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The retirement age and the hiring of senior workers[☆]

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ABSTRACT

This paper examines the effect of decreasing the distance to the minimum legal retirement age on the hiring of senior workers. Using variation brought about by a comprehensive old-age public pension reform in Norway in 2011, I identify a positive effect on the hiring of senior workers. The results suggest that the increase was mainly caused by a positive shift in the demand for “risky” senior workers (workers with prior sick-leave and blue-collar workers). This lends support to the notion that risk-averse firms are more willing to hire senior workers when the minimum legal retirement age is decreased.

1. Introduction

In light of increased longevity and the need to extend working lives, reforms of old-age public pension systems are on the agenda for many OECD countries (OECD, 2015). In reforming these systems, policymakers can (somewhat simplified) choose from a limited policy menu consisting of: (i) the flexibility approach, allowing for flexible pension benefit claiming while also strengthening the incentives for continued work, and (ii) the prescription approach, increasing the minimum legal retirement age (LRA).[1] The main contribution of this paper is to illustrate that implementing the flexibility approach can lead to a positive effect on the demand for senior workers through increased hiring. To show this, I consider a comprehensive reform of the Norwegian old-age public pension system in 2011. The Norwegian reform followed the flexibility approach, and the policy discontinuity induced by reducing the minimum LRA from age 67 to age 62 in private sector firms without early retirement schemes allows me to identify a (positive) causal effect of decreasing the distance to the minimum LRA on the hiring of senior workers.

One possible explanation why decreasing the distance to the minimum LRA leads to increased hiring of senior workers relates to the argument made by Lazear (1979). Any firm considering a senior applicant to a vacancy must take into account both the applicant’s perceived current productivity and the expected productivity in the future. While these two factors may be accurately probed at the time of hire, there may be a high downside risk that the senior worker will be subject to a negative productivity shock after the time of hire. The combination of strict employee protection legislation and the limited outside labor market opportunities for senior workers means that the firm may then be left with an “expensive” worker in terms of the wage-productivity differential. It is only when the worker qualifies for pension benefits and can withdraw from the labor market that the firm is relieved of this negative lock-in effect. Lowering the minimum LRA (as in the Norwegian reform) facilitates for the senior worker to withdraw from the labor market earlier than before. As such, reducing the distance to the minimum LRA reduces the firm’s exposure to the downside risk of hiring a senior worker and could therefore increase its propensity to hire senior workers. In addition, we would expect that lowering the minimum LRA would affect different groups of senior workers differently according to the risk associated with hiring them. At the same time, the lowering of the minimum LRA in the Norwegian reform was accompanied by an actuarial adjustment of pension benefits. This means that while there was now a possibility for workers to retire earlier than before (e.g. due to a negative health shock), there were no changes in the economic incentives to do so. For the hiring firm, the reform therefore reduced the risk of a negative lock-in effect while keeping the expected length of the employment relationship the same.

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[1] By the minimum legal retirement age (LRA), I mean the earliest age at which a senior worker can claim old-age public pension benefits conditional on having sufficient individual pension benefit accrual.

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The estimation results accord well with this story. A difference-in-difference approach indicates that the reform increased the probability of private sector firms hiring a senior worker (defined as a worker aged 50 to 61 at the time of hire) by 0.81 percentage points or 16.3% $(100 \times (0.0081/0.0497) = 16.3)$. Estimations using age-specific difference-in-difference specifications show a clear age-differential effect on the probability of a new senior hire, with the effect being lower for higher ages. In light of this, I exploit the reform-induced discontinuity in the distance to the minimum LRA in order to assess the effect of decreasing the distance to the minimum LRA. Using a linear probability model, the results indicate that a one-year decrease in the distance to the minimum LRA increases the probability of a new senior hire by 4.2% $(100 \times (0.0021/0.0497) = 4.2)$. Furthermore, the increase in probability of a new senior hire is concentrated on the extensive margin (with the firm going from no senior hire to one senior hire). To examine the mechanism behind these results, I consider the effect on senior workers with a prior history of sick-leave and blue-collar senior workers. These are two worker characteristics that may indicate the risk of hiring for the firm. Restricting attention to senior workers with a prior history of sick-leave, the effect of decreasing the distance to the minimum LRA by one year increases the probability of a new senior hire by 4.6% $(100 \times (0.0011/0.0234) = 4.6)$. For blue-collar senior workers, the effect is 7.4% $(100 \times (0.0015/0.0201) = 7.4)$. In short, the increase in the hiring of senior workers from decreasing the distance to the minimum LRA is mainly driven by firms being more willing to hire "risky" senior workers than before, as we would expect from the argument made by Lazear (1979).

As traditional pay-as-you-go public pension systems are reformed, increasing the minimum LRA is considered an effective policy measure to lengthen working lives. The results in this paper suggest that doing so may have an unintended detrimental effect on the labor market opportunities of displaced senior workers or senior workers wanting to change jobs to facilitate partial retirement (Kantarci, 2013). In terms of the policy menu, the results therefore point to the flexibility approach being the better choice in terms of accommodating for labor market mobility of senior workers. This paper adds to the existing literature on the effect of changing the minimum LRA, which has mostly focused on labor supply responses and program substitution (see e.g. Börsch-Supan et al., 2017; Geyer and Welteke, 2017; Manoli and Weber, 2016; Staubli and Zweimüller, 2013; Vestad, 2013), by examining the consequences for the demand for senior workers.

The remainder of the paper is structured as follows: Section 2 summarizes some of the existing literature on senior worker labor market mobility. Section 3 introduces the institutional changes brought about by the Norwegian old-age public pension reform in 2011, and motivates the quasi-experimental methods used later in the paper. Section 4 discusses how the reduction in the minimum LRA as a result of the reform can affect both (i) the firm's propensity to hire senior workers and (ii) the (latent) labor supply of senior workers. Section 5 describes the data used, and presents summary statistics of this data. Section 6 introduces the baseline methodology used to estimate the effect of the reform on the hiring of senior workers, and discusses the estimation results. Section 7 sets out the extended methodology to estimate the effect of a one-year decrease in the distance to the minimum LRA on the hiring of senior workers, and discusses the estimation results. Section 8 applies the extended methodology on a subsample of senior worker hires to investigate the possible mechanisms behind the distance-to-retirement effect. Section 9 concludes.

2. Literature

Saint-Paul (2009) suggests three stylized facts that distinguish senior workers from junior workers: (i) their remaining career time is short, (ii) their human capital is more specific to their current job, and (iii) their labor productivity is falling. Taken together, these three characteristics are likely to adversely affect both the supply of senior workers and the demand for senior workers.

On the demand side of the labor market, the use of deferred compensation structures to alleviate monitoring problems reduces the hiring rate of senior applicants as the short employment horizon for these workers reduces the effectiveness of this strategy (Daniel and Heywood, 2007; Heywood et al., 2010; Lazear, 2011). The fixed costs the firm faces when hiring also reduces the hiring rate, especially as senior workers have a greater stock of human capital which needs to be updated upon starting in a new job (Hutchens, 1986; Montizaan and Fouarge, 2016). Firms are also risk-averse with respect to the expected productivity of the applicant worker, which is often declining with age, and therefore leads firms to hire fewer senior workers than junior workers (Heyma et al., 2014; Skirbekk, 2004). The essence of this risk-aversion argument is that hiring a senior applicant who is paid a high wage but only has a potentially high productivity level is less appealing to the firm than to hire a junior applicant with a lower wage and lower productivity.[2]

On the supply side of the labor market, the (latent) labor supply of searching senior workers (either currently employed or unemployed) will be low due to (i) the potential wage loss upon changing employer (owing to loss of firm-specific human capital), (ii) institutional rules (such as the accrual rules of defined benefit occupational pension schemes) and (iii) constraints on outside labor market opportunities acting as a lock-in effect at the current employer (Hurd, 1996; Stier and Endeweld, 2015). Technological age-bias in job search efficiency (such as online job portals) owing to a lower command of new job search techniques leads to the job search efficiency of senior workers to decline with age, which results in a low optimal job search effort (Legendre and Sabatier, 2017). Furthermore, the closer an unemployed senior individual is to the retirement eligibility age, the lower the optimal job search effort becomes through the decrease in the expected additional gain in pension benefits of searching (Hairault et al., 2010).

In terms of labor market reforms, Heyman and Skedinger (2011) study the Swedish reform of reducing notice periods for employer-initiated separations for senior workers, and find a positive effect on firms' propensity to hire senior workers. Behaghel et al. (2008) use a legislative change exempting firms from layoff taxation for workers hired after the age of 50, and find that the transition rate from unemployment to employment (interpreted as the hiring rate) for workers over the age of 50 increased compared to workers under the age of 50. More closely related to the topic of this paper are reforms to the minimum LRA. Ilmakunnas and Ilmakunnas (2015) exploit a Finnish reform increasing eligibility ages for unemployment pension benefits and individual early retirement pension benefits and find a positive effect on the hiring of senior workers. The Finnish reform had direct implications for firms' labor costs (through pension contributions and liabilities), while this paper studies a reform that did not directly affect the labor costs of the firm. Brunello and Langella (2012) study the change in the early retirement age for several European countries, and find that in Northern Europe a one-year increase in the early retirement age increases the probability of senior workers taking up a bridge job, whereas in Southern Europe the same change does not change the probability of taking on a bridge job. Martins et al. (2009) study an increase in the minimum LRA for women in Portugal, and find a reduction in the hiring of junior workers

[2] There may also be a large upside risk of hiring a senior worker (the worker having a higher-than-expected productivity level) although the downside risk generally outweighs the upside risk, at least in terms of physically demanding jobs (Ilmarinen, 2002). To the extent that the variance of productivity levels is greater for senior workers compared to junior workers, statistical discrimination could explain firms' low propensity to hire senior workers (Heyma et al., 2014). In terms of modeling statistical discrimination in the hiring of senior workers, Manger (2014) shows how labor market frictions endogenously lead to older unemployed workers facing lower hiring opportunities, even for age-independent productivity levels and no training costs upon being hired.

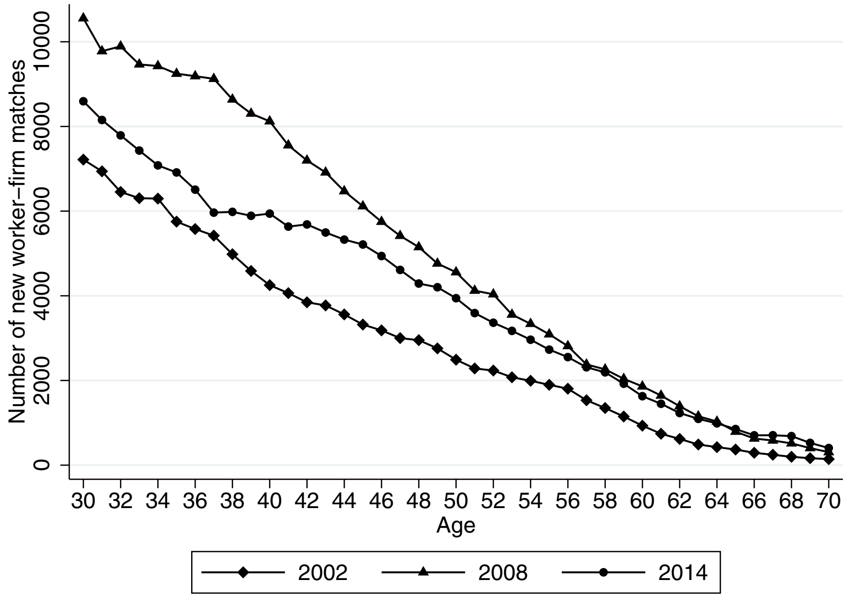

Fig. 1. Number of new worker-firm matches by age, 2002, 2008 and 2014. Source: Author's own calculations using data from Statistics Norway. Note: Number of new worker-firm matches in 2002, 2008 and 2014 in Norway by worker age. The sample is restricted to new worker-firm matches with workers aged 30 to 70 at the time of hire. See Section 5 for the definition of a new worker-firm match.

in the affected firms as a result of senior female workers being retained in the firm for longer.

Putting together these stylized facts about the demand for senior workers and the supply of senior workers, the low labor market mobility of senior workers is well understood. As in most Western European countries, the persistently low labor market mobility of senior workers is also manifest in Norway, as illustrated by the number of new worker-firm matches by age in Fig. 1. In this paper, I aim to investigate if reducing the minimum LRA changes the extent to which senior workers are hired, using the reform of the Norwegian old-age public pension system in 2011 as a case study.

3. Institutional setting

In order to study the effect of reducing the minimum LRA on the hiring of senior workers, I consider the case of the Norwegian old-age public pension system and the associated reform in 2011.[3] In this Section, I provide a brief overview of the Norwegian old-age public pension system, and how the reform changed the system differentially for different groups of workers depending on their firm affiliation. I focus on the earnings model for pension benefits and the eligibility criteria for pension claiming, and relate the differential reform elements to the quasi-experimental setup used later in the paper.

3.1. Before the reform

I restrict attention to senior workers in two types of firms: private sector firms without early retirement schemes (henceforth denoted as private sector firms), and public sector firms with early retirement schemes (henceforth denoted as public sector firms). Prior to the reform, the minimum LRA for old-age public pension benefits, henceforth denoted as FT (*folketrygd*) benefits, was age 67 for all workers conditional on eligibility. The annual FT pension benefits granted upon claiming were calculated based on the sum of (i) a basic pension (*grunnpensjon*), and (ii) an earnings-related pension (*tilleggspensjon*) based on a scaled measure of the sum of pension points from the 20 years of highest annual labor income (Kudrna, 2017).

The full basic pension required a 40-year period of residence (counting from age 17 until the claiming age), and equaled between 0.85 and 1 base amount (1 base amount (BA) in 2016 = 2016-NOK 92,576 ≈ 2016-USD 11,109) depending on the marital status and spousal income.[4] For each year of residence less than 40 (but more than three), the basic pension was proportionally reduced. For single individuals, and letting $R \geq 3$ denote the number of residence years, the annual basic pension (*GP*) measured in BA was:

$$GP = \frac{\min\{R, 40\}}{40}$$

The earnings-related pension was based on pension points earned each year (P_t) accumulated by the following scheme, where $Y_t > 0$ denotes pre-tax pensionable income in year t measured in BA:[5]

$$P_{t\in\{1967,1991\}} = \min\left\{\max\left\{Y_t - 1, 0\right\}, 7\right\}$$
$$+ \max\left\{\min\left\{\frac{1}{3}\left[Y_t - 1 - 7\right], \frac{4}{3}\right\}, 0\right\}$$
$$P_{t\in\{1992,\ldots,\}} = \min\left\{\max\left\{Y_t - 1, 0\right\}, 5\right\}$$
$$+ \max\left\{\min\left\{\frac{1}{3}\left[Y_t - 1 - 5\right], 2\right\}, 0\right\}$$

For ages 17 and up to the age of pension claiming, the pension points accumulated in the years of highest pre-tax pensionable income (up to 20 years) were summed and divided by the number of years ($K \leq 20$) to obtain the final pension point count (*sluttpoengtallet*), denoted *SLP*:

$$SLP = \frac{1}{K}\sum_{j=1}^{K} P_j$$

The earnings-related pension (*TP*), measured in BA, was calculated as a scaled sum of the final pension point count (*SLP*), where *A* denotes the

[3] The exposition of the institutional details of the old-age public pension system in Norway and the changes brought about by the pension reform of 2011 in this Section borrows heavily from Hernæs et al. (2016); Kudrna (2017) and Andersen (2018).

[4] The nominal value of the base amount (BA) measure is adjusted each year to follow the average wage growth in the economy.

[5] In addition to pension point accrual based on pensionable income, individuals taking care of others (primarily young children) are awarded caretaker points (*omsorgspoeng*) as a replacement of foregone earnings. See Andersen (2018) for more details.

number of years with positive pension point accumulation (i.e. years with more than one BA of pre-tax pensionable income), as follows:

$$TP = \frac{SLP \times 0.45 \times \max\left\{A_{t\in\{1967,1991\}}, 40\right\}}{40} + \frac{SLP \times 0.42 \times \min\left\{A_{t\in\{1992,\ldots,\}}, 40 - A_{t\in\{1967,1991\}}\right\}}{40}$$

Before the reform, the annual FT pension benefits ($FTP = GP + TP$) were not adjusted and therefore constituted a fixed lifetime annual entitlement upon claiming. There was also a minimum annual FT pension benefit threshold (*minstepensjon*), varying between 1.59 BA and 3 BA depending on marital status and spousal income. If the annual FT pension benefits fell below the minimum annual FT pension benefit threshold (*MP*) due to low earnings-related pensions (*TP*), the individual was granted supplemental pension benefits (*særtillegg*) equal to the difference between *MP* and the annual FT pension benefits (*FTP*):

$$ST = \max\{MP - FTP, 0\}$$

As a result, the total annual pension benefits for workers in the private sector with no early retirement scheme equaled $FTPT = FTP + ST$. For the majority of senior workers in the private sector with no early retirement scheme prior to the reform, the age limit of 67 was the binding eligibility constraint in terms of accessing their full old-age pension benefits, due to sufficient number of accumulation years ($A \geq 40$) and being past the peak of their age-earnings profile in terms of their 20 best years.[6]

However, senior workers in the public sector firms had access to a supplementary early retirement scheme, henceforth denoted as the AFP (*avtalefestet pensjon*) scheme, where the minimum LRA was age 62 conditional on eligibility. The eligibility criteria for accessing the AFP benefits at age 62 were (i) current employment in a public sector firm, (ii) at least 3 years tenure with the present employer, (iii) at least 10 years of pensionable income above one BA since age 50, and (iv) an average of the 10 highest annual incomes after 1966 exceeding two BA (Hernæs et al., 2016). The size of the annual AFP benefits were typically comparable to the total annual FT pension benefits (*FTPT*). The fixed annual AFP benefits, measured in BA, were calculated as the total FT pension benefits that *would* have been awarded given continued work up to the age of 66 (denote this $FTPT_{\hat{P}}$) plus a nominal annual supplemental pension of NOK 20,400 ($\approx$ 2016-USD 2,450) divided by the nominal value of BA in year t (G_t):

$$AFP = FTPT_{\hat{P}} + \frac{20,400}{G_t}$$

The AFP benefits were subject to a strict earnings test upon claiming that was effective from low levels of labor income (NOK 15,000 $\approx$ 2016-USD 1,800) with no actuarial adjustment of benefits (meaning it constituted a fixed annual entitlement), implying a very high implicit tax rate on continued work after the age of 62.[7] As a result of the age limits and the lack of actuarial adjustments of the old-age pension benefits, the observed worker retirement ages prior to the reform were largely concentrated at age 62 and 63 for public sector workers and at age 67 for private sector workers without AFP. The distribution of retirement ages for workers from the public and private sector firms prior to the reform is illustrated in Fig. 2.

3.2. After the reform

The old-age public pension reform, which was implemented in January 2011, reduced the minimum LRA for FT pension benefits from 67 to 62, allowed for flexible claiming of pension benefits (either fully or partially) between ages 62–75, and introduced actuarial adjustment of the pension benefits. This meant that in contrast to the pre-reform period the annual old-age public pension entitlements were not fixed regardless of the time of claiming, but rather depended on the individual's age at claiming and birth cohort.

3.2.1. Cohorts 1943–1953

After the reform, the model for calculating FT pension benefits was kept largely unchanged for the cohorts 1943–1953. However, to actuarially adjust the annual pension benefits, the two components of the FT benefits (the basic pension *GP* and the earnings-related pension *TP*) are adjusted to take into account (i) the cohort-specific longevity (relative to the 1943 cohort) and (ii) the age at time of claiming. More specifically, for cohort c claiming pension benefits from age $62 \leq T \leq 75$ the two components are divided by an adjustment factor (***forholdstall***) to adjust for this (denote this adjustment factor by $FHT_{c,T}$):

$$GP_{c,T} = \frac{GP}{FHT_{c,T}}$$
$$TP_{c,T} = \frac{TP}{FHT_{c,T}}$$

The sum of these two components then constitutes the FT pension benefits for cohort c when claiming benefits from age $62 \leq T \leq 75$ (denote this $FTP_{c,T}$), measured in BA:

$$FTP_{c,T} = \frac{GP}{FHT_{c,T}} + \frac{TP}{FHT_{c,T}} = GP_{c,T} + TP_{c,T}$$

Similar to before, if the FT pension benefits (before adjusting for cohort-specific longevity and age at claiming) falls below the product of the new minimum annual FT pension benefit (*minste pensjonsnivå*) guaranteed at age 67 ($MPN_{c,67}$) and the adjustment factor evaluated at age 67 ($FHT_{c,67}$), the difference is granted as supplemental pension benefits (*pensjonstillegg*) divided by the adjustment factor ($FHT_{c,T}$):

$$PT_{c,T} = \frac{\max\left\{\left(FHT_{c,67} \times MPN_{c,67}\right) - \left(FHT_{c,T} \times FTP_{c,T}\right), 0\right\}}{FHT_{c,T}}$$

The total annual pension benefits at age a for cohort $1943 \leq c \leq 1953$ when claiming from age $62 \leq T \leq 75$ (denote this $FTPT_{c,T,a}$) after the reform is adjusted by wage growth minus 0.75%:

$$FTPT_{c,T,a} = \left[FTP_{c,T} + PT_{c,T}\right](1 - 0.0075)^{a-T}$$

The criterion for being eligible to claim FT pension benefits prior to age 67 for these cohorts is based on the notion of sufficient accrual. This consists of having enough FT pension benefit entitlements (*FTP*) and supplemental pension benefits (*PT*) at the age of pension benefit claiming to achieve a hypothetical level of total pension benefits at age 67 larger than the new (predicted) cohort-specific minimum annual FT pension benefit guaranteed at the age of 67 ($MPN_{c,67}$), adjusting for wage growth minus 0.75% up to age 67. Denoting the claiming age by $A < 67$, the condition for sufficient accrual can be expressed as:

$$\left[FTP_{c,A} + PT_{c,A}\right](1 - 0.0075)^{67-A} \geq MPN_{c,67}$$

For public sector workers in the cohorts 1943–1953, the reform did not change the structure of the AFP scheme, meaning these workers still faced the same disincentive to work past the age of 62 as before and no change in the minimum LRA. After the reform, eligible workers in the public sector can however claim FT pension benefits from age 62 instead of age 67, but would then have to disavow their AFP benefits. The main change for these workers is the adjustment of the AFP benefits by the wage growth minus 0.75%, such that the annual AFP benefits at age a when claiming from age T ($AFP_{c,T,a}$) becomes:

$$AFP_{c,T,a} = FTPT_{\hat{P}}(1 - 0.0075)^{a-T} + \frac{20,400}{G_t}$$

[6] For cohorts 1938–1943 employed in the private sector with no early retirement scheme (and not on disability insurance at age 66), 99.9% of the sample had at least three years of residence and 97.8% had 40 or more years of residence at age 66.

[7] For cohorts 1942–1947 employed in the public sector (and not on disability insurance at age 61), the age limit of 62 was the binding eligibility constraint at age 61 for 86.0% of the sample.

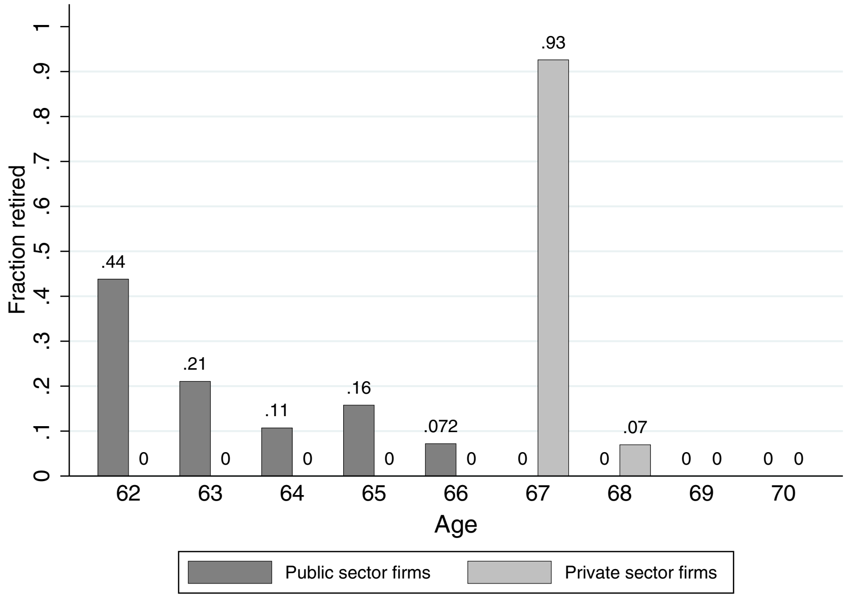

Fig. 2. Distribution of retirement ages, public sector firms and private sector firms. Source: Author's own calculations using data from the Statistics Norway. Note: Fraction of sample observed claiming old-age public pension for each age $a = 62, \ldots, 70$ and that are not working at age $a + 1$ (defined as the annual labor income at age $a + 1$ being less than 1/2 of annual labor income at age a) prior to 2011. The sample consists of the pre-reform worker birth cohorts 1938–1943, and claiming decisions are observed up to and including 2010. The distribution is not conditional on workers having sufficient pension benefit accrual to retire at the earliest possible age.

3.2.2. Cohorts 1954 and onwards

For the cohorts 1954–1962, the share of pension benefits calculated by the old earnings model described in Section 3.1 decreases by 1/10 for each successive cohort. This means that for the 1954 cohort, 1/10 of the benefits are calculated according to a new earnings model (described in more detail below) and 9/10 of the benefits are calculated according to the old earnings model. The new earnings model for calculating pension benefits applies fully for the 1963 cohort and onwards.

Under the new earnings model for calculating pension benefits, pension wealth is accumulated for any positive pensionable income (up to 7.1 BA) each year between the ages 17–75 at an accumulation rate of 18.1%. More specifically, pension points earned in year t (P_t) are calculated according to the following formula, where $Y_t > 0$ denotes pre-tax pensionable income in year t measured in BA:

$$P_t = \min\left\{Y_t, 7.1\right\}$$

The sum of pension points are then converted into earnings-related pension wealth (*pensjonsbeholdning*) earned by age $a \leq 75$ (PB_a) by an accumulation rate of 18.1%:

$$PB_a = 0.181 \times \sum_{j=17}^{a} P_j$$

Based on the earnings-related pension wealth, the FT pension entitlement (*inntektspensjon*) for an individual from cohort c claiming from age $62 \leq T \leq 75$ ($IP_{c,T}$) is found by dividing the pension wealth accumulated by age T by an adjustment factor (*delingstall*) to account for cohort-specific life expectancy and age at time of claiming (denote this adjustment factor $DT_{c,T}$):

$$IP_{c,T} = \frac{PB_T}{DT_{c,T}}$$

If 80% of the total FT pension wealth earned by age $62 \leq T \leq 75$ falls below the product of the new minimum FT pension benefits (*garantipensjon*) guaranteed at age 67 ($GAP_{c,67}$) and the adjustment factor evaluated at claiming age 67 ($DT_{c,67}$), the difference is granted as new supplemental pension benefits (*garantitillegg*) divided by the adjustment factor evaluated at age $62 \leq T \leq 75$ ($DT_{c,T}$):[8]

$$GAT_{c,T} = \frac{\max\left\{\left(DT_{c,67} \times GAP_{c,67}\right) - 0.8 \times PB_T, 0\right\}}{DT_{c,T}}$$

The total annual pension benefits at age a for cohort $c \geq 1963$ when claiming benefits from age $62 \leq T \leq 75$ ($IPT_{c,T,a}$) after the reform is adjusted by wage growth minus 0.75%:

$$IPT_{c,T,a} = \left[IP_{c,T} + GAT_{c,T}\right](1 - 0.0075)^{a-T}$$

The condition for claiming FT pension benefits before age 67 for the pure new-system cohorts (cohorts 1963 and onwards) is based on the notion of sufficient accrual, which consists of having enough FT pension benefit entitlements (*IP*) and supplemental pension benefits (*GAT*) at the age of claiming to achieve a hypothetical level of total FT pension benefits at age 67 greater than the new (predicted) cohort-specific minimum annual FT pension benefit guaranteed at the age of 67 ($GAP_{c,67}$), adjusting by wage growth minus 0.75% up to age 67. Denoting the claiming age by $A < 67$, the eligibility condition can be expressed as:

$$\left[IP_{c,A} + GAT_{c,A}\right](1 - 0.0075)^{67-A} \geq GAP_{c,67}$$

For the hybrid cohorts 1954–1962, the calculation of annual pension benefits and the eligibility condition is a weighted combination of the old pension system and the new pension system. Similar to the system applied to the cohorts 1943–1953 as described in Section 3.2.1, the reform did not change the structure of the AFP scheme for public sector workers from the cohorts 1954 and onwards. Similar to the cohorts 1943–1953, the main change was the adjustment of the AFP benefits by the wage growth minus 0.75%, such that the annual AFP benefits at age a when claiming benefits from age T ($AFP_{c,T,a}$) becomes:

$$AFP_{c,T,a} = IPT_{\hat{P}}(1 - 0.0075)^{a-T} + \frac{20,400}{G_t}$$

[8] Similar to the pre-reform basic pension (*grunnpensjon*), the full new minimum FT pension benefits requires a 40-year period of residence (counting from age 17) and is proportionally reduced for each year of residence less than 40 (but more than three).

Table 1
Summary of the Norwegian old-age public pension reform.

	Changes from the reform
Private sector, no AFP scheme	No change in return to work Reduction of minimum LRA
Public sector, AFP scheme	No change in return to work No change in minimum LRA

Source: Column 1 in Table 1 in Hernæs et al. (2016, p. 41).

3.3. Summary of the reform

While the reform's change of the earnings-related part of the public pension benefit accrual may in the long run improve the work incentives of the cohorts 1954 and onwards (as all positive pensionable earnings count for pension benefit accrual, and not only pensionable earnings above one BA as under the old system), the main effect was to increase the private sector workers' liquidity with respect to the possibility of claiming FT pension benefits earlier than before.[9] Based on the preceding discussion, the main reform elements for the workers in private sector firms and the workers in public sector firms are summarized in Table 1.

Table 1 illustrates that the reform closely resembled a quasi-experimental setup, where the treatment group consists of workers in the private sector (where the minimum LRA was reduced from age 67 to age 62) and the control group consists of workers in the public sector (where the minimum LRA was kept constant at age 62). As expected from Table 1, the reform did not lead to economically significant changes in the labor supply of neither the workers in the private sector firms nor of the workers in the public sector firms (Brinch et al., 2015; Hernæs et al., 2016).[10] The reform did however lead to increased hiring of senior workers, as illustrated by the number of new senior hires over the period 2007–2014 in Fig. 3. The level difference of new senior hires was substantial prior to the reform (average annual difference of 7,211) while it converges after the implementation of the reform (average annual difference of 4,222), mostly as a result of an increase in the number of senior hires in the private sector firms. This could suggest that the reform's liquidity element affected the hiring of senior workers positively.

3.4. Challenges to the quasi-experimental setup

However, the pension reform also brought about changes to a third group of workers, namely workers in private sector firms with the AFP scheme (constituting around 50% of all workers in the private sector). Prior to the reform, the eligibility criteria for and the earnings test of the private sector AFP benefits were largely the same as in the public sector firms. More specifically, the worker had to be employed in an AFP-affiliated private sector firm at the time of benefit claiming and either (i) be employed by the same firm for the last three years or (ii) be employed by a private sector firm with AFP for the last five years. The annual nominal supplemental pension was approximately NOK 19,000 (≈ 2016-USD 2,280). However, after the reform the AFP benefits were subject to an actuarial adjustment upon claiming and the earnings test was removed, which spurred a substantial increase in the labor supply of senior workers in these firms (Brinch et al., 2015; Hernæs et al., 2016).[11] As such, there may be differential spillover effects from these firms that could invalidate the quasi-experimental setup outlined in Table 1. For instance, the increased labor supply of workers in private sector firms with AFP after the reform could differentially change the pool of job applicants available to the public sector firms and the private sector firms without AFP. If we consider the rate at which senior workers transfer between the three types of firms in the pre-reform and post-reform period in Table 2, there seems to have been no large changes in the aggregate rates over time. More importantly, the transition rates suggest that there were no differential spillover effects from private sector firms with AFP to the treatment group (private sector firms without AFP) and the control group (public sector firms).

Table 2
New senior hires from private sector firms with AFP, 2007–2014.

	Fraction of new senior hires from private sector firms with AFP:	
Year	Private sector firms without AFP	Public sector firms
2007	0.192	0.059
2008	0.192	0.070
2009	0.194	0.074
2010	0.179	0.079
2011	0.166	0.069
2012	0.167	0.067
2013	0.162	0.068
2014	0.188	0.066

Source: Author's own calculations using data from Statistics Norway.
Note: Fraction of new senior hires in (i) private sector firms without AFP (treatment group) and (ii) public sector firms (control group) coming from private sector firms with AFP, 2007–2014. A senior worker is defined as a worker aged 50 to 61 at time of hire. The sample is restricted to firms more than 5 full-time equivalents in the year prior to hiring.

In terms of determining whether public sector firms serve as good controls for the private sector firms in a quasi-experimental setup, Schumann (2017) points to the stable unit treatment value assumption (SUTVA) and the associated trade-off between (i) independence of entities and (ii) comparability in terms of hiring processes. The public sector firms and private sector firms generally do not overlap much in terms of firms' industry classification codes, as illustrated by the distribution of firms in the sample across industries over the period 2007–2014 using the first two digits of the SN2007 industry classification system in Table 3.[12] This could suggest that the public and private sector firms may differ in terms of pool of potential job candidates when hiring, so that increased hiring from private sector firms as result of the reform did not affect the potential for public sector firms to hire. The lack of overlap between private sector firms and public sector firms is also evident when considering the job classification codes of the newly hired senior workers, further substantiating the claim that the SUTVA is satisfied (see the Online Appendix for the distribution of new senior hires across the different job classification codes using the STYRK98 job classification system). As such, the private sector firms and the public sector firms are fairly independent entities, suggesting that the SUTVA may well be sat-

[9] The main sample used later in the paper consists of workers aged 50–61 and covers the years 2007–2014 (in other words the cohorts 1946–1964). Considering the gradual implementation of the new earnings-related part of the pension benefit accrual after the reform for cohorts 1954–1962, we can expect the labor supply responses as a result of this to be negligible in this application.

[10] A small positive extensive margin response was recorded for public sector workers, indicating that some found continued work after the age of 62 attractive despite that this implied them disavowing their AFP pension benefits (Hernæs et al., 2015). Furthermore, a small and marginally significant negative extensive margin response was found for workers in private sector firms without AFP, despite no explicit changes in the work incentives as a result of the reform (Hernæs et al., 2016).

[11] For the sake of brevity, I exclude the details on the changes in the eligibility criteria, pension benefit accrual rules and the earnings test for the private sector AFP benefits. See Kudrna (2017) and Andersen (2018) for more details.

[12] The estimation results reported later in Section 6.2 and Section 7.2 do not change much when including sector-specific linear time trends to mitigate the lack of sectoral overlap. The complete estimation results are relegated to the Online Appendix.

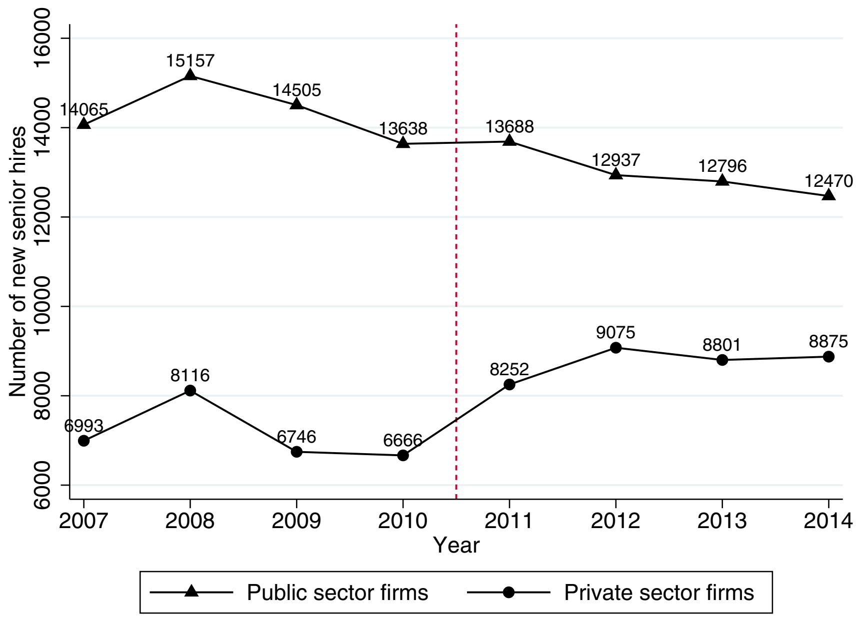

Fig. 3. Number of new senior worker hires, 2007–2014. Source: Author's own calculations using data from Statistics Norway. Note: Number of new senior worker hires (in private sector firms and public sector firms) with more than 5 full-time equivalents in the year prior to hiring, 2007–2014. A senior worker is defined as a worker aged 50 to 61 at time of hire. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011.

Table 3
Distribution of firms across industries.

Industry classification	Private sector firms	Public sector firms
Primary sector	0.0208	0.0025
Manufacturing	0.1003	0.0063
Construction	0.0823	0.0072
Wholesale retail and trade	0.1533	0.0230
Transportation and storage	0.1745	0.0160
Accommodation and food service	0.0280	0.0008
Information and finance	0.0776	0.0136
Professional, scientific and legal activities	0.1774	0.1789
Administration	0.0428	0.2267
Education, human health and social work	0.0852	0.4760
Other	0.0578	0.0489

Source: Author's own calculations using data from Statistics Norway.
Note: Fraction of firms in each industry category (by firm type), based on the two first digits of the four-digit SN2007 industry classification as follows: primary sector $\in$ {01, 02, 03, 05, 06, 07, 08, 09}; manufacturing $\in$ {$1x$, $2x$, 30, 31, 32, 33, 35, 36, 37, 38, 39} for $x = 0, 1, \ldots, 9$; construction $\in$ {41, 42, 43}; wholesale retail and trade $\in$ {45, 46, 47}; transportation and storage $\in$ {49, 50, 51, 52, 53}; accommodation and food service $\in \{55, 56\}$; information and finance $\in$ {58, 59, 60, 61, 62, 63, 64, 65, 66}; professional, scientific and legal activities $\in$ {69, 70, 71, 72, 73, 74, 75}; administration $\in$ {77, 78, 79, 80, 81, 82, 84}; education, human health and social work $\in$ {85, 86, 87, 88, 89}; other = none of the above. See Statistics Norway (2008) for a comprehensive overview of the SN2007 industry classification standard.

isfied.[13] To assess whether the lack of sectoral overlap could challenge the economic comparability of the hiring process in the private sector firms and public sector firms, I now consider the mechanisms behind how a reduction in the minimum LRA can affect both the firms' hiring of senior workers and the (latent) labor supply of senior workers.

[13] To the extent that there are potential spillover effects, it would reduce the public sector firms' propensity to hire senior workers. In this case, the estimated effect of the reform on the hiring of senior workers would therefore be overestimated.

4. Mechanisms

In this Section, I discuss how changes in the distance to retirement as a result of reducing the minimum LRA could potentially affect both the firms' hiring of senior workers and the senior workers' job search behavior, with reference to the elements of the old-age public pension reform sketched out in Section 3.

4.1. Hiring of senior workers

When a firm chooses to hire a senior applicant for a vacant position, this is equivalent to the firm making a complex, and in light of strict employee protection legislation (henceforth denoted as EPL), a potentially irreversible investment decision. As argued by Hairault et al. (2010, p. 1037) in the context of firms hiring senior workers:

> [...] retirement age is private and asymmetric information. Firms must solve a potentially difficult problem to infer the expected retirement age for each individual, especially when the SS [social security] system is actuarially fair, that is, when retirement age relies on individual preferences for leisure.

Of course, the same complex investment problem of deriving the expected length of the employment relationship arises for a firm hiring a worker at any age. However, what distinguishes a senior worker from other workers is their clear outside option of claiming old-age pension benefits and withdrawing from the labor market upon reaching eligibility. This means that, roughly speaking, a senior worker who has reached eligibility has two opportunities from the firm's point of view: (i) continue working (with or without claiming old-age pension benefits), or (ii) retire and claim old-age pension benefits. Since a senior worker's extensive margin choice essentially only concerns continued employment within the same firm (the supply side factors are sketched out in greater detail in the literature review in Section 2), the horizon of the employment relationship is more salient for firms when hiring senior workers compared to hiring junior workers.[14] For firms hiring senior workers,

[14] Surveys of Norwegian employers over the period 2008–2015 indicate that this is a salient consideration for firms, with around 30% of respondents (man-

the horizon of the employment relationship will therefore in general coincide with the distance to the worker's expected retirement age.

The expected distance to retirement for a new senior hire matters to the hiring firm for two reasons, which we can denote as (i) the "horizon" effect and (ii) the "lock-in" effect. The first effect concerns the time period over which the firm can amortize any fixed costs of hiring such as vacancy posting costs or training costs (Mühlemann and Leiser, 2015; Oi, 1962). The second effect concerns the possibility of the firm at a later point in time ending up with a senior employee with a low productivity relative to the perceived productivity at the time of hire (due to for example a negative health shock, poor match quality or general productivity depreciation) who it cannot dismiss due to strict EPL and whose wage may not unilaterally be adjusted downwards.[15] The longer the expected distance to retirement, the longer is the time period over which the fixed costs of hiring can be amortized (a positive effect on firm surplus). However, the greater is also the cumulative risk of ending up with a low-productive employee (a negative effect on firm surplus). In short, the horizon effect increases the firm's propensity to hire senior workers, whereas the lock-in effect reduces the firm's propensity to hire senior workers.

When arguing that the expected distance to retirement matters to the firm when hiring a senior worker, we can distinguish between two metrics the firm cares about: (i) the minimum LRA, and (ii) the age at which the worker actually retires and claims old-age pension benefits. Under the Norwegian pre-reform regime, the two metrics would largely coincide for both the public sector workers (at age 62) and the private sector workers (at age 67), as illustrated in Fig. 2. Under the Norwegian post-reform regime, the minimum LRA constitutes a rule-based lower bound on the actual retirement age. For private sector workers, this lower bound is disentangled from the actual retirement age, which is separately determined by individual characteristics (e.g. preferences, wealth, family situation, health, etc.). For public sector workers, the two metrics continued to coincide due to the work (dis)incentives after age 62 being the same before and after the reform.

Under the pre-reform regime, the high minimum LRA for workers in private sector firms often meant a large spillover into the disability insurance (DI) programme. Under screening conditions related to health status and associated reduced work capacity, a senior worker could claim DI pension benefits prior to the minimum LRA. This meant that for some eligible senior workers the DI programme essentially constituted a conditional-on-health early retirement scheme. This health-dependent retirement pathway was extensively used both prior to and after the reform, as illustrated by the age distribution of new DI benefit recipients in Fig. 4. Under the pre-reform regime, a private sector firm who employed a senior worker subject to an adverse health shock was therefore not exposed to the full extent of the negative lock-in effect, as the worker could enroll in the DI programme conditional on passing the screening process.

However, for senior workers who became low-productive because of either a less-than-adverse health shock (where the worker in question did not pass the screening process for DI benefits), general productivity depreciation (and consequent lower command of job tasks), or who simply turned out to be a poor worker-firm match after the time of hire, the lock-in effect was still present for private sector firms.[16] If the interests of the firm and the worker were non-aligned, meaning the firm wanted to dismiss the worker but the worker wanted continued employment, the firm could not simply dismiss the worker at its own behest due to strict EPL.[17] There could also be stigma costs associated with being on DI, deterring individuals from applying (Hernandez et al., 2007). Of course, the interests of the firm and the senior worker could very well be aligned: the firm wanting to dismiss the worker and the worker wanting to leave the firm. However, the combination of poor outside employment opportunities for the senior worker (Dorn and Sousa-Poza, 2010) and the firm facing potentially high costs of severance payments to compensate the worker for the lack of good outside opportunities (Hallberg and Eklöf, 2010; Huber et al., 2016) led to the mutually undesirable lock-in effect.

While the private sector firms and the public sector firms operate in different sectors and post vacancies that differ in terms of job classification codes (see Section 3.4), both types of firms are likely to want to avoid this general lock-in effect when considering job applicants. With public firms mostly operating within education, human health and social work and private sector firms within transportation, manufacturing and retail, both types of firms are exposed to workers having strenuous tasks and the subsequent risk of the negative lock-in effect. To illustrate that this lock-in effect could very well be a important consideration for employers in general when hiring, Fig. 5 shows the average transition rate from good to bad health (based on the self-reported health status from the Norwegian Labor Force Survey) by age and education. While the transition rate from good to bad health is generally increasing by age for all education groups, the increase is especially pronounced for workers with only completed compulsory education.

Under the post-reform regime, the lock-in effect in the case of aligned interests was reduced for both senior workers and firms since the minimum LRA was reduced. For the workers, the option of retirement was made available at an earlier stage for those wanting to leave the firm, for example due to poorer health. There could be a host of other so-called "jump factors" leading to senior workers wanting to retire at an early age, but there is evidence pointing to poor health being the key factor explaining why senior workers choose to retire early (see e.g. (Bråthen and Bakken, 2013; Dwyer and Mitchell, 1999; García-Gómez, 2011)). Indeed, the distribution of retirement ages for workers coming from the private sector firms became more evenly distributed after the reform (illustrated in Fig. 6), which indicates that the lock-in effect was mitigated as a result of the reduction in the minimum LRA. For firms, the possibility for senior workers to claim benefits earlier could mean that the associated potential costs of severance payments were reduced, at least within a behavioral interpretation of worker behavior.[18] This then implies that the reduction in the minimum LRA could be a salient feature for private sector firms when hiring, as a reduction of this threshold reduced the potential negative lock-in effect. The reduction in the minimum LRA could therefore be thought of as giving firms access to a "safety valve" when hiring a senior worker, with the downside risk

agers) reporting that a worker who is to be hired should have at least 10 years left of his or her career. Only 15–20% of the respondents indicate that the distance to retirement does not affect their hiring decision (Dalen, 2015).

[15] Collective wage bargaining in Norway follows a two-tier system, with the setting of central level "tariff wages" and allowing for local level "wage drifts" (Barth et al., 2014). The collective wage agreements cover around 2/3 of all employees and lead to a compressed wage distribution (IMF, 2016), such that the scope for unilateral downwards adjustment of wages by the firm is limited.

[16] According to Galaasen et al. (2012) the rejection rates for DI benefit applicants are the highest for workers with complex musculoskeletal diagnoses, as they are on average regarded as less incapable of working (compared to the well-defined musculoskeletal diagnoses) by the National Insurance administration. With these conditions often stemming from repetitive manual work or psychosocial strain, senior workers with long working lives behind them could be especially exposed to DI application rejection.

[17] Norway is ranked as a country with a fairly high degree of EPL, and firms can generally only dismiss workers if the firm is making a loss or in cases of a worker's material breach of the employment contract (OECD, 2013, p. 81–83). There are no formal rules on how downsizing should be carried out, although the "last in, first out"-rule is a strong norm (Huttunen et al., 2011). Upon job loss workers can apply for unemployment benefits, with a requirement of income of at least 1.5 BA in the past calendar year.

[18] An economically rational worker would demand the same severance pay as before the reform, as the annual old-age public pension benefits after the reform were subject to actuarially fair readjustments (see Section 3.2). For individuals with hyperbolic discounting or who exhibit partial inattention behavior (see e.g. Gabaix, 2016), this need not be the case.

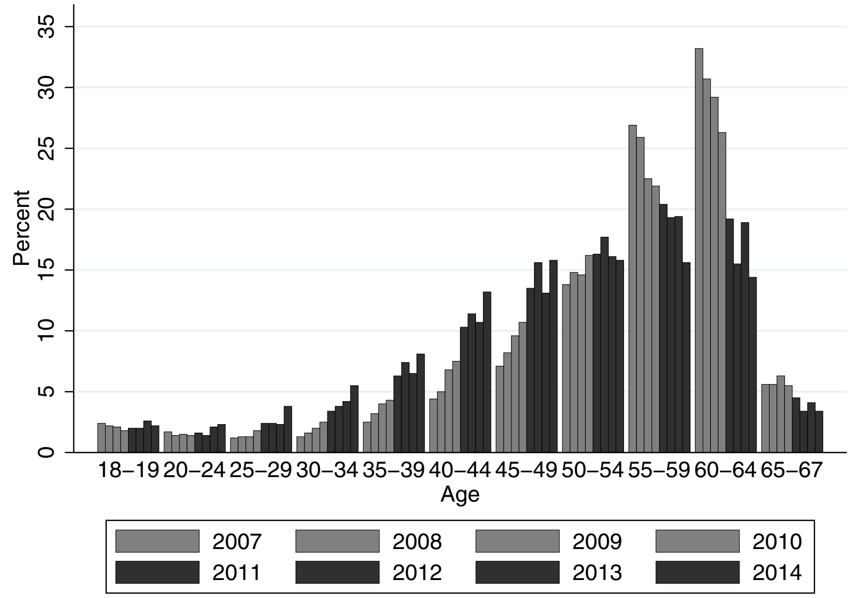

Fig. 4. Distribution of new DI benefit recipients by age, 2007–2014. Source: Data from Norwegian Labour and Welfare Administration (2016). Note: Percent of new disability insurance (DI) benefit recipients within each five-year age bin (ages 18–67). Light gray bars indicate period under the pre-reform old-age public pension system (2007–2010), while dark gray bars indicate period under the post-reform old-age public pension system (2011–2014).

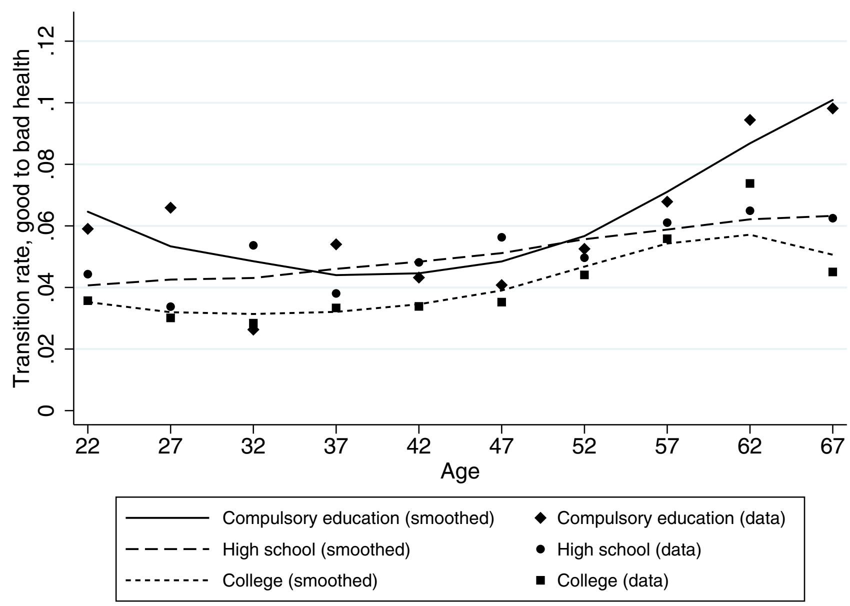

Fig. 5. Transition rates from good to bad health, by age and education. Source: Author's own calculations using data from the Norwegian Labor Force Survey (covering the years 2002, 2004–2008 and 2010–2014). Note: Average rate of transition from good to bad health (self-reported) for employed individuals by age (collapsed into 5-year bins) and education level (compulsory education, high school and college) for the years 2002, 2004–2008 and 2010–2014. The smoothing of the data points is done using locally weighted linear regression of the average transition rate on age, with bandwidth = 0.8.

of hiring a senior worker potentially being greater than the upside risk. As such, we should expect a positive shift in private sector firms' hiring of senior workers, resulting from the reduction of the minimum LRA and the subsequent reduction of the potential lock-in effect. Due to the actuarial adjustment of the FT pension benefits, the incentives to keep working for longer (or shorter) than before were not changed as a result of the reform. While the worker could now leave the firm earlier than before, the incentives to do so did not change and thus there was no change in the uncertainty of "losing" productive workers. Therefore we should expect that there was no change in private sector firms' hiring due to changes to the horizon effect, with the only change coming from the reduced lock-in effect. For public sector firms, we should expect no change in hiring as a result of the reform. This is because the incentives to work past the unchanged minimum LRA did not change, meaning there was no change in neither the horizon effect nor the lock-in effect.

4.2. Senior workers' job search

One of the potential consequences of a reduced minimum LRA on the latent labor supply of searching senior workers (the job search intensity) was a reduction in the need for bridge jobs and thus a reduction in the number of senior job-to-job transitions. Bridge jobs are jobs that senior workers who want to perform gradual retirement in the form of reducing the work burden (either through reducing the number of

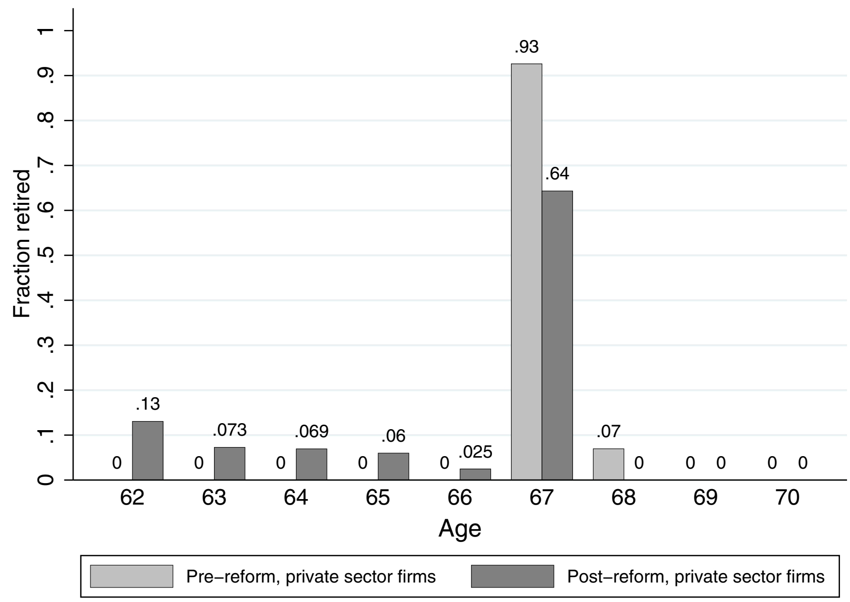

Fig. 6. Distribution of retirement ages, private sector firms. Source: Author's own calculations using data from the Statistics Norway. Note: Fraction of sample observed claiming old-age public pension for each age $a = 62, \ldots, 70$ and that are not working at age $a + 1$ (defined as the annual labor income at age $a + 1$ being less than 1/2 of annual labor income at age a) prior to 2011. The pre-reform sample consists of worker cohorts 1938–1943 (claiming decisions are observed up to and including 2010), while the post-reform sample consists of worker cohorts 1947–1951 (claiming decisions are observed during the period 2011–2014).

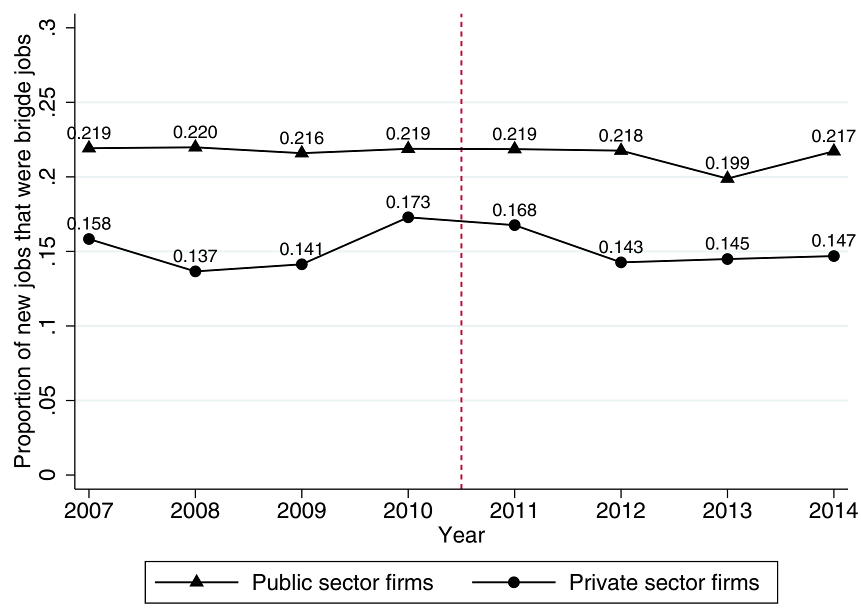

Fig. 7. Bridge jobs as proportion of new senior hires, 2007–2014. Source: Author's own calculations using data from Statistics Norway. Note: Proportion of new senior worker job-to-job transitions that were bridge jobs in private sector firms and public sector firms, 2007–2014. A bridge job is defined as a job with a reduction in both the number of working hours and the annual total wage income by 25% or more compared to previous job. A senior worker is defined as a worker aged 50 to 61 at time of hire. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011.

working hours or changing work tasks) take on by changing from their career employer to a short-term job in the years prior to full retirement (Brunello and Langella, 2012; Kantarci, 2013). Why the reform should affect this is perhaps best illustrated by a hypothetical example. Consider a worker of age 60 who worked in a physically demanding job (e.g. industrial processing) in a private sector firm. Prior to the reform this worker would have 7 more years left until the minimum LRA, and would therefore perhaps want to change to a less demanding job to ease the work burden. After the reform, the same worker only had 2 years left until the minimum LRA and would therefore perhaps not have the same need for a bridge job to reduce the work burden, as the time remaining was shorter. Fig. 7 illustrates the proportion of the senior job-to-job transitions where both the number of working hours and the annual wage income in the new job was reduced by 25% or more compared to the previous job. As is evident, there was a slight reduction from 2010 to 2011 in the proportion of new senior worker job-to-job transitions that were bridge jobs for the private sector firms and seemingly a trend reversal, while the proportion for the public sector firms stayed fairly stable. This lends some support to the aforementioned hypothesis.

The reduced minimum LRA could also potentially reduce the incentives for job-to-job movements from private sector firms to public sector firms. This is also best illustrated by a hypothetical example. Consider a worker of age 55 in a private sector firm who wanted to retire as early as possible, but did not qualify for DI benefits. The worker could then apply for a job in a public sector firm, fulfill the eligibility criteria for public AFP benefits and thus retire at age 62 instead of age 67. After the reform, the same worker did not have the same incentive to search for jobs in the public sector due to the reduced minimum LRA in the private sector firms. While this could change the pool of applicants public sector jobs and thus challenge the quasi-experimental setup (as discussed in Section 3.4), it is worthwhile noting that there is a lack of sectoral overlap for the public sector and private sector firms (see Table 3 for the distribution of firms across industries). The lack of sectoral overlap in conjunction with largely non-overlapping job codes (see the Online Appendix) suggests that the work tasks of the employees in these firms do not overlap, which will limit the extent to which a senior worker can easily change employment across private and public sector firms. As such, the change in the incentives to perform job-to-job movements as a result of the reform seems limited.

5. Data and summary statistics

For the remainder of the paper, I consider the population of private sector firms and public sector firms as units of analysis, restricting the sample to firms with 5 or more full-time equivalents (FTEs) to avoid the hiring decisions of sole proprietorships and smaller firms. A new hire in year t is defined as an individual registered as an employee in a firm j in year t who in year $t-1$ was (i) registered as employed in another firm $k \neq j$, (ii) received some form of social security benefits, or (iii) was outside the labor force. As noted by Johansen (2013), the number of new hires in any given year will differ substantially depending on the definition of "another firm". I define another firm as a firm with a different establishment identifier (but not necessarily a different enterprise identifier) to avoid possible misclassification of job-to-job transitions (Johansen, 2013, p. 9).[19] An additional restriction imposed is that an establishment is defined as a hiring firm in year t only if less than 50% of the new hires in the establishment in year t come from the same enterprise (Lien, 2014). If a firm hires several new workers, but where more than 50% of these workers come from a different establishment identifier but with identical enterprise identifier, the firm is defined as a non-hiring firm and the new workers in the firm are not considered new hires.

In the empirical analysis I restrict attention to new senior hires, defined as workers between ages 50 to 61 at time of hire. This means that I only consider new senior hires where the distance to the minimum LRA is one year or more, regardless of the old-age public pension regime in place and the firm type in question (public or private) at the time of hire. The period of analysis is restricted to the period 2007 to 2014. The decision to analyze this period relates to the timing of the reform. The reform's basic principles were agreed upon in Parliament in May 2005 (Norwegian Ministry of Finance, 2005), the law was passed by Parliament in May 2009 (Norwegian Ministry of Labour and Social Inclusion, 2009) and it was implemented in January 2011. However, a law on mandatory occupational pension schemes was implemented in 2006 (Norwegian Ministry of Finance, 2006), which may serve to confound hiring propensities before 2007. A detailed description of how the data set used in the empirical analysis is constructed is deferred to the Online Appendix.

The resulting data set indicates that the public sector firms are larger than the private sector firms in terms of the number of full-time equivalents over the sample period, illustrated by the firm size distributions in Panel A in Fig. 8. The differences in firm sizes are also reflected in the average number of new hires for private sector firms and public sector firms, illustrated by the distributions of average number of new hires each year in Panel B in Fig. 8. In the empirical specifications in Sections 6 and 7, I take into account the substantial level differences in the number of new hires between the private sector firms and the public sector firms by weighting the regressions by the number of full-time equivalents in the year prior to hire.

[19] The results in Section 6.2 and Section 7.2 when restricting the definition of "another firm" to both a different establishment identifier and a different enterprise identifier are qualitatively the same. The results are available from the author upon request.

Table 4
Summary statistics of new hires.

	Private sector firms	Public sector firms
Number of new hires in total	491,145	627,010
Number of unique firms	33,679	18,961
Prop. of new hires that are senior	0.129	0.174
	[0.336]	[0.379]
Education level		
Compulsory	0.432	0.291
High school	0.287	0.203
College	0.281	0.506
Pre-hire state		
Sick-leave benefits	0.064	0.082
Work assessment allowance	0.043	0.074
Disability benefits	0.040	0.070
Financial assistance program	0.012	0.010
Unemployment benefits	0.102	0.058
Outside labor force	0.198	0.115
Employed	0.541	0.590

Source: Author's own calculations using data from Statistics Norway.
Note: Summary statistics of new hires in private sector firms and public sector firms with more than 5 FTEs in the year prior to hiring, 2007–2014. Numbers in square brackets denote the standard deviations. Education level and pre-hire state refers to the new senior hires (workers aged 50–61 at time of hire). If the senior worker is not employed in year $t-1$, the state in year $t-1$ is defined as the last observed state in the three months prior to the employment start date in year t. Summary statistics by pre- and post-reform period, more detailed worker-level summary statistics and the evolution of mean real hourly wages are deferred to the Online Appendix.

From Table 4, we find that there were more new hires in the public sector firms than in the private sector firms. At the same time, there are about half as many public sector firms (18,961) as there are private sector firms (33,679), reflecting the pattern illustrated in Panel B in Fig. 8. Of the new hires, public sector firms have a greater proportion that are senior compared to the private sector firms (17.4% versus 12.9%). 50.6% of the senior workers hired in public sector firms have college education while only 28.1% of senior workers hired in private sector firms have college education, mirroring the distribution of private and public sector firms across industries in Table 3. A greater proportion of the senior workers hired in private sector firms come from unemployment benefits or were outside the labor force in the year prior to hire compared to the senior workers hired in public sector firms (10.2 and 19.8% versus 5.8 and 11.5%). Common for both types of firms is that most of the new senior hires are job-to-job transitions (54.1% for private sector firms and 59.0% for public sector firms).

6. Baseline methodology

In this Section, I introduce the baseline difference-in-difference linear probability models used to estimate the effect of the old-age public pension reform on the hiring of senior workers. The estimation results from the baseline specification are presented, and several specification tests are carried out.

6.1. Empirical specification

The ideal unit of analysis to identify the effect of changing the minimum LRA on the hiring of senior workers would be detailed information

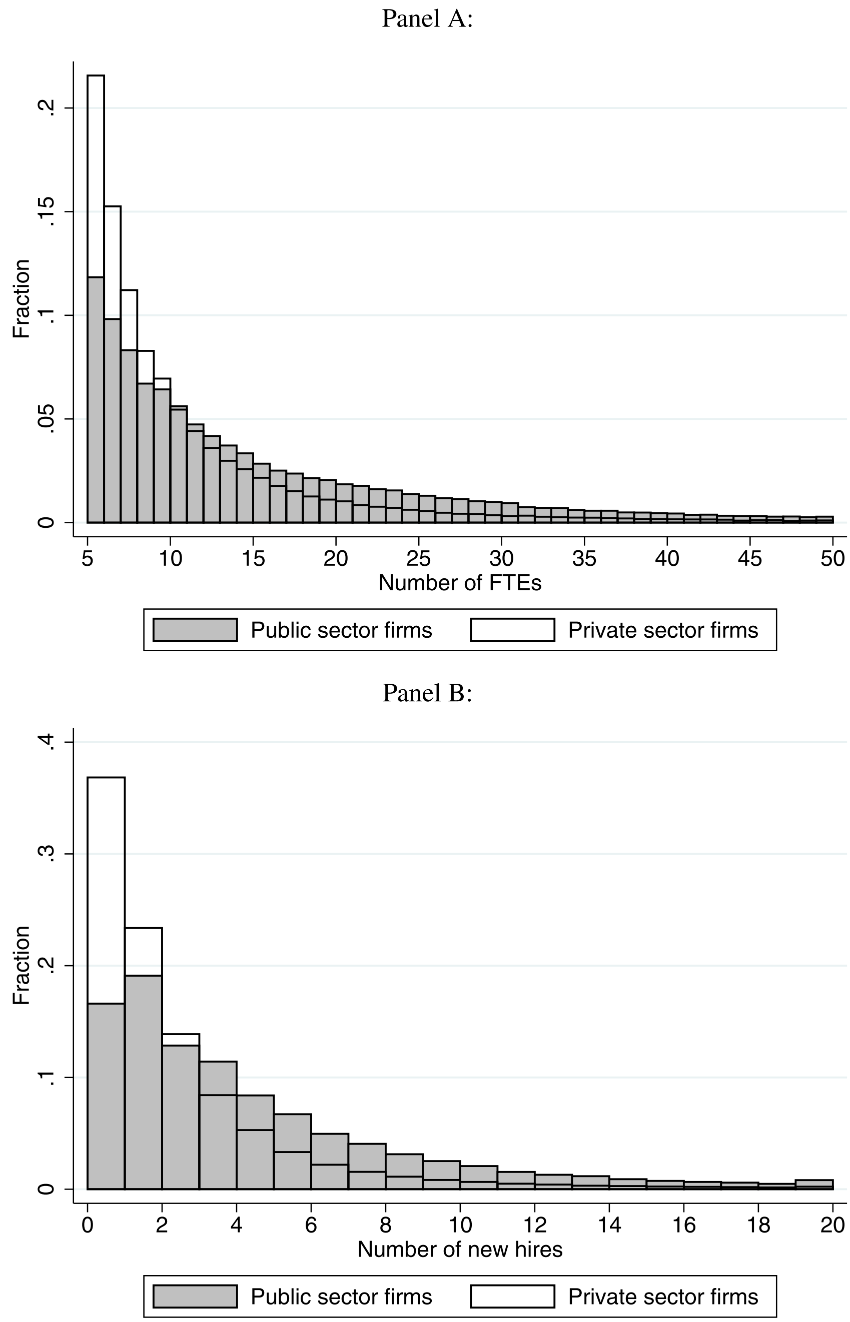

Fig. 8. Distribution of full-time equivalents (FTEs) and new hires, 2007–2014. Source: Author's own calculations using data from Statistics Norway.
Note: Number of full-time equivalents (Panel A) and new hires (Panel B) in private sector firms (hollow bars) and public sector firms (gray bars), 2007–2014. The distribution of full-time equivalents (Panel A) is here bounded over the interval [5,50] for ease of exposition, while the distribution of average number of new hires (Panel B) is bounded over the interval [0,20] for ease of exposition. See the Online Appendix for further details on how FTEs are measured. The sample of new hires is restricted to workers aged 30 to 75 at time of hire.

about applicants to a posted vacancy, as we could then observe both the realized firm-worker match and the discarded applicants. However, this data is not readily available. As a result, the primary unit of analysis in the empirical model is a firm (either a private sector firm or a public sector firm) that either hires or does not hire a senior worker. As a starting point, I estimate the following conventional difference-in-difference linear probability model:

$$Y_{jt} = \alpha_j + \beta T_j R_t + \delta T_j + \lambda R_t + \varepsilon_{jt} \quad (1)$$

In Eq. (1), the outcome variable Y_{jt} is a dummy variable equal to one if firm j hires at least one worker of age $a \in [50, 61]$ in year t, and zero otherwise. T_j is a treatment variable equal to one if firm j is a private sector firm (treated firms), and zero for public sector firms (non-treated firms). R_t is an indicator variable equal to one if year t is in the post-reform period (2011–2014) and zero in the pre-reform period (2007–2010). Furthermore, α_j is a fixed effect for firm j. The difference-in-difference coefficient β measures the reform effect of decreasing the minimum LRA by five years on the (unconditional) probability of at least one new senior hire (measured in percentage points). The resulting point estimate for the aggregate reform effect ($\hat{\beta} = 0.0081$, SE($\hat{\beta}$) = 0.0008) indicates that the reform lead to a 0.81 percentage point increase in the probability of hiring a senior worker (or a 16.3% increase in the probability of hiring a senior worker relative to the pre-reform mean of 0.0497). In terms of the reform's effect on the employment/non-employment margin of senior workers, I estimate Eq. (1) with the *number* of new senior hires coming from one of the non-employment states (sick-leave benefits, work assessment allowance, disability benefits, financial assistance program, unemployment benefits or outside labor force) as the dependent variable. The estimated aggregate reform effect ($\hat{\beta} = 0.0037$, $\text{SE}(\hat{\beta}) = 0.0008$) suggests that the reform led to a 0.37 percentage point increase in the employment of senior workers (equivalently a 14.3% increase relative to the pre-reform mean of 0.0258) due to increased hiring of senior workers from the non-employment states.

To investigate if there was an age-differential effect of the reform, as we may expect if the firm cares about the distance to retirement when hiring senior workers (see Section 4), I estimate the following age-specific difference-in-difference specification, again formulated as a linear probability model:

$$Y^a_{jt} = \alpha_j + \beta T_j R_t + \delta T_j + \lambda R_t + \epsilon_{jt} \tag{2}$$

In Eq. (2), the outcome variable Y^a_{jt} is a dummy variable equal to one if firm *j* hires at least one worker of age $a \in \{50, \ldots, 61\}$ in year *t*, and zero otherwise. T_j is a treatment variable equal to one if firm *j* is a private sector firm (treated firms), and zero for public sector firms (non-treated firms). R_t is an indicator variable equal to one if year *t* is in the post-reform period (2011–2014) and zero in the pre-reform period (2007–2010). Furthermore, α_j is a fixed effect for firm *j*. The difference-in-difference coefficient β now measures the reform effect of decreasing the minimum LRA by five years on the (unconditional) probability of at least one new senior hire of age *a* (measured in percentage points).

6.2. Age-specific estimation results

The results of estimating the baseline difference-in-difference specification in Eq. (2) separately for ages $a \in \{50, \ldots, 61\}$ are illustrated in Panel A of Fig. 9. The results indicate that the reform had a positive effect on the probability of a new senior hire, but that this positive effect was mainly restricted to workers aged $a \in \{50, \ldots, 57\}$ and generally decreasing with age. For workers aged 50, the increase in the probability of a new worker-firm match due to the reform is estimated to be 1.1 percentage points, whereas for workers aged 54 the reform effect is estimated to be 0.9 percentage points.

Due to the large differences in firm sizes and levels of hiring for the treated and the non-treated firms (illustrated in Panel A and Panel B in Fig. 8), I also estimate the same baseline difference-in-difference specification in Eq. (2) separately for ages $a \in \{50, \ldots, 61\}$ where I weight the regressions by the lagged number of employees (measured as full-time equivalents). In short, this procedure weights up the hiring decisions of the larger firms relative to the hiring decisions of the smaller firms (see the Online Appendix for the weighting formula used). The same age-differential pattern of reform effects emerges, but the estimated reform effects and their associated standard errors change, indicating that the firm size affects the propensity to hire senior workers. Notably, weighting the empirical specification in Eq. (2) leads to less significant effects compared to the unweighted specification, with the statistically significant positive reform effect now mainly restricted to workers aged $a \in \{51, \ldots, 55\}$. This is illustrated in Panel B of Fig. 9.[20]

Extending the baseline specification in Eq. (2) to include ages $a \in \{62, \ldots, 66\}$ allows us to investigate whether the reform had any effect on the hiring of workers who could retire immediately after the implementation of the reform. The estimation results for the unweighted and the weighted version of Eq. (2) for the ages $a \in \{50, \ldots, 66\}$ are also shown in Panel A and Panel B of Fig. 9, respectively. The results indicate that the reform generally had a small (and statistically insignificant) effect on the hiring probabilities for ages 62 to 66. The lack of reform effects for workers aged 62–66 could owe to the horizon effect (see Section 4) being a salient consideration for firms when hiring workers in this age range, as the amortization period for these workers is fairly short. As such, the choice of age 61 as upper age bound on the main sample of senior workers in the empirical analysis is not wholly inappropriate.[21]

These results, in combination with the reduced proportion of job-to-job transitions that are bridge job transitions illustrated in Fig. 7, should imply that the absolute size of the senior worker labor force in the treated firms (private sector firms) increases as a result of the reform. The increase in the size of the senior worker labor force (measured both by the number of senior workers and the number of senior full-time equivalents) in private sector firms from 2010 to 2011 illustrated in Fig. 10 confirms this.

From Fig. 9, it is evident that the pension reform had a differential effect on hiring across age groups, which indicates that the *distance* to the minimum LRA (and not only the *level* of the minimum LRA) is important in understanding the change in the firms' propensity to hire senior workers. However, it is important to emphasize that the illustrated age-specific reform effects are not necessarily orthogonal to each other, which may complicate the otherwise clear age-differential reform effect pattern shown in Fig. 9. For instance, the (strong) reform effect for age 52 is not necessarily independent of the (weaker) reform effect for age 53. This is because a firm's decision to hire a worker of age 52 is not necessarily independent of the decision not to hire a worker of age 53. Going beyond the assumption of age-orthogonality of the reform effects and investigating whether the distance to the minimum LRA matters for the probability of a new senior hire is studied using the extended empirical specification in Section 7.

6.3. Specification checks

6.3.1. Parallel trends

A critical assumption behind the difference-in-difference approach is the parallel trends assumption. The parallel trends assumption states that in the absence of the reform, the difference in the probability of a new senior hire between private sector firms (treated) and public sector firms (control) would have been the same in both the pre-reform period and the post-reform period (Angrist and Pischke, 2009, p. 230). To visually investigate if this counter-factual plausibly holds for the empirical specification in Eq. (2), Fig. 11 shows the (unweighted) mean of the outcome variable Y^a_{jt} in the pre-reform period (2007–2010) and the post-reform period (2011–2014) for all ages $a \in \{50, \ldots, 61\}$ pooled together.[22] From Fig. 11, the assumption of parallel trends in the outcome variable prior to the reform seems fairly reasonable. However, from 2009 the probability of a new senior hire in private sector firms seems to follow a slightly different trend than that of the public sector firms. This change in the trend for the treatment group could indicate the presence of an announcement effect in 2009, which would concord well with the passing of the old-age public pension reform in Parliament in May 2009 (Norwegian Ministry of Labour and Social Inclusion, 2009).

In order to test if there was an announcement effect as Fig. 11 suggests, I estimate the specification in Eq. (2) but (counter-factually) impose that the reform took place in 2009. For this placebo model, I therefore redefine R_t in Eq. (2) to be an indicator variable equal to one if year *t* is in the placebo post-reform period (2009–2014) and zero in the

[20] The estimation results from an extended version of the baseline specification in Eq. (2) where time-fixed effects are added are deferred to the Online Appendix. For both the unweighted and the weighted version of Eq. (2) with time-fixed effects, the estimation results reveal (qualitatively) the same age-differential pattern as illustrated in Panel A and Panel B in Fig. 9.

[21] The results of estimating the baseline specification in Eq. (2) for younger workers (aged $a \in \{41, \ldots, 49\}$) are deferred to the Online Appendix. The estimation results from the unweighted version of Eq. (2) show a generally upward-sloping reform effect over the age interval $a \in \{44, \ldots, 49\}$, while the reform effects are statistically insignificant in the weighted version.

[22] Figures showing the trends in the (unweighted) mean of the outcome variable (Y^a_{jt}) for the ages $a \in \{50, \ldots, 61\}$ separately in the pre-reform period (2007–2010) and the post-reform period (2011–2014) are available from the author upon request.

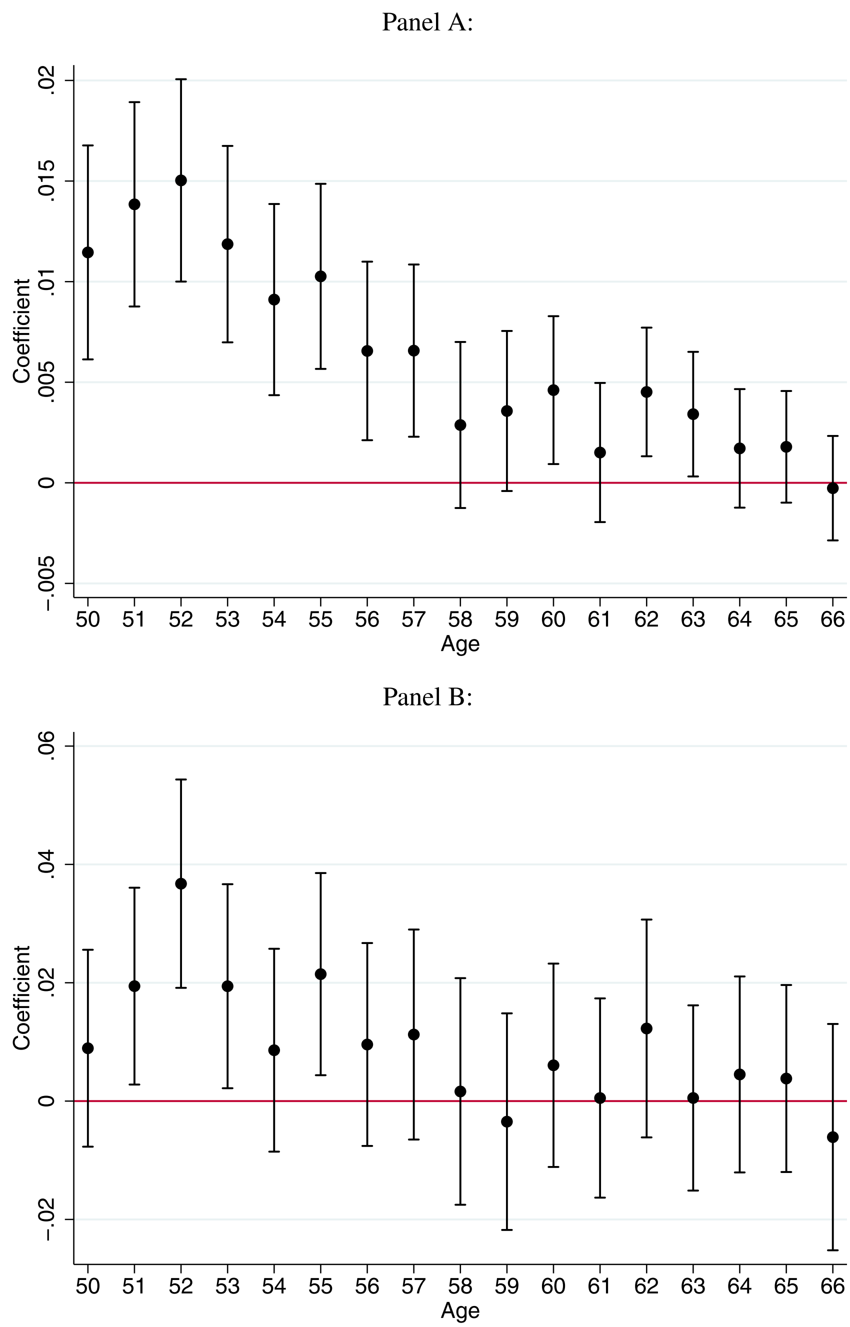

Fig. 9. Estimation results, baseline specification. Source: Author's own calculations using data from Statistics Norway. Note: Difference-in-difference estimates ($\hat{\beta}$) and the associated 95% confidence intervals (caped lines) from estimation of the baseline difference-in-difference specification in Eq. (2) separately for ages $a \in \{50, \dots, 66\}$. Panel A shows unweighted estimation results, Panel B shows weighted estimation results. Standard errors (robust) are clustered on the firm level. Point estimates and associated standard errors are relegated to the Online Appendix.

placebo pre-reform period (2007–2008). In other words, I pretend the reform took place in 2009 (the year of Parliament passing the reform) instead of 2011 (the year of the actual implementation) and estimate the age-specific reform effects. The estimation results from the placebo version of Eq. (2) shown in Panel A (unweighted) and Panel B (weighted) in Fig. 12 largely confirm the absence of statistically significant announcement effects (except for ages $a = 52$ and $a = 55$ in the unweighted version). The absence of statistically significant pre-reform effects lends support to the causal interpretation of the estimation results reported in Fig. 9.[23]

6.3.2. Linear time trends

Another specification check involves investigating if the reform correlates with other time trends within the treatment and control group, which may serve as a threat to disentangling the causal effect of the reform from other underlying (linear) time trends. To investigate this, I estimate a version of Eq. (2) where I also include a linear group-specific time trend:

$$Y^a_{jt} = \alpha_j + \beta T_j R_t + \delta T_j + \lambda R_t + T_j \times t + \varepsilon_{jt} \tag{3}$$

[23] The same placebo exercise when counter-factually imposing that the reform took place in 2008 (the year prior to Parliament passing the reform) instead of 2011 (the year of implementation) yields statistically insignificant reform effects for all ages $a \in \{50, \dots, 61\}$ in both the unweighted and weighted version. The estimation results are deferred to the Online Appendix.

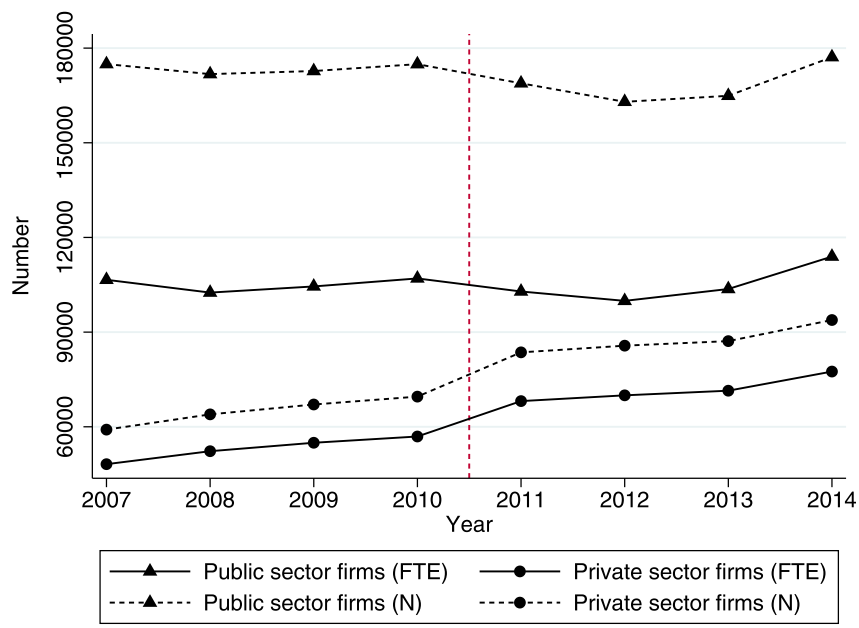

Fig. 10. Number of senior workers and senior full-time equivalents, 2007–2014. Source: Author's own calculations using data from Statistics Norway. Note: Total number of workers (N) and total number of full-time equivalents (FTE) in the age range 50–61 in private sector firms and public sector firms with more than 5 FTEs, 2007–2014. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011. Full-time equivalents are defined in the Online Appendix.

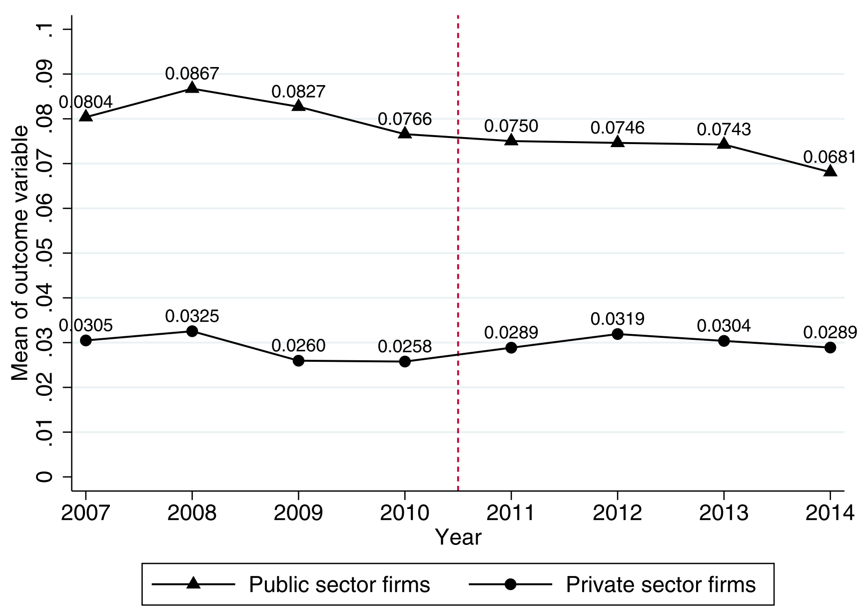

Fig. 11. Parallel trends assumption, all ages pooled. Source: Author's own calculations using data from Statistics Norway. Note: Mean of the outcome variable Y^a_{jt} (unweighted) in private sector firms (treated) and public sector firms (control) for workers in all ages $a \in \{50, \ldots, 61\}$ pooled together, 2007–2014. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011.

In Eq. (3), the variables Y^a_{jt}, T_j, R_t and α_j are defined as in the baseline specification in Eq. (2). The addition of the interaction between the treatment status and a linear time trend ($T_j \times t$) controls for other (linear) time trends within each group that may confound the reform effect. If the results are driven by different trends across the groups, then including the parametric time trend should change the estimated reform effects. The results from estimating Eq. (3) are illustrated in Panel A (unweighted) and Panel B (weighted) in Fig. 13. The estimation results generally show only slightly changed reform effects compared to the estimation results in Fig. 9 (but with increased standard errors), supporting the notion that the results in Fig. 9 could reflect the reform effect on the age-specific probability of a new senior hire.

7. Extended methodology

The estimation results from Section 6 indicate that the distance to the minimum LRA matters for the hiring of senior workers. In this Section, I introduce the extended specification used to estimate the effect of decreasing the distance to the minimum LRA by one year on the hiring of senior workers. The estimation results are presented and specification tests are carried out. I also investigate if the effect differs by the pre-hire

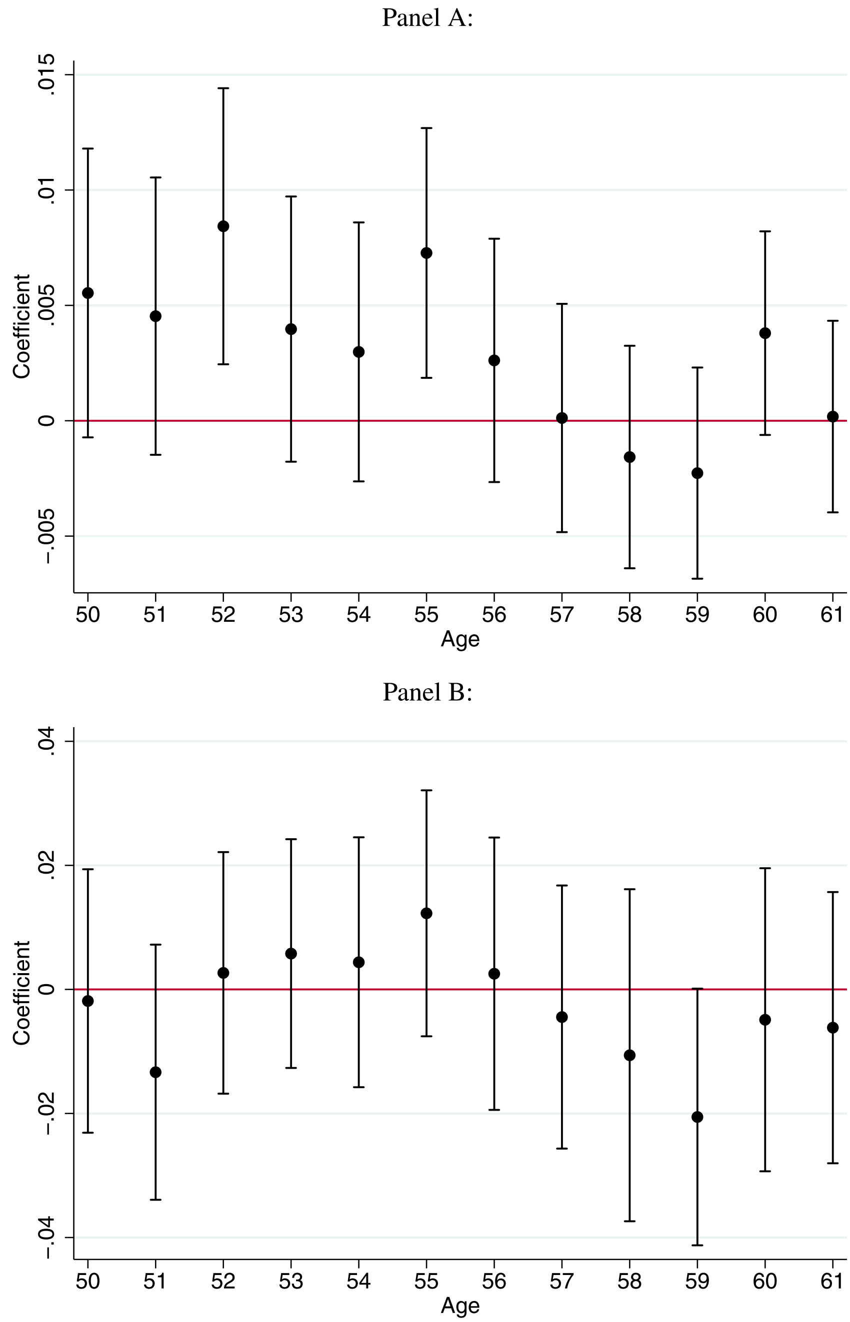

Fig. 12. Estimation results, placebo regression. Source: Author's own calculations using data from Statistics Norway. Note: Difference-in-difference estimates ($\hat{\beta}$) and the associated 95% confidence intervals (caped lines) from estimation of the placebo specification in Eq. (2) separately for ages $a \in \{50, \ldots, 61\}$, where $R_t = 0$ if $t \in \{2007, 2008\}$ and $R_t = 1$ if $t \in \{2009, \ldots, 2014\}$. Panel A shows unweighted estimation results, Panel B shows weighted estimation results. Standard errors (robust) are clustered on the firm level. Point estimates and associated standard errors are relegated to the Online Appendix.

status of the senior worker, and estimate the specification with different measures of hiring.

7.1. Empirical specification

The fact that the difference-in-difference estimates in Fig. 9 indicate that there was an differential effect of the reform across ages lends support to the hypothesis that the distance to the minimum LRA is indeed a relevant metric when trying to understand the effect of changing the minimum LRA on the hiring of senior workers, as argued in Section 4. In order to go beyond estimating only the age-specific effect of decreasing the minimum LRA on the probability of a new senior hire, we can construct an age-year-specific function (D_{ajt}) that measures the distance to the minimum LRA. This function can then be used to estimate the effect of a one-year *increase* in the distance to the minimum LRA on the probability of a new senior hire. While the specification gives the effect of *increasing* the distance to the minimum LRA, I will interpret the results in terms of *decreasing* the distance to the minimum LRA as the data used comes from a reform decreasing the minimum LRA. Following the setup of Ilmakunnas and Ilmakunnas (2015), the extended specification for estimating the distance-to-retirement effect is the following linear probability model:

$$Y_{ajt} = \alpha_j + \beta D_{ajt} + \delta_a + \pi_t + \varepsilon_{ajt} \tag{4}$$

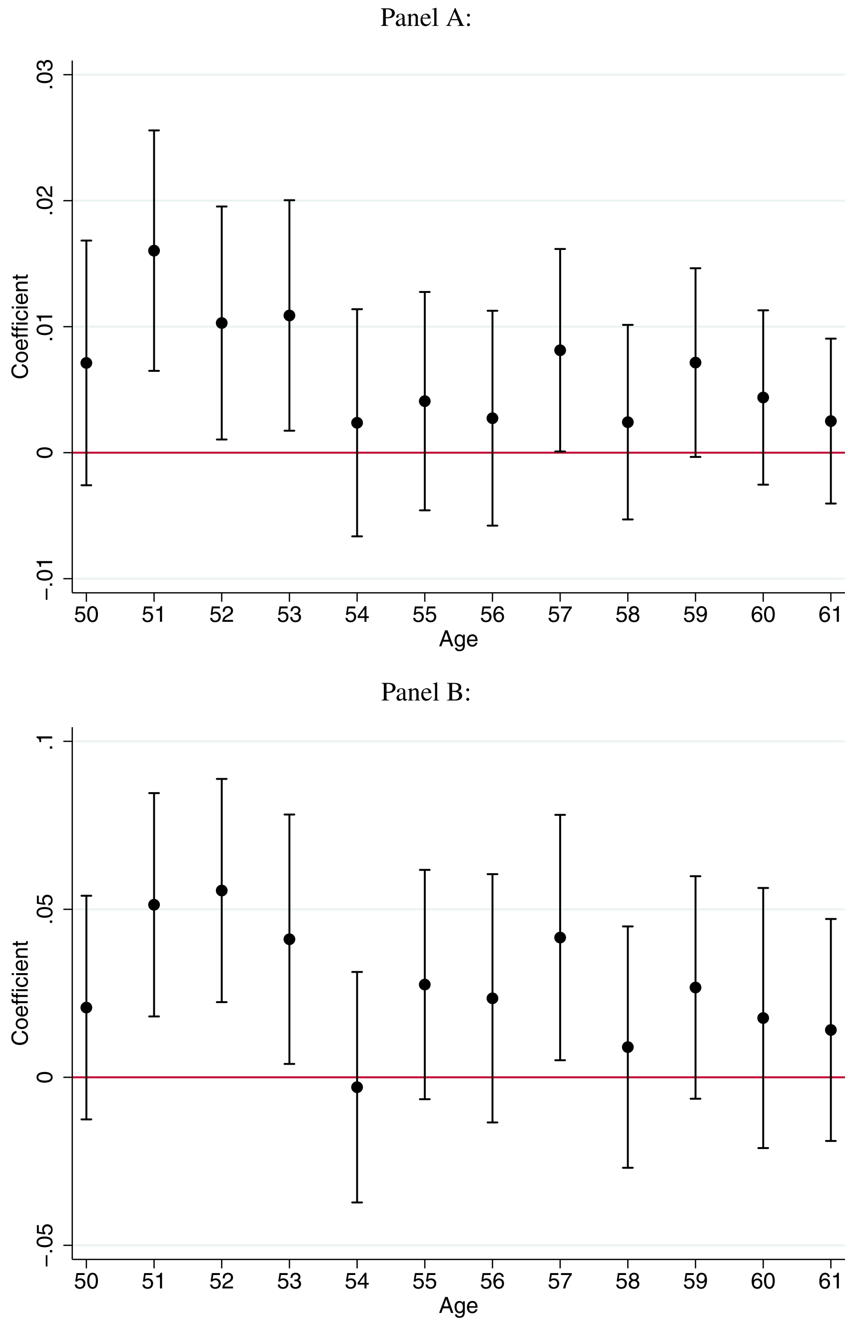

Fig. 13. Estimation results, linear time trend. Source: Author's own calculations using data from Statistics Norway. Note: Difference-in-difference estimates ($\hat{\beta}$) and the associated 95% confidence intervals (caped lines) from estimation of the baseline specification in Eq. (3) separately for ages $a \in \{50, \dots, 61\}$. Panel A shows unweighted estimation results, Panel B shows weighted estimation results. Standard errors (robust) are clustered on the firm level. Point estimates and associated standard errors are relegated to the Online Appendix.

In Eq. (4), the outcome variable Y_{ajt} is a dummy variable equal to one if firm j hires at least one worker in the age range $a \in [50, 61]$ in year t, and zero otherwise. D_{ajt} is a function measuring the distance to the minimum LRA for a worker of age a hired by firm j in year t, and is specified in Eq. (T1). δ_a is a fixed effect for age a, and π_t is a fixed effect for year t. Finally, α_j is a fixed effect for firm j.[24]

To establish that the coefficient β in Eq. (4) captures the causal effect of a one-year increase in the distance to the minimum LRA on the probability of a new senior hire, we need to identify the exogenous policy discontinuity in the function measuring the distance to the minimum LRA (D_{ajt}). This identification strategy is perhaps best illustrated using an example. Consider a private sector firm hiring a 58 year old worker. If the firm did so by hiring a worker born from a birth cohort $c \leq 1944$

[24] While using the linear probability model imposes a strong functional form assumption (especially the assumption of a constant marginal effect of the distance to the minimum LRA), it can provide reasonable estimates of the average effects (Wooldridge, 2010, p. 608). In addition, non-linear binary response models with fixed effects have limited practical applicability. For instance, deriving the average marginal effect of a one-year increase in the distance to the minimum LRA from a fixed-effect conditional logit model without encountering the incidental parameters problem requires very strong assumptions about the distribution of the fixed effect.

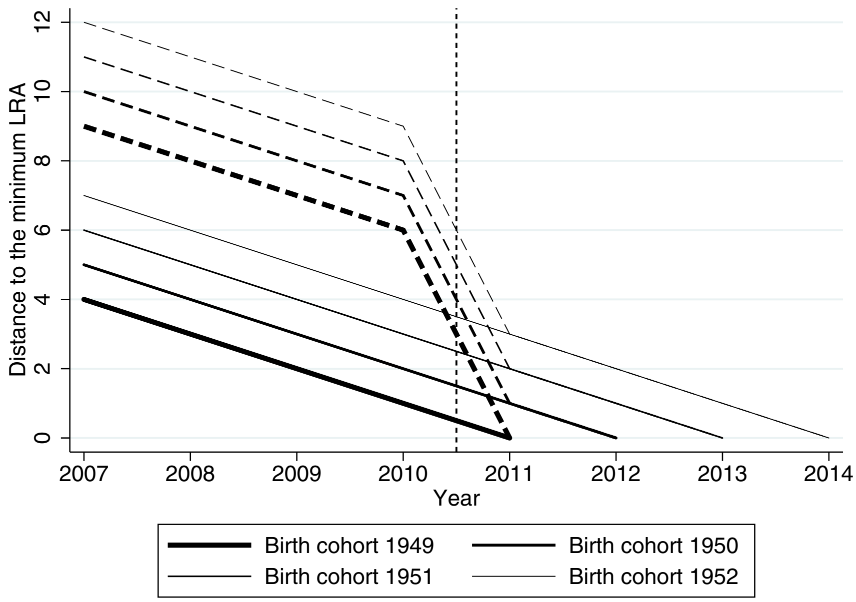

Fig. 14. Distance-to-retirement functions (D_{ajt}), 2007–2014. Source: Author's own calculations. Note: Distance to the minimum LRA (D_{ajt}) for workers in private sector firms (dashed line) and public sector firms (solid line) over the period 2007–2014, separately for the worker birth cohorts 1949–1952. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011. Examples of the distance-to-retirement function D_{ajt} for various worker birth cohort-year combinations in private sector firms and public sector firms are shown in the Online Appendix.

in year t, the earliest access to the old-age public pension benefits was 9 years later, regardless of the time of hire (t). If the firm did so by hiring a worker from a birth cohort $c \in \{1945, \dots, 1949\}$ in year t, the worker's earliest access was $2011 - t$ years later (meaning the distance to the minimum LRA was dependent on the distance to the reform). If the firm instead hired a 58 year old worker from a birth cohort $c \geq 1950$ the earliest access was 4 years later, regardless of the time of hire (t). A 58 year old worker hired in an public sector firm would in contrast have earliest access to the AFP benefits 4 years later, regardless of the birth cohort of the worker (c) and the time of hire (t).[25] The function D_{ajt}, which we can call the "distance-to-retirement" function, is therefore defined as follows for the private and public sector firms:

	$t < 2011$	$t \geq 2011$
Private sector firms	$\max\{0, 67 - a\}$	0 if $c < 1950$ $\max\{0, 62 - a\}$ if $c \geq 1950$
Public sector firms	$\max\{0, 62 - a\}$	$\max\{0, 62 - a\}$

(T1)

To illustrate, Fig. 14 shows the distance-to-retirement functions (D_{ajt}) for private sector firms and public sector firms for the worker birth cohorts 1949–1952. The exogenous variation in D_{ajt} brought about by the reform discontinuity for private sector firms in Fig. 14 demonstrates that the coefficient β is an estimator for the effect of a unit increase in the distance to the minimum LRA on the probability of a new senior hire.[26] In summary, the identification strategy used here entails isolating the effect of the discontinuity in the distance-to-retirement function (D_{ajt}) on the probability of a new senior hire.

7.2. Estimation results

The results from estimating the empirical specification in Eq. (4) are reported in Column (1) and Column (2) in Table 5. In line with the results presented in Section 6.2, a decrease in the distance to the minimum LRA increases the probability of a new senior hire. Comparing the non-weighted results in Column (1) to the weighted results in Column (2), the magnitude of the estimated effect of decreasing the distance to the minimum LRA by one year on the probability of a new senior hire changes from 2.9% ($100 \times (0.0015/0.0497) = 2.9$) to 3.2% ($100 \times (0.0016/0.0497) = 3.2$).[27] As expected from the differences in the firm sizes and the average number of new hires shown in Panel A and Panel B in Fig. 8, weighting the estimation yields much higher explanatory power of the specification in Eq. (4). Assuming a linear effect of the distance to the minimum LRA, these results imply that the reform led to an increase in the probability of a new senior hire of 14.5 to 16.0% (calculated as $2.9\% \times 5$ years $= 14.5\%$ and $3.2\% \times 5$ years $= 16.0\%$, respectively). Including the hiring of senior workers aged 62–66 in the estimation of Eq. (4) shifts up the estimated effects relative to those reported in Table 5 (see the Online Appendix for complete estimation results). Seemingly in contrast to the generally insignificant estimation results for these ages reported in Section 6.2, this could be a result of the age-specific effects not being independent of each other (but note however that including ages 62–66 in the estimation zero-inflates the distance-to-retirement function in Eq. (T1), which could explain some of the difference).

The effect of the distance to the minimum LRA could potentially correlate with other (linear) time trends within the treatment group (private sector firms) and the control group (public sector firms), which

[25] This example assumes that the senior worker in question has sufficient old-age public pension benefit accrual to retire at the minimum LRA (see Section 3.1 and 3.2 for details on the eligibility criteria). As individual old-age public pension benefit accrual is (generally) private information, the firm will (generally) have to proxy the actual eligibility age for each individual hire.

[26] A comparison of the coefficients from the baseline specifications in Section 6.2 and the coefficients from the extended specifications in Section 7.2 (using a comparable metric) is deferred to the Online Appendix.

[27] Estimating the extended specification in Eq. (4) for worker ages $a \in [41, 49]$ yields much lower estimated relative effects (and statistically insignificant for the weighted specification) of decreasing the distance to the minimum LRA by one year (with the estimated relative effects ranging from 0.6 to 1.8%). The results concord well with the argument presented in Section 4, namely that the horizon of the employment relationship until the minimum LRA is not a salient consideration for firms when hiring more junior workers. The complete estimation results are available from the author upon request.

Table 5
Estimation results, extended specification.

	(1)	(2)	(3)	(4)
Distance to minimum legal retirement age (D_{ajt})	−0.0015***	−0.0016**	−0.0016***	−0.0021***
	(0.0001)	(0.0007)	(0.0002)	(0.0007)
Weighted by number of FTEs (lagged)	×	✓	×	✓
Group-specific linear time trend included	×	×	✓	✓
Number of unique firms (n)	52,640	52,640	52,640	52,640
Total size of panel ($n \times T$)	252,585	252,585	252,585	252,585
Number of obs. ($n \times T \times 12$)	3,031,020	3,031,020	3,031,020	3,031,020
Adjusted R^2	0.1149	0.3215	0.1149	0.3216
Pre-reform mean of Y_{ajt}	0.0497	0.0497	0.0497	0.0497

Source: Author's own calculations using data from Statistics Norway.
Note: Results from estimation of Eq. (4) and Eq. (5). The dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$, and zero otherwise. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T1). Estimation sample covers firms (private sector and public sector) with 5 FTEs or more in the prior year, over the period 2007–2014. Reported robust standard errors (in parentheses) are clustered on the firm level.
*** $p < .01$, ** $p < .05$.

may serve as a threat to disentangling the causal effect of the distance to the minimum LRA from other underlying time trends. Controlling for a group-specific linear time trend, the specification in Eq. (4) becomes:

$$Y_{ajt} = \alpha_j + \beta D_{ajt} + \delta_a + \pi_t + T_j \times t + \varepsilon_{ajt} \quad (5)$$

The variables in Eq. (5) are defined as in Eq. (4), but with the added interaction between treatment status and a linear time trend ($T_j \times t$). The estimation results from this exercise are presented in Column (3) and Column (4) in Table 5. Compared to Column (1) and Column (2), the estimates do not change very much when including the linear time trend, with the relative effect in Column (3) and Column (4) being 3.1% and 4.2%, respectively. This suggests that the reform did not significantly interact with other (linear) time trends within the treatment group and the control group, and strengthens the interpretation that the previous results show the effect of decreasing the distance to the minimum LRA by one year.

7.3. Heterogeneity by pre-hire status

We may expect that the effect of decreasing the distance to the minimum LRA has a heterogeneous effect depending on the status of the senior worker prior to being hired. For instance, we may expect the effect to be different for workers who are performing a job-to-job transfers versus workers who are on the fringe of the labor market, with the distance to the minimum LRA being a more salient feature for firms when hiring from the latter group. Performing the analysis separately for each of the pre-hire states (which is either sick-leave benefits, work assessment allowance, disability benefits, financial assistance program, unemployment benefits, outside labor force or employed) confirms this. The effect effect is the largest for workers coming from either the financial assistance program or from disability benefits, both in the unweighted and the weighted version of Eq. (5). The estimation results are shown in Panel A (unweighted) and Panel B (weighted) in Fig. 15, where the relative effects (measured as percentage changes relative to pre-reform mean) express the effect of *increasing* the distance to the minimum LRA by one year. While the results suggest that there is some heterogeneity in hiring of senior workers from different pre-hire states, the estimated effects should be interpreted with some caution as the number of hired workers from each state is often rather small and the scope for valid statistical inference is limited (see the Online Appendix for number of new senior hires by pre-hire status over the period 2007–2014).

7.4. Other measures of hiring

Estimating the linear probability model in Eq. (5) means that a great deal of information on the outcome of interest (the extent of new senior hires) is potentially discarded. To examine if the effect on the hiring of senior workers is concentrated on the extensive or intensive margin, I estimate Eq. (5) with the total number of new hires of age $a \in [50, 61]$ in firm j in year t (N_{ajt}) as the outcome variable:

$$N_{ajt} = \alpha_j + \beta D_{ajt} + \delta_a + \pi_t + T_j \times t + \varepsilon_{ajt} \quad (6)$$

In Eq. (6), the right-hand side variables are the same as in Eq. (5). The results from estimating the specification in Eq. (6) are reported in Column (1) in Table 6. I do not weight the specification in Eq. (6) by the number of FTEs in year $t-1$, as it would mechanically confound the effect of interest. The magnitude of the estimated effect relative to the pre-reform mean of the dependent variable is 2.9%, meaning that decreasing the distance to the minimum LRA by one year led to an increase in the number of new senior hires by 2.9%. The similarity between this estimate and the linear probability model estimate reported in Column (3) in Table 5 reflects that most of the effect of decreasing the distance to the minimum LRA concerns firms now hiring one senior worker instead of hiring no senior workers. In short, the effect of decreasing the distance to the minimum LRA is concentrated mainly on the extensive margin of hiring.

To take into account the substantial level differences in firm size and new hires illustrated in Panel A and Panel B in Fig. 8 while still using the unweighted linear count model in Eq. (6), I replace the outcome variable in Eq. (5) with the number of new hires of age $a \in [50, 61]$ in firm j in year t (N_{ajt}) relative to the total size of the firm (measured as the number of FTEs) in year $t-1$ ($N^{tot}_{j,t-1}$):

$$\frac{N_{ajt}}{N^{tot}_{j,t-1}} = \alpha_j + \beta D_{ajt} + \delta_a + \pi_t + T_j \times t + \varepsilon_{ajt} \quad (7)$$

In Eq. (7), the right-hand side variables are the same as in Eq. (5). The results from estimating the specification in Eq. (7) are reported in Column (2) in Table 6, and shows a relative effect of 2.4%. This means that decreasing the distance to the minimum LRA by one year led to an increase in the size-relative expansion of the firm with new senior workers by 2.4%. In summary, the results from Tables 5 and 6 show that decreasing the distance to the minimum LRA (i) leads to an increased probability of a new senior hire, (ii) that most of this effect is concentrated on the extensive margin of hiring, and (iii) that the effect is present for all firm sizes, but more so for smaller firms.[28] Given that the effect on hiring is mostly concentrated on the extensive margin of

[28] Estimating the baseline specification in Eq. (2) with the count outcomes yields the same qualitative conclusions. Figures of the estimation results (and associated tables with coefficients and standard errors) are deferred to the Online Appendix.

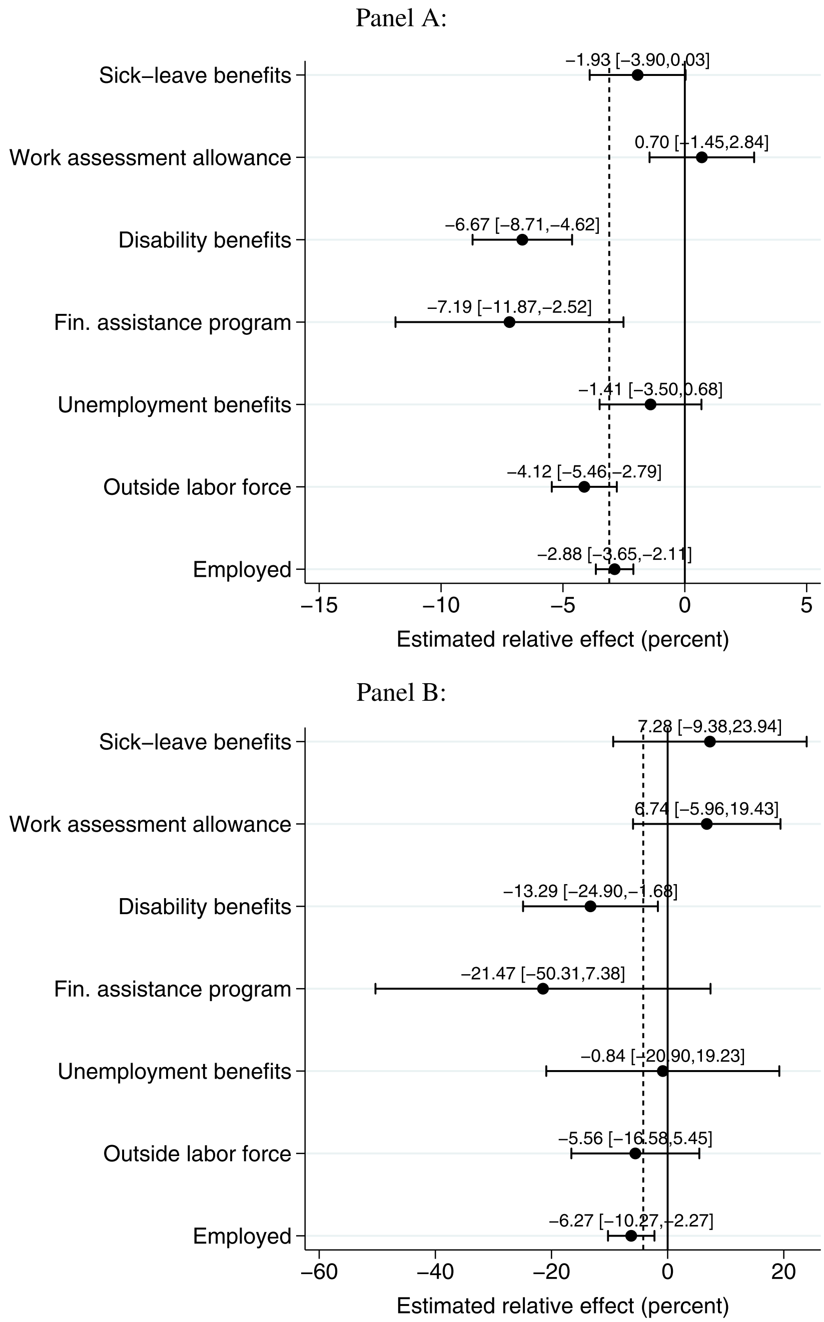

Fig. 15. Estimation results, heterogeneous effects. Source: Author's own calculations using data from Statistics Norway. Note: Results from estimating Eq. (5) separately for each pre-hire status (denote this k). The dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$ coming from pre-hire status k, and zero otherwise. Panel A shows unweighted estimation results, Panel B shows weighted estimation results. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T1). Estimation sample covers firms (private sector and public sector) with 5 FTEs or more in the prior year, over the period 2007–2014. The estimated effects and 95% confidence intervals (using robust standard errors clustered on firm level) are re-scaled to percentage changes relative to pre-reform mean, and is reported for each marker. The dashed vertical lines indicate the estimated effect for the full sample.

hiring, I maintain the linear probability model in Eq. (5) as the preferred specification.

7.5. Specification checks

To ensure that using the discontinuity in the distance to the minimum LRA (D_{ajt}) as the identifying assumption does not merely pick up a secular trend in firms' hiring propensities or is confounded by other institutional changes, a placebo test of the specification in Eq. (5) can be carried out. In order to do so, I impose the same discontinuity in the distance to the minimum LRA as described in Section 7.1, but I shift the age range of the new hires down from $a \in [50, 61]$ to $a \in [30, 41]$, while at the same time imposing the same reform as before. In other words, this placebo exercise entails counter-factually imposing the pension reform on younger workers. The strategy behind this exercise is best illustrated with an example. For workers in private sector firms, the earliest age at which they could access their old-age public pension benefits was decreased from age 47 to age 42 as a result of the reform. For workers in public sector firms, the earliest age at which they could access their AFP benefits stayed constant at age 42. For this exercise,

Table 6
Estimation results, extended specification, count outcome.

	(1)	(2)
Distance to minimum legal retirement age (D_{ajt})	−0.0017***	−0.00009***
	(0.0003)	(0.00002)
Type of dependent variable	Count	Relative count
	(N_{ajt})	($N_{ajt}/N^{tot}_{j,t-1}$)
Number of unique firms (n)	52,640	52,640
Total size of panel ($n \times T$)	252,585	252,585
Number of obs. ($n \times T \times 12$)	3,031,020	3,031,020
R^2 (adjusted)	0.2187	0.0471
Pre-reform mean of dep. variable	0.0590	0.0036

Source: Author's own calculations using data from Statistics Norway.
Note: Results from estimation of Eq. (6) and Eq. (7). In Column (1), the dependent variable (N_{ajt}) is the total number of new hires of age $a \in [50, 61]$ in firm j in year t. In Column (2), the dependent variable ($N_{ajt}/N^{tot}_{j,t-1}$) is the total number of new hires of age $a \in [50, 61]$ in firm j in year t relative to the size (measured as the number of FTEs) of firm j in year $t-1$. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T1). Estimation sample covers firms (private sector and public sector) with 5 FTEs or more in the prior year, over the period 2007–2014. Reported robust standard errors (in parentheses) are clustered on the firm level. *** $p < .01$.

Table 7
Estimation results, extended specification, placebo analysis.

	(1)	(2)
Distance to minimum legal retirement age (D_{ajt})	−0.0003	0.0003
	(0.0007)	(0.0009)
Type of placebo test	Sample of younger workers	Implementation in pre-reform year
Weighted by number of FTEs (lagged)	✓	✓
Number of unique firms (n)	52,640	52,640
Total size of panel ($n \times T$)	252,585	252,585
Number of obs. ($n \times T \times 12$)	3,031,020	3,031,020
Adjusted R^2	0.3839	0.3215
Pre-reform mean of Y_{ajt}	0.1271	0.0530

Source: Author's own calculations using data from Statistics Norway.
Note: Results from estimation of Eq. (5). In Column (1), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [30, 41]$, and zero otherwise. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T2). In Column (2), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$, and zero otherwise. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T1), but with the discontinuity occurring in 2009. Estimation sample covers firms (private sector and public sector) with 5 FTEs or more in the prior year, over the period 2007–2014. Reported robust standard errors (in parentheses) are clustered on the firm level.

the distance-to-retirement function (D_{ajt}) in Eq. (T1) is redefined as:

$$\begin{array}{l|cc} & t < 2011 & t \geq 2011 \\ \hline \text{Private sector firms} & \max\{0, 47 - age\} & \begin{cases} 0 \text{ if } c < 1970 \\ \max\{0, 42 - age\} \text{ if } c \geq 1970 \end{cases} \\ \hline \text{Public sector firms} & \max\{0, 42 - age\} & \max\{0, 42 - age\} \end{array} \quad \text{(T2)}$$

If the empirical strategy sufficiently isolates the effect of the reform, we should expect that estimating Eq. (5) on the sample of junior workers does not yield a statistically significant effect of decreasing the distance to the minimum LRA. The estimation results from this placebo exercise are shown in Column (1) in Table 7. The results show a statistically insignificant effect of decreasing the distance to the minimum LRA on the probability of a new junior hire (p-value = 0.725). Thus, the results from the placebo version of Eq. (5) reported in Column (1) in Table 7 lend support to the previous estimation results showing the effect of decreasing the minimum LRA on the hiring of senior workers.

Another test of the validity of using the reform to identify the effect of decreasing the distance to the minimum LRA by one year entails counter-factually imposing the reform discontinuity in a pre-reform year. In line with the placebo exercise from Section 6.3.1, I estimate the specification in Eq. (5) but impose that the reform took place in 2009 (and not in 2011). As no institutional change was implemented by that stage (only the passing of the reform in Parliament), the estimation results should yield non-significant effects of the distance to the minimum LRA. The estimation results from this placebo exercise are shown in Column (2) in Table 7. The statistically insignificant effect of decreasing the distance to the minimum LRA (p-value = 0.739) further supports the notion that the previous estimation results reflect the effect of decreasing the distance to the minimum LRA.[29,30]

[29] Estimating the extended specification in Eq. (5) using a more matched sample of treated and non-treated firms (matching on the proportion of workers in the firm that are senior workers, the proportion of workers with higher education, and the size of the firm measured in full-time equivalents) generally yields slightly larger effect magnitudes compared to the estimation results reported in Table 5. A full description of the matching procedure and complete estimation results are deferred to the Online Appendix.

[30] The same placebo exercise with the reform implemented in 2010 yields only a marginally significant effect (p-value = 0.098), suggesting that announcement effects in 2010 were also negligible. Imposing the pension reform in May 2009 (the time of passing of the reform in Parliament) yields a statistically insignificant effect (p-value = 0.582). Complete estimation results from these two placebo exercises are available from the author upon request.

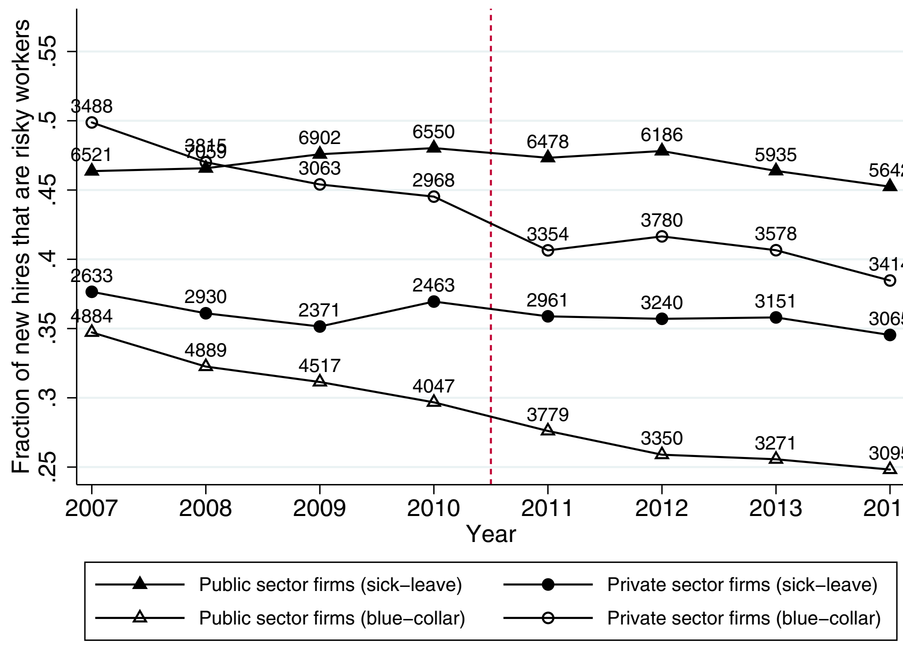

Fig. 16. Fraction and number of new "risky" senior hires, 2007–2014. Source: Author's own calculations using data from Statistics Norway. Note: The fraction (vertical axis) and the total number (marker) of senior hires that were "risky" workers (senior workers with sick-leave histories and/or blue-collar senior workers) in private sector firms and public sector firms with more than 5 FTEs in year prior to hire, 2007–2014. A senior worker is defined as a worker aged 50 to 61 at time of hire. The dashed vertical line indicates the implementation of the Norwegian old-age public pension reform in 2011.

Table 8
Estimation results, extended specification, risky workers.

	History of sick-leave		Blue-collar	
	(1)	(2)	(3)	(4)
Distance to minimum legal retirement age (D_{ajt})	−0.0012*	−0.0011**	−0.0011*	−0.0015**
	(0.0007)	(0.0005)	(0.0006)	(0.0007)
Weighted by number of FTEs (lagged)	✓	✓	✓	✓
History of sick-leave prior to hire	×	✓	–	–
Blue-collar worker	–	–	×	✓
Number of unique firms (n)	52,640	52,640	52,640	52,640
Total size of panel ($n \times T$)	252,585	252,585	252,585	252,585
Number of obs. ($n \times T \times 12$)	3,031,020	3,031,020	3,031,020	3,031,020
Adjusted R^2	0.2779	0.2538	0.3143	0.1720
Pre-reform mean of Y_{ajt}	0.0296	0.0234	0.0323	0.0201

Source: Author's own calculations using data from Statistics Norway. Note: Results from estimation of Eq. (5). In Column (1), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$ who had no recorded months of sick-leave in the three years prior to hire, and zero otherwise. In Column (2), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$ who had at least one month of sick-leave in the three years prior to hire, and zero otherwise. In Column (3), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$ who had completed high school or more, and zero otherwise. In Column (4), the dependent variable (Y_{ajt}) is an indicator variable equal to one in year t if firm j hired a worker in the age range $a \in [50, 61]$ who had only completed compulsory education, and zero otherwise. The distance-to-retirement function (D_{ajt}) is defined in Eq. (T1). Estimation sample covers firms (private sector and public sector) with 5 FTEs or more in the prior year, over the period 2007–2014. Reported robust standard errors (in parentheses) are clustered on the firm level.
** $p < .05$, * $p < .10$.

8. Mechanisms

To the extent that firms are risk-averse when making the decision to hire, the results in Table 5 indicate that decreasing the minimum LRA whilst at the same time imposing actuarial adjustment of pension benefits does act a safety valve for firms when hiring senior workers, subsequently increasing the propensity of firms to hire senior workers. As noted in Section 4, the hypothesized mechanism behind the increased hiring is that by reducing the minimum LRA one facilitates for selective attrition from employment (Boeri and van Ours, 2013, chap. 6), while the actuarial adjustment of benefits keeps the expected length of the employment relationship the same as before. While the estimation results point to that the reform generally stimulated the hiring of senior workers, it is also worth investigating if firms' risk aversion with respect to the potential lock-in effect could be a driving mechanism behind the increased hiring, as hypothesized in Section 4. To examine this mechanism, I consider the change in hiring of two groups of “risky” senior workers where the firm may be apprehensive of the potential negative lock-in effect. These two groups are: (i) senior workers with sick-leave histories and (ii) blue-collar senior workers. The first group of workers are workers with a history of sick-leave during the last three years prior to being hired, while the second group of workers are defined as workers that have completed only compulsory education. The fraction of and the number of new senior hires from these two groups in private sector and public sector firms over the period 2007–2014 are shown in Fig. 16.

To examine the effect of decreasing the distance to the minimum LRA on the hiring of workers from these two groups, I estimate the specification in Eq. (5) for each group separately. The dependent variable Y_{ajt} in Eq. (5) is now only equal to one if firm j hires a “risky” worker in the age range $a \in [50, 61]$ in year t, and zero otherwise. The results from estimation indicate that the firms’ risk-aversion could be the driving mechanism behind the estimation results in Section 7, since the positive effect of decreasing the distance to the minimum LRA is greatest for the hiring of senior workers with sick-leave histories and blue-collar senior workers – the “risky” senior workers. The estimation results for the first group of workers are shown in Table 8, where Column (1) shows the effect on hiring senior workers with no sick-leave history and Column (2) shows the effect on hiring senior workers with a sick-leave history. The relative effect for the “risky” workers is 4.6%, which is larger than the relative effect reported for the similar specification in Column (4) in Table 5. The relative effect for the “non-risky” workers is 3.9% and only marginally significant (p-value = 0.099), as well as being smaller than the relative effect reported for the similar specification in Column (4) in Table 5. The estimation results for the second group of workers are shown in Column (3) and Column (4) in Table 8, where Column (3) shows the effect on hiring white-collar senior workers (completed high school or more) and Column (4) shows the effect on hiring blue-collar senior workers (only compulsory education). The magnitude of the relative effect on the probability of a new senior blue-collar hire is substantially higher than the relative effect reported in Column (4) in Table 5, with the estimated effect in Column (4) in Table 8 being 7.4%. The relative effect on the hiring of white-collar senior workers is on the other hand estimated at 3.3%, a lower magnitude than for the sample as a whole.

The results in Table 8 therefore lend support to the idea that the increased hiring of senior workers when decreasing the distance to the minimum LRA is driven by firms being more willing to hire “risky” senior workers, with the largest effect on the hiring of blue-collar senior workers. To the extent that firms are aware of the educational gradient in the transition rate from good to bad health at older ages illustrated in Fig. 5, these results accord well with the argument (made in Section 4) that firms are risk-averse when hiring senior workers due to the potential negative lock-in effects.

9. Conclusion

If the intention of reforming old-age public pension systems is to extend the working lives of senior workers, policymakers can choose from a limited policy menu: (i) strengthen the incentives of continued work while simultaneously allowing for flexible pension benefit claiming, and (ii) increase the minimum legal retirement age (LRA). While several papers have studied the effect of changing the minimum LRA on individual labor supply (see e.g. Charni, 2016; Sánchez-Martin et al., 2014), there have been few studies of the effect on firms’ hiring behavior. The aim of this paper has been to investigate the consequences of decreasing the distance to the minimum LRA on the hiring of senior workers. By using a reduction of the minimum LRA from age 67 to age 62 as a result of the Norwegian old-age public pension reform in 2011, I am able to estimate the effect of decreasing the distance to the minimum LRA on the propensity of firms to hire senior workers. As a first pass, a difference-in-difference approach indicates that the reform increased the probability of hiring a senior worker (defined as a worker aged 50 to 61 at the time of hire) by 0.81 percentage points or 16.3% ($100 \times (0.0081/0.0497) = 16.3$). Using the discontinuity in the distance to the minimum LRA brought about by the reform, further estimation results show that, as a whole, the probability of hiring a senior worker increased by 4.2% ($100 \times (0.0021/0.0497) = 4.2$) when decreasing the distance to the minimum LRA by one year. Most of the effect is found to be on the extensive margin of hiring. Furthermore, in terms of the potential mechanisms behind the change in the propensity to hire senior workers, the relative effects of decreasing the distance to the minimum LRA are strongest for (i) senior workers with a prior history of sick-leave with an increase of 4.6% ($100 \times (0.0011/0.0234) = 4.6$), and (ii) blue-collar senior workers with an increase of 7.4% ($100 \times (0.0015/0.0201) = 7.4$). This lends support to the notion that risk-averse firms become more inclined to hire “risky” senior workers when the potential risk of doing so decreases (see e.g. Heywood and Jirjahn (2016)). In terms of the policy menu, increasing the minimum LRA could therefore have an unintended negative effect on the extent to which senior workers are hired. With a negative effect on hiring, the room for senior worker labor market mobility is subsequently reduced. While increasing the minimum LRA may lengthen the working lives of the senior workers, the side effect of reducing the firms’ propensity to hire could increase the extent of labor misallocation at older ages. As longevity increases and working lives become longer, the negative effect of such a policy on labor misallocation (and subsequently on productivity) may become ever more pronounced.

Supplementary material

Supplementary material associated with this article can be found, in the online version, at 10.1016/j.labeco.2018.01.004

References

Andersen, A.G., 2018. Technical note - pension. Working Paper, Ragnar Frisch Centre for Economic Research (forthcoming)
Angrist, J.D., Pischke, J.-S., 2009. Mostly Harmless Econometrics–An Empiricist' Companion. Princeton University Press, Princeton.
Barth, E., Moene, K.O., Willumsen, F., 2014. The Scandinavian model–an interpretation. J. Publ. Econ. 117, 60–72.
Behaghel, L., Crépon, B., Sédillot, B., 2008. The perverse effect of partial employment protection reform: the case of French older workers. J. Publ. Econ. 92, 696–721.
Boeri, T., van Ours, J., 2013. The Economics of Imperfect Labor Markets. Princeton University Press, Princeton.
Brinch, C.N., Vestad, O.L., Zweimüller, J., 2015. Excess Early Retirement? Evidence from the Norwegian 2011 Pension Reform.
Brunello, G., Langella, M., 2012. Bridge Jobs in Europe. IZA Discussion Paper No. 6938, Institute for the Study of Labor.
Bråthen, M., Bakken, F., 2013. Arbeid eller alderspensjon? [Work or old-age pension?]. Report 2013:1, Norwegian Labour and Welfare Service.
Börsch-Supan, A., Bucher-Koenen, T., Kutlu-Koc, V., Goll, N., 2017. Dangerous Flexibility–Retirement Reforms Reconsidered. Munich Center for the Economics of Aging.
Charni, K., 2016. The French Pension Reforms and their Impact on Unemployed Older Workers. Aix-Marseille School of Economics.
Dalen, E., 2015. Norsk Seniorpolitisk Barometer: Ledere The Norwegian Senior Policy Barometer Survey: Managers. Ipsos MMI.
Daniel, K., Heywood, J.S., 2007. The determinants of hiring older workers: UK evidence. Labour Econ. 14, 35–51.
Dorn, D., Sousa-Poza, A., 2010. ‘Voluntary’ and ‘Involuntary’ early retirement: an international analysis. Appl. Econ. 42, 427–438.
Dwyer, D.S., Mitchell, O.S., 1999. Health problems as determinants of retirement: are self-rated measures endogenous? J. Health Econ. 18, 173–193.
Gabaix, X., 2016. Behavioral Macroeconomics Via Sparse Dynamic Programming. NBER Working Paper No. 21848
Galaasen, A.M., Bruusgaard, D., Claussen, B., 2012. Excluded from social security: rejections of disability pension applications in Norway 1998–2004. Scand. J. Publ. Health 40, 142–149.
García-Gómez, P., 2011. Institutions, health shocks and labour market outcomes across Europe. J. Health Econ. 30, 200–213.
Geyer, J., Welteke, C., 2017. Closing Routes to Retirement: How Do People Respond?. IZA Discussion Paper No. 10681, Institute for the Study of Labor.
Hairault, J.-O., Sopraseuthm, T., Langot, F., 2010. Distance to retirement and older workers' employment: the case for delaying the retirement age. J. Eur. Econ. Assoc. 8 (5), 1034–1076.
Hallberg, D., Eklöf, M., 2010. Do buy-outs of older workers matter? Int. J. Manpower 31 (3), 337–359.
Hernæs, E., Markussen, S., Piggott, J., Røed, K., 2015. Pension Reform and Labor Supply: Flexibility vs. Prescription. NBER.

Hernæs, E., Markussen, S., Piggott, J., Røed, K., 2016. Pension reform and labor supply. J. Publ. Econ. 142, 39–55.
Hernandez, M., Pudney, S., Hancock, R., 2007. The welfare cost of means-testing: pensioner participation in income support. J. Appl. Econ. 22, 581–598.
Heyma, A., van der Werff, S., Nauta, A., van Sloten, G., 2014. What makes older job-seekers attractive to employers? Economist (Leiden) 162, 397–414.
Heyman, F., Skedinger, P., 2011. Employment Protection Reform, Enforcement in Collective Agreements and Worker Flows. Research Institute of Industrial Economics. IFN Working Paper No. 876
Heywood, J.S., Jirjahn, U., 2016. The hiring and employment of older workers in Germany: a comparative perspective. J. Lab. Market Res. 49, 349–366.
Heywood, J.S., Jirjahn, U., Tsertsvardze, G., 2010. Hiring older workers and employing older workers: German evidence. J. Popul. Econ. 23, 595–615.
Huber, M., Lechner, M., Wunsch, C., 2016. The effect of firms' phased retirement policies on the labor market outcomes of their employees. ILR Rev. 69 (5), 1216–1248.
Hurd, M.D., 1996. The effect of labor market rigidities on the labor force rigidities on the labor force behavior of older workers. In: Wise, D.A. (Ed.), Advances in the Economics of Aging. University of Chicago Press, pp. 11–60.
Hutchens, R., 1986. Delayed payment contracts and a firm's propensity to hire older workers. J. Lab. Econ. 4 (4), 439–457.
Huttunen, K., Møen, J., Salvanes, K.G., 2011. How destructive is creative destruction? Effects of job loss on job mobility, withdrawal and income. J. Eur. Econ. Assoc. 9 (5), 840–870.
Ilmakunnas, P., Ilmakunnas, S., 2015. Hiring older employees: do the age limits of early retirement and the contribution rates of firms matter? Scand. J. Econ. 117 (1), 164–194.
Ilmarinen, J., 2002. Physical requirements associated with the work of aging workers in the European union. Exp. Aging Res. 21 (1), 7–23.
IMF, 2016. Norway–Selected Issues. Technical Report, IMF Country Report No. 16/215. International Monetary Fund.
Johansen, I., 2013. Jobbskifter blant eldre arbeidstakere [Job Changes Among Older Employees]. Rapporter 60/2013, Statistics Norway.
Kantarci, T., 2013. Factors Limiting the Opportunities for Partial Retirement. Working Paper Series, Paper No: 2013-009, CESR-Schaeffer.
Kudrna, G., 2017. The Norwegian Pension Reform: An External Perspective. CEPAR Working Paper 2017/07, CEPAR..
Lazear, E.P., 1979. Why is there mandatory retirement? J. Polit. Econ. 87 (6), 1261–1284.
Lazear, E.P., 2011. Wages, productivity, and retirement. Int. Tax Publ. Finance 18, 17–35.
Legendre, B., Sabatier, M., 2017. The puzzle of older workers' employment: distance to retirement and health effects. Int. J. Manpower 38 (1), 45–61.
Lien, O.C., 2014. Få bytter jobb etter fylte 50 år [Few change jobs after the age of 50]. Arbeid og velferd 1, 37–50.
Manger, C., 2014. Endogenous age discrimination. J. Popul. Econ. 27, 1087–1106.
Manoli, D.S., Weber, A., 2016. The Effects of the Early Retirement Age on Retirement Decisions. IZA Discussion Paper No. 10154, Institute for the Study of Labor.
Martins, P.S., Novo, A., Portugal, P., 2009. Increasing the Legal Retirement Age: The Impact on Wages, Worker Flows and Firm Performance. IZA Discussion Paper No. 4187, Institute for the Study of Labor.
Montizaan, R., Fouarge, D., 2016. The Impact of Employer's Characteristics on the Willingness to Hire Older Workers: Evidence from a Stated Preferences Experiment. Netspar Discussion Paper 01/2016-054
Mühlemann, S., Leiser, M.S., 2015. Ten Facts You Need To Know About Hiring.
Norwegian Labour and Welfare Administration, 2016. Uføretrygd [Disability]. https://www.nav.no/no/NAV+og+samfunn/Statistikk/AAP+nedsatt+arbeidsevne+og+uforetrygd+-+statistikk/Uforetrygd
Norwegian Ministry of Finance, 2005. Pensjonsreform - trygghet for pensjonene [Pension Reform - Securing Pensions]. St.Meld. nr. 12 (2004–2005), St.Meld.
Norwegian Ministry of Finance, 2006. Om lov om obligatorisk tjenestepensjon [On The Mandatory Occupational Pensions Act]. Ot.prop. nr. 10 (2005–2006), Ot.prop.
Norwegian Ministry of Labour and Social Inclusion, 2009. Om lov om endringer i folketrygdloven (ny alderspensjon) [On Changing The National Insurance Act (New Old-Age Pension)]. Ot.prop. nr. 37 (2008–2009), Ot.prop.
OECD, 2013. OECD Employment Outlook 2013. Technical Report. Organisation for Economic Co-operation and Development.
OECD, 2015. Pensions at a Glance 2015. Technical Report. Organisation for Economic Co-operation and Development.
Oi, W.Y., 1962. Labor as a quasi-Fixed factor. J. Polit. Econ. 70 (6), 538–555.
Saint-Paul, G., 2009. Does the Welfare State Make Older Workers Unemployable? IZA Discussion Paper No. 4440, Institute for the Study of Labor.
Schumann, M., 2017. The effects of minimum wages on firm-financed apprenticeship training. Lab. Econ. 47, 163–181.
Skirbekk, V., 2004. Age and individual productivity: a literature survey. Vienna Yearb. Popul. Res. 2, 133–153.
Statistics Norway, 2008. Standard for næringsgruppering [Standard Industrial Classification]. Official Statistics of Norway D383, Statistics Norway.
Staubli, S., Zweimüller, J., 2013. Does raising the early retirement age increase employment of older workers? J. Publ. Econ. 108, 17–32.
Stier, H., Endeweld, M., 2015. Employment transitions and labor market exits: age and gender in the Israeli labor market. Res. Soc. Stratif. Mob. 41, 92–102.
Sánchez-Martin, A.R., García-Pérez, J.I., Jiménez-Martín, S., 2014. Delaying the normal and early retirement ages in spain: behavioural and welfare consequences for employed and unemployed workers. Economist (Leiden) 162, 341–375.
Vestad, O.L., 2013. Labour supply effects of early retirement provision. Lab. Econ. 25, 98–109.
Wooldridge, J.M., 2010. Econometric Analysis of Cross Section and Panel Data. MIT Press, Cambridge.