Retirements, Vacancy Chains, and the Secular Decline in Worker Reallocation*

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Abstract

The pace of worker reallocation in the U.S., measured by employer-to-employer (EE) transitions, has been declining since the early 1990's. This paper explores the impact of older workers' decisions to delay their retirement from the labor force on the pace of worker reallocation. Using variation in the age composition of firms' workforces and shifts in workers' retirement rates caused by two Social Security rule changes, I measure the effect of a worker's voluntary separation on a firm's replacement hiring rate. After documenting that the average worker retires almost two years later for cohorts reaching early retirement age in 2008 versus 1990, I use a vacancy chain framework to evaluate the effect of workers' delayed retirements on the overall pace of worker reallocation. A retiring worker may generate a chain of vacancies (and EE transitions) if the employer replaces the retiring worker with an already employed worker. This new hire must then quit their old job, thus creating a new vacancies and perpetuating the vacancy chain. I find that approximately 30% of the secular decline in the rate of EE transitions from 1990 to 2015 can be explained by the delaying of retirement by older workers.

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1 Introduction

According to a broad range of measures, the dynamism of the U.S. labor market has been declining since the early 1990's.[1] This paper examines one measure of labor market dynamism - the rate at which workers switch jobs from employer to employer (EE transitions) - and explores the extent to which changes in workers' retirement decisions contribute to fluctuations in the EE transition rate. Since job switching promotes the quality of match-specific productivity, worker bargaining power, and worker lifetime earnings growth; a long-term decline in the EE transition rate could have important macroeconomic implications for productivity and earnings inequality.[2] Previous research has shown that the aging of the workforce can explain approximately 20% of the secular decline in the EE transition rate since the mid-1990's simply because older workers tend to switch jobs less frequently (Hyatt and Spletzer (2013); Molloy, Smith, Trezzi and Wozniak (2016)). In this paper, I propose an additional mechanism by which older workers affect the EE transition rate: the timing of their retirement from the labor force.

A worker's decision to retire from employment can generate many EE transitions by initiating a vacancy chain - where a firm chooses to fill the job vacated by the retiring worker with an already employed worker, who must, in turn, quit their previous job, thus creating a chain of additional vacancies (Akerlof, Rose, Yellen, Ball and Hall (1988)).[3] When workers choose to delay their retirement from the labor force, the number of EE transitions falls as fewer vacancy chains are created. Older workers' shift towards later retirement ages has been dramatic, with the median employed worker who reaches early retirement age (62) in 2008 choosing to retire nearly two years later than the similar median worker in 1990. Furthermore, the largest increases in older workers' employment rate have occurred among workers who are high on the job ladder at age 62.[4] The fact that workers high on the job ladder exhibit the largest retirement delays has important implications

[1] This declining dynamism has been documented for job creation and destruction (Davis, Faberman, Haltiwanger, Jarmin and Miranda (2010); Decker, Haltiwanger, Jarmin and Miranda (2014); Engbom (2017)), interstate migration (Kaplan and Schulhofer-Wohl (2012), Molloy, Smith and Wozniak (2014)), job tenure (Hyatt and Spletzer (2016)), and EE transitions (Hyatt and Spletzer (2013); Davis and Haltiwanger (2014); Molloy, Smith, Trezzi and Wozniak (2016)).

[2] Studies that show job switching promotes the quality of match-specific productivity include Mendes, Van Den Berg and Lindeboom (2010); Hagedorn and Manovskii (2013); Card, Cardoso, Heining and Kline (2018); and Hagedorn, Law and Manovskii (2017). Cahuc, Postel-Vinay and Robin (2006) examines the effect of job switching on worker bargaining power. Topel and Ward (1992) and Haltiwanger, Hyatt, Kahn and McEntarfer (2018) demonstrate the effect of job switching on workers' lifetime earnings growth.

[3] More recently, Mercan and Schoefer (2019) and Elsby, Michaels and Ratner (2019) explore the implications of vacancy chains for cyclical fluctuations as workers' employer-to-employer transitions generate job vacancies and replacement hiring at their previous employers.

[4] I define the job ladder by ranking firms into employment-weighted quintiles according to each firm's workers' average full-quarter earnings.

for worker reallocation since I find that retirements high on the job ladder generate longer vacancy chains and thus more worker reallocation.

Approximately 1 million workers retire from the labor force each quarter, but these are only a fraction of the 5-8 million quarterly EE transitions. Given these level differences, changes in older workers' retirement behavior will have a substantial effect on the EE transition rate only if vacancy chains are relatively long - where each delayed retirement would have generated many EE transitions. The length of a typical vacancy chain is determined by three parameters: 1) firms' replacement hiring rate in response to a worker retirement, 2) the probability that the replacement hire is poached from another firm (generating an EE transition), and 3) firms' replacement hiring rate in response to quits (which perpetuates the vacancy chain).

I use a 4-state sample of the employer-employee linked Longitudinal Employer-Household Dynamics (LEHD) dataset from 1990-2015 to show that the new hire poaching probability increases with firm rank. Firms in the top quintile of the job ladder are almost twice as likely to poach new hires from other firms relative to firms in the lowest quintile. To estimate firms' replacement hiring rate in response to retirements and worker poaching, I examine the response of firms' 4-quarter change in hires-per-employee to changes in their retiring or poached-workers-per-employee.[5] To address both measurement error and omitted variable bias, I construct two share-shift (Bartik-style) instrumental variables that approximate distinct components of the change in retirements at a given firm. The first IV exploits the substantial variation in retirement rates across ages within a given time period. The second IV exploits variation across time in workers' retirement rates at each specific age. I construct similar instruments for poached workers at a firm using variation in the quit rate both over time and across ages and tenures. I find that the average retirement (poaching) generates approximately 2.1 (1.5) new hires over the course of the year following the retirement (poaching).[6]

The validity of these share-shift instrumental variables hinges on the exogeneity of both the share and the shift components. With regards to the share component of the IV, the age-sex composition of a firm's workforce, the concern is that the age of workers at the firm may be correlated with

[5]Elsby, Michaels and Ratner (2019) also explore firms tendency to conduct replacement hiring. In their paper, they examine the consistent pattern of firms with zero quarter-over-quarter employment change but significant within quarter hires and separations.

[6]Jäger and Heining (2019) examine firm's replacement hiring after an unexpected death of an employee. They find that in the period immediately following a worker's death, firms hire on average 0.417 new workers. This is very similar to my IV estimates that, in the quarter in which a worker retires or quits, firms increase their hiring by approximately 0.5 hires on average.

unobserved factors that affect the firm's replacement hiring rate. For instance, if younger workers are more likely to depart firms that face declining product demand, then firms with a larger share of older workers may be less likely to replace retiring workers. To address this concern, I control for the share of older workers at the firm. This is equivalent to a uniform weighting of older workers by age, and thus ensures that the instruments exploit only variation due to differences in retirement rates across age and time.

For the shift component of the two instrumental variables (either differences in the retirement rate by age (within a given period) or by time (for a given age), the validity concern is that the factors driving changes in older worker's retirement behavior are also affecting firms' replacement hiring rates. For instance, if the worker-firm match quality has improved over time, then this would induce both lower replacement hiring rates (as firms need to churn through fewer new hires to find a good worker) and delayed retirements of older workers (because older workers receive more value from their jobs). To rule out this endogeneity of the changes in retirement behavior to firms' replacement hiring rates, I demonstrate that two Social Security Old-Age and Survivors Insurance (OASI) rule changes were responsible for most of the delayed retirement of older workers.

The first OASI rule change that I examine is the gradual increase in the Delayed Retirement Credit (DRC) legislated by the Social Security Amendments of 1983. The DRC is the percent increase in an individual's monthly OASI benefit for every year (prorated) that the individual waits to claim OASI after their Full Retirement Age (FRA) (up to age 70). Prior to 1990, the DRC was 3.0%. Starting in 1990, the DRC was increased for individuals reaching their FRA in the given year by 0.5pp increments every two years until it reached 8.0% in 2008. To identify the effect of the DRC increase on older workers' retirement behavior, I use a regression discontinuity analysis on individuals born immediately before versus after each of the DRC increase cutoff dates. Due to sample size limitations, earlier studies of the impact of the DRC increase on older workers' employment rates have pooled together workers into broad age categories (e.g. ages 65-69 in Pingle (2006) or ages 55-69 in blau2010can). With administrative data it is possible to examine more precise age-specific responses to the DRC increase. I find that there is significant heterogeneity across ages in the response of employment to the DRC increases, with the greatest employment rate changes occurring for 66 year-olds (with each 0.5pp increase generating a 2pp increase in the employment rate), whereas individuals under the age of 64 and over the age of 69 exhibit little or no change in employment rates.

The second OASI rule change that I examine is the 2000 elimination of the Earnings Test for workers over their Full Retirement Age. Prior to 2000, if individuals age 65-69 had claimed their OASI benefits and earned above an exemption threshold, then their monthly benefits were reduced by 33 cents for every dollar they earned above the threshold. In March of 2000, the Earnings Test was eliminated for individuals over their FRA (a more stringent Earnings Test still applies to individuals between ages 62 and their FRA). To identify the effect of this change, I use a difference-in-difference strategy similar to that of Gelber, Jones, Sacks and Song (2017), where the control group is composed of workers whose calendar year earnings at age 63 are below the anticipated exemption threshold that would apply at age 65. The results indicate that the 2000 elimination of the Earnings Test: i) caused treated workers turning 65 in 2000 to increase their employment rate at age 66 by 2.5pp relative to similar workers in 1998 (before any announcement of legislation to eliminate the Earnings Test); ii) that this effect persisted through at least age 68; and iii) that workers age 63 or 64 in 2000 also increased their employment rate in anticipation that the Earnings Test would no longer bind at age 65, such that workers age 63 in 2000 had a 4.5pp higher employment rate at age 66 relative to similar workers in 1998. This last finding may explain why some earlier studies of the affect of the elimination of the Earnings Test (Song and Manchester (2007)) found little effect of the Earnings Test on the extensive margin of labor supply since they used 62-64 year-old workers as the control group in their difference-in-differences estimations.

Having established the validity of the IV estimates of firms' replacement hiring rates, I then use these replacement hiring rates to estimate the length of the vacancy chains that are generated by retirements at different rungs of the job ladder. I find that vacancy chains tend to be longer when workers retire from jobs high on the job ladder because firms that are higher up the job ladder are more likely to poach their new hires away from other firms (versus hiring workers from nonemployment). For instance, the average retirement from a firm at the bottom of the job ladder generates a vacancy chain with 1.6 EE transitions, whereas retirements from the top of the job ladder generate more than twice as many EE transitions.

By combining the estimated lengths of vacancy chains with an estimate of the number of "missing" retirement that would have occurred if the age-specific retirement rates had remained at the rate observed in 1990, I estimate the EE transition rate in the counterfactual scenario where older workers continued to retire at their 1990 rates. These estimates imply that approximately 30% of the secular decline in the EE transition rate can be explained by older workers choosing to delay retirement.

This paper proceeds as follows: Section 2 describes the subsample of the Longitudinal Employer-Household Dynamics (LEHD) dataset that I use in this paper. Section 3 documents the increased employment rate of older workers across the job ladder and estimates the causal impact on older worker's retirement behavior of two OASI rule changes: i) the gradual increase in the Delayed Retirement Credit from 3.0% in 1989 to 8.0% in 2008, and ii) the 2000 elimination of the Social Security Earnings Test for workers over their Full Retirement Age (FRA).[7] Section 4 describes the vacancy chain framework and the estimation of the key parameters affecting the length of vacancy chains. In particular, this section documents how the new hire poaching probability varies over the job ladder and estimates firms' replacement hiring rates. The section concludes with a simple counterfactual exercise that estimates the EE transition rate if older workers' retirement rates were unchanged since 1990.

2 Data

The primary dataset used in this paper is the Census Bureau's Longitudinal Employer-Household Dynamics (LEHD) database - an employer-employee linked dataset with quarterly earnings for approximately 96% of all employment in a state. The LEHD database is derived from the Quarterly Census of Employment and Wages survey and Unemployment Insurance filings. The primary sample of the paper is a four-state sample (IL, MD, WA, WI) covering the period from 1990:Q1 to 2015:Q4.[8] For the analysis of vacancy chains over the job ladder, I use a secondary sample that includes 32-states, but only for the period from 1998:Q1 to 2015:Q4. The LEHD dataset includes a set of worker characteristics, such as date of birth, sex, race, and education, which are derived from both the Social Security Numident File and various U.S. Census surveys. The LEHD dataset also has a set of firm characteristics, such as industry, firm age, and firm size, where the firm identifier is at the level of a state employer identification number (SEIN).[9]

By combining workers' dates of birth with their quarterly earnings, each worker's age-specific employment status is identified as either employed (E) or nonemployed (N) - where each worker's

[7] I have also examined the impact of the extension of the FRA from age 65 to 66, but found little impact on the timing of workers' retirement from the labor force at ages other than 65. I do find that much of the previously observed spike in the retirement hazard rate at age 65 shifts outward in lock-step with the two-month increases in the FRA that occurred between 2003-2008.

[8] Much of the analysis in this paper involves transitions between employment states, so the sample necessarily becomes truncated by one quarter depending on the type of transitions. Separations truncate the last quarter, whereas hires truncate the first quarter.

[9] Note that a single firm may have multiple SEINs, but I will generally refer to each SEIN as a firm.

age is measured in quarters since birth. Per Hyatt, McEntarfer, McKinney, Tibbets and Walton (2014), if a worker has multiple employers in the same quarter, then the worker's dominant employer is identified as whichever firm paid the worker the highest average quarterly real earnings in quarters t and $t-1$.

I also determine whether the worker transitioned between employment states in the given quarter, focusing on the following three types of employment state transitions. The first employment transition is "no earnings-gap" employer-to-employer transitions ($EE^{\text{no gap}}$) - which serve as a proxy for quits. Following Haltiwanger, Hyatt, Kahn and McEntarfer (2018), no earnings-gap employer-to-employer transitions are defined as instances where an individual transitions between dominant employers and the estimated gap in earnings between jobs is less than one month of the workers' average earnings in adjacent time periods. The second employment transition is employment-to-nonemployment transitions where the non-employment spell lasts at least 4-quarters (EN^{4+}). For older workers, EN^{4+} serve as a proxy for retirements. The EN^{4+} transitions of younger workers serve as a useful control variable for these involuntary transitions of older workers to persistent nonemployment. The third employment transition used is hires (H), where the hire may be an individual transitioning from non-employment or employment at another firm (a poaching). This hiring measure excludes two types of hires: 1) "recall" hires where the individual had worked at the hiring firm within the last four quarters, and 2) short-duration hires where the hired individual was no longer employed at the firm at the end of the hiring quarter.

3 Changing Retirement Behavior and OASI Rule Changes

Older workers' retirement from the labor force is strongly correlated with their age. Figure 1 shows the EN^{4+} transition rate of workers in 1990 and 2014 by age in quarter increments from 55-70. This figure evidences two facts of worker retirement. First, the empirical hazard rate of older workers' exit from the labor force is strongly correlated with their age, with spikes in the retirement rate when workers become eligible for i) OASI benefits at 62 (their OASI early retirement age), ii) Medicare benefits at age 65, and iii) their OASI Full Retirement Age (65 in 1990, 66 in 2014). Second, the retirement rate of older workers has declined dramatically between 1990 and 2014.

This large decline in older workers' retirement rates corresponds to substantial increases in the employment rate of older workers. Figure 2 shows the change (relative to 1990) in the age-specific employment rate of individuals who turned 62 in the given year and were employed at age 61.75.

The obvious take-away from this figure is that workers of all ages between 62-70 have increased their employment rate, with the increase being particularly large (14-16pp) for workers between the ages of 63-65. More subtly, the figure reveals three distinct periods of changes in the employment rate of older workers. During the 1990's, workers of all ages (63-72) appear to have had similar employment rate changes, with their employment rates rising between 5-8pp. Between 1999 and 2005, there was another dramatic change in older workers' employment rates, but these changes were concentrated among workers aged 63 to 66, who increased their employment rate by approximately 6pp. And since 2005, there have been only small changes in the employment rate of older workers.

By grouping together all workers of a particular birth cohort, Figure 2 masks substantial heterogeneity in the changing employment rate of workers based on the quality of the job that the worker held at age 61.75. The three panels of Figure 3 also plot the change (relative to 1990) in the employment rate of workers age 62-72, but differentiates workers by the rank of the workers' employers at age 61.75.[10] These three panels show that individuals working at low quality firms at age 61.75 had relatively small changes in their employment rate (they were already more likely to work until later ages), whereas individuals working at the highest quality firms at age 61.75 have increased their employment rate by between 20-25pp between ages 63-66. The fact that workers at high quality jobs evidence the largest increases in their employment rates has important implications for the EE transition rate, since, as discussed in Section 4, retirements from jobs high on the job ladder generate longer vacancy chains and more EE transitions.

In Section 4.1, I use the substantial variation across ages (within time) and over time (within age) to construct instrumental variables for estimating firms' replacement hiring rate. In order to exploit the changes over time in older workers' retirement behavior, it must be the case that these changes are exogenous to non-retirement factors that affect firms' replacement hiring rates. Thus, showing that two OASI rule changes have caused most of the change in older workers' employment rates helps demonstrate the validity of these instruments. Section 3.1 considers the impact of the increases in the Delayed Retirement Credit that occurred between 1990-2008 on older workers' employment rates. Section 3.2 examines the effect of the 2000 elimination of the Earnings Test for workers over their Full Retirement Age.

[10] I rank firms into employment-weighted quintiles according to the average full-quarter real earnings that they pay their workers.

3.1 Rule Change 1: Delayed Retirement Credit Increase

For every year after their FRA (up to age 70) that individuals wait to claim OASI, their monthly OASI benefit increases by the DRC amount. Prior to 1990, the DRC was 3.0%. Starting in 1990, the DRC was increased for individuals reaching their FRA in the given year by 0.5pp increments every two years until it reached 8.0% in 2008. While earlier studies have examined the impact of the increase in the DRC on the employment rate of workers in the 65-69 age category (Pingle (2006), Blau and Goodstein (2010)), the instrumental variable estimation of firms' replacement rates in Section 4.1 requires knowing how the increase in the DRC affects the employment rate of workers at each specific age (in quarters) between ages 62-70. To identify the effect of the increase in the DRC, I use a regression discontinuity analysis on individuals born immediately before versus after the DRC increase cutoff dates.

To estimate the effect of a 0.5pp increase in the DRC on employment rates at specific ages, I construct 15 distinct groups of individuals who were born in December or January for each of the 15 calendar year transitions between 1993 and 2008 (eight with DRC increases, and seven with no change in the DRC). The effect is identified by comparing the employment rate at each age a of individuals within each group who were born in the 30 days immediately before and after a jump in the DRC (which occurs between January 1st and 2nd). The regression discontinuity analysis allows for group-specific linear pre and post trends in the distance of an individual's day of birth from January 1st (pre) and 2nd (post). Including the years without any DRC increase controls for differences in the employment rate of individuals born in January versus December.

More specifically, I estimate a linear probability model for every age $a \in [62, 72]$ where the outcome variable is an indicator variable equal to one if individual i is employed at age a. For simplicity, this specification assumes that the effect of each 0.5pp increase in the DRC is the same,[11] and thus the explanatory variable of interest is the indicator variable (I_i) that equals one if the individual was born in January into one of the eight birth cohorts where the DRC increased at the calendar year transition. I estimate the following model for every age (in quarters) between 62 and 72:

[11] This assumption is unlikely to hold since the increases in the Full Retirement Age from 65 in 2002 to 66 in 2008 would have decreased the value of the higher DRC. Thus, I would expect that the DRC increases to have a larger effect on employment rates in the 1990's.

$$\mathbb{1}\left[\text{employed}\right]_i^a = \boldsymbol{\gamma}^{\boldsymbol{a}} I_i + \beta_1^a J_i + \sum_{g=1993}^{2008} \left(\beta_2^{g,a} d_i^g + \beta_3^{g,a} d_i^g R_i^{pre} + \beta_4^{g,a} d_i^g R_i^{post} \right) + \epsilon_i^a \tag{1}$$

The indicator variable J_i equals one if the individual was born in January. It controls for systematic differences in the employment rate of individuals born in January versus December and is identified by the seven calendar year transitions in which there was no increase in the DRC. The group-specific indicator variable (d_i^g) controls for changes in the age-specific employment that affect all individuals born around a given calendar year transition. Following Imbens and Lemieux (2008), the specification includes two local linear variables for each group. The pre-cutoff running variable, R_i^{pre}, is the distance of individual i's birthday from the DRC increase cutoff date (January 1st) if the individual was born on or before January 1st and zero if born after January 1st. The post running variable, R_i^{post}, is similarly the distance of individual i's date of birth from January 2nd ($DOB-$Jan 2^{nd}) if born on or after January 2nd, and zero otherwise.

The results of this regression discontinuity analysis are shown in Figure 4. Each 0.5 percentage point increase in the DRC caused the employment rate of workers between the ages of 64-69 to increase by between 0.75 to 2.0pp. The fact that there were nine such 0.5pp increases in the DRC between 1990 and 2008 suggests that the increase in the DRC is responsible for much of the increase in the employment rate of older workers.

3.2 Rule Change 2: Elimination of the Earnings Test For Older Workers

Prior to 2000, if individuals age 65-69 had claimed their OASI benefits and earned above an exemption threshold, then their monthly benefits were reduced by 33 cents for every dollar they earned above the threshold. In March of 2000, the Earnings Test was eliminated for individuals over their FRA (a more stringent Earnings Test still applies to individuals between ages 62 to their FRA).

To identify the effect of this change, I use a difference-in-difference strategy similar to that of Gelber, Jones, Sacks and Song (2017), where the control group is comprised of workers whose calendar year earnings at age 63 are below the anticipated exemption threshold that would apply at age 65. Because these individuals' earnings at age 63 are below the Earnings Test exemption threshold, it is unlikely that these individuals would have been constrained by the Earnings Test at age 65 (even if the Earnings Test had remained in effect, these individuals could have claimed OASI benefits without penalty). Thus, the elimination of the Earnings Test in 2000 is unlikely to have affected these individuals' employment decisions. Assuming that these individuals are unaffected by the elimination of the Earnings Test, they will serve as an effective control group to the extent that they respond to labor market shocks in a similar fashion as higher earning workers (who are the treated group).

More specifically, I estimate a set of age-specific linear probability models, for ages $a \in [63, 70]$, where the outcome variable is an indicator variable equal to one if individual i is employed at age a. There are two explanatory variables. One, an indicator variable if the individuals' earnings at age 63 are above the (anticipated) Earnings Test threshold that would (or did) apply to the worker at age 65 (Y_i^{above}). And two, an age-specific treatment indicator variable T_i^a that equals one if either i) the earnings test no longer applies to individuals of age a (if $a >= 65$), or ii) an individual of age a knows that the Earnings Test post FRA has been eliminated (for $a < 65$). With these variables, I estimate the following linear probability model for each age in quarters $a \in [63, 70]$:

$$\mathbb{1}\left[employed\right]_i^a = \beta_0^a + \beta_1^a Y_i^{above} + \beta_2^a T_i^a + \boldsymbol{\theta^a} T_i^a Y_i^{above} + \epsilon_i^a \tag{2}$$

The coefficient of interest is θ^a. Figure 5 shows the results of this estimation when individuals turning age 65 in 2000 (when the Earnings Test was eliminated) are compared against workers turning 65 in 1997, 1998, and 1999. Note that the first discussions of eliminating the Earnings Test occurred in January 1999 when President Clinton proposed it in his State of the Union address. I draw two conclusions from this figure regarding the impact of the 2000 elimination of the Earnings Test on older workers' employment: first, the elimination caused individuals turning 65 in 2000 to increase their employment rate at age 66 by 2.5pp relative to similar workers in 1998 (before any announcement of legislation to eliminate the Earnings Test); and second, this effect persisted through at least age 68 (shown from the comparison with workers turning age 65 in 1997).

The elimination of the Earnings Test for workers over their Full Retirement Age (65 years old at the time), also affected the employment decisions of workers who had yet to reach their Full Retirement Age. That younger workers increased their employment rate in anticipation of not being constrained by the Earnings Test at age 65 can be seen in Figure 6. This figure compares workers who were 63, 64, or 65 years old in 2000 against similar workers reaching age 65 in 1998. Anticipating that the Earnings Test would no longer bind at age 65, workers turning age 63 in 2000 had a 4.5pp higher employment rate at age 66 relative to similar workers in 1998. This last finding may explain why some earlier difference-in-differences studies found little effect of the Earnings

Test on the extensive margin of labor supply since they used 62-64 year-old workers as the control group in their D-in-D estimations (Song and Manchester (2007)).

4 Vacancy Chains, Replacement Hiring, and New Hire Poaching

In order to explore the effect of retirements on EE transitions, I utilize a vacancy chain model similar to that described in Akerlof, Rose, Yellen, Ball and Hall (1988). A worker's decision to retire from employment can generate many EE transitions by initiating a vacancy chain. If a firm chooses to fill the job vacated by the retiring worker with an already employed worker, then this both generates an EE transition and creates a vacancy at the new hire's previous employer. This vacancy chain continues until either a firm hires a replacement worker from non-employment or a firm chooses not to replace a retiring/quitting worker.

In a simple vacancy chain model that ignores differences across the job ladder, the length of a typical vacancy chain is determined by three parameters: i) firms' replacement hiring rate in response to a worker retirement (ρ_R), ii) the probability that the replacement hire is poached from another firm (π), and iii) firms' replacement hiring rate in response to quits (ρ_Q) - which perpetuates the vacancy chain. The number of hires generated by a retirement can be expressed as the following infinite series:

$$H^R = \underbrace{\rho^R}_{H_1} + \overbrace{\underbrace{\rho^R \pi}_{EE_1} \rho^Q}^{H_2} + \overbrace{\underbrace{\rho^R \rho^Q \pi^2}_{EE_2} \rho^Q}^{H_3} + ... \tag{3}$$

For a vacancy chain that is generated by a retiring worker, H_1 denotes the first stage of the vacancy chain, namely the number of replacements that a firm hires in order to fill the retiring worker's position. The second stage of the vacancy chains comes from these replacement hires generating EE transitions (EE_1), the number of which are determined by the new hire poaching probability π. The previous employers of these poached workers will replace these poached workers at the rate ρ_Q, which determines the number of replacement hires that occur at the second stage of the vacancy chain (H_2). Some number of these second stage replacement hires will be poached from other employers, thus perpetuating the vacancy chain. As with any infinite series, so long as $\rho_Q \pi < 1$, the number of hires generated by a retirement can be simply expressed as:

$$H^R = \frac{\rho_R}{1 - \rho_Q \pi} \tag{4}$$

And the number of EE^R transitions generated by a retiring worker is simply:

$$EE^R = \pi H^R = \frac{\pi \rho_R}{1 - \rho_Q \pi} \tag{5}$$

In a more complicated vacancy chain model that allows for the parameters to differ across the job ladder, a similar infinite series representation of the number of hires and EE transitions exists. Consider a job ladder with n distinct ranks of firms. Let $\boldsymbol{r}$ be a $1 \times n$ vector with a value of one in the position associated with the rank of the retiring worker's firm and zeros in all other positions. The replacement hiring rates become $n \times n$ diagonal matrices ($\boldsymbol{\rho_R}$ and $\boldsymbol{\rho_Q}$), with the firm rank-specific replacement hiring rates along the diagonal. Lastly, the poaching probability becomes a $n \times n$ matrix ($\boldsymbol{\Pi}$) where the j^{th} row of the matrix denote the probabilities that a firm of rank j fills a vacancy by poaching a worker from a firm of rank k, where k corresponds to the column. The resulting infinite series expression of the number of hires generated by a retirement is:

$$\boldsymbol{H^R}(\boldsymbol{r}) = \boldsymbol{r}\boldsymbol{\rho_R}\left(\boldsymbol{I} + \boldsymbol{\rho_Q}\boldsymbol{\Pi} + \boldsymbol{\rho_Q^2}\boldsymbol{\Pi^2}\right) \tag{6}$$

And the number of EE Transitions generated by a retirement is:

$$\boldsymbol{EE^R}(\boldsymbol{r}) = \boldsymbol{H^R}(\boldsymbol{r})\boldsymbol{\Pi} = \boldsymbol{r}\boldsymbol{\rho_R}\left(\boldsymbol{I} + \boldsymbol{\rho_Q}\boldsymbol{\Pi} + \boldsymbol{\rho_Q^2}\boldsymbol{\Pi^2}\right)\boldsymbol{\Pi} \tag{7}$$

While there is no simple algebraic expression for this infinite series of matrices, the number of hires (and thus EE transitions) generated by a retirement from a given firm rank can still be estimated by recursively calculating the above infinite series until it converges. The next three subsections describe the estimation of the three key parameters that are needed in order to calculate the length of the vacancy chains generated by worker retirements.

4.1 Estimating Replacement Hiring Rates

Calculating the length of vacancy chains requires estimates of the replacement hiring rates - the number of new workers that a firm hires in response to a retirement or a quit of an existing worker. To demonstrate the strong relationship between quits, retirements, and hires, Figure 7 plots the $EE^{\text{no gap}}$ and EN^{4+} transition rates by the age of the worker at single worker firms. These rates are then compared against the hiring rate at the single worker firms in the next two periods. This figure shows the strong correlation between the age-specific quit rate and the subsequent hiring rate (particularly at younger ages). For retirements, the age-specific spikes in the EN^{4+} transition rate at 62 and 65 are clearly mirrored in the subsequent hiring rate at the firms.

To estimate firms' replacement hiring rate over the x quarters following a worker's retirement or poaching, I examine the response of firms' 4-quarter change in no-recall hires-per-employee ($\Delta_4 H_{k,t,t+x}$) to changes in the retirements-per-worker ($\Delta_4 R_{kt}$) or quits-per-worker ($\Delta_4 Q_{kt}$). Using the 4-quarter change controls for firm-specific seasonality in hiring, quits, and retirements. The replacement hiring rates for retirements ($\boldsymbol{\rho_R}$) and quits ($\boldsymbol{\rho_Q}$) can be estimated using the following model of the change in the hiring rate at firm k:

$$\Delta_4 H_{k,t,t+x} = \Delta_4 R_{kt} \boldsymbol{\rho_R} + \Delta_4 Q_{kt} \boldsymbol{\rho_Q} + X_{kt}\beta + \Delta_4 \epsilon_{kt} \tag{8}$$

where X_{kt} is a set of control variables that include the change in EN^{4+} transitions of workers age 25-59, the share of employees who are age 60-80, and time-period specific fixed effects for each 2-digit NAICS industry. Since the LEHD data does not identify voluntary retirements and quits at a firm, I must instead proxy for these variables using older workers (age 60-80) transitions from employment to persistent nonemployment (EN^{4+}) for retirements and workers' no-earnings gap EE transitions to other employers ($EE^{\text{no gap}}$) for quits. Thus, I estimate:

$$\Delta_4 H_{k,t,t+x} = \Delta_4 EN^{4+}_{kt} \rho_R + \Delta_4 EE^{\text{no gap}}_{kt} \rho_Q + X_{kt}\beta + \Delta_4 \epsilon_{kt} \tag{9}$$

A direct OLS estimation of the model described in Equation 9 could not be interpreted as a causal estimate of ρ_R and ρ_Q for two key reasons. First, these estimates would be subject to significant bias since I do not directly observe voluntary retirements (R_{kt}) and quits (Q_{kt}), but instead must proxy for them using the EN^{4+} and $EE^{\text{no gap}}$ transition rates. These proxies will fail to identify some voluntary quits and retirements - generating attenuation bias due to classical measurement error. Even more critically, these proxies will identify some involuntary layoffs as quits and retirements. The inclusion of these involuntary layoffs generates a negative bias because involuntary layoffs are negatively correlated with unobserved job creation, and the unobserved job creation is positively correlated with hires.

The second issue with a direct OLS estimation of Equation 9 is that changes in workers' retirement and quit behavior at firm k are likely to be correlated with unobserved components of the error term. Most obviously, changes in job creation at the firm are included in the error term.

Since workers' quit and retirement decisions can be affected by the firm's growth prospects (which drives job creation), the observed quits and retirements are correlated with the error term. This generates bias of an indeterminant sign.

To address both measurement error and omitted variable bias, I construct two share-shift (Bartik-style) instrumental variables that approximate distinct components of the change in firm retirements (quits). The first instrumental variable (R^{IV-A}) exploits variation across time in the retirement (quit) rate of each specific age. The second instrumental variable (R^{IV-B}) exploits the substantial variation in retirement (quit) rates across ages within a given time period. The relevance of these two instrumental variables can seen by decomposing the retirements at the firm into the sum of the interactions between the firm's start-of-period employment level for every age (a) and sex (s) combination ($E_{k,t}^{a,s}$), and the firm-specific retirement rate of that age-sex group ($r_{k,t}^{a,s}$). More specifically, the following identity holds:

$$\Delta_4 R_{kt} = \sum_{a=60*4}^{80*4} \sum_{s=M,F} E_{k,t}^{a,s} r_{k,t}^{a,s} - \sum_{a=60*4}^{80*4} \sum_{s=M,F} E_{k,t-4}^{a,s} r_{k,t-4}^{a,s} \tag{10}$$

After some algebra, this can be rewritten as the sum of two distinct components:

$$\Delta_4 R_{kt} = \underbrace{\sum_a \sum_s E_{k,t-4}^{a,s} \left(r_{k,t}^{a,s} - r_{k,t-4}^{a,s} \right)}_{A} + \underbrace{\sum_a \sum_s \left(E_{k,t}^{a,s} - E_{k,t-4}^{a,s} \right) r_{k,t}^{a,s}}_{B} \tag{11}$$

The two instrumental variables, R^{IV-A} and R^{IV-B}, are designed to directly relate to the A and B components of the change in retirements. Considering the A component first, if the firm-specific retirement rates ($r_{k,t}^{a,s}$) are replaced with the leave-one-out national retirement rate for a specific age-sex combination ($r_{\boldsymbol{n},t}^{a,s}$), then this becomes the standard Bartik-style share-shift instrumental variable:

$$R_{kt}^{IV-A} = \sum_a \sum_s E_{k,t-4}^{a,s} \left(r_{\boldsymbol{n},t}^{a,s} - r_{\boldsymbol{n},t-4}^{a,s} \right) \tag{12}$$

The validity of R^{IV-A} relies on the the error term in Equation 9, $\Delta_4 \epsilon_{kt}$, being uncorrelated with both year-over-year changes in the national retirement rate and the one-year lag in the level of employment for each age-sex combination. The analysis in Section 3 of the impact of OASI rule changes on older workers' retirement rates demonstrates that the changes in the national retirement rates are plausibly exogenous to the error component of firms' replacement hiring rates. To address the concern that the one-year lag of the employment level of each age-sex combination is correlated with the error term, I include the share of older workers at the firm as a control variable. This is equivalent to a uniform weighting of older workers by age, and thus ensures that IV^A exploits only variation due to differences in retirement rates across age and time.

The instrumental variable corresponding to B, the second component of the change in the retirement rate, uses the fact that workers of age a in period $t-4$ will be age $a+4$ in period t. Thus, the firms' employment level of a particular age a and sex s in period t can be approximated by the employment level of workers of age $a-4$ in period $t-4$.[12] Using this approximation and performing some algebra, the B component of the 4-quarter change in retirements can be approximated by:

$$B \approx \sum_{a=60*4}^{79*4} \sum_{s} E_{k,t-4}^{a,s} \left(r_{k,t}^{a+4,s} - r_{k,t}^{a,s} \right) + \sum_{a=59*4}^{59.75*4} \sum_{s} E_{k,t-4}^{a,s} r_{k,t}^{a+4,s} - \sum_{a=79.25*4}^{80*4} \sum_{s} E_{k,t-4}^{a,s} r_{k,t}^{a,s} \tag{13}$$

The instrumental variable R^{IV-B} simply replaces the firm-specific retirement rates for a given age-sex combination with the leave-one-out national retirement rate for the same age-sex combination. Namely,

$$\Delta R_{k,t}^{IV-B} = \sum_{a=60*4}^{79*4} \sum_{s} E_{k,t-4}^{a,s} \left(r_{\boldsymbol{n},t}^{a+4,s} - r_{\boldsymbol{n},t}^{a,s} \right) + \sum_{a=59*4}^{59.75*4} \sum_{s} E_{k,t-4}^{a,s} r_{\boldsymbol{n},t}^{a+4,s} - \sum_{a=79.25*4}^{80*4} \sum_{s} E_{k,t-4}^{a,s} r_{\boldsymbol{n},t}^{a,s} \tag{14}$$

As can be seen from Equation 14, the R^{IV-B} instrumental variable exploits the variation in the retirement rate across ages within a given period that is shown in Figure 1. Two notes regarding this variation. First, the difference in the age-specific retirement rate from its one-year older value is only significant at ages immediately around workers' early retirement age (62), their Medicaid eligibility age (65), and their full retirement age (65-66). Given these ages are fixed by policy, it is plausible that differences in the age-specific retirement rates are exogenous to the error component of firms' replacement hiring rate. Second, the fact the variation across ages is negligible for other ages means that the R^{IV-B} instrumental primarily identifies a local average treatment effect for the effect on firm hiring of workers who choose to retire at ages 62, 65, or 66.

For the validity of this instrumental variable, the same concern discussed for R^{IV-A} regarding the exogeneity of the lagged age-sex employment levels applies to R^{IV-B}. Thus, including the firm-specific employment share of older workers (age 60-80) as a control variable ensures that the

[12]This approximation will be inaccurate to the extent that new workers are hired in periods $t-3$ to $t-1$ who will be age a in period t or workers of age $a-4$ in period $t-4$ separate from the firm.

instrumental variable only exploits variation in the retirement rate across ages.

A similar pair of instrumental variables can be constructed for the change in quits ($\Delta_4 EE^{\text{no gap}}$) using worker ages and tenures. For these instrumental variables, I differentiate employment by both age and tenure at the firm and then use national level variation in quit rates to construct the share-shift instruments. Note that the there is more reason to be concerned about the validity of the instrumental variables for the change in quits since national level fluctuations in the age-specific quit rate are more likely to be correlated with firm replacement hiring rates (due to cyclicality) and the tenure of a firm's workers is also likely to be correlated with the error term (which includes job creation).

I estimate two separate 2SLS procedures, one for the causal estimate of the change in retirements and the other for the change in quits. The first-stage results of these estimates are shown in Table 1. For both quits and retirements, the instrumental variables are strong instruments and the coefficient estimates are of the correct sign. Furthermore, the two instrumental variables appear to be exploiting different variation in retirement/quit rates as there is not an issue with multicolinearity between the instruments.

Figure 8 plots the coefficient estimates for both the OLS and 2SLS estimates of the effect of retirements and quits on firm's replacement hiring. The figure plots the no-recall hiring rate over $t + x$ periods where $x \in [0, 3]$. The OLS coefficient estimates for both quits and retirements are statistically insignificant and near zero, whereas the 2SLS estimates indicate that the average retirement (poached worker) generates approximately 2.1 (1.5) new hires over the course of the year following the retirement (poaching). The dramatic difference in the OLS and 2SLS estimates implies that using proxies for the quit and retirement rates generates severe attenuation and negative bias in the OLS coefficient estimates.

4.2 New Hiring Poaching Rates

The third parameter that is necessary to estimate the length of the vacancy chains generated by a retirement is the probability that a new hire is poached from another employer. Using the $EE^{\text{no gap}}$ transitions as the measure of poaching, the LEHD data allows for a simple calculation of the new hire poaching probability. I find that the new hire poaching probability differs significantly across the job ladder, increasing almost linearly from 33% at firms ranked in the lowest quintile to 53% at firms in the top quintile.

I also find that there are significant differences across the job ladder in where poached workers are hired from. As shown in Figure 9, firms are most likely to poach workers from similarly ranked firms and they rarely poach workers from higher ranked firms. This has important implications for the length of vacancy chains since it implies that vacancy chains will be longer high up the job ladder where firms are more likely to hire employed workers and these employed workers are more likely to come from firms high on the job ladder. Workers retiring from low ranked firms are much more likely to be replace by a nonemployed worker or a worker at a low ranked firm (where the vacancy chain is less likely to persist).

4.3 Length of Vacancy Chains Across the Job Ladder

Using the formula for the length of vacancy chains described in Equation 7 and the parameter estimates from Sections 4.1 and 4.2, I calculate the length of vacancy chains generated by retirements across the job ladder. The implied vacancy chain lengths increase almost linearly from 1.6 $EE^{\text{no gap}}$ transitions generated by a retirement at a low ranked firm to 3.4 $EE^{\text{no gap}}$ transitions generated at firms at the top of the job ladder.

There are also differences over the job ladder in where the EE transitions are generated. As shown in Figure 10, retirements at top ranked firms both generate more $EE^{\text{no gap}}$ transitions and these $EE^{\text{no gap}}$ transitions are more equally distributed across the job ladder. Retirements from the lowest ranked firms, on the other hand, create fewer $EE^{\text{no gap}}$ transitions and these are concentrated among similarly low ranked firms. Table 2 shows the differences in the number of $EE^{\text{no gap}}$ transitions at firms of given origin firm rank generated by a retirement from a firm of a give destination rank.

4.4 Changing Retirements and the Secular Decline in Worker Reallocation

To evaluate the impact of older workers' decisions to delay their retirement from the labor on the aggregate EE transition rate of younger workers (age 25-59), I first estimate the number of missing retirements that would have occurred at each level of the job ladder if the age-sex retirement rates at each rung of the job ladder had held constant at their 1990 levels.[13] The number of "missing" $EE^{\text{no gap}}$ transitions is calculated by multiplying the number of "missing"retirements at each firm rank by the estimates from Section 4.3 of the number of $EE^{\text{no gap}}$ transitions generated by a retirement. These "missing" $EE^{\text{no gap}}$ transitions are then added to the actual $EE^{\text{no gap}}$

[13]The calculation of the missing retirements adjusts the actual age-sex-firm rank employment level in each period to account for the fact that the employment level would have been lower if workers had retired in previous periods at the higher 1990 retirement rates.

transitions for the period. Figure 11 shows the comparison of the actual $EE^{\text{no gap}}$ transitions to the counterfactual $EE^{\text{no gap}}$ transitions if retirement rates had remained at their 1990 levels. This estimate implies that approximately 30% of the secular decline in the EE transition rate can be explained by the delaying of older workers' retirement from the labor force.

5 Conclusion

This paper documents the increasing employment rate of older workers since 1990 and explores the implications of this delayed retirement trend for the employer-to-employer transition rate, a measure of worker reallocation. The median 62 year-old worker in 2008 retired nearly two years later relative to the similar median worker in 1990. This paper provides evidence that much of this trend towards working at older ages was caused by two rule changes in the Social Security OASI program: the gradual increase in the Delayed Retirement Credit from 1990 to 2008 and the 2000 elimination of the Earnings Test for workers over their Full Retirement Age.

Older workers' decisions to delay their retirement from the labor force have a multiplier effect on the EE transition rate because of vacancy chains - firms' tendency to fill job vacancies with already employed workers, who must quit their previous job, thus creating additional vacancies. The three key parameters determining the length of vacancy chains are i) firms' replacement hiring rate in response to a worker retirement, ii) the probability that these replacement hires are poached from other employers, and iii) firms' replacement hiring rate in response to a quitting worker. Using a set of share-shift instrumental variables, I estimate the causal effect of worker retirements (quits) on firm hiring, finding that the average retirement (poached worker) generates approximately 2.1 (1.5) new hires over the course of the year following the retirement (poaching). Combining these estimates with the fact that firms higher up the job ladder are more likely to poach their new hires away from other firms (versus hiring workers from nonemployment), I show that the average retirement from a firm at the bottom of the job ladder generates a vacancy chain with 1.6 EE transitions, whereas retirements from the top of the job ladder generate twice as many EE transitions. These estimates imply that approximately 30% of the secular decline in the EE transition rate can be explained by older workers choosing to delay their retirement from employment.

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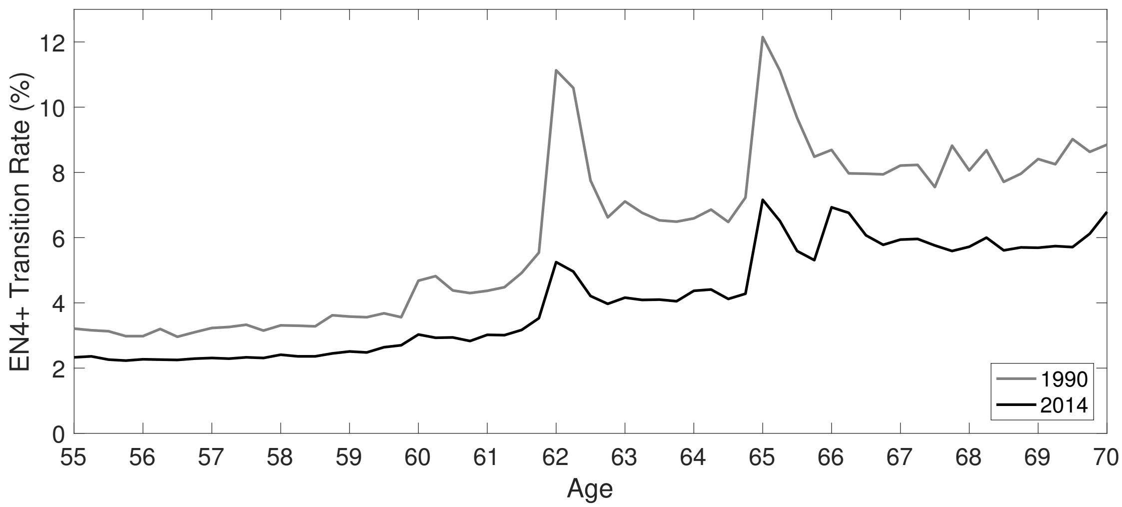
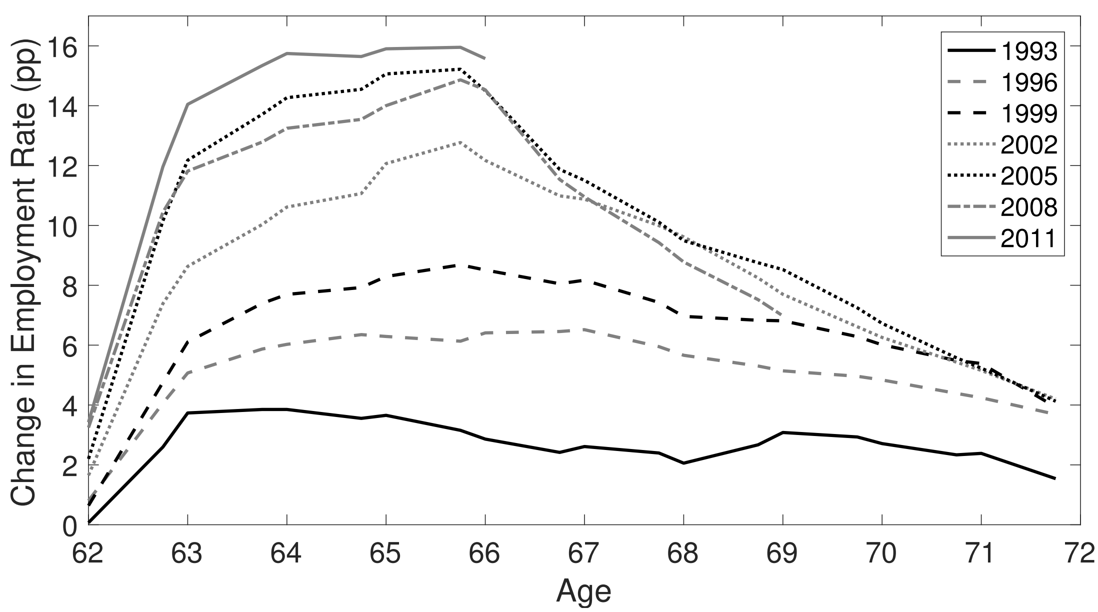
Figures

Figure 1: EN^{4+} Transition Rate by Age in Quarters



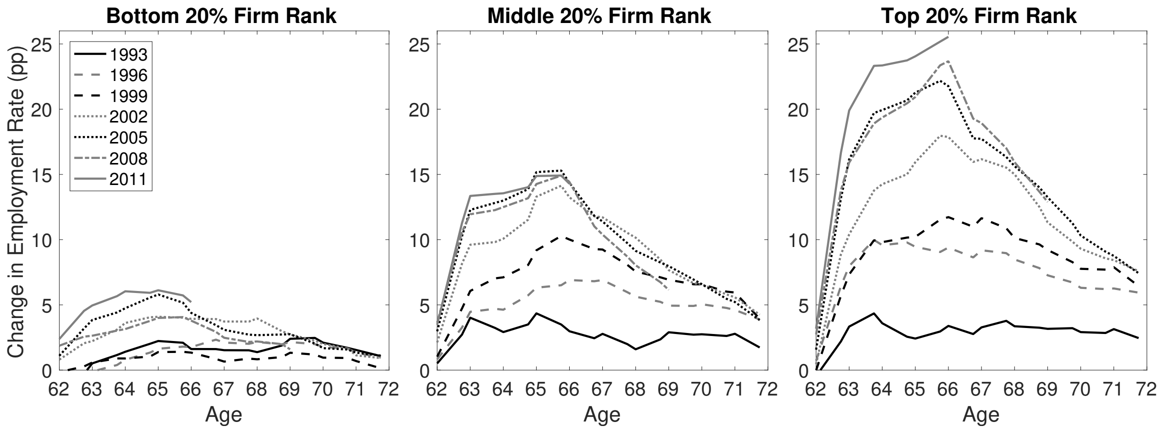
Notes: Employment to persistent nonemployment transition rate of workers of given age in 1990 and 2014. The nonemployment spell must last at least 4 quarters. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 2: Change in the Conditional Age-Specific Employment Rate



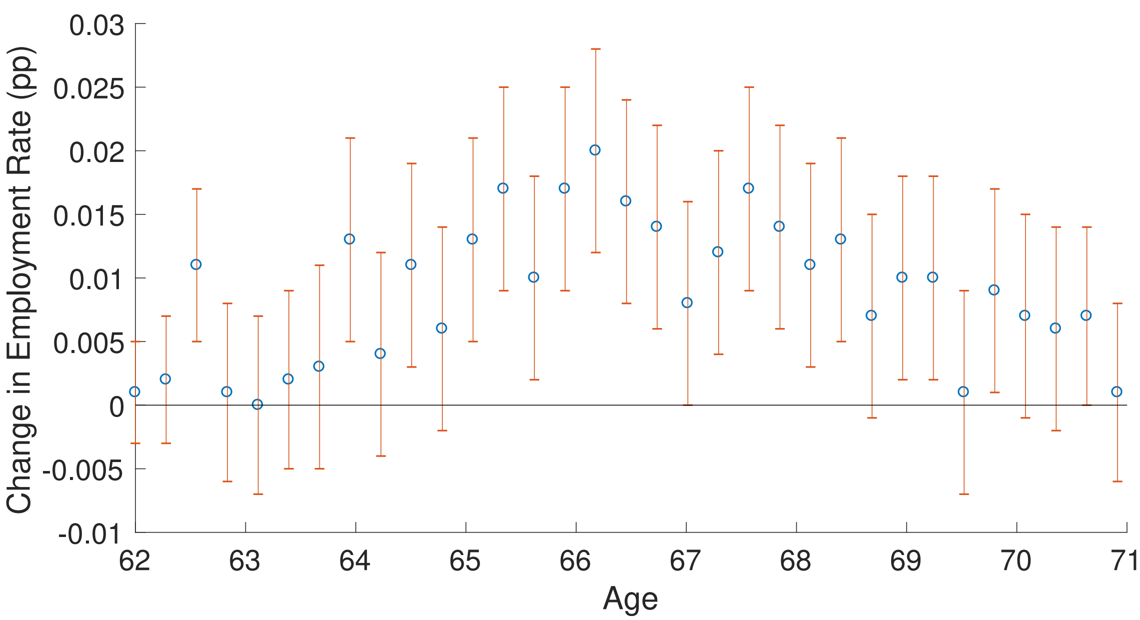
Notes: Change relative to the 1990 cohort in the probability that a worker turning 62 in a particular year is still employed at the given age, conditional on the worker being employed at age 61.75 years old. Displayed are linear spline estimates for the five quarters between and including every workers' birthday between ages 62-72. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 3: Change in the Conditional Age-Specific Employment Rate By Firm Rank at age 61.75



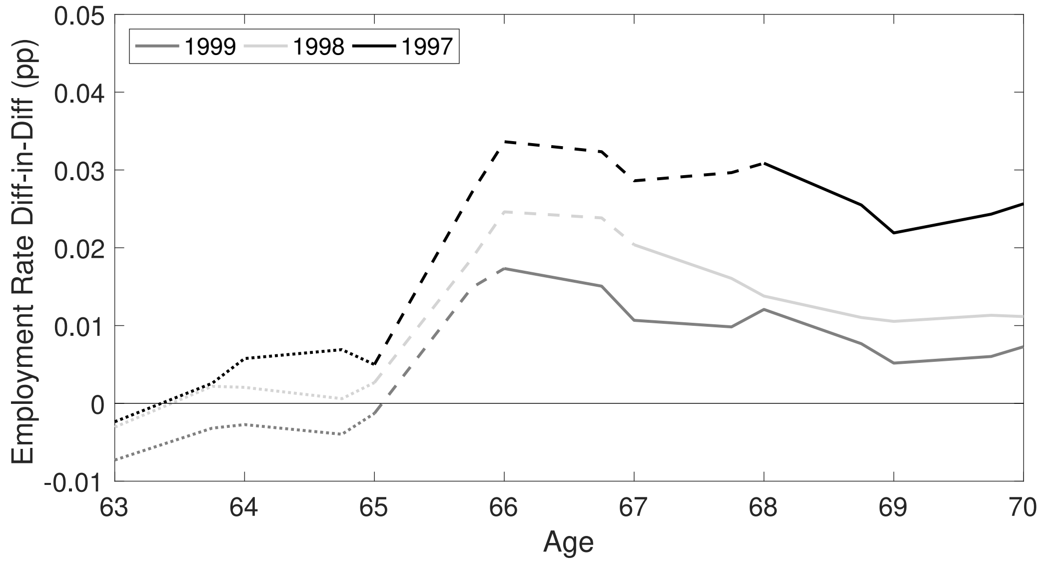
Notes: Change relative to the 1990 cohort in the probability that a worker turning 62 in a particular year is still employed at the given age, conditional on the worker being employed at a firm of the given rank at age 61.75 years old. Firm rank is measured by the employment-weighted quintile of the average full-quarter real earnings paid by the firm. Displayed are linear spline estimates for the five quarters between and including every workers' birthday between ages 62-72. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 4: Change in Employment Rate Caused by 0.5pp Increase in Delayed Retirement Credit



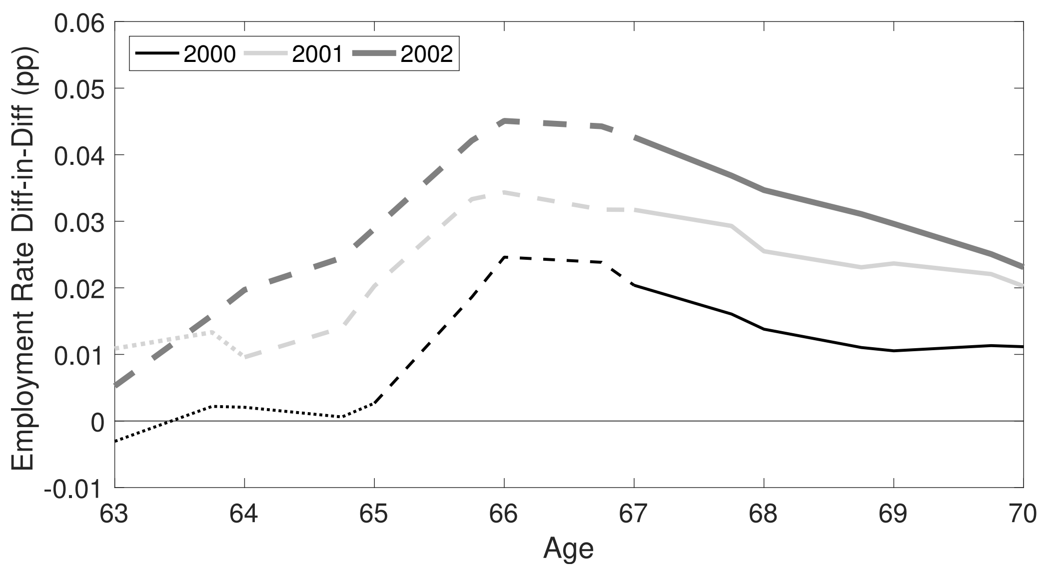
Notes: Difference-in-differences regression discontinuity estimates of γ^a, the effect of a 0.5 percentage point increase in the Delayed Retirement Credit on the probability that a worker who was employed at age 61.75 remains employed at a given age a. Circles are point estimates. Bars are 95% confidence intervals. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 5: Change in 65+ Employment Rate Caused by Elimination of Earnings Test



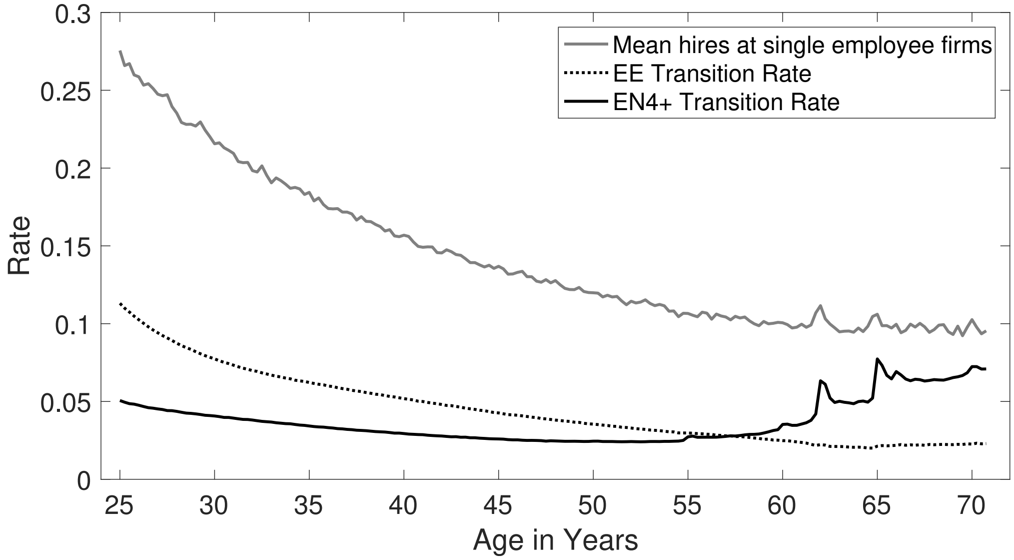
Notes: Difference-in-differences estimate. The first difference is between the employment rate of individuals who had earnings above versus below the earnings test threshold at age 63. The second difference is between cohorts turning age 65 in year 2000 (coincident with the Earnings Test elimination) versus years 1999, 1998, and 1997. The dotted lines indicated the pre-treatment period. The dashed lines indicate the period in which only the 2000 cohort is treated. The solid lines indicate the period in which no cohorts are affected by the Earnings Test. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 6: Change in 63+ Employment Rate Caused by Elimination of Earnings Test



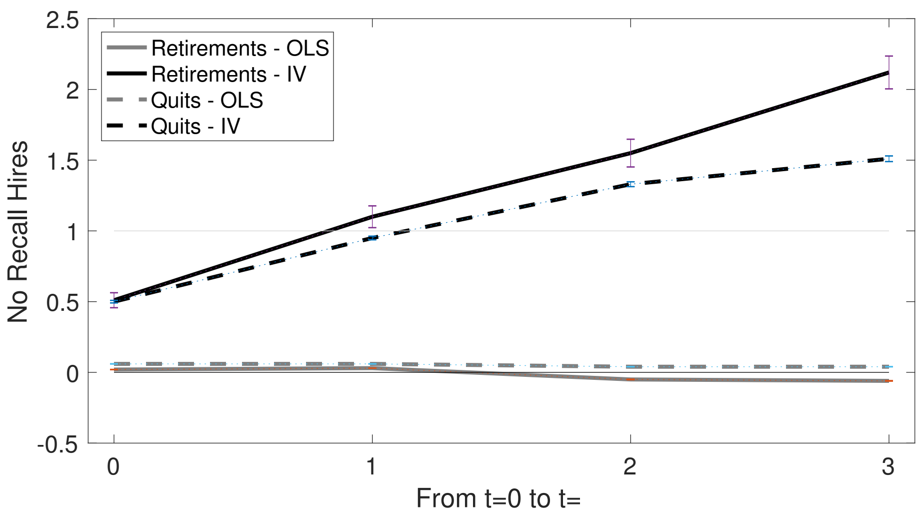
Notes: Difference-in-differences estimate. The first difference is between the employment rate of individuals who had earnings above versus below the earnings test threshold at age 63. The second difference is between cohorts turning age 65 in years 2000, 2001, or 2002 versus years 1998. The dotted lines indicated the pre-treatment period. The dashed lines indicate the period in which the 1998 cohort is untreated. The solid lines indicate the period in which no cohorts are affected by the Earnings Test. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 7: Age-Specific $EE^{\text{no gap}}$, EN^{4+}, and Hiring Rates at Single-Worker Firms



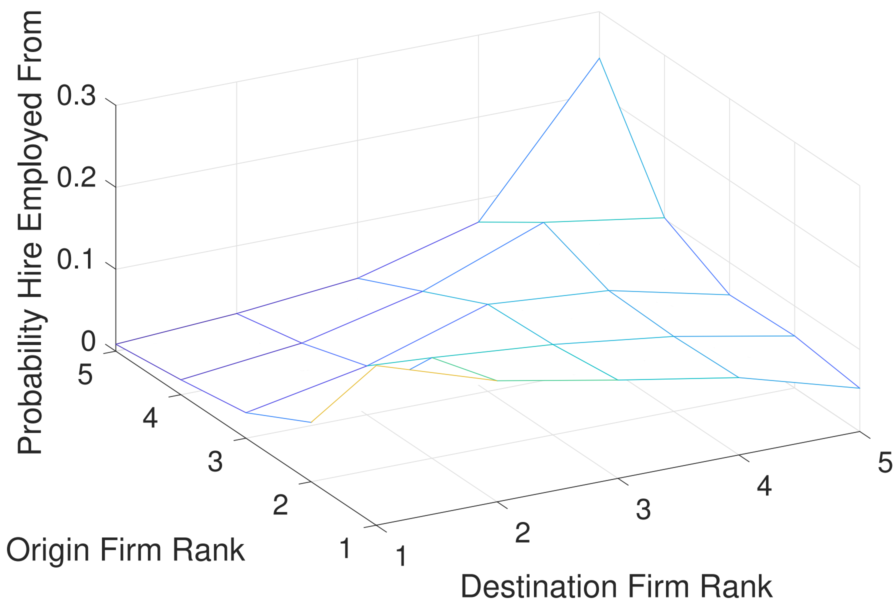
Notes: Shown are the transition rates at single-worker firms based on the age of the worker. The three transition rate measures are: 1) total hires per employee at the worker's employer in period t and $t+1$; 2) no earnings gap EE transitions in period t (($EE^{\text{no gap}}$); and 3) employment-to-persistent nonemployment transitions (EN^{4+}). Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 8: 2SLS and OLS Estimates of Replacement Hiring Coefficients ρ_R and ρ_Q



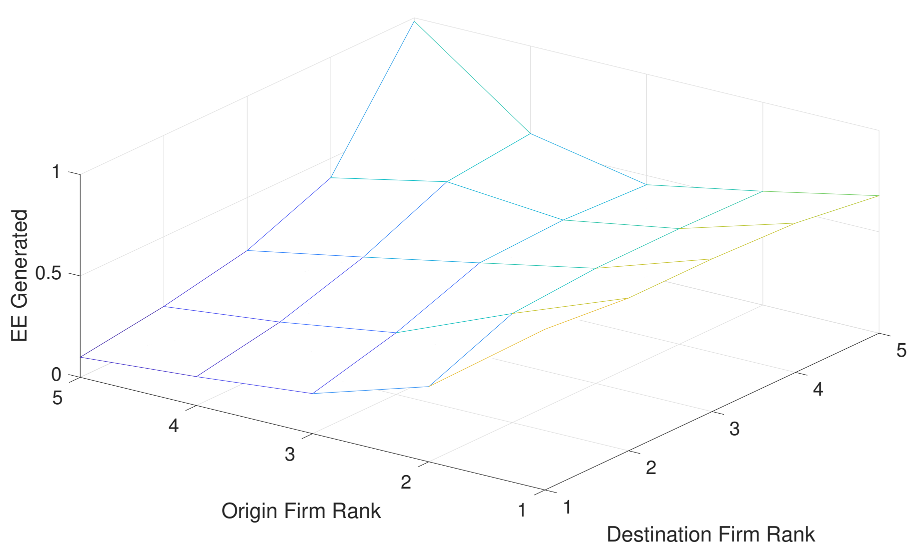
Notes: OLS and 2SLS estimation results regressing the change in no-recall hires in period $t = 0$ to x on the change in employment-to-persistent nonemployment (EN^{4+}) transitions of workers age 60-80 and no earnings-gap EE transitions ($EE^{\text{no gap}}$. The IV estimates use both IV-A and IV-B instruments for the endogenous variable of interest (quits or retirements), but separate models are estimated for quits and for retirements. The control variables include share of old workers at the firm, the EN^{4+} transition rate of younger workers at the firm, and industry-specific time period fixed effects. Estimation performed on 25.5M SEIN-quarter observations, with observations weighted by total start-of-quarter employment at the SEIN in period $t-4$. All reported 2SLS coefficient estimates are statistically significant at the .001% level when using robust standard errors clustered at the SEIN-level. None of the reported OLS coefficient estimates are statistically significant at the 0.5% level. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Figure 9: Probability New Hire at Destination Firm Rank is Poached from Origin Firm Rank



Notes: Probability new hire by firm with destination firm rank poaches worker from a firm with origin firm rank. Source: LEHD microdata 32 state sample from 1998:Q1-2015:Q3.

Figure 10: Vacancy Chain Length Over the Job Ladder

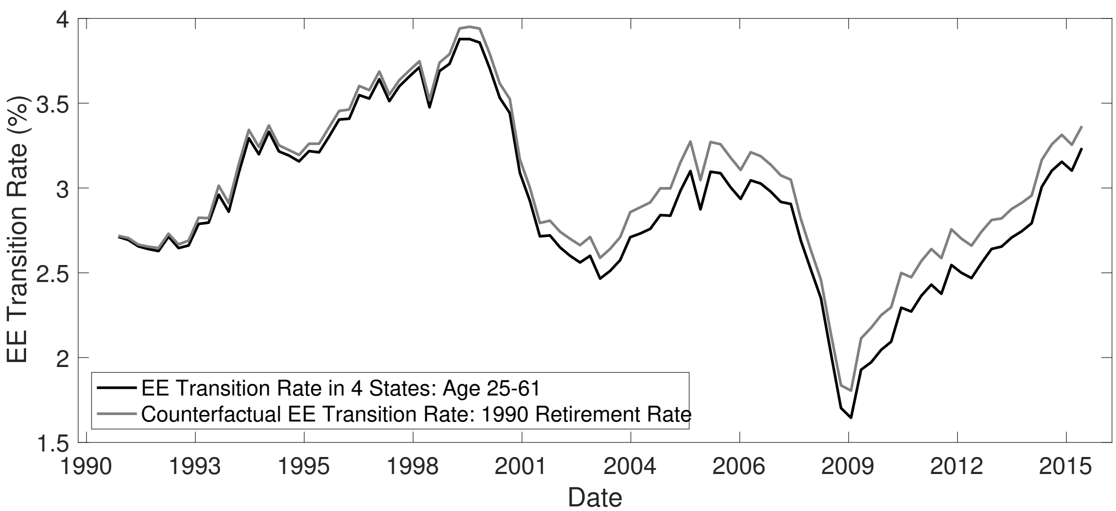


Notes: Calculated based on the vacancy chain model of:

$$\boldsymbol{EE^R(r) = r\rho_R \left(I + \rho_Q \Pi + \rho_Q^2 \Pi^2\right) \Pi}$$

With the $\boldsymbol{\rho_R}$, $\boldsymbol{\rho_Q}$, and $\boldsymbol{\Pi}$ parameters calculated as described in Sections 4.1 and 4.2. Source: LEHD microdata 32 state sample from 1998:Q1-2015:Q3.

Figure 11: Counterfactual EE Transition Rate at 1990 Retirement Rates



Notes: Comparison of actual $EE^{\text{no gap}}$ transition rate and the counterfactual transition rate if older workers' retirement rates had remained at their 1990 level.

Tables

Table 1: **First-Stage Estimation Results**

	$EN4+_{kt}$	$EE_{kt}^{SQ+AQ,nogap}$
IV-A	0.39***	0.54***
	(0.004)	(0.003)
IV-B	0.48***	0.56***
	(0.008)	(0.005)
F-Statistics	5853	16515

First-stage estimation results regressing either the change in employment-to-persistent nonemployment (EN^{4+}) transitions of workers age 60-80 or the change in no earnings-gap employer-to-employer transitions of workers ($EE^{\text{no gap}}$)on the instrumental variables and control variables. The IV estimates use both IV-A and IV-B instruments for the endogenous variable of interest (quits or retirements), but separate models are estimated for quits and for retirements. The control variables include the share of old workers at the firm, the EN^{4+} transition rate of younger workers (age 25-59) at the firm, and industry-specific time period fixed effects. Estimation performed on 25.5M SEIN-quarter observations, with observations weighted by total start-of-quarter employment at the SEIN in period $t-4$. All reported coefficient estimates are statistically significant at the .001 level when using robust standard errors clustered at the SEIN-level. Source LEHD microdata - 4 state sample from 1990:Q1-2015:Q3.

Table 2: **Comparison of EE Generated Over the Job Ladder**

	Origin Rank				
Destination Rank	**1**	**2**	**3**	**4**	**5**
1	**0.79**	0.37	0.20	0.14	0.10
	Difference from Retirement at Rank 1 Firm				
3	-0.04	0.20	**0.26**	0.20	0.14
5	-0.11	0.19	0.25	0.43	**0.88**

Number of $EE^{\text{no gap}}$ transitions generated at firms of origin rank by a retirement at a firm of given destination rank. The Destination Rank 3 and 5 values are differences from the $EE^{\text{no gap}}$ generated by a Rank 1 firm.