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Examining differences in bunching at statutory retirement ages in Australia and the Netherlands

Exploring the role of firms and executives

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Academic paper

Examining differences in bunching at statutory retirement ages in Australia and the Netherlands: Exploring the role of firms and executives *

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Abstract

In a cross-country analysis, we document stark differences in the retirement hazard at the Statutory Retirement Age (SRA) for Dutch workers (65%) and Australian workers (5%). To understand the role of firms and executives in explaining this disparity, we use comprehensive administrative data for both countries to construct monthly worker-firm-executive datasets and examine job separation rates at the SRA. Focusing on firms/executives employing multiple older workers, we perform a variance decomposition and estimate a series of regressions to quantify firm and executive effects. Consistently, the results show that firms and executives matter in both countries, though to a much greater extent in the Netherlands. Our estimates imply that moving a worker from a bottom quartile firm to a top quartile firm, in terms of SRA separation propensities, would raise the worker's separation probability at the SRA by 28 percentage points in the Netherlands, while the comparative figure in Australia is just 5 percentage points. We discuss possible institutional reasons for the differences across countries, including the system of automatic job termination at the SRA in the Netherlands and the more targeted public pension in Australia.

JEL Codes: H55, L22, J23, J26,
Keywords: Retirement, Labor demand, Managers, Firm heterogeneityCross-country comparison

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1 Introduction

Reducing barriers towards old-age employment is an important policy goal to address the demographic challenges associated with population aging. Many countries have implemented policies to incentivize workers to delay retirement, such as raising the pension age and scaling back early retirement pathways. While existing studies find that these supply-side interventions are effective in increasing the employment rates of older workers on average (see e.g., Manoli and Weber (2016); Staubli and Zweimüller (2013) among others), there are substantial differences in the responses to those reforms. Employers can influence rates of involuntary retirement (Dorn and Sousa-Poza, 2010), enforcement of mandatory retirement (Rabaté, 2019; Morris and Dostie, 2023) and the effectiveness of a pension reform (Bello and Galasso, 2020; Geyer et al., 2022). Despite the importance of demand-side interventions aiming at boosting old-age employment, little is known about the causal impact of firms on retirement outcomes.

In this paper, we study the phenomenon of retirement bunching in the form of a spike in job separations and retirement at the statutory retirement age (SRA) (Seibold, 2021; Lalive et al., 2023; Gudgeon et al., 2023; Rabaté et al., 2024). In particular, we study the phenomenon by comparing the amount of bunching in two countries: the Netherlands and Australia. This is motivated by the striking differences between these countries. In the Netherlands, the retirement hazard at SRA is approximately 70% (Rabaté et al., 2024), which translates into one of the highest incidence of bunching in the world. In contrast, there appears to be only a moderate amount of bunching at the corresponding threshold in Australia (Morris, 2022). Our analysis shows that in the month prior to the SRA, the estimated increase in the retirement hazard is 8.3 percentage points in the Netherlands compared to 1.0 percentage point in Australia. This disparity is even larger in the SRA month: 43.4 percentage points (Netherlands) and 3.4 percentage points (Australia). Our bunching estimates indicate that around 68% of Dutch workers who are still working near the SRA retire within one year of attaining eligibility for the public pension, while this figure is closer to 9% in Australia.

In addition, we document large variations in job separation rates at the SRA across firms in the Netherlands and Australia. We compare the distribution of firm-specific retirement hazards at SRA with a pseudo-distribution when we randomly reshuffling workers across the firms in our sample. This exercise indicates that, in the Netherlands, there is 150% more variation in the data than would be expected by chance. The difference is much smaller — around 25% — in Australia. To understand how much of this variation comes from industry-specific retirement hazards, we compare it with another pseudo-distribution when we randomly reshuffle workers across firms within the same industry. While industry-specific factors clearly explain some of the variation across firms, about half remains unexplained. Moreover, firms in the private sector exhibit especially large variations in job separation rates at the SRA.

We then ask the questions: What are the main drivers of the heterogeneity in retirement bunching? How much of a role do firms and managers play in determining retirement outcomes? Using high-quality population-wide linked employer-employee datasets, we first assess whether the extent of retirement bunching within individual firms can be predicted by a host of worker- and firm-level covariates, thus evaluating the relative importance of these attributes in explaining the variation in individual retirement behavior. We use a variance decomposition approach to quantify how worker, firm, sector, and executive factors explain variation in separation outcomes at the SRA, finding that firm and executive fixed effects account for significantly more variance than observable characteristics — especially in the Netherlands. In the Dutch private sector, executive fixed effects alone explain more than 20% of the variance, while in Australia all observable and unobservable components combined explain less than 10%. These results suggest that firms and executives play a much larger role in shaping retirement behavior in the Netherlands than in Australia.

Furthermore, using regression analyses, we show a robust correlation between a worker's job separation probability at the SRA and the leave-one-out mean of other workers at the firm, suggesting the presence of firm-specific effects. In the Netherlands, a 10 percentage point increase in coworkers' separation rates is associated with a 5.9 pp increase in an individual's own probability of separating at the SRA, compared to 2.7 pp in Australia. These results remain robust after accounting for observable controls, selection on unobservables (via Oster bounds), and potential measurement error, which tends to attenuate the estimates. Our estimates imply that moving a worker from a bottom-quartile firm to a top -quartile firm, in terms of SRA separation propensities, would raise the worker's separation probability at the SRA by 28 percentage points in the Netherlands, while the comparative figure in Australia is 5 percentage points.

Lastly, we show that the SRA separation outcomes are strongly correlated both within firms and across workers who reach the SRA under the same executive. We show that a 10 percentage point increase in the separation rates of coworkers' under the same executive is associated with a 3 pp increase in an individual's own probability of separating at the SRA, compared to 2.5pp in Australia. This result indicates that HR policies — such as those around mandatory retirement — are influenced by individual leaders and not fixed at the firm level. In both the Netherlands and Australia, executive-level effects are statistically significant, suggesting that retirement practices can vary meaningfully depending on who is in charge, but there appears to be much more scope for executives to matter in the Dutch context.

This paper connects to two strands of literature. First, it speaks to the growing literature studying the retirement decisions, with the new focus on the role of firms and managers. Most of the existing studies focus on the supply side determinants of retirement decision (see, e.g., Manoli and Weber (2016); Seibold (2021); Lalive et al. (2023)). By focusing on firm-level variations in retirement bunching at statutory retirement age, we complement existing work on firms' influence in workers' retirement decisions (e.g. Frimmel et al. (2018); Rabaté (2019); Deshpande et al. (2024); Morris and Dostie (2023); Saez et al. (2024)) and a broader literature that emphasizes the importance of firms on dismissal and unemployment outcomes (e.g., Van Doornik et al. (2023); Gudgeon et al. (2023); Carry and Schoefer (2024)).

Second, this paper links to the emerging studies on the role of CEOs and managers. Previous research finds that managers affect productivity (Bandiera et al., 2020), personnel choices (Hoffman and Tadelis, 2021) and wages (He and le Maire, 2022). Especially, recent papers have shown the importance of CEO and directors by exploring characteristics such as leadership quality (Lazear et al., 2015; Black, 2019), gender (Flabbi et al., 2019), managerial talent (Fenizia, 2022) and business education (Acemoglu et al., 2023). The dimension we study zooms in the dismissal and retention of older employees. We categorize directors according to their revealed preferences for older workers, proxied by the amount of job separations at the SRA. This allows us to provide a new data-driven way to examine the how managers shape retirement outcomes.

The remainder of the paper proceeds as follows: Section 2 describes the pension system in the Netherlands and Australia and the data sources. Section 3 compares the retirement bunching at SRA between the Netherlands and Australia. Section 4 studies firm's role in explaining the SRA bunching and compares the firm's impacts in these two countries. Section 5 concludes.

2 Background on Dutch and Australian pension systems

Dutch Pension System. The Dutch pension system consists of three pillars. The first pillar is a flat-rate public pension (*AOW, Algemene Ouderdomswet*), which is indexed to the minimum wage and financed by pay-as-you-go social insurance contributions. In 2024, this pension amounted to €1,497.77 per month for singles (€17,973 per annum) and €2,049.24 per month for couples (€24,591 per annum). All Dutch residents can receive the first-pillar pension once they reach their birth cohort-specific "AOW age" or Statutory Retirement Age (SRA). The SRA was 65 years old before 2013. It increased by one month per year in 2013–15, three months per year in 2016–18 and four months in 2019. For cohorts born between January 1948 and December 1955, the SRA increased from 65 to 67 (Figure 1a). Early claiming of the first-pillar pension is not possible. Moreover, individuals lose eligibility for unemployment and disability benefits after reaching the SRA.

The SRA constitutes the end point of most employment contracts, which means that individuals who wish to continue working past the SRA have to be offered a new contract. Similar "mandatory retirement" age thresholds are common policies in many OECD countries (Van Vuuren, 2014; Saez et al., 2024), particularly those with more stringent employment protection legislation, such as Germany, France and the Netherlands.

The second pillar consists of mandatory employer contribution schemes, which are collective pension schemes connected to a specific industry or company, capital-funded and managed by pension funds. Through the second-pillar schemes, employees can retire earlier (i.e., before the SRA). The third pillar, which consists of individual savings for retirement, is less important than in other developed countries (Grip et al., 2012).

Australian Pension System. The Australian pension system consists of roughly three pillars as well. The first pillar is the *Age Pension*, which is available to Australian residents (for at least 10 years) who satisfy a means test based on household income and assets. Eligible residents can claim the Age Pension at the Age Pension Age (APA), which we simply refer to as the Statutory Retirement Age (SRA). Before 1994, the SRA was 60 for women and 65 for men. The 1994 pension reform legislated that women's SRA would gradually increase to 65, rising by 6 months every 2 years from July 1995 to July 2013. For female cohorts born between July 1935 to January 1949, the SRA increased from 60 to 65. The 2009 reform raised the SRA again. This time for both men and women. The SRA was increased from 65 to 67 by 6 months every 2 years from July 1, 2017 to July 1, 2023. In other words, for cohorts born between July 1952 and December 1956, the SRA increased from 65 to 67 (Figure 1 b). After reaching the Age Pension age, eligible individuals cannot claim unemployment and disability benefits any more.

Compared with the Netherlands, Australia has relatively weak employment protection and no longer has mandatory retirement laws. Firms can dismiss workers after a notification period, which has a maximum period of five weeks for people older than 45 with more than 5 years of tenure. Therefore, SRA is not necessarily tied with end of employment contracts.

Most Australians still partially or fully fund their retirement via the Age Pension (Morris, 2022). In 2021, around two-thirds of people above the pension age received income from the Age Pension. Rates are higher among women and older cohorts, and around 50% of people claiming the Age Pension immediately when they reach the SRA (Oguzoglu et al., 2020). In 2024, the maximum basic payment was 1,149.00 AUD per fortnight for single pensioners (29,976 AUD per annum, or €17,032 based on 2025 exchange rates) and 1,732.20 AUD per fortnight for couples (45,192 AUD per annum or €25,677).

Thus, the maximum Age Pension payment is similar to the universal rate in the Dutch public pension. In addition, the access ages are now 67 in both countries — recently rising from 65 — and both ages signal the end of eligibility for unemployment and disability benefits. However, unlike the Dutch pension, Australia's Age Pension is means-tested based on the current income and assets of the household. Pensioners receive the payment specified by either the income test or the assets test, whichever is lower. Each of these tests has an initial phase-out threshold. If a household's income and assets are below the relevant thresholds, pensioners receive the maximum payment. Otherwise, payments are gradually reduced to zero. In 2016, around 60% of Age Pensioners received the maximum payment (Department of Social Services, 2016).[1]

The second pillar consists of a private mandatory defined-contribution scheme known as superannuation, which is assessable under the Age Pension means test. Since 1992, superannuation contributions for employees have been mandatory, with their employers currently required to contribute at least 12% on top of their pre-tax earnings into the employee's superannuation account. The mandatory contribution rate has risen over time, starting from 3%, which means that average balances have risen considerably over time, particularly among wealthier households. In June 2021, among 60–64-year-olds, the average balance was $402,838 among men (median = $211,996) and $318,203 among women (median = $158,806).[2] The minimum age to access the superannuation is between 55 and 60, depending on the birth cohorts.[3] Superannuation benefits are paid as either a lump-sum or as an income stream. The third pillar is voluntary savings.

Overall, the Age Pension and superannuation have similar importance in funding the retirements of Australians, but they serve very different segments of the population. The Age Pension primarily benefits low-income households, providing little or no benefit to high-income households due to the means test. Meanwhile, low-income households have little saved in superannuation, while high-income households typically have large superannuation balances. The cross-over point, where superannuation becomes more important than the Age Pension, is at about the 50th percentile.[4]

[1] The income test includes earned income (above a modest, allowed amount) and the assets test excludes the value of the primary residence. The taper rate of both tests is high: for each $1 of non-pension income, Age Pension payments are reduced by $0.50, implying a 50% implicit tax rate; for each $1,000 of assessable assets, payments are reduced by $3 per fortnight (equivalent to $78 per annum). The phase-out thresholds depend on whether the person is single or partnered, and a homeowner or renter in the case of the assets test. For a single person, the phase-out threshold is $314,000 for the assets test if the person is a homeowner ($566,000 for a non-homeowner) and $212 per fortnight ($5,500) for the income test. For couples, the phase-out thresholds are around 50% higher for the assets test and 75% higher for the income test.

[2] https://www.superannuation.asn.au/wp-content/uploads/2024/01/2311_An_update_on_superannuation_account_balances_Paper_V2.pdf

[3] It is 55 for people born before July 1st 1960, and 60 for people born after June 30th 1964.

[4] https://treasury.gov.au/sites/default/files/2021-02/p2020-100554-ud00b_key_obs.pdf

3 Aggregate bunching at the SRA

3.1 Data

Our empirical analyses explore administrative data provided by Statistics Netherlands (2006–24) and the Australian Bureau of Statistics (2020–24). Both agencies maintain linked-employer-employee datasets that can be used to the role of firms and executives in influencing retirement bunching. For Australia, we also draw on a high-quality longitudinal survey, which allows us to quantify the overall retirement bunching at the SRA in Australia over a longer time span (2000–23).

Dutch sample. The Dutch sample comes from several linked population registers maintained by Statistics Netherlands. The cornerstone of our dataset is the SPOLISBUS register, which is a payroll register that tracks the full population of workers living in the Netherlands from January 2006 to June 2024. The key information is monthly records of workers' earnings and work hours (both regular and overtime) and a set of hashed IDs (individual and firm). Individual IDs are used to link relevant information from other population registers, such as worker's gender and month-year of birth.

To study retirement bunching at the SRA, we start with the universe of Dutch residents born between 1941 and 1957, who reach the SRA within the span of our data. Over this period, the SRA increased from 65 to 67 (Figure 1).

Our analysis of aggregate retirement bunching will focus on workers' transitions to retirement in the 12 months around the SRA. We define retirement as an event that occurs where the individual stops working for at least 6 months. That is, worker i retires in month t if $\text{employed}_{i,t} = 1$ and $\text{employed}_{i,t+k} = 0$ for $k \in [1, 6]$.

Australia sample. To quantify the amount of bunching at the SRA in Australia, we use a mixture of survey and administrative data sources. First, we use data from waves 1–23 of the Household Income and Labour Dynamics in Australia (HILDA) survey. This nationally representative, longitudinal dataset allows us to construct a monthly employment series for older Australians spanning July 2000 to June 2023.[5] Second, we use administrative data from the Australian Tax Office's Single Touch Payroll (STP), which contains payslip-level data from January 2020 to June 2024 for nearly all businesses and their employees. In both datasets, we define retirement in the same way as in the Dutch data and consider the same sample (i.e., workers within 12 months of their SRA).

[5] Starting in 2001, individuals are interviewed annually, typically between August and October. They are required to report their start- and end-dates for each job from July of the previous year, as well as any periods of unemployment or time out of the labor force. This allows us to define employment at a monthly frequency.

This means that we can define the monthly retirement hazard for individuals in HILDA from July 2000 to December 2022 and in the STP from January 2020 to December 2023.[6]

Both datasets have their strengths and weaknesses but are complementary. On the one hand, the survey data has a much longer panel, which allows us to consider more cohorts approaching the SRA (1939–1955 for women and 1935–1955 for men, compared to 1954–57 in the administrative data). On the other hand, the information in the survey data relies on individuals reporting when each of their job spells start and finish (for each fiscal year), which means that the monthly employment series may contain substantial measurement error. Furthermore, the sample size is low compared to the administrative data, and the survey may not be perfectly representative of the Australian population. The administrative data is likely to be of higher quality, as it contains (nearly) the full population of employees and is based on data automatically transmitted to the tax office each time an employee is paid, and result in more precise estimates. However, the panel is shorter and it overlaps with the COVID-19 pandemic.

3.2 Graphical evidence

Figure 2 shows the hazard of exiting the labor market at ages 12 months around the SRA for the Netherlands and Australia. Panel a shows the average hazard for Dutch workers, based on the administrative monthly payslip data described above, while panels b and c show the average hazard for Australian workers based on the HILDA survey data (panel b) and the administrative data (panel c). In both countries, there is a clear increase in the retirement hazard in the three months immediately around the SRA, with the largest increase in the exact month the worker reaches the SRA. However, the magnitude of these effects differ markedly across the two countries. In the Netherlands, the retirement hazard is nearly 50% in the SRA month (and elevated in the previous and subsequent month as well), while it is only around 4–5% in Australia. Reassuringly, there is strong alignment between the two sets of results for Australia from different data sources. Beyond the SRA, the retirement hazard appears to remain elevated, particularly when looking at the administrative data.

3.3 Regression evidence

We quantify the bunching at the SRA in the retirement hazard using two regressions. First, we use an event-study model, where we set the reference period to be 2 months before the SRA (event-time $= -2$). This choice of reference period reflects the fact that individuals may choose to retire

[6]In HILDA, we also exclude individuals who are unlikely to satisfy the residency condition for the Age Pension (born outside Australia and moved to Australia less than ten years ago). Such information is not available in the STP, though only 0.5% of individuals in the HILDA sample are excluded due to the residency condition.

in the month before they reach the SRA, particularly if they will reach the SRA early in the following month (e.g., a person may retire in December if they will reach the SRA in early January). Our event-study model includes time fixed effects (month-year of observation) and worker age fixed effects.[7] The former controls for business cycle fluctuations and seasonality in retirement transitions, while the latter controls for preferences to retire at specific ages (e.g., at particular birthdays like age 65). We estimate the following regressions:

$$\text{retires}_{it} = \alpha + \sum_{\substack{k=-12 \\ k \neq -2}}^{11} \beta_k \mathbf{1}\left(\text{age}_{it} - \text{SRA}_i = k\right) + \text{age}_{it} + \gamma_t + \epsilon_{it} \tag{1}$$

where i denotes a worker and t time (the month-year of observation); retires_{it} is a dummy variable equal to 1 if the worker retires in month t (0 if they remain employed in month $t+1$); α is the constant term; γ_t is a set of time fixed effects; age_{it} is a set of age-in-months fixed effects; and ϵ_{it} is an idiosyncratic error term.[8] The sample is an unbalanced panel of workers because workers drop out of the sample as they retire. The β_k terms are the coefficients of interest, denoting the difference in the retirement hazard of workers in event-month k and the reference period. We expect the β_k coefficients to be zero for $k < -2$ and positive for $k = 0$ (and possibly $k = -1$ and $k = 1$). If this is the case, we would conclude that we have robust evidence of bunching at the SRA. The β_k coefficients may also be positive for $k > 1$, if workers are more likely to retire at all ages once they are eligible for a pension (i.e., not just at the first possible age).

We present the estimates of equation (1) in Figure 3. For both countries, the patterns are qualitatively similar but quantitatively very different. The estimates for both countries show minimal evidence of any trend in the retirement hazard in the lead up to the SRA, with the coefficients hovering around zero. The retirement hazard then increases in the month before the SRA and peaks in the SRA month, before declining back towards baseline levels. In the month prior to the SRA, the estimated increase in the retirement hazard is 8.3 percentage points (pp) in the Netherlands compared to 0.3–1.0pp in Australia. This disparity is even larger in the SRA month: 43.4pp (Netherlands) and 1.6–3.4pp (Australia).

The retirement hazard remains partially elevated thereafter in both countries, with some of the subsequent coefficients attaining statistical significance. Thus, we observe strong evidence of bunching at the SRA and some evidence that the retirement hazard remains elevated afterwards,

[7] We include fixed effects for the worker's age in months in the Dutch data. Estimating such a rich specification is not possible with the Australian policy variation, which features larger SRA increments, due to multicollinearity. Instead, we include more parsimonious age controls for the Australian samples (fixed effects for age in half-year bins in the HILDA survey and a linear age term with the administrative data).

[8] We cluster standard errors at the level of randomization: birth year and month (Netherlands) and birth year and month by gender (Australia). This results in 215 clusters in the Dutch sample and 428 in the Australian sample.

particularly in the administrative data.

We quantify the amount of bunching at and following the SRA by estimating a slightly modified version of equation (1), where we aggregate the β_k coefficients for $k > -2$ into two groups (around-SRA: event-months -1 to 1 and post-SRA: event-months 2 to 11). Specifically, we estimate the following regressions:

$$\text{retires}_{it} = \alpha + \delta_1 \frac{\mathbf{1}\left((\text{age}_{it} - \text{SRA}_i) \in \{-1, 0, 1\}\right)}{3} + \delta_2 \frac{\mathbf{1}\left((\text{age}_{it} - \text{SRA}_i) \geq 2\right)}{10} + \text{age}_{it} + \gamma_t + \epsilon_{it} \quad (2)$$

where δ_1 quantifies the cumulative amount of bunching in the retirement hazard immediately around the SRA, relative to the year before the SRA, and δ_2 captures the cumulative increase in the retirement hazard in the subsequent 10 months.

We quantify the cumulative amount of bunching immediately around the SRA and subsequently using equation (2). As shown in Table 1, in the three months around the SRA, we estimate a 64.9-percentage-point increase in the retirement hazard in the Netherlands and a 4.2–4.6-percentage-point increase in Australia (both $p < 0.01$). These estimated increases in the cumulative hazard over these months amounts to a 819% increase in the Netherlands and a 69% increase in Australia. In the subsequent 10 months, we estimate a further 7.5-percentage-point increase in the retirement hazard in the Netherlands ($p < 0.01$) and a 1.6–4.7-percentage-point increase in Australia (the latter estimate, based on the administrative data, is statistically significant at the 1% level). Overall, our estimates suggest that around 68% of Dutch workers who are still working near the SRA retire within one year of attaining eligibility for the public pension, while this figure is closer to 9% in Australia.[9] The magnitude of both estimates is broadly consistent with earlier studies that focused on a narrower set of birth cohorts and/or a single gender (Rabaté et al., 2024; Morris, 2022).[10]

4 Firms' Role in SRA Bunching

What role do firms play in the much higher level of retirement bunching at the SRA observed in the Netherlands? We investigate this question by examining how the average job separation rate at the SRA varies across firms in the Netherlands and Australia using our linked employer-employee administrative data. We focus on separations because they are the main margin that individual firms can affect workers' retirement timing (since firms cannot prevent workers from

[9] $68\% = 0.649 + (1 - 0.649) \times 0.075$. $9\% = 0.046 + (1 - 0.046) \times 0.047$

[10] Rabaté et al. (2024) examine the effects of raising the SRA in the Netherlands from 65 to 66 years and 4 months. They estimate large employment effects, with a retirement hazard at the SRA of 72%. Morris (2022) uses an earlier version of the HILDA data (2001–14) to estimate the employment effects of increasing the SRA for women from 61.5 to 65. He estimates an increase of 2.7pp ($p = 0.121$) off a base employment rate of 30.3%, which is a 9% increase.

finding employment elsewhere).[11] While we would expect some variation in separation rates across firms due to random factors, such as differences in workers' retirement preferences, the extent of variation across firms is informative about the role of firms in the retirement process. Our preferred definition of job separations at the SRA considers individuals who are employed by a firm in event-month -1 and assesses whether they still work for the firm in event-month 2.[12] That is, $\text{separates}_{if} = 1$ if worker i is no longer employed by firm f in event-month 2 (otherwise, $\text{separates}_{if} = 0$). The advantage of this approach is that we have a single observation for each older worker and allow separations in any of the three months around the SRA where separations are elevated.

Firms may influence SRA separation rates in several ways. First, the nature of the work or management practices may affect the likelihood of workers choosing to retire voluntarily. Second, workers may respond to peer effects — if a worker observes all of their coworkers retiring at a specific age, they may also be more likely to do so. Third, and perhaps most importantly in the Dutch setting, firms may effectively force their older workers into retirement by choosing not to renew their contract when it ends at the SRA.

For this analysis, we restrict the set of older workers and firms in two ways. First, we restrict the set of 'focal' workers to have at least 12 months consecutive employment in the firm by event-month -1. Second, we focus on firms that had at least five focal workers, since the average SRA separation rate is uninformative about the role of firms if the firms have few workers reaching that age.[13] In total, we retain 57.3% of all Dutch workers who reach the SRA and 31.9% of Australian workers (see Table A1). The main reason for this difference is that the Australian panel is much shorter, which (i) makes the tenure requirement more restrictive (effectively, we lose all workers reaching the SRA in 2020) and (ii) limits the set of firms with at least five workers reaching the SRA.

We summarize the characteristics of older workers and firms in the two country samples in Table 2. Among older workers, a similar proportion are female (40.1% in the Netherlands vs 45.0% in Australia, and the average SRA is only 0.6 years different (65.9 in the Netherlands vs 66.5 in Australia). The average monthly earnings among the older workers are also similar (€3,421 in the Netherlands and $6,424 in Australia, which is equivalent to €3,659 at 2025 exchange rates). Among older workers in the two countries, the most notable difference is the separation rate in the

[11] Indeed, a very high proportion of job separations at the SRA coincide with workers exiting employment and retiring.

[12] Figures 1 and 2 show that the spike in workers' retirement hazard around the SRA starts in event-month -1 and continues until event-month 1, so we allow for job separations in the three months around the worker's SRA.

[13] For example, firms with a single worker reaching the SRA will either have an average retirement hazard of 0% or 100%. Moreover, these workers are often the owners of the business.

three months around the SRA, which is 76.7% in the Netherlands and 10.9% in Australia.[14]

Firms in the Australian sample are larger on average (1,460 workers compared to 489), which likely is due to the combination of our five-older-worker requirement and the shorter nature of the panel (i.e., Australian firms have to be larger to make it into the sample). Other firm characteristics are quite similar, in terms of the average ages and earnings of workers. Wage-age gradients are similar across countries but slightly higher in the Netherlands: on average, an additional year of age is associated with 1.2% higher earnings in the Netherlands and 1.0% higher earnings in Australia.

The last panel of Table 2 shows the characteristics of the highest-earning worker in the firm at the time, whom we call the executive. On average, highest-earners are predominantly male in both countries, although to a larger extent in the Netherlands (82.6% vs 63.6%). In both countries, executives are aged around 52–55 years on average and earn six- to nine-times the average salary within the firm. We validate our interpretation of highest-earners as executives by linking a subset of highest-earners in the Dutch sample to their contemporaneous responses in the Dutch Labour Force Survey (LFS).[15] Based on the International Standard Classification of Occupations, 79% of individuals are classified as a manager, and, of those asked, 93% report managing others, with 70% of these individuals managing at least 20 people. Moreover, 96% report managing personnel decisions and 95% report managing strategy decisions. Overall, these responses suggest that highest-earners have influential positions within their firms and are likely to be in a position to influence retirement policies.[16]

4.1 Variations in Retirement Bunching Across Firms

Figure 4 shows the distribution of firm-specific SRA separation rates faced by older workers. We weight firms by the number of older workers (i.e., the denominator of their average separation rate), which means larger firms and firms with more older workers matter more.[17] We present the distribution for Dutch employers on the left-hand side and Australian employers on the right. In

[14] Another notable difference is the share of immigrants. In the Australian sample, nearly half of workers are born outside of Australia (44.6%), compared to just one in ten in the Netherlands (9.9%). Most of these people immigrated to Australia at much younger ages and therefore satisfy the residency condition for the Age Pension (living in Australia for at least 10 years). Our analysis of the HILDA data suggests that less than 1% of individuals who are working at the SRA do not satisfy the residency condition.

[15] The LFS is a nationally representative quarterly survey with a rotating panel — individuals are re-interviewed in the panel for up to 5 quarters — that contains detailed information on individuals' occupation and job characteristics. We manage to link 1,939 observations on 748 highest-earners, which is approximately 1.5% of highest-earners in the sample.

[16] Previous studies have also shown that top-earners are usually top executives, partners, or top-level managers. For example, Bakija et al. (2010) shows that 40% of US taxpayers in the top 0.1% of the income distribution are executives, managers or supervisors. Song et al. (2019) assume that top-earners are the chief executive.

[17] We can thus interpret this distribution as the distribution of firm-specific SRA separation probabilities faced by the population of older workers (who work in sufficiently large firms).

panel a, we show the distribution for all employers, while the other panels focus on firms in the private sector (panel b) and public sector (panel c). In the Dutch sample, we observe considerable variation across firms, with a standard deviation of 0.19 around the mean SRA separation rate of 0.78. For example, approximately 15% of workers work in a firm with a separation rate of at least 0.95, but a similar proportion have a rate below 0.6 (15%). There is also a large mass of older workers in firms with a separation rate of 1 (5%). In Australia, there is a much lower mean (0.109) and considerably less dispersion, with a standard deviation of 0.084.

How large is the variation across firms in the two countries? One useful benchmark to consider is the case where the separation outcomes of coworkers are independent. In this case, we may expect less variation in the firm-specific separations because we explicitly rule out relevant firm-level factors, such as correlations in coworkers' job characteristics, peer effects and employer decision-making. To implement this comparison, we create a pseudo-distribution of firm-specific separation rates after randomly reshuffling workers across the firms in our sample.[18] As shown in Figure 4 (in dark gray), this pseudo-distribution has a bell-shape and is much less dispersed in the Netherlands, with a standard deviation that is 42% as large as the observed distribution. Thus, there is nearly 2.5-times as much variation in the Dutch data than would be expected by chance. In the Australian sample, we also observe slightly more dispersion in the true distribution than the one after workers are reshuffled across firms, but the difference is much smaller (around 25% instead of 150%). This suggests that firms have a much stronger influence on worker separations at the SRA in the Netherlands than in Australia.

A valid critique of this analysis is that we are not controlling for other factors that may vary between workers in different firms, such as the type of work involved. Workers in some industries, for example, may be more likely to voluntarily elect to retire at the SRA than other workers. Moreover, collective agreements in the Netherlands, which are typically set at the industry level, often dictate rules around mandatory retirement. However, we show in Figure 4 (in light gray) that the excess variation in firms' SRA separation rates persists within industries. To implement this comparison, we create a pseudo-distribution of firm-specific separation rates after randomly reshuffling workers across firms in the same industry. While industry-specific factors clearly explain some of the excess variation across Dutch firms, about half is unexplained, suggesting that "firm effects" matter in addition to "industry effects".[19] This highlights the influence of firm-specific factors on workers' retirement timing, especially in the Dutch context. The pseudo-distribution when reshuffling workers across firms in the same industry is similar to the pseudo-distribution when reshuffling

[18] We maintain the distribution of firm sizes (i.e., the number of older workers in each firm) to ensure comparability across the two distributions.

[19] The standard deviation of the within-industry pseudo-distribution is approximately half way between the standard deviation of the observed distribution and the other pseudo-distribution.

workers across the firms, suggesting that industry plays a smaller role in Australia.

When we examine public and private sectors separately, we observe much more dispersion, in private sector (panel b) with a much larger standard deviation (0.232) compared to the public sector (0.123) in the Netherlands. The two pseudo-distributions fail to capture this dispersion, particularly at the top and bottom of the distribution. For example, it is twice as common to observe private-sector firms with a separation rate of 100% than would be expected by chance, and 17-times more common to observe private-sector firms with a separation rate below 30%. In the Australian sample, there is much less difference between the public and private sectors.

4.2 What influences separation rates at the SRA?

To answer this question, we first use a variance decomposition method similar to Bertrand and Schoar (2003) to show how different components, such as worker characteristics, firm characteristics, sector and executive characteristics, and unobserved firm-level and executive-level factors explain individual separation outcomes at the SRA. Rows 1 to 5 of Table 3 report the adjusted R^2 when we regress separates$_{if}$ on different sets of controls, including (1) a rich set of worker characteristics; (2) firm characteristics; (3) industry and sector dummies; (4) the characteristics of executives / high-earners within the firm; and (5) time fixed effects.[20] Panel A shows the results for the Netherlands. Individually, these controls explain only a relatively small share of the variance in workers' SRA separation outcomes, ranging from 1% to 9% when we do not condition on sector. Industry and sector dummies are the most important, explaining 9%, while worker and firm characteristics explain 4% and 5% respectively, while time fixed effects explain less than 1% of the variation. Combined, these five types of controls (with 29 different variables and five dimensions of fixed effects) explain 13% of the variance in the full sample. Interestingly, columns (2) and (3) show that the adjusted R^2 values are consistently about twice as high for workers in the private sector. For example, the full list of controls explains 15.8% of the variance in the private sector and only 8.5% in the public sector. Panel B shows the results for Australia. In contrast to the

[20] We attempt to make the list of controls as uniform as possible for both countries. For the Dutch analysis, the full list of controls is as follows: (1) worker characteristics —- a female dummy; SRA-cohort dummies; dummies for being an immigrant; pre-SRA monthly earnings, hours/days worked and hourly/daily wage (and the worker's percentile ranks within the firm for each); a dummy for being in the top 5% of earners within the firm; a dummy for being the highest earner; and a dummy for being a director or major shareholder of the firm; (2) firm characteristics — dummies for each decile of firm size; a dummy for having multiple establishments; the number of establishments; the average age, earnings, hours worked and hourly wages of workers in the firm; the share who are female, immigrants; the wage-age gradient in the firm; and the average separation rate of older workers in the firm in the three years prior to the SRA; (3) industry dummies at the two-digit level; (4) time fixed effects at the calendar month-year level; and (5) executive characteristics — a dummy for having a director / major shareholder observed in the data; the number of directors / major shareholders; average age, earnings and female share of the directors; age, earnings and a female dummy of the highest earner in the firm; average age, earnings and female share of the directors of the top 5% earners within the firm.

Netherlands, worker characteristics explain the largest component of the variance (1.8%). Other factors, such as firm characteristics, sector fixed effects and executive characteristics, account for only 0.1% to 0.4%. The full list of controls explains just 3.2% of the variance and unlike in the Netherlands, there is no meaningful difference between public and private sectors.

In row (7), we demonstrate the importance of unobservable differences across firms: by themselves, firm fixed effects have an adjusted R^2 of 0.17 for all employers and 0.21 in the private sector in the Netherlands. These values are approximately 35% higher than the full list of controls. In the public sector, the adjusted R^2 terms are also higher with firm fixed effects, but the difference is smaller. Across all columns, the adjusted R^2 values increase further when we augment our firm fixed effects model with the full list of controls. In Australia, firm fixed effects explain 3% of the variance, which is similar to the adjusted R^2 when adding the full list of controls.

In row (8), we show that we achieve higher adjusted R^2 terms if we control for executive fixed effects. Our richest model, with executive fixed effects and the full list of controls, attains an adjusted R^2 value of 0.22 (0.09) for all employers and 0.27 (0.07) in the private sector in the Netherlands (Australia).

Overall, this exercise highlights a number of interesting patterns in the data. First, firm effects appear important in determining workers' separation outcomes at the SRA, particularly in the Netherlands: firm fixed effects explain five times as much of the variation in workers' outcomes as workers' own characteristics, and more than twice as much as industry and sector fixed effects. Second, differences in HR practices across executives may be an important component of the firm fixed effects. Third, the comparison between the Netherlands and Australia suggests that the role of firms and executives in shaping retirement behavior is more pronounced in the Dutch context.

4.3 Quantifying firm effects

While the descriptive analysis above highlights the importance of firms in explaining the variation in individuals' SRA separation outcomes, it is subject to several limitations. First, since older workers only reach the SRA once in their lives (and typically when employed by a single firm), it is difficult to disentangle unobserved individual and firm effects. Second, the R^2 terms with firm fixed effects may (slightly) overstate the importance of firms effects, since a worker's own outcome partly determines the mean separation rate in his/her firm, mechanically inflating the R^2 terms. Although we use adjusted R^2 terms for this reason, our analysis may still be subject to overfitting. Third, the mechanisms underpinning the firms effects remain unclear.

To address these limitations, we quantify the importance of firm effects using a series of regressions. Our basic approach is to estimate the relationship between an individual's SRA separation outcome and those of their coworkers using the following regressions:

$$\text{separates}_{ift} = \alpha + \beta \left(\frac{\sum_{j \in f, j \neq i} \text{separates}_{jf}}{N_f - 1} \right) + X_{ift} + \epsilon_{ift} \tag{3}$$

where i denotes an individual older worker in firm f who is just about to reach the SRA at time t; the term in parentheses is the leave-one-out (L1O) mean SRA separation rate of older workers in the firm (with N_f the total number of older workers in the firm observed over the sample period); X_{ift} is a set of individual-level, firm-level and time controls (see footnote 20 for the full list); and ϵ_{ift} is an idiosyncratic error term. We cluster standard errors by firm. The coefficient of interest is β, which shows the relationship between workers' SRA separation outcomes and those of other workers in the same firm: $\beta = 1$ would imply a perfect correlation, while $\beta = 0$ would imply that there is no correlation. Positive values of β suggest that there is a common factor — the firm — that affects the SRA separation rates of workers. As expected, placebo regressions that rule out this common factor — by randomizing workers across firms and recalculating L1O separation rates — yield precise-zero estimates of β in both the country samples.

We present the estimates in Table 4. In all models, we include industry and sector dummies, which is important to account for mandatory retirement provisions in collective agreements within the Netherlands, which differ by industry. The estimate of β in column 1, which includes no other controls, is 0.662 and highly statistically significant ($p < 0.001$). This means that a 10pp increase in the average SRA separation rate of coworkers is associated with a 6.6pp increase in the worker's own separation probability, even after controlling for differences in separation probabilities across industries and sectors. Nonetheless, this relationship may not be causal evidence of 'firm effects', since coworkers' SRA separation rates may be positively correlated for other reasons . Thus, to get closer to causality, we progressively add detailed controls at multiple levels in columns 2–5. We see that the estimates fall with both the individual- and firm-level controls, while controlling for the characteristics of the executive(s) and adding time fixed effects barely moves the coefficient. In all models, the estimated β coefficient remains large and highly statistically significant. Our preferred estimate, with the inclusion of all controls, indicates that a 10pp increase in the average SRA separation rate of coworkers is associated with a 5.9pp increase in the worker's own SRA separation probability in the Netherlands and a 2.7pp increase in Australia. These estimates imply that moving a worker from a bottom quartile (Q1) firm to a top quartile (Q4) firm, in terms of SRA separation propensities, would raise a Dutch worker's SRA separation propensity by 27.5pp and an Australian worker's SRA separation propensity by 5.3pp.

To what extent could omitted variable bias (i.e., spurious correlations in coworkers' SRA separation propensities) explain our positive estimates of β? To assess this question, we perform an

Oster (2019) test of coefficient stability using the associated Stata command `psacalc`. This test uses information on (i) the stability of treatment coefficients when additional controls are included and (ii) R^2 movements to determine bounds for β that are robust to selection on unobservables. Following the recommendations of Oster (2019), we impose an equal selection assumption (i.e., selection on observables and unobservables are equally important) and assume that the R^2 on the hypothetical regression of all observable and unobservable controls would be 30% higher than the R^2 with all observable controls.[21] Under these assumptions, we estimate a β of 0.337 in the Dutch sample and 0.244 in the Australian sample. Evidently, both estimates remain well above 0, indicating that the positive estimates of β are robust to selection on unobservables.

We also acknowledge that our estimates of β are likely to be attenuated due to measurement error. Ideally, we would like to measure the association between a worker's SRA separation rate and their coworkers' separation *probabilities*. Although the L1O coworker mean provides an estimate of this separation probability, it may be a noisy measure, particularly for firms with a relatively small number of older workers. In Appendix Table A2, we show that our estimates of β increase as we impose more stringent restrictions on the minimum value of N_f. For example, the estimates of β increase in the Dutch sample from 0.590 with $N_f \geq 5$ to 0.709 with $N_f \geq 10$ and 0.781 with $N_f \geq 20$. In the Australian sample, the increases in the estimates are even more noticeable: from 0.274 with $N_f \geq 5$ to 0.445 with $N_f \geq 10$ and 0.592 with $N_f \geq 20$. This pattern is consistent with classical measurement error, which would attenuate the β estimates toward zero.

4.4 The Role of Executives

To what extent are the differences in SRA separation rates across firms driven by different HR policies (e.g., regarding mandatory retirement)? And are these policies immutable or can they be influenced by leaders within the firm? To assess these questions, we study how workers' SRA separation outcomes are correlated when they reach the SRA under the same executive (highest-earning worker in the firm). Specifically, we define the L1O SRA separation rate of older workers who reach the SRA under the same executive (either within the same firm or another firm), $\left(\frac{\sum_{j\in e, j\neq i} \text{separates}_{jf}}{N_e - 1}\right)$, where i indexes the older worker and e the executive. Then, we re-estimate equation (3), where we augment the L1O firm mean with the L1O executive mean.[22] Specifically, we estimate the following model

$$\text{separates}_{ift} = \alpha + \beta_f \left(\frac{\sum_{j\in f, j\neq i} \text{separates}_{jf}}{N_f - 1}\right) + \beta_e \left(\frac{\sum_{j\in e, j\neq i} \text{separates}_{je}}{N_e - 1}\right) + X_{ift} + \epsilon_{ift} \quad (4)$$

[21] For this test, we assume the 'uncontrolled' regression contains industry and sector dummies (Table 4, column 1) but no other controls.

[22] We use a slightly modified sample, excluding observations where individual executives are linked to less than five older workers (to mirror our sample restriction on firms). This reduces the size of the sample by about 20%.

We present the estimates in Table 5. In the Dutch sample, the estimates of β_f (0.364) and β_e (0.306) are both positive and highly statistically significant. This suggests that we have sufficient variation in separation rates across executives within the same firm, and firms' SRA separation propensities are not immutable. To illustrate the magnitude of these effects, consider an older worker i in a firm that has two executives over its history: executive A and executive B. Suppose that, among i's coworkers, executive A always enforced mandatory retirement and never re-employed older workers (100% separation rate), while executive B re-employed 50%. The estimates suggest that the older worker will be 15pp more likely to separate from the firm at the SRA if they reach the SRA under executive A instead of executive B. In the Australian sample, the estimates of β_f (0.200) and β_e (0.246) are also positive and highly statistically significant. However, although these coefficients are only slightly smaller than in the Dutch sample, the R^2 terms are much smaller (about one-quarter of the size), indicating that there is much less variation in the SRA separation outcomes across firms and executives in the Australian context. This suggests that firms' practices and managerial discretion has a much smaller influence on workers' retirement outcomes in the Australian context.

5 Discussion and Conclusion

This paper provides new evidence on the role of firms and executives in shaping retirement behavior at the statutory retirement age (SRA), using rich administrative data from the Netherlands and Australia. We document striking differences in the degree of retirement bunching across the two countries and show that these differences are partly explained by the influence of firms and their leaders. In the Netherlands, firm and executive fixed effects explain a substantial share of the variation in SRA separation outcomes—much more so than worker or industry characteristics—while in Australia, the role of firms is notably smaller.

Using leave-one-out methods, we demonstrate that older workers' separation decisions are strongly correlated with those of their coworkers in the same firm and coworkers under the same executive, suggesting that retirement policies and practices are shaped by workplace norms and managerial styles. Importantly, this finding implies that managerial policies on retirement are not immutable and can vary depending on who is in charge. This paper highlights the importance of the demand side influences (firms and firm leaders) when designing policies aimed at extending working lives, especially in institutional contexts where employer discretion over retirement is greater.

The observed differences in retirement behavior around the statutory retirement age between Australia and the Netherlands could be partly related to the different institutional setting. In particular, employment contracts end automatically at the statutory retirement age (SRA/AOW) in the

Netherlands. This feature effectively leads to mandatory retirement at the SRA. Moreover, the margin of responses heavily rely on how firms and executives effectively rehire older workers by initiating a new contract. In contrast, Australia operates a more flexible system with a public pension scheme that is means-tested and not universally granted at a fixed age. This system introduces incentives for continued labor force participation among older individuals, especially those with insufficient private savings. However, Appendix Figure A1 shows that even when we restrict the sample to workers who we know are eligible for the Age Pension, as they receive pension income within 12 months of reaching the SRA, the bunching at SRA is still much smaller in Australia, less than 12% of separation rate at SRA. This suggests that the differences in retirement behavior around the statutory retirement age between Australia and the Netherlands is not only caused by the fact that Australia public pension is means-tested. Another possible factor could be Australia's relatively weak employment protection, which makes it easier for firms to fire and rehire older workers. These institutional differences likely shape not only when individuals retire, but also how firms play a role in determining the retirement pattern.

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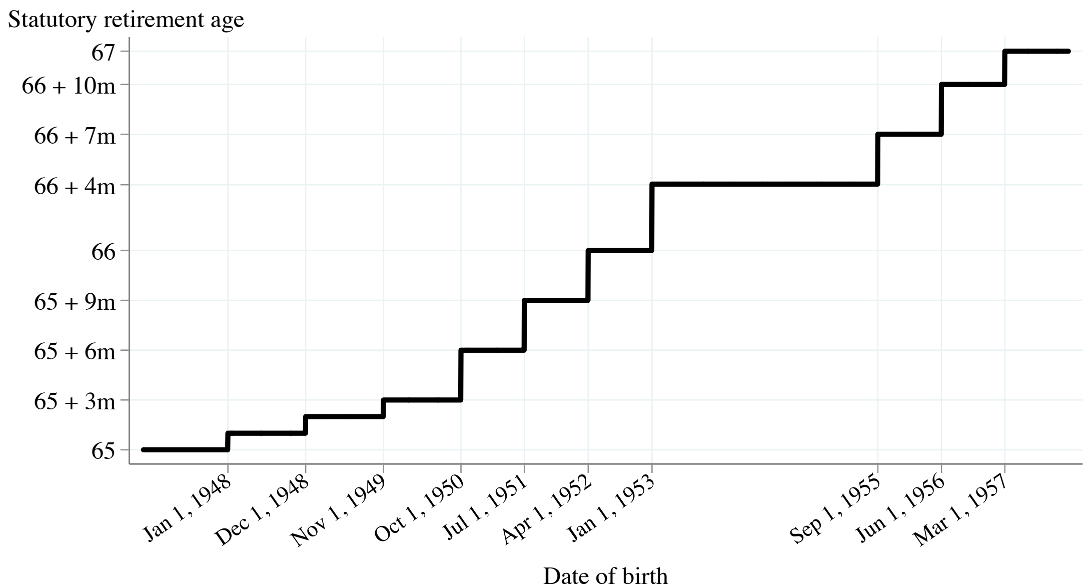
Figure 1: Statutory retirement ages in Australia and the Netherlands

(a) The Netherlands



(b) Australia

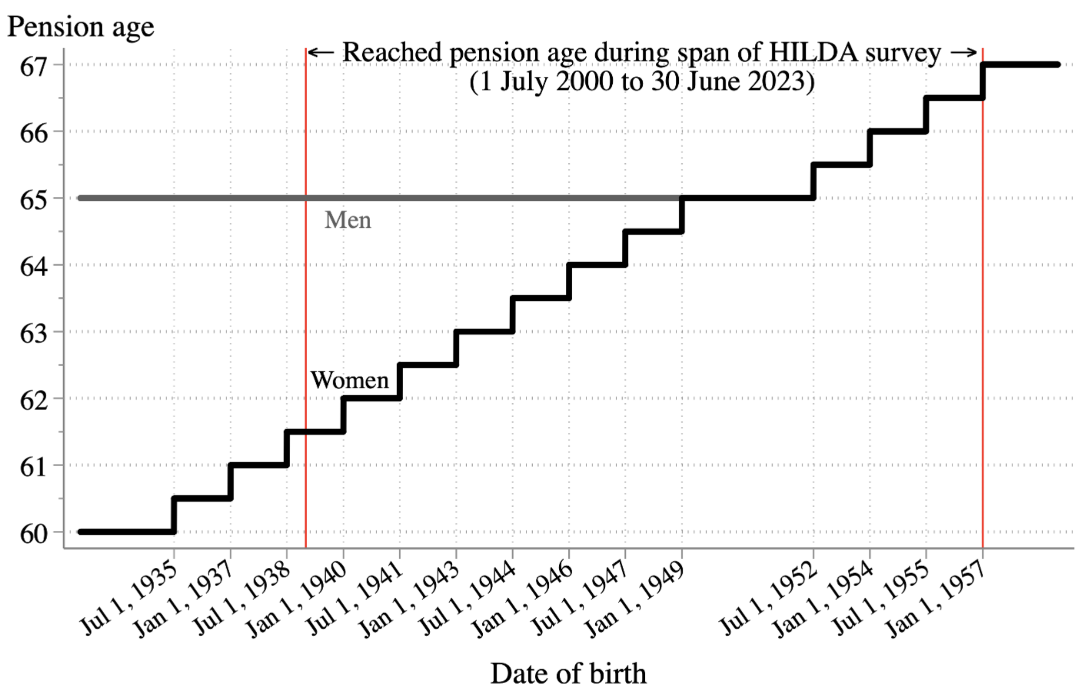
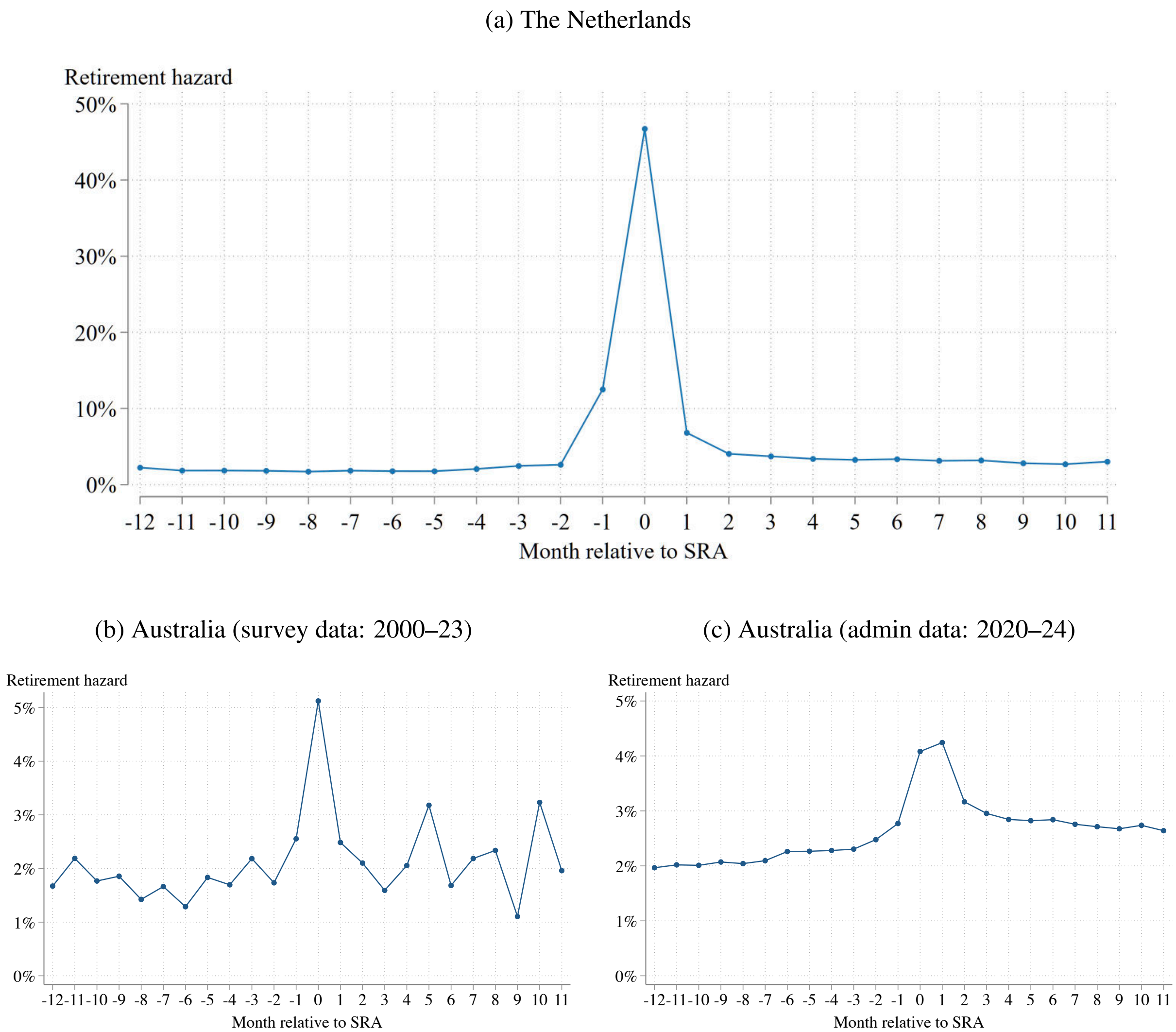


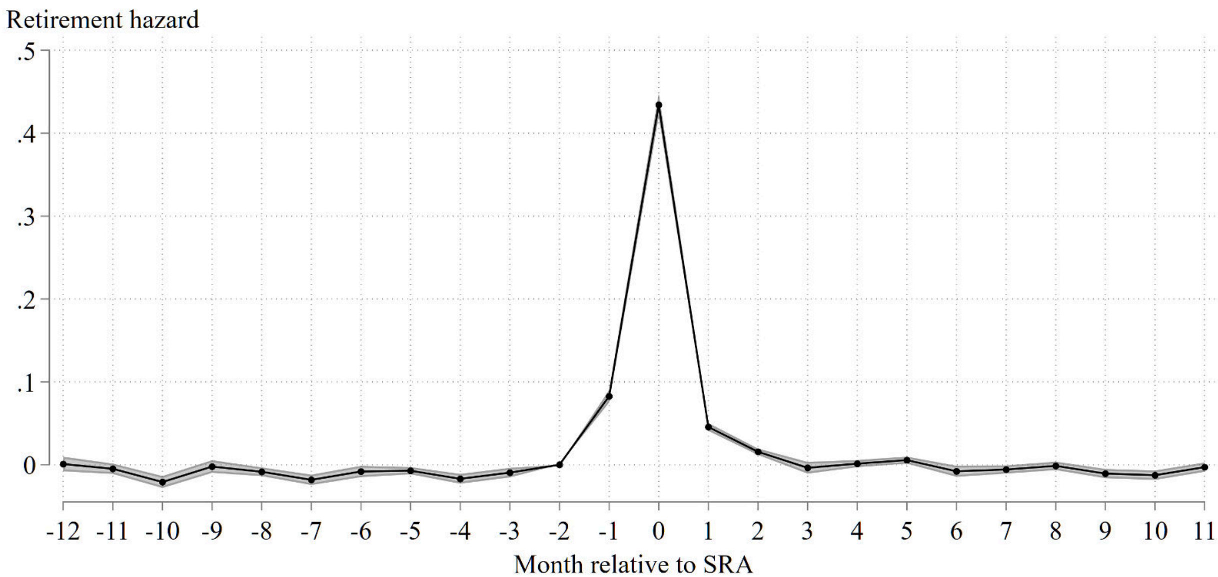
Figure 2: Transitions into retirement around the statutory retirement age



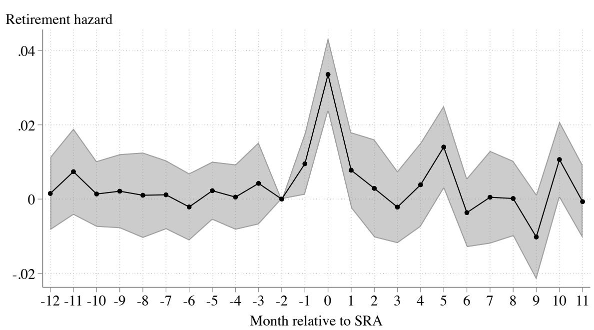
Notes: These figures plot the average retirement hazard of Dutch and Australian workers by the number of months until the SRA.
Source: CBS and ABS administrative data; HILDA survey data.

Figure 3: Event-study estimates of bunching at SRA

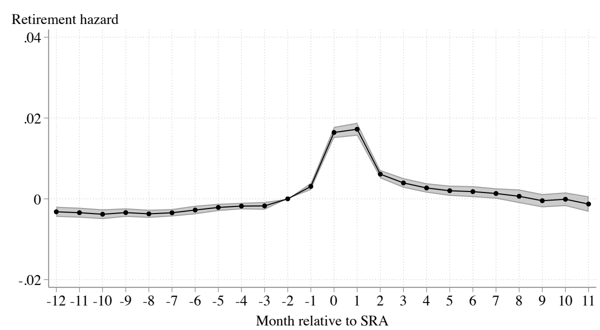
(a) The Netherlands



(b) Australia (survey data: 2000–23)



(c) Australia (admin data: 2020–24)

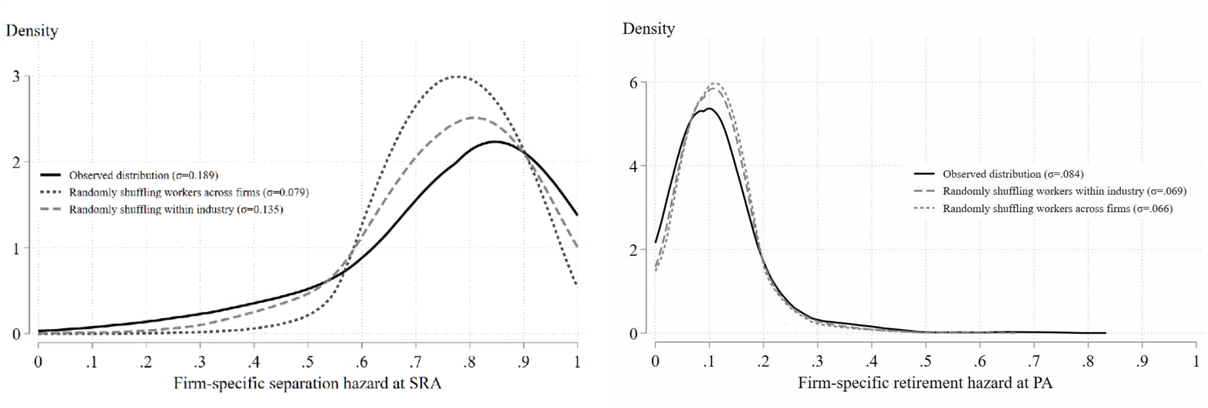


Notes: These figures plot the regression-adjusted retirement hazard around the SRA for the Netherlands (Figure a) and Australia (Figure b). The event-study estimates are relative to the reference period (two months prior to the SRA) and are based on equation (1). We present 95% confidence intervals with standard errors clustered by month of birth (Netherlands) and gender by month of birth (Australia).
Source: CBS and ABS administrative data; HILDA survey data.

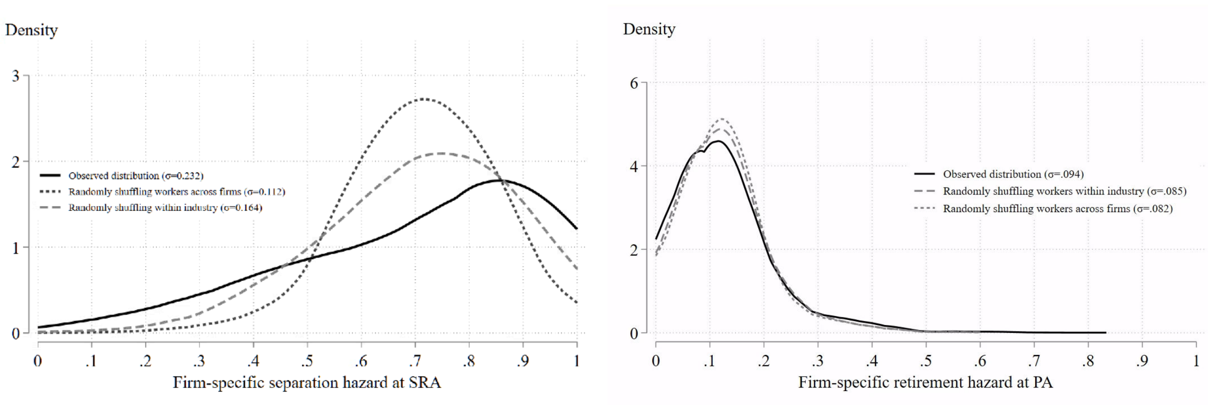
Figure 4: Distribution of SRA retirement hazards across employers, the Netherlands

Netherlands | Australia

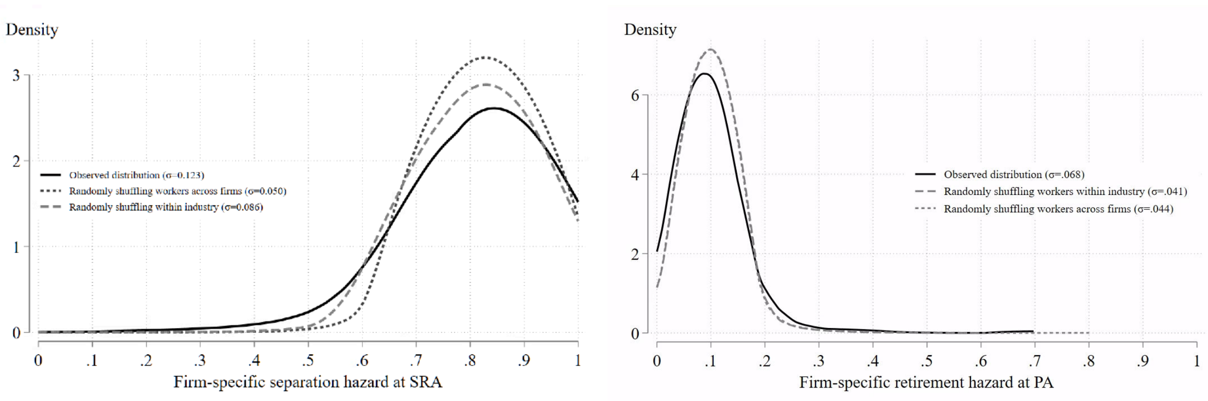
(a) All employers



(b) Private sector



(c) Public sector



Notes: These figures compare the distribution of firm-level incidences of bunching at statutory retirement age (solid black line) with the distribution when randomly shuffle workers across firms (dotted line) and randomly shuffle within sectors (dashed line) for all firms (Figure a), private sector firms (Figure b), public sector firms (Figure c) and subsidized sectors firms (Figure d), respectively.
Source: CBS and ABS administrative data.

Table 1: Comparison of SRA bunching in Netherlands and Australia

	(1)	(2)	(3)
	Netherlands	Australia	
	Admin data (2006–24)	Survey data (2000–23)	Admin data (2020–24)
Around SRA (months -1 to 1)	0.649***	0.042***	0.046***
	(0.007)	(0.013)	(0.001)
Implied % increase	819%	69%	71%
Post SRA (months 2 to 11)	0.075***	0.016	0.047***
	(0.025)	(0.029)	(0.005)
Implied % increase	30%	8%	20%
Number of observations	12,732,742	28,829	5,502,118
Number of workers	956,135	1,629	433,675

*** denotes p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses are clustered by month of birth.
Source: CBS and ABS administrative data; HILDA survey data.

Table 2: Descriptive statistics for older workers, firms and executives

	Netherlands		Australia	
	Mean	Std Dev	Mean	Std Dev
Older worker characteristics				
Separates at SRA	76.7%		10.9%	
Share female	40.1%		45.0%	
1st gen immigrant	9.9%		44.6%	
Private sector	45.4%		56.2%	
SRA in years	65.9	(0.6)	66.5	(0.3)
Monthly earnings pre SRA	€3,421	(€2,255)	$6,424	($5,457)
Percentile rank in firm	0.52	(0.30)	0.48	(0.28)
Monthly hours/Days employed	125.4	(46.5)	29.4	(5.2)
Percentile rank in firm	0.48	(0.30)	0.54	(0.27)
Hourly/Daily wage pre SRA	€26.5	(€12.9)	$227	($339)
Percentile rank in firm	0.57	(0.29)	0.46	(0.28)
Number of older workers	384,775		88,656	
Firm characteristics				
Number of employees	489	(2,041)	1,460	(5,629)
Average age of workers (years)	45.0	(5.1)	44.3	(4.7)
Average monthly earnings	€3,281	(€1,235)	$6,468	($2,835)
Share of female workers	41.7%	(0.286)	50.3%	(0.255)
Share of 1st gen immigrants	10.6%	(0.126)	39.0%	(0.166)
Wage-age gradient (log pts)	0.012	(0.009)	0.010	(0.002)
Number of firms	13,936		3,905	
Executive (highest-earner) characteristics				
Share female	17.4%		36.4%	
Age in years	54.9	(7.4)	51.9	(10.6)
Monthly earnings	€19,117	(€16,307)	$57,387	($24,355)
Number of observations	384,775		88,656	

Notes: This table presents descriptive statistics of older workers, firms and executives. We present the means for all variables and standard deviations for continuous variables. At 2025 exchange rates, €1 = AU$1.76. Observations are measured two months before the older worker reaches the SRA, with one observation per worker. For the firm characteristics in the middle panel, we take the mean of each variable across all observations within the given firm before taking the average across firms.
Source: CBS and ABS administrative data.

Table 3: Share of variance in SRA separation rates explained by different controls, based on adjusted R^2 values

	(1)	(2)	(3)
		By sector	
	All employers	Private	Public
Panel A: Netherlands			
(1) Own worker characteristics	0.041	0.089	0.043
(2) Firm characteristics	0.053	0.070	0.043
(3) Industry & sector fixed effects	0.088	0.094	0.050
(4) Executive characteristics	0.029	0.036	0.009
(5) Time fixed effects	0.007	0.011	0.006
(6): All controls, (1) to (5)	0.131	0.158	0.085
(7): Firm fixed effects	0.173	0.214	0.093
(7a): (7) + (6)	0.203	0.247	0.127
(8): Executive fixed effects	0.190	0.237	0.110
(8a): (8) + (6)	0.219	0.267	0.142
Observations	384,775	174,530	210,245
Panel B: Australia			
(1) Own worker characteristics	0.018	0.020	0.020
(2) Firm characteristics	0.003	0.001	0.006
(3) Industry & sector fixed effects	0.004	0.004	0.001
(4) Executive characteristics	0.001	0.001	0.002
(5) Time fixed effects	0.009	0.009	0.009
(6): All controls, (1) to (5)	0.032	0.033	0.035
(7): Firm fixed effects	0.030	0.023	0.034
(7a): (7) + (6)	0.057	0.052	0.061
(8): Executive fixed effects	0.067	0.047	0.071
(8a): (8) + (6)	0.089	0.073	0.092
Observations	88,656	49,820	38,836

Source: CBS and ABS administrative data.

Table 4: Quantifying firm effects with different controls

	(1)	(2)	(3)	(4)	(5)
			Panel A: Netherlands		
L1O firm mean	0.662***	0.621***	0.597***	0.590***	0.590***
	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)
Controls					
Industry & sector dummies	✓	✓	✓	✓	✓
Worker characteristics		✓	✓	✓	✓
Firm characteristics			✓	✓	✓
Executive characteristics				✓	✓
Time fixed effects					✓
R-squared	0.147	0.168	0.169	0.170	0.171
Mean SRA hazard			0.767		
Observations			384,775		
			Panel B: Australia		
L1O firm mean	0.301***	0.294***	0.282***	0.277***	0.274***
	(0.073)	(0.074)	(0.073)	(0.072)	(0.072)
Controls					
Industry & sector dummies	✓	✓	✓	✓	✓
Worker characteristics		✓	✓	✓	✓
Firm characteristics			✓	✓	✓
Executive characteristics				✓	✓
Time fixed effects					✓
R-squared	0.011	0.030	0.030	0.031	0.039
Mean SRA hazard			0.109		
Observations			88,656		

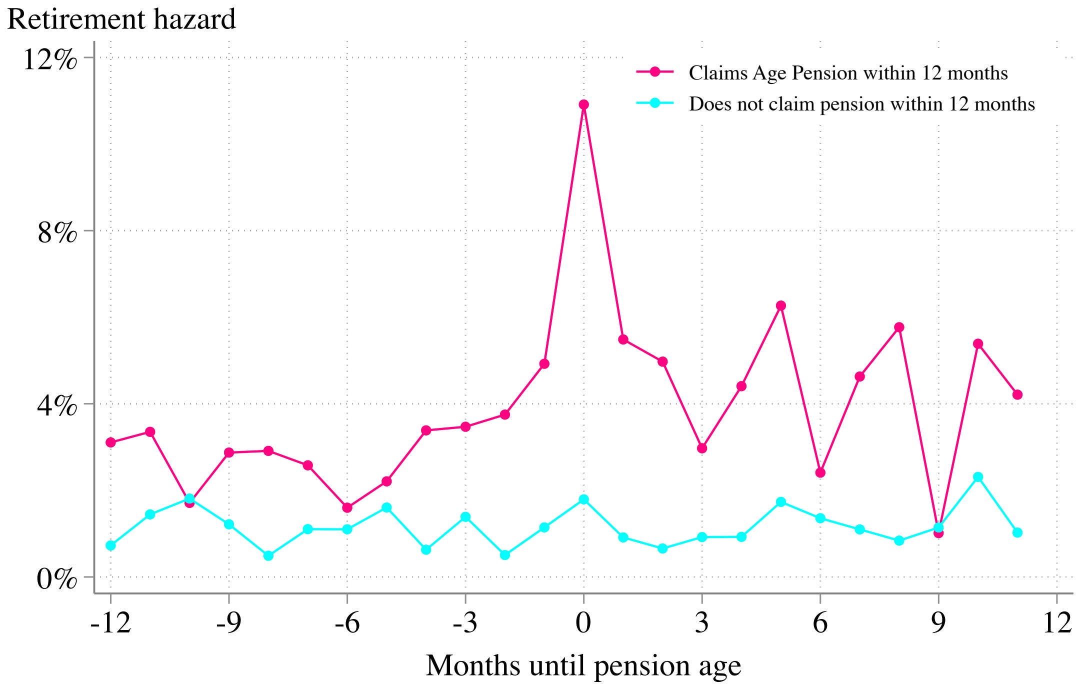
*** denotes p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses are clustered by firm.
Source: CBS and ABS administrative data.

Table 5: Mechanisms: Examining the role of executives

	Netherlands		Australia	
	(1)	(2)	(3)	(4)
L1O mean in firm	0.650***	0.365***	0.437***	0.200***
	(0.010)	(0.018)	(0.125)	(0.076)
L1O mean of executive		0.306***		0.246***
		(0.020)		(0.070)
All controls	✓	✓	✓	✓
R-squared	0.165	0.169	0.045	0.049
Mean SRA hazard	0.782		0.109	
Observations	323,512		56,716	

*** denotes p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses are clustered by firm.
Source: CBS and ABS administrative data.

Figure A1: Heterogeneity in Retirement Bunching at the Statutory Retirement Age in Australia



Notes: This figure compares the extent of retirement bunching at the statutory retirement age (SRA) in Australia between two groups: those who may not be eligible for the Age Pension (i.e., individuals who do not claim pension benefits within 12 months of reaching the SRA) and those who are definitely eligible (i.e., individuals who begin receiving pension income within 12 months of reaching the SRA).
Source: HILDA.

Table A1: Impact of sample restrictions

	Dutch sample		Australian sample	
	Workers	Firms	Workers	Firms
Employees aged one month below SRA	671,495	190,219	277,877	176,451
Continuous tenure of at least 12 months	564,872	135,744	160,702	62,011
At least 5 such workers ever observed in the firm	384,775	13,936	88,656	3,905

Source: CBS and ABS administrative data.

Table A2: Sensitivity of estimates to measurement error in coworker separation probability and heterogeneity by sector

	(1)	(2)	(3)
	Minimum observations per firm		
	5	10	20
Panel A: Netherlands			
L1O mean in firm	0.590***	0.708***	0.781***
	(0.009)	(0.009)	(0.010)
R-squared	0.171	0.163	0.159
Observations	384,775	338,739	293,359
Panel B: Australia			
L1O mean in firm	0.274***	0.445***	0.592***
	(0.072)	(0.102)	(0.123)
R-squared	0.039	0.044	0.048
Observations	88,656	74,632	62,608

*** denotes p<0.01, ** p<0.05, * p<0.1. Standard errors in parentheses are clustered by firm.
Source: CBS and ABS administrative data.