in Section 4. Estimations are run separately for private and public sector workers. By comparing affiliated workers only, one avoids problems related to selection into ER affiliated firms, at the cost of some external validity.[12] The treatment indicator is now a function of year of birth only, which is arguably exogenously determined from each individual's point of view. Hence, there is no need to worry about non-randomness of the treatment indicator in this particular setting. Comparing individuals of different birth cohorts makes room for another source of bias, however, namely that any estimated differences in labour market outcomes between the cohorts might be contaminated by different labour market conditions. We will return to this issue below.

Figure 6 and 7 show ATEs under no selection for the outcomes 'work', 'disability', 'sickness leave', 'unemployment' and 'other', for private and public sector workers, respectively. Starting with the outcome 'work', the difference ranges from about -20 percentage points to about -40 percentage points between age 62 and 64. From age 64.5 the difference is reduced, but the cohort with age limit 64 remains more likely to be working than the cohort with age limit 62 also after the lower age limit is reached for the older cohort (at 64 years). Private sector workers with age limit 64 are also more likely to receive unemployment benefits, and slightly more likely to receive disability pensions as of age 65 than are those with age limit 62, whereas public sector workers with age limit 64 are *less* likely to receive such pensions between age 61 and 64 than those with age limit 62. For both private and public sector workers, we note that those with age limit 62 are more likely than workers with age limit 64 to be on long-term sickness leave until they reach the lower age limit, while the difference turns negative around age 62. This may well reflect a response to the incentives provided by the ER scheme to remain in the labour market until the lower age limit is reached. The state 'other' is generally more populated for those with age limit 64.

Table 5 first gives the observed outcomes at age 63 for the two birth cohorts

[12]This statement is true only if the selection of workers into ER affiliated firms, that is, into firms taking part in the central tariff agreements, has been constant over time. According to Nergaard and Stokke (2010), there were only minor fluctuations in the share of employees working in firms covered by the central tariff agreements from the 1960s and until the late 1990s.

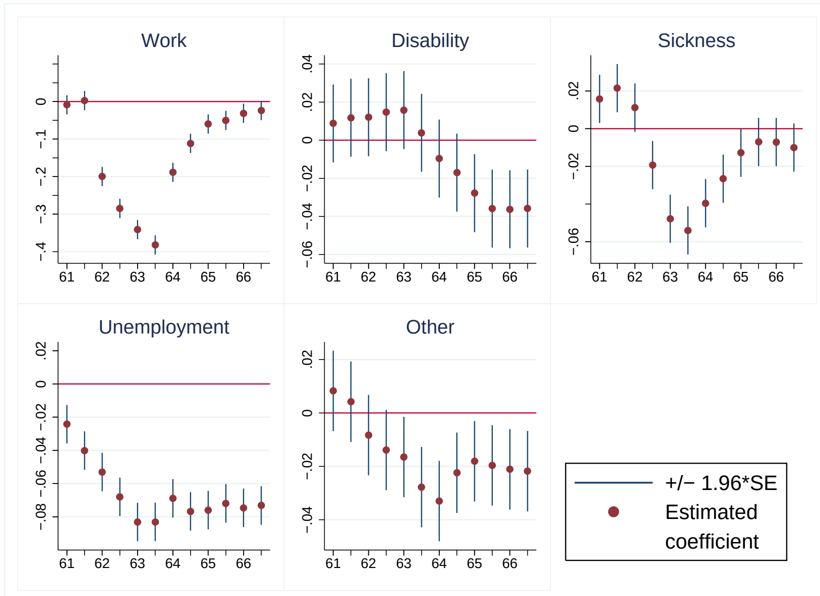

Figure 6: ATEs under no selection from OLS on equation (5) with $\delta_i = 1$ if age limit $= 62$. Private sector workers at different ages.

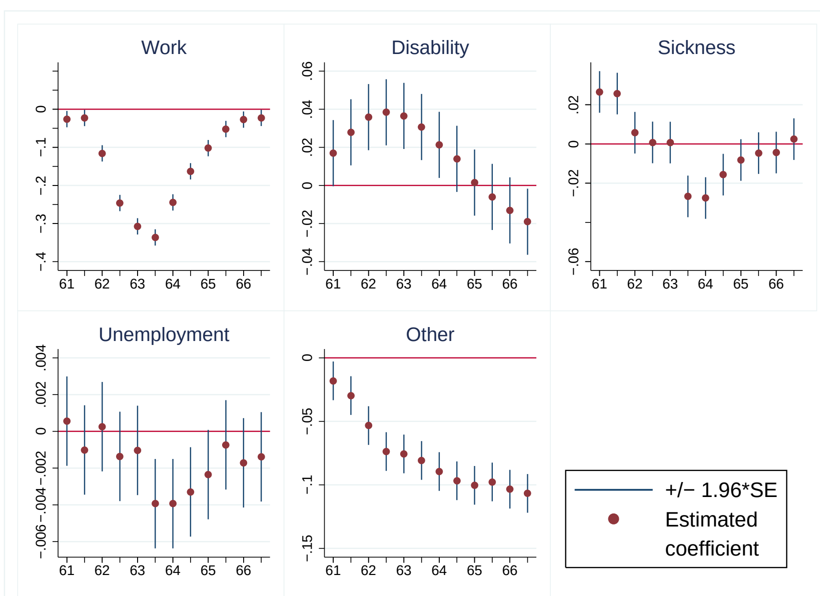

Figure 7: ATEs under no selection from OLS on equation (5) with $\delta_i = 1$ if age limit $= 62$. Public sector workers at different ages.

and then compares the ATEs under no selection to ATEs under selection on observables, where the covariates in X are the same as those in Section 5.2 (namely years of experience, days on sick leave and receipt of unemployment benefits during the year prior to the base year (1992/1996), and dummies for full time work, educational attainment, gender, industry, geographical location and firm size (base year)). There are only minor changes in the estimated ATEs of age limit 62 on the propensities to be in each of the four states when covariates are added. The total substitution from other exit routes now adds up to -12 and -4 percentage points for private and public sector workers, respectively.

Table 5: ATEs of age limit 62 on outcomes at 63

	Observed (%)		Estimated ATEs (%)			
	62	64	No selection		Selection on observables[a]	
			Private sector workers			
Work	31.0	65.1	-34.1	[-37.1,-31.2]	-34.6	[-37.7,-31.5]
Disability	12.1	10.5	1.6	[-0.4,3.6]	1.2	[-0.8,3.2]
Sick leave	3.0	7.8	-4.8	[-6.2,-3.4]	-3.8	[-5.4,-2.3]
Unemployment	0.8	9.1	-8.3	[-9.6,-7.0]	-7.6	[-9.0,-6.2]
Other	5.8	7.4	-1.6	[-3.2,-0.1]	-1.7	[-3.3,-0.1]
ER	39.1	0.0				
Work and ER	8.2	0.0				
Observations	2225	1732				
			Public sector workers			
Work	37.4	68.1	-30.7	[-33.1,-28.4]	-29.5	[-31.9,-27.1]
Disability	14.8	11.2	3.6	[2.0,5.3]	3.5	[1.8,5.1]
Sick leave	5.9	5.8	0.1	[-1.1,1.2]	0.4	[-0.8,1.6]
Unemployment	0.2	0.2	-0.1	[-0.3,0.1]	-0.2	[-0.4,0.0]
Other	7.0	14.6	-7.6	[-9.1,-6.0]	-7.7	[-9.3,-6.2]
ER	20.4	0.0				
Work and ER	14.3	0.0				
Observations	3183	3068				

95% confidence intervals in brackets. The first two columns give observed relative frequencies for the cohorts with age limit 62 and 64, respectively. The unweighted differences are $\hat{\gamma}_t^k$ from equation (5), while "OLS w/controls" refers to $\hat{\gamma}_t^k$ from equation (6), for $t = 63$, and with the treatment indicator $\delta_i = 1$ for individuals with age limit 62.

[a] Included controls are years of experience, days on sick leave during 1992, a dummy for receipt of unemployment benefits in 1992, and dummies for full time work, educational attainment, gender, geographical location, industry and firm size in 1993 (the two latter for private sector workers only).

To investigate whether important heterogeneities are masked by the average treatment effects, I have estimated versions of equation (6) where the treatment indicator is interacted with each of the variables Pension points 1993/1997, Male, Married, Hours per week ≥ 30, Educational attainment (three dummies), and Firm size (five dummies). For private sector workers, this method reveals no significant differences for the outcomes 'work' and 'other'. The positive treatment effect for the outcome 'disability' appears to be driven by women, as the treatment effect for men is negative and non-significant. The negative effect on the outcome 'sick leave' is driven by individuals with low levels of education and by workers in the smallest firms, whereas the propensity to be unemployed decreases most for those with low levels of education and for workers in large firms.

For public sector workers the interaction terms reveal no significant differences for the outcomes 'work' and 'unemployment'. The positive treatment effect for the outcome 'disability' appears to be driven by those who earned relatively few pension points the year they turned 60 (1993/1997). A lower age limit increases the propensity to be on sickness leave for the non-married, but not for the married, and decreases the same propensity slightly for those with high levels of education, whereas the treatment effect for the outcome 'other' is mostly driven by men. (Results not reported.)

6.2 Difference-in-differences estimation

It may be that the differences in Figure 6 and 7 and Table 5 are not only due to the different age limits; they might be biased due to the fact that the cohorts are observed at different times and under different labour market conditions. The fact that private sector workers with age limit 62 are less likely to be unemployed than are those with age limit 64 even before they turn 62 gives a particular reason to suspect such biases. One way of coping with these is to rely on difference-in-differences (DD) or triple differences (DDD) strategies. Instead of requiring that the observed outcomes *at each age* are influenced by the year of birth only through the lower age limit for early retirement, the identifying assumption for DD and DDD estimation relates to differences in outcomes *over time.*

Table 6 illustrates DD and DDD estimation of the impact of a lower age limit at 62 relative to one at 64 on the propensity to be classified as still working at the age of 63.[13] The top panel compares the change in work propensities for affiliated private sector workers with age limit 62 to the change for those with

[13] Both the notation and the framing of the DD and DDD strategies of this section are borrowed from Gruber (1994).

age limit 64. There was a 52.5 percentage point fall in the propensity to work between age 61 and 63 for workers with age limit 62, compared to only 19.3 percentage points for those with age limit 64, which constitutes a 33 percentage points relative fall in the work propensity between age 61 and 63 among ER affiliated workers. This is the difference-in-differences average treatment effect estimate of a two-years reduction in the lower age limit, and its magnitude is very close to the ATEs reported in Table 5.

The possibility of bias due to different labour market conditions is examined in panel B, where the same exercise as the one in panel A is performed for a control group consisting of non-affiliated private sector workers. It turns out that individuals belonging to the cohort with age limit 64 (the older cohort) are more likely to be working at both ages than those with age limit 62 (the younger cohort), even among the non-affiliated: The relative difference in work propensities is a significant -6.1 percentage points.

The triple difference estimate (the difference between panel A and B) is at -27.1 percentage points, that is, there was a 27 percentage points fall in the relative work propensities of affiliated workers with age limit 62, compared to the change in relative propensities among those with age limit 64, between age 61 and 63. This estimate indicates that more than two out of three (-27.1/39.1) ER pensioners with age limit 62 would still have been working at the age of 63 if the lower age limit was instead at 64, which corresponds to roughly twice the magnitude of the effects of increased early retirement age in Austria, as documented by Staubli and Zweimüller (2012).

One may want to investigate whether the DDD estimate in Table 6 is sensitive to conditioning on other observables that affect the propensity to continue working, and a way of doing so is to include a set of covariates in a regression equation like the following:

$$\begin{aligned} Y_{ijt} = \alpha &+ \beta_1 X_{ijt} + \beta_2 \tau_t + \beta_3 \delta_j + \beta_4 \mathrm{D}_i \\ &+ \beta_5 (\delta_j \times \tau_t) + \beta_6 (\mathrm{D}_i \times \tau_t) + \beta_7 (\mathrm{D}_i \times \delta_j) \\ &+ \beta_8 (\mathrm{D}_i \times \delta_j \times \tau_t) , \end{aligned} \tag{7}$$

where i indexes individuals, j indexes age limits (1 if age limit 62, 0 if 64), and t indexes time/age (1 if age 63 ("after"), 0 if 61 ("before")). Y_{ijt} equals 1 if individual i is observed working at age t, and zero otherwise, and X is a vector of observable characteristics, δ_j is a fixed cohort (or age limit) effect, τ_t is a fixed age effect, and D_i takes the value 1 if individual i is affiliated with the ER scheme, and zero otherwise. The fixed effects control for the effects of age (β_2), time-invariant differences between the two cohorts (β_3), and time-invariant characteristics common to affiliated workers (β_4). The second-level

Table 6: Work propensities and DD and DDD estimates of reduced lower age limit

	Before (Age 61)		**After** (Age 63)	**After-Before**
A. Treatment individuals: Affiliated private sector workers				
Age limit 62	83.5		31.0	-52.5
	[2,225]		[2,225]	(1.3)
Age limit 64	84.4		65.1	-19.3
	[1,732]		[1,732]	(1.4)
Diff. between cohorts at same age:	-0.9		-34.1	
	(1.2)		(1.5)	
Difference-in-differences:		-33.2		
		(1.9)		
B. Control group: Non-affiliated private sector workers				
Age limit 62	81.2		59.9	-21.3
	[1,323]		[1,323]	(1.7)
Age limit 64	81.9		66.7	-15.2
	[1,488]		[1,488]	(1.6)
Diff. between cohorts at same age:	-0.7		-6.9	
	(1.5)		(1.8)	
Difference-in-differences:		-6.1		
		(2.3)		
Triple difference:		-27.1		
		(3.0)		

Standard errors in parentheses, number of individuals in square brackets. Observed relative frequencies and point estimates in percentages.

interactions control for changes over time for the cohort with age limit 62 (β_5), changes over time for affiliated workers (β_6), and time-invariant characteristics common to affiliated workers with age limit 62 (β_7). The third level interaction (β_8) captures all variation in the work propensity specific to affiliated workers (relative to non-affiliated) with age limit 62 (relative to those with age limit 64) at age 63 (relative to age 61). This is the DDD estimator of age limit 62 on the propensity to be working at age 63. Neither the DD nor the DDD estimates were altered in any significant way when the same individual characteristics as those in Table 5 were included in X.

Although the estimates in Table 6 are robust with respect to the inclusion of controls, they will for the same reasons as in Section 5.2 be subject to contamination bias. Following the same logic as above we compute that $3.7/0.473 = 7.8$ percent of the workers with age limit 62 classified as non-affiliated were actually eligible at the age of 63 (numbers taken from Table 4). Dividing the estimated propensity to work at the age of 63 for non-affiliated workers with age limit 62 by the estimated fraction of non-affiliated workers, we arrive at an adjusted estimate of $59.9/(1 - 0.078) = 65.0$. The DD estimate in panel B is thus reduced to -1.0 and the DDD estimate increases in magnitude to -32.2, which would imply that more than four out of five (32.2/39.1) of the ER pensioners would still have been working at the age of 63 had the age limit been 64 rather than 62.

Both unadjusted and adjusted DD and DDD estimates of age limit 62 on the propensity to be in each of the five outcome states are given in Table 7. We note that the DDD estimates are generally very close to the treatment effects reported in Table 5.

Figure 8 shows nation-wide unemployment rates along with lower age limits for ER affiliated workers, and how unemployment was higher in 1994 (5.2 percent) and 1996 (4.2 percent), when the outcomes of the older cohort are observed, than in 1998 (2.4 percent) and 2000 (2.7 percent), when the outcomes of the younger cohort are observed. In Table 6 and 7 we saw how individuals in the older cohort are more likely to be working than those in the younger cohort, also among the non-affiliated. This may be taken as an indication that norms regarding what is an acceptable or normal retirement age have changed as the ER programme has become more settled. The link between benefit eligibility ages and social norms for retirement is also put forward by Gruber and Wise (2004), along with limited private saving and liquidity constraints, as a possible explanation for the widely observed spikes in retirement rates at benefit eligibility ages.

Table 7: DD and DDD estimates of age limit 62 on outcomes at age 63

	DD				**DDD**	
	ER		No ER			
A. Unadjusted effects (%)						
Work	-33.2	(1.9)	-6.1	(2.3)	-27.1	(3.0)
Disability	0.7	(1.2)	2.8	(1.4)	-2.1	(1.8)
Sick leave	-6.4	(1.1)	0.5	(1.3)	-6.8	(1.7)
Unemployment	-5.9	(0.8)	0.7	(0.9)	-6.6	(1.2)
Other	-2.5	(1.1)	-1.5	(1.3)	-0.9	(1.7)
B. Adjusted effects (%)						
Work	-33.2		-1.0		-32.2	
Disability	0.7		4.1		-3.4	
Sick leave	-6.4		1.2		-7.6	
Unemployment	-5.9		0.9		-6.8	
Other	-2.5		-0.6		-1.9	

Standard errors in parentheses. The adjusted effects in panel B are point estimates adjusted for contamination bias as described in the text.

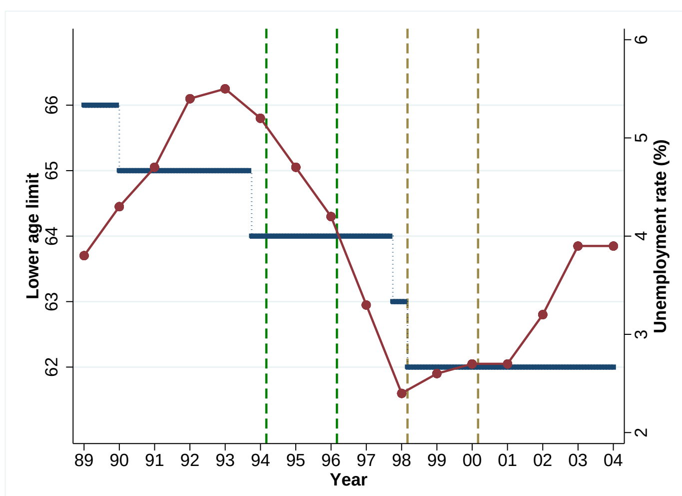

Figure 8: Register based unemployment rates from the Norwegian Labour and Welfare Administration (NAV) along with lower age limits for ER affiliated workers. Green (brown) lines indicate *before* and *after* for the "old" ("young") cohort.

7 Conclusion

The purpose of this paper has been to estimate causal effects of the early retirement programme AFP on labour market outcomes among elderly workers, including induced retirement effects and substitution from other exit routes. We have seen that differences between public and private sector workers, in terms of observed characteristics and retirement behaviour, makes a good case for making separate analyses for the two groups, which has typically not been done in earlier studies of this topic. Moreover, we have found indications that estimates based on comparisons between workers in affiliated and non-affiliated firms and relying on the conditional independence assumption may be subject to selection bias. Nevertheless, average treatment effects indicate that at least one out of three ER pensioners would still be working at the age of 66.5 if early retirement was not an option, and this magnitude is almost identical for two cohorts faced with different lower age limits. The results also indicate rather substantial substitution from disability pensions. Exploiting a reduction in the lower age limit as a source of exogenous variation in individual eligibility we have obtained robust difference-in-differences and triple differences estimates indicating that more than two out of three ER pensioners would be working at the age of 63 had the age limit been 64 rather than 62.

Although previous trends towards early retirement appear to have come to an end in many industrialised countries, the financial sustainability of public pension systems is still under pressure due to rapid increases in life expectancy. Policy reforms intended to increase labour force participation among elderly workers are thus likely to be called for also in the years to come. The magnitude of the average treatment effects reported in this paper clearly suggest that there may be much to gain in terms of increased labour supply from restricting access to early retirement schemes.

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A Appendix: Technical details

A.1 On the classification of workers into labour force states

The following procedure is used for the classification of individuals into one of the seven labour force states (work, part-time work and ER benefits, ER, disability, long term sick leave, unemployment, and 'other') at each age for those registered with more than one type of benefits: ER pensions are given priority over disability benefits, and disability benefits over sick leave and unemployment benefits, and only those who did not receive either of these benefits may be classified into the state 'work'. An individual is classified into the state 'part-time work and ER benefits' if she received ER pensions the relevant month, had at least one employment record with termination date after the relevant month, and had an annualized wage exceeding 2BA. Finally, those not registered as benefit recipients and for whom no active employment record is found for the relevant month are classified into the state 'other', which would include individuals retired through private firm-provided retirement schemes or public sector early retirement schemes other than AFP, self-employed, and individuals out of the labour force for other reasons.

A.2 On the classification of firms according to ER affiliation

There is no direct information on ER affiliation at the firm level, and prior to 1999 there is no direct way of separating private from public sector firms. To get around these limitations, I have identified all recipients of public and private ER pension benefits and tried to track down the last job prior to pension take-up for each of these individuals. As a general rule, a firm is classified as public if at least one of its former employees received public ER pensions. In the relatively few cases where the same firm has some former employees receiving public ER benefits and some receiving private ER benefits, the firm is classified as public if the number of public ER benefits recipients exceed the number of private ER benefit recipients and the first take-up of public ER benefits was prior to the first take-up of private ER benefits, or if the number of private ER benefits recipients is very small compared to the number of public ER benefits recipients. A private sector firm is classified as ER affiliated in the base year (1993/1997) if the first take-up of private ER benefits of a former employee happened within the first five years after the base year.

B

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