Career Spillovers in Internal Labor Markets*

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Abstract

This paper studies career spillovers across workers, which arise in firms with limited promotion opportunities. We exploit a 2011 Italian pension reform that unexpectedly tightened eligibility criteria for the public pension, leading to sudden, substantial, and heterogeneous retirement delays. Using administrative data on Italian private-sector workers, the analysis leverages cross-firm variation to isolate the effect of retirement delays among soon-to-retire workers on the wage growth and promotions of their colleagues. We find evidence of spillover patterns consistent with older workers blocking the careers of their younger colleagues. These effects are present only in firms with limited promotion opportunities.

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1 Introduction

Workers of every generation fear that their elders are holding their careers back. Millennials worry that their careers are "stalled because older employees are staying in the workplace longer,"[1] and Gen Xers similarly complain about "boomers blocking their way to the top as older workers delay retirement."[2] When older workers linger in their positions, the thinking goes, it has a negative spillover effect on the careers of younger workers. These career spillovers are not only important for younger workers but also for their employers. If employers attract, retain, and motivate workers by promising them careers rather than jobs, they need to design personnel policies and make strategic decisions that enable them to deliver on these promises.

Despite the popular attention these career spillovers receive, there is no systematic evidence that they actually matter. The vast empirical literature on internal labor markets, dating back to at least Baker, Gibbs, and Holmström (1994), neglects these spillovers, as it treats workers' careers independently.[3] This empirical neglect is not an oversight but results from standard economic reasoning. When one qualified worker's career appears to be blocked because a higher-level position is already occupied, the firm always has the option of creating another higher-level position. And, even if the firm cannot do so, the worker can always move to another firm that can. According to this logic, workers' careers should therefore be determined only by their individual characteristics, such as their human capital, and by broad market-level factors.[4] However, if it is difficult for firms to create positions and for workers to switch to another firm, then career spillovers should matter: one worker's career success may come at the expense of his or her coworkers'.

In this paper, we show that career spillovers matter by providing evidence that retirement delays among older workers negatively impact the career progressions of their younger coworkers. An ideal test for such career spillovers would randomly prevent older workers in one firm from retiring while allowing older workers in another firm to retire, and it would compare the career progressions of younger workers between these two firms. While such a test is not feasible, we argue that a recent reform to the Italian pension system created a reasonably close approximation of this ideal. The Fornero reform, which was swiftly implemented in December of 2011 to contain public expenditures, led to an overall increase

[1] https://www.hrdive.com/news/millennials-feel-boomer-and-gen-x-bosses-are-blocking-their-progress/504129/

[2] http://www.bbc.com/capital/story/20130710-the-forgotten-generation

[3] For a rare exception, see Friebel and Panova (2008), which uses personnel records from a large heavy-industry firm in Russia following privatization reforms and finds evidence that reduced turnover at the top led to blocked promotions for younger workers.

[4] See Gibbons and Waldman (1999), Rubinstein and Weiss (2006), Lazear and Oyer (2013), and Waldman (2013) for surveys on standard approaches to analyzing workers' careers.

in the minimum retirement-eligibility age. Grandfather clauses were limited, and the reform unexpectedly caused retirement delays among senior employees who were slated to retire soon after December, 2011. Moreover, the change to the eligibility criteria led otherwise similar workers to face significantly different retirement delays based on small differences in their ages and years of contribution to social security.

The unanticipated nature of the reform and the differential treatment of otherwise similar employees provide a clean empirical setting to study the effect of retirement delays among senior workers on the careers of younger workers. Our identification strategy compares changes in wage growth and internal promotions of younger employees across firms experiencing different average retirement delays for senior workers, both before and after the reform. We measure the exposure of each firm to the pension reform as the average change in retirement eligibility caused by the reform among senior workers close to retirement. By controlling for differences in age and gender distributions between firms, we exploit the variation in treatment that does not stem from cross-firm differences in broad demographic compositions, which could affect internal career trajectories through other channels. Rather, our analysis reflects idiosyncratic differences in gender, age, and years of contribution to social security among workers close to retirement.

We leverage two sources of data, both provided by the Italian Social Security Institute (INPS). First, we use a panel of matched employer–employee records for all private, non-agricultural firms with 10 to 200 workers in the first quarter of 2009. Drawing on these records, we are able to compute monthly average contractual wage growth as well as categorical promotions within the company between 2009 and 2015.[5] Second, we use the complete contribution histories for all workers employed in these firms. These data allow us to compute the retirement delays among workers who were slated to retire within three years of 2011.

Our main finding is that career spillovers exist: longer retirement delays among older workers cause larger decreases in wage growth for younger workers. A one-standard-deviation increase in the average retirement delay among workers who were close to retirement before the reform decreases the wage growth of their younger colleagues by 2.5 percent per year with respect to the pre-reform baseline growth, and these effects persist throughout the four years of the treatment period.

To better understand the underlying mechanism behind these career spillovers, we develop a model that describes when career spillovers are likely to matter most. The model allows us to ask, and answer, the following three questions: First, does it matter whose retirement

[5] The contractual wage is the wage written into a worker's labor contract, not their take-home pay. As we explain in Section 2.2, contractual wages are closely related to job titles, which is a unique feature of the Italian labor market.

is delayed? Second, are career spillovers stronger in firms that have more limited promotion opportunities? Finally, do retirement delays affect the careers of different workers differently?

Our findings show that, as the model predicts, retirement delays among older workers reduce the promotion rates of younger workers, but only if the older worker is in the higher-level position. Next, we show that career spillovers are most relevant for workers in slow-growth firms. We divide firms into tertiles of pre-period employment growth and look at how the effects of the treatment differ by the growth rate of the firms. The decrease in wage growth from a one-standard-deviation increase in our treatment variable is 8 percent for workers in bottom-tertile firms, which are all shrinking in size, and approximately zero for workers in top-tertile firms, which are all expanding their ranks. Similarly, we show that career spillovers are concentrated among firms with larger spans, measured as the pre-period fraction of jobs in the firm that are relatively highly paid. Finally, we find that retirement delays among older workers have a bigger impact on the careers of their coworkers who are 55 years or older than on their younger coworkers. This finding is consistent with firms using seniority as one of the criteria to assign promotions.

Our model also generates specific predictions about workers' and firms' extensive-margin responses to retirement delays. For workers, our model predicts that, even if retirement delays reduce promotion opportunities, we would not expect to see workers leave for other firms where they may have to restart the career ladder. Consistent with this prediction, we do not observe younger workers responding to retirement delays among their older coworkers by voluntarily leaving the firm. For firms, our model predicts that they will respond to retirement delays by laying off existing workers and hiring fewer new workers. Empirically, we find that a one-standard-deviation increase in retirement delays leads firms to increase layoffs by 12 percent and reduce hiring by 2 percent.

Are the career spillovers we document large? The reduced wage growth of younger workers due to a one-standard-deviation increase in retirement delays amounts to a monetary loss of up to €718 over the course of four years. As another way of interpreting this magnitude, we can benchmark it against other drivers of wage growth studied in the literature. Our main estimate is similar in magnitude to the effect of moving to a metro area with $800,000$ fewer residents (Wheeler, 2006) or to the effects of receiving six fewer months' worth of on-the-job full-time training (Bartel, 2002).

A natural alternative way to measure the firm-level short-term effect of the reform would be to compute the share of workers who had to delay their retirement during the first years after 2011. Unlike our preferred treatment variable, however, the share of affected workers is correlated with the firm-level age distribution. Using this variable as our treatment, therefore, would induce a comparison between career progressions for workers in firms with older and younger workers, which would confound the effects of the reform. To overcome this issue, we estimate an instrumental variable specification in which we instrument the share of close-to-retirement workers whose retirement-eligibility age increased by at least one year due to the reform with our baseline treatment variable, the average retirement delay among close-to-retirement workers. The findings of these IV specifications are qualitatively similar to the main OLS coefficients.

We conclude the analysis by evaluating the extent to which our findings are consistent with other career-spillover channels. For example, financially constrained firms that face retirement delays may simply be unable to afford to promote their workers. We find evidence consistent with financial-constraint-driven career spillovers, but financial constraints alone cannot account for the full range of our findings.

Our paper contributes to the literature on the theory of internal labor markets by highlighting the empirical relevance of slot constraints in determining a worker's career progression. Slot constraints, defined as limits on available slots for internal promotions and the inability to easily add positions to the organization, have featured prominently in the literature in sociology and organizational theory.[6] Most leading models of internal labor markets in economics, however, abstract from slot constraints, focusing instead on individual factors, such as human capital acquisition, learning, insurance, signaling, and incentives (see Gibbons and Waldman (1999) and Waldman (2013) for surveys of the theoretical literature on internal labor markets). As a result, most of the empirical work on internal labor markets has focused on these worker-level factors.[7] Our findings suggest the importance of incorporating firm-level factors, such as slot constraints for understanding workers' career dynamics (for early theoretical work in this direction, see Lazear and Rosen (1981), and for more recent work, see DeVaro and Morita (2013), Ke, Li, and Powell (2018), and Li, Powell, and Ke (2019)).

Our paper documents the impact of career spillovers on workers due to blocked promotion opportunities. Other papers in the literature establish a number of other channels through which there may be career spillovers. Hayes, Oyer, and Schaefer (2006) and Jäger and Heining (2019) show, for example, that career spillovers can arise because of team production. There are several papers that emphasize the role of limited career opportunities, but they focus on jobs with strict institutional features that give rise to rigid slot constraints, such as bureaucracies (Bertrand et al., 2018), academia (Borjas and Doran, 2012), sports (Brown, 2011; Gong, Sun, and Wei, 2017), and firms in transitioning economies (Friebel and

[6] For early conceptual work, see Simon (1951) and White (1970). See Stewman and Konda (1983) and Stewman (1986) for surveys, and see Bidwell and Keller (2014) for recent empirical evidence on the importance of available slots for the firm's decision about whether to hire externally.

[7] Chiappori, Salanié, and Valentin (1999) focuses on learning; DeVaro and Waldman (2012) focuses on asymmetric information and signaling; and Benson, Li, and Shue (2019) focuses on job performance.

Panova, 2008).[8] Our paper shows that scarce career opportunities lead to career spillovers in representative private-sector firms in which there are no obvious institutional constraints to creating additional positions. Finally, we focus on the effect of limited career opportunities for the career advancement of workers who have already been hired. Several recent papers examine the implications of limited opportunities on whether workers are hired to begin with and which occupations they decide to pursue (Liang, Wang, and Lazear, 2018; Lazear, Shaw, and Stanton, 2018).

We also contribute to the growing body of literature that shows how workers' careers are shaped by luck. There are many studies documenting how labor market conditions at the time a worker is hired affect his or her entire career trajectory (Von Wachter and Bender, 2006; Oyer, 2006; Kahn, 2010; Schmieder and Von Wachter, 2010; Shu, 2012). Lazear, Shaw, and Stanton (2018) shows that idiosyncratic luck at the time of hiring can also play an important role. We complement these findings by showing that luck matters throughout a worker's career. Even after being hired, a worker's career progression depends on whether senior workers happen to leave their positions and open up advancement opportunities.

Finally, we provide new evidence on the consequences of the Fornero reform, arguably the most important Italian reform of the last decade. In a contemporaneous paper, Boeri, Garibaldi, and Moen (2017) also uses firm-level variation in exposure to the reform to bring granular evidence to bear on an important question in labor economics: How do pension reforms affect youth hiring and unemployment (Gruber and Wise, 2010)? Our primary focus is instead on career progression inside firms, which requires us to use more detailed individual-level data both for identification purposes and for analyzing the underlying mechanisms. On the identification side, these data enable us to precisely measure individual retirement delays. This is necessary for isolating the variation in retirement delays that is driven by small differences in demographics from the variation that is driven by other firm-level differences. In terms of mechanisms, the data enable us to decompose firm-level average retirement delays into multiple firm-level shocks that occur at different parts of the wage distribution and for different categories of workers. Moreover, they enable us to look at the effect of retirement delays on different categories of workers—namely younger, middle-tenure, and older workers.

The rest of the paper is organized as follows. Section 2 describes the institutional details and the data. Section 3 introduces a stylized theoretical model and develops several predictions. Section 4 lays out the identification strategy. Section 5 presents the main results and several robustness checks. Section 6 discusses alternative mechanisms, and Section 7 concludes the paper.

[8] The natural experiment we use could also be interpreted as an unexpected shock to labor supply. In contrast to other papers that study shocks to labor supply stemming from large influxes of outside workers (e.g., Card, 1990; Ottaviano and Peri, 2012; Dustmann, Schönberg, and Stuhler, 2017), our paper studies an increase in labor supply stemming from workers who were already employed by local firms.

2 Institutional Background and Data

2.1 The 2011 Reform of the Italian Pension System

On December 6, 2011, the Italian government enacted a reform of the pension system—known as the Fornero reform—as part of a larger package of interventions called the "Save Italy" decree.[9] The reform became fully effective on January 1, 2012, only 26 days after its presentation to the Parliament (Figure A1). The goal of the reform was to quickly reduce public spending by raising the eligibility requirements for public pensions.

The Fornero reform had three characteristics that are important for our empirical analysis. First, many workers experienced a substantial increase in their retirement-eligibility age. Most workers in private-sector firms retire as soon as they become eligible for a public pension (88 percent in our sample), so this increase in the retirement-eligibility age led to retirement delays. In Italy, private-sector employees become eligible to claim full pension benefits based on one of two sets of criteria. One is based on age alone (age-based criteria) and the other is based on a combination of age and years of contribution to social security (seniority-based criteria). The Fornero reform raised the requirements to become eligible under both sets of criteria. In the case of the age-based criteria, the minimum retirement age was immediately increased by one year for men and two years for women (Figure A3, panel A). In the case of seniority-based pensions, the minimum number of years of contribution required for eligibility increased by two to seven years for men and one to six years for women (Figure A3, panel B). Appendix B includes a more thorough description of the changes induced by the Fornero reform.

The second important feature of the reform is that grandfather clauses were very limited. They only applied to workers who were eligible to claim a pension under the old rules by December 31, 2011, and to a couple other specific categories.[10] The lack of grandfather clauses meant the reform had an immediate effect on the retirement decisions of most Italian workers.

Finally, workers and firms could not have anticipated the detailed provisions of the reform. Even though Italy had been facing increasing financial difficulties prior to December 2011, the political events that led to the reform happened in rapid succession.[11] The reform

[9] The pension reform was the central component of the decree. Other interventions mainly increased taxation on real estate, cars, and consumption. The whole text of the law can be accessed at https://www.gazzettaufficiale.it/gunewsletter/dettaglio.jsp?service=1&datagu=2011-12-06&task=dettaglio&numgu=284&redaz=011G0247&tmstp=1323252589195.

[10] We list these rare exceptions in Appendix B.

[11] The government lost its parliamentary majority on November 8, Prime Minister Berlusconi resigned four days later on November 12, and a new technocratic government took office without general elections on November 16.

was presented only 20 days after the appointment of a new technocratic government and started being enforced 26 days after its presentation.[12] Stock markets responded sharply on December 6, when the reform was officially presented, suggesting that at least some aspects of the reform were not anticipated (Figure A2). We can therefore consider these increases in the retirement-eligibility age as largely unexpected shocks to firms' internal labor markets.

The changes introduced by the reform provide a clean empirical setting to study career trajectories within private-sector firms. Small differences in observable characteristics generated large differences in retirement delays (Figure A4). For instance, consider a group of male workers born in 1951 and 1952, who started working at 23 and contributed to social security without interruption. In spite of being born only one year apart, the 1951 cohort became eligible for a seniority-based pension in 2011 under the old rules, while the 1952 cohort faced a 4-year and 7-month delay in retirement (Appendix B.3).

To summarize, the reform represents an unexpected and substantial shock to the minimum requirements for public pension eligibility. Moreover, small demographic differences led to large differences in retirement delays for individuals. The reform, therefore, could have very different effects across firms with similar demographic characteristics among their workforces. Our empirical analysis will exploit cross-firm differences in the retirement delays of older workers that stem from individual variation in gender, age, and years of contribution but that are not correlated with other firm-level determinants of career trajectories.

2.2 Data

Our empirical analysis uses confidential administrative data provided by the Italian Social Security Institute (INPS). Specifically, we use seven years of matched employer–employee data to build firm-level measures of career progression, and we use a separate dataset containing the complete working history of workers to compute individual-level reform-induced retirement delays.

The first dataset consists of matched employer–employee records for all private-sector, non-agricultural firms with at least one salaried employee. The dataset combines individual-level information about workers, such as demographic characteristics, wage, type of contract (full-time vs. part-time, open-ended vs. fixed-term), and position within the firm (blue-collar, white-collar, and manager), with information about the firm, such as sector, location, and age. In this dataset, we restrict our analysis to workers who were not eligible in 2011 to retire within the following three years. These are individuals not immediately affected by changes brought about by the reform because they were relatively far from retiring at the

[12] Moreover, the technocratic cabinet implemented the reform using the legal instrument of the "decree-law," which does not require a public discussion in the Parliament.

time of the reform.[13] We further focus on full-time permanent employees because we want to study the career trajectories of workers who are central to firm activities.

We use this information to construct several measures of career progression. First, we compute the average monthly contractual wage growth—an indirect measure of promotions—for each firm and year in the sample. To do so, we use the monthly contractual wage for each worker instead of the more commonly available take-home pay. The contractual wage is the monthly wage that each employee should receive based on his or her labor contract. Unlike take-home pay, it is not affected by transitory shocks, such as leaves of absence (maternity, injury, sick) and bonuses. Rather, it is closely related to job titles, which we do not observe in the data. Assigning a new job title to an employee, in fact, often requires by law a modification of the contractual wage to reflect the different responsibilities attached to the new position (Art. 2103 c.c.). In summary, our measure of monthly wage growth likely captures more permanent changes in job titles instead of transitory shocks to hours worked or bonuses.

Second, we create two direct measures of categorical promotions by computing the number of workers moving from blue- to white-collar jobs or from blue/white-collar jobs to managerial positions for each firm and year. These variables capture substantial leaps within the firm's hierarchy. The combination of contractual wage growth and categorical promotions should provide a relatively complete description of internal promotions within the private-sector firms in our dataset.

The second dataset consists of the complete contribution histories of individuals who, between 2009 and 2015, worked in private-sector, non-agricultural firms that employed between 10 and 200 employees in the first quarter of 2009.[14] In this dataset, the unit of observation is an event that generated a contribution to the pension system. Available information includes the type of event associated with the contribution (e.g., paid work, sick leave, or maternity leave), its monetary value, and its duration. This rich dataset is essential for identifying senior workers close to retirement under pre-reform rules and precisely determining the firm-level shock to the retirement decisions of older employees, which Section 4 discusses in greater detail.

2.3 Sample

We restrict the sample to firms that employed between 10 and 200 workers in the first quarter of 2009. We impose the upper bound to comply with INPS's request to limit the

[13] The results are robust to focusing on workers who were eligible in 2011 to retire within the following two, four, or five years (Table A7).

[14] The restriction on firm-size is due to constraints on the number of contribution histories that could be extracted by INPS.

size of the data extraction. Moreover, we set the lower bound to remove very small firms with organizational structures that are too simple to properly study career spillovers. Even with these constraints in place, the sample is highly representative of the Italian productive landscape, which is mostly populated by small to medium-large firms. Indeed, only 0.08 percent of firms have more than 250 employees.[15] Furthermore, we only consider firms that operated in all years between 2009 and 2015 and employed at least one full-time permanent worker in each year in order to have a balanced sample.

Table A2 (columns 1 and 2) shows the main characteristics of the master sample, which comprises $104,182$ firms, at the beginning of the sample period in 2009.[16] The average firm employed 26 workers and had been operational for 18 years. The majority of firms operated in the service sector. The majority of workers were between 35 and 55 years old. Of all employees, 59 percent were in blue-collar jobs, 33 percent held white-collar positions, 2 percent were managers, and the rest were apprentices. The vast majority of workers were permanent and full-time.

Firm-level summary statistics also indicate that the turnover of older workers, as well as the wage growth and number of categorical promotions of younger workers, decreased after 2011 (Table 1). The number of workers retiring at a given firm and in a given year decreased by 18 percent in the post-reform period. The number of vacancies, measured as the number of all workers leaving the firm (due to retirement, voluntary or involuntary turnover), shows a similar percentage drop. Together with a decrease in turnover, we observe a decline in average wage growth and in the number of promotions from blue-collar to white-collar jobs and from blue/white-collar jobs to managerial positions. As discussed above, these last three career outcomes are computed without including workers who were within three years of retirement in 2011. Of course, the comparison of pre- and post-reform averages does not identify the causal effect of retirement delays among senior workers on the career trajectories of younger coworkers. In fact, many other factors—including macroeconomic conditions—might have changed between the two periods. In Section 4, we outline the empirical strategy we employ to isolate the effect of the reform.

3 A Stylized Model of Career Spillovers

Before analyzing the effects of retirement delays on the career progression of younger workers, we provide a conceptual framework to explore how constraints on a firm's career capacity—its

[15] Data between 2012 and 2016 are available from Istat at http://dati.istat.it/Index.aspx?DataSetCode=DICA_ASIAUE1P.

[16] In addition to the constraints discussed in the previous paragraph, we limit the sample to firms that have non-missing values for all measures of career progression. This step reduces the number of firms from 104,924 to 104,182.

ability to provide advancement opportunities to qualified workers—affect the career progression of its employees.

Our conceptual framework is related to the models of internal labor markets of Gibbons and Waldman (1999), Ke, Li, and Powell (2018), and Li, Powell, and Ke (2019). The contribution of our analysis is to incorporate the idea of limited career capacity into the Gibbons and Waldman (1999) framework, which gives rise to career spillovers across workers.

Our analysis yields eight empirical predictions that describe how retirement delays among older workers affect the career progression of younger workers. We summarize these predictions at the end of this section.

3.1 Model Setup

A firm operates for two periods and in each period requires workers to perform two different jobs, job 1 and job 2. Worker productivity depends on their effort, their innate ability, and the job they are assigned. The worker either exerts effort, $e_i = 1$, or shirks, $e_i = 0$, and their effort costs depend on which job they are assigned: if they are assigned to job j, their effort costs are c_j, where $c_2 > c_1$. Effort is not directly observed—if a worker shirks in a given period, the firm observes this with probability q_j if they are assigned to job j, where $q_1 > q_2$. Job 1 is therefore easier to do and easier to monitor. Workers are heterogeneous, and their innate ability, $\theta_i = \theta_L, \theta_H$, is initially unknown to all parties. Workers have high ability with probability λ, and their ability is revealed at the end of their first period of employment. This innate ability affects their productivity in job 2 but not in job 1. All parties discount future payoffs with discount factor $\delta < 1$.

Production. If worker i is assigned to job j in period t, and they shirk, their output is 0, and if they exert effort, then they produce

$$Y_{j,t} = f_j + h_j \theta_i.$$

We assume that $h_1 = 0$, so their output in job 1 does not depend on their ability. We also assume that $f_1 > 0 > f_2$ and $0 > (1-\lambda)(f_2 + h_2\theta_L) + \lambda(f_2 + h_2\theta_H)$, so if the worker's ability is unknown, their expected productivity is negative if they are assigned to job 2. Finally, we assume that $f_2 + h_2\theta_H > f_1$, so if the worker is known to be high ability, they are more productive in job 2 than in job 1. The firm is capacity constrained and can assign up to $\overline{N}_{j,t}$ workers to job j in period t, and if it assigns $N_{j,t} \leq \overline{N}_{j,t}$ workers who all exert effort to job j in period t, then it receives revenues $N_{j,t}Y_{j,t}$. Throughout, we will also assume that in the first period the firm is endowed with $\overline{N}_{2,1}$ high-ability workers, which it assigns to job 2, reflecting the idea that the results of the first period reflect past optimizing behavior on the part of the firm. We will refer to such workers as legacy workers.

Personnel Policies. To motivate workers to exert effort, the firm has three instruments at its disposal. First, the firm pays non-negative wages $w_{j,t}$ to a worker assigned to job j at the end of period t if they are not caught shirking. If the worker is caught shirking, we assume without loss of generality that the worker will be paid 0 and will be terminated. Next, the firm chooses reassignment probabilities $p_{k,j}(\theta)$ between period 1 and period 2, where $p_{k,j}(\theta)$ is the probability that a worker of type θ assigned to job k in period 1 is assigned to job j in period 2 if they have not been caught shirking. Finally, if the firm hires new workers, it has to decide what job to assign them in their first period of employment.

Timing. The timing of the game is as follows. In each period t, the firm chooses the number of workers to assign to each job $N_{j,t}$. The firm then offers each worker assigned to job j a contract that specifies a nonnegative wage $w_{j,t} \geq 0$ that the worker will receive if they are not caught shirking as well as a next-period assignment $p_{k,j}(\theta)$ if they continue their employment at the firm. The worker then decides whether to accept the contract or reject it in favor of an outside opportunity that yields a payoff of 0. If they accept the contract, they choose whether to exert effort or to shirk, which the firm observes with noise. The firm then makes payments to workers according to the contract. The worker's ability θ is then observed by both the firm and the worker, and the worker departs the firm for exogenous reasons with probability d_j.

The Firm's Problem. The firm's problem is to choose the number of workers it assigns to each job in each period, $(N_{j,t})_{j,t}$, its wage policy $(w_{j,t})_{j,t}$, its promotion policy $(p_{k,j})_{k,j}$, and its second-period hiring policy $(H_j)_j$ to maximize its profits

$$N_{1,1}(Y_{1,1} - w_{1,1}) + N_{1,2}(Y_{1,2} - w_{1,2}) + \delta(N_{2,1}(Y_{2,1} - w_{2,1}) + N_{2,2}(Y_{2,2} - w_{2,2})),$$

subject to the constraint that each worker has the incentives to exert effort in each period and to three additional sets of constraints. We detail these constraints below.

Incentive Constraints. The firm needs to motivate its workers to exert effort in both the first and second periods. In the second period, workers assigned to job j need to prefer to exert effort, in which case they receive $w_{j,2} - c_j$, rather than to shirk, in which case they do not incur their effort cost, and with probability $1 - q_j$ they are not caught and therefore are still paid $w_{j,2}$.

In the first period, workers' incentives to exert effort depend on the probabilities with which they will be assigned to each of the two jobs in the second period. If they remain at the firm, the job they will be assigned in the second period depends on their ability and on the firm's promotion policy. A worker who is found to have high ability, which occurs with probability λ, will be assigned to job k in the next period with probability $p_{j,k}(\theta_H)$, and a low-ability worker will be assigned to job k in the next period with probability $p_{j,k}(\theta_L)$. Hence, an unknown-ability worker will receive an expected payoff of

$$V_j = \lambda(p_{j,1}(\theta_H)v_{1,2} + p_{j,2}(\theta_H)v_{2,2}) + (1-\lambda)(p_{j,1}(\theta_L)v_{1,2} + p_{j,2}(\theta_L)v_{2,2})$$

in the second period, where $v_{k,2}$ is the utility they will receive in period 2 if they are assigned to job k. Workers will therefore prefer to exert effort in the first period if

$$w_{j,1} - c_j + \delta(1-d_j)V_j \geq (1-q_j)[w_{j,1} + \delta(1-d_j)V_j].$$

That is, they will prefer to exert effort if their expected discounted payoffs are higher if they work than if they shirk.

Other Constraints. In addition to satisfying workers' incentive constraints, the firm also has to satisfy three additional sets of constraints: participation, flow, and slot constraints. The participation constraints require that, in each period, each worker prefers to work at the firm rather than to take their outside option.

The flow constraints ensure that, in period 2, the number of workers assigned to job j is equal to the sum of the number of new hires into job j, H_j, and the number of workers who were assigned to job k in period 1, who did not leave the firm exogenously, and who were assigned to job j in period 2. That is, for $j = 1, 2$,

$$N_{j,2} = H_j + N_{1,1}(1-d_1)(\lambda p_{1,j}(\theta_H) + (1-\lambda)p_{1,j}(\theta_L)) + N_{2,1}(1-d_2)p_{2,j}(\theta_H),$$

where H_j is the number of workers the firm hires in period 2 and which it assigns to job j. Finally, the firm has to satisfy slot constraints, $N_{j,t} \leq \overline{N}_{j,t}$ for each job j and in each period t.

3.2 Optimal Personnel Policies

The model is stylized, but the assumptions are empirically motivated. In particular, the two jobs correspond to blue-collar and white-collar jobs. Workers' effort in blue-collar jobs is often easier to monitor than in white-collar jobs. The lowest wage that workers can be paid is set to zero for simplicity, and the analysis can easily accommodate any other value for the lowest wage.

In this model, optimal personnel policies resemble an internal labor market. There is a port of entry in the sense that, except for legacy workers, new workers are assigned to job 1. Optimal personnel policies also feature a well-defined career path. Workers are motivated by a combination of wages in their current job and, if they turn out to be high-ability, the prospect of promotion to job 2, which is coupled with an additional wage increase. In addition, workers are never demoted.

The following proposition describes the firm's hiring policies and the expected wage growth for workers assigned to job 1 in period 1 and shows that it depends on the promotion rate. For ease of exposition, we will assume that, in terms of the firm's capacity, its organizational span, $\overline{N}_{1,t}/\overline{N}_{2,t}$, is fixed and equal to s. Denote the firm's growth rate by $g = (\overline{N}_{2,2} - \overline{N}_{2,1})/\overline{N}_{2,1}$, and define the variable $R_i = (1 - q_i)c_i/q_i$, which is a measure of the amount of rents required to motivate a worker assigned to job i. We also assume that the output job-1 workers produce, f_1, is greater than $c_1 + R_1$, so job-1 workers in the second period produce strictly positive profits for the firm. Proofs are in Appendix D.

Proposition 1. Suppose $f_1 > c_1 + R_1$. A worker assigned to job 1 in period 1 will receive an expected wage increase of

$$\Delta w^* = w^*_{1,2} - w^*_{1,1} + \lambda p^*_{1,2}(w^*_{2,2} - w^*_{1,2}),$$

where

$$p^*_{1,2} = min\left\{\frac{g + d_2}{(1 - d_1)\lambda s}, 1\right\}.$$

Moreover, the number of new hires in the second period satisfies $H^*_1 = N^*_{1,2} + N^*_{2.2} - (1 - d_1)N^*_{1,1} - (1 - d_2)N^*_{2,1}$.

The expression for wage growth in Proposition 1 describes the two sources of wage growth. The wage growth within job 1 is given by $w^*_{1,2} - w^*_{1,1}$, and the promotion premium is given by $w^*_{2,2} - w^*_{1,2}$. The key result of Proposition 1 is that workers' promotion rates are determined by $p^*_{1,2}$, which is governed by two regimes. In particular, when $p^*_{1,2} = 1$, the firm has abundant career capacity, so all high-ability workers are promoted in a given period. When $p^*_{1,2} < 1$, the firm has limited career capacity, so not all high-ability workers are promoted.

Which of the two regimes prevails depends, in part, on the firm's growth rate and its span. A firm that grows quickly or has a low span will have abundant career capacity, while a firm that grows slowly or has a high span will have limited career capacity.

In firms with abundant career capacity, a change in the exogenous departure rate for workers in job 2 has no effect on the promotion probability and therefore no effect on the expected wage growth for workers in job 1. In contrast, in firms with limited career capacity, a reduction in the departure rate for workers in job 2 means that fewer slots are freed up for workers in job 1, which reduces their promotion probability. As a result, their expected wage growth will also be lower. The same is true for within-job-1 wage growth.[17]

Finally, the proposition shows that the firm always hires directly into the bottom job. The number of new hires is equal to the total number of positions minus the number of workers from the previous period who have not departed.

Proposition 1 therefore allows us to make predictions regarding how expected wage growth and promotion rates for younger workers will be affected by the pension reform. If we think of the pension reform as primarily reducing the exogenous departure rate for certain workers, then our model shows how the reform will affect workers' wage growth and promotion rates within firms. Our model delivers several predictions, which we describe in the following corollary. We assume that $g + d_1 + d_2 < 1$ because it is the empirically relevant case.

Corollary 1. Suppose $g + d_1 + d_2 < 1$. Then, the following are true:

(i.) Δw^* and $p^*_{1,2}$ are increasing in d_1 and d_2;

(ii.) $\partial p^*_{1,2}/\partial d_2 > \partial p^*_{1,2}/\partial d_1$;

(iii.) $\partial \Delta w^*/\partial d_1$ and $\partial \Delta w^*/\partial d_2$ are decreasing in g and increasing in s;

(iv.) H^*_1 is increasing in d_1 and d_2.

The first part of Corollary 1 shows that the expected wage growth and promotion rate for younger workers are decreasing in retirement delays, as measured by a reduction in d_1 and d_2. The second part shows that the impact of retirement delays on promotion rates is higher if the workers whose retirements are being delayed are in job 2. The third part shows that the effect of retirement delays on expected wage growth is more pronounced in slow-growing firms and firms with larger spans. The last part shows that retirement delays lead the firm to reduce hiring in the second period.

The model we analyzed above is a two-period model in which parties' learning about worker qualifications is immediate. The model can be extended to allow firms to be long-lived and workers to live for a finite number of periods, with their ability gradually revealed over time. Such an extension preserves the predictions of Corollary 1, and it allows us to

[17] In the model, second-period wages are determined by the worker's incentive constraint and are $c_i + R_i$ in job i. The promotion premium therefore does not depend on the departure rate d_2. Wage growth in job 1 does, however, because promotions and current wages, which act like bonuses, are substitutes (see, for example, Ekinci, Kauhanen, and Waldman, 2019): A reduction in d_2 raises the wage that has to be paid to motivate job-1 workers in the first period and therefore reduces within-job-1 wage growth.

make an additional prediction. In particular, when learning is gradual, workers who have recently been hired at the firm may not have had the opportunity to demonstrate that they are qualified for job 2. In this case, when a position in job 2 is freed up, it is more likely to be filled by someone who has longer tenure in job 1. As a result, retirement delays will have a bigger impact on relatively more senior workers in job 1.

In addition, we have only explored the effects of turnover of older workers on wages and promotions for younger workers, but we might also expect it to affect their voluntary departure decisions: if older workers are less likely to leave, then younger workers are less likely to get promoted, and they may seek alternative opportunities. In the model, however, the need for the firm to provide incentives implies that workers receive rents that are increasing over time, so they strictly value their current job over their next best alternatives. This feature of internal labor markets in our setting—a feature shared with the models of Ghosh and Ray (1996) and Kranton (1996)—implies that turnover among older workers is likely to have a limited effect on turnover decisions among younger workers.

Next, we discuss how retirement delays affect involuntary turnover. When promotion opportunities are used to motivate employees, vacancies created through layoffs can have beneficial incentive effects for younger workers. These incentive effects are larger when the firm has more limited career capacity. Firms may therefore lay off workers in order to create more promotion opportunities (see, for example, Ke, Li, and Powell (2018)). More generally, if worker ability is heterogeneous, then the firm will retain its higher-ability workers, and this retention threshold will be higher when the firm has more limited career capacity. This result, in turn, leads to more layoffs in both jobs in response to retirement delays.

Finally, the model can be extended to allow for hiring directly into job 2 by incorporating skill heterogeneity as in Ke, Li, and Powell (2018). In such an extension, firms may fill job 2 vacancies with outside hires, but they are biased towards filling them with internal candidates. This is because hiring outside candidates reduces the advancement opportunities for workers lower in the organization. The degree of this insider bias is greater, and there is therefore less hiring into the top job, when firms have more limited career capacity.

3.3 Empirical Predictions

Our model illustrates how career spillovers can result when retirement delays block younger workers' promotion prospects. Career spillovers are stronger in firms with limited career capacity, where workers' promotion prospects are already low. These observations give rise to a host of empirical predictions regarding the pattern of the resulting career spillovers. In the subsequent sections, we test eight key predictions:

(1) the wage growth of young workers decreases in response to retirement delays;

(2) promotion rates are reduced more by retirement delays in higher-level positions;

(3) the effect of retirement delays on wage growth is larger in slow-growing firms;

(4) the effect of retirement delays on wage growth is larger for firms with larger spans;

(5) the effects of retirement delays on wage growth are larger for more-senior workers;

(6) the voluntary departure rate is independent of retirement delays;

(7) the number of layoffs rises in response to retirement delays;

(8) the number of new hires falls in response to retirement delays.

Each of these theoretical predictions receives empirical support. We also discuss alternative interpretations of our empirical results in Section 6.

4 Empirical Strategy

4.1 The Treatment Variable

This section describes how we isolate the effect of retirement delays among senior employees on the career progression of their younger coworkers. The desired treatment variable should measure the reform-induced retirement delays in each firm. To construct this variable, we focus on senior workers, who we henceforth refer to as CTR (close-to-retirement) workers, to isolate the short-term effect of the reform. We classify a worker as a CTR worker if they are a full-time permanent employee who, in December 2011, would have become eligible to retire by December 2014 under the pre-reform rules. When compared to other employees, CTR workers are older, more experienced, and have a longer tenure at the firm. They also earn a higher wage (Table A3).

To identify CTR workers, we use data on gender, age, and years of contribution at the time of the reform that is contained in the contribution histories provided by INPS. We use this information to compute the retirement-eligibility date under the pre-reform rules, had they remained in place, for each employee in the sample.[18] We also compute the retirement-eligibility date under the post-reform rules. We define the worker-level retirement delay as the difference between the post- and pre-reform retirement-eligibility dates:

$$D_\psi = \text{Years until retirement}^{\text{post}} \quad - \quad \text{Years until retirement}^{\text{pre}},$$

where ψ represents the worker's group, which depends on their gender, age, and years of contribution in December, 2011. Even though all CTR workers were similarly close to retirement under pre-reform rules, there is substantial variation in their reform-induced retirement delays (Figure 1, panel A). The variable D_ψ has a mean of 1.36 years and

[18] Appendix C includes more details on how the retirement dates are computed.

standard deviation of 1.42 years. As discussed in Section 2, these individual-level differences in retirement delays arise from small variations in demographic characteristics (Figure A4).

To construct the firm-level treatment, we weight the retirement delay for each worker group by the share of CTR workers belonging to that group. Specifically, we compute:

$$\begin{aligned} \text{Delay}_i &= \sum_{\psi} \pi_{\psi,i} \times D_{\psi} \\ \pi_{\psi,i} &= \frac{\#\text{CTR workers}_{\psi,i}}{\#\text{CTR workers}_i}. \end{aligned} \tag{1}$$

Our treatment Delay_i therefore measures the weighted average retirement delay of CTR workers at firm i. Throughout the rest of the paper, we will refer to the weighted average retirement delay among CTR workers at firm i simply as the "retirement delay" or the "firm-level retirement delay." As with the worker-level variable D_{ψ}, there is substantial variation in the firm-level retirement delay (Figure 1, panel B). The average retirement delay is 0.44 years, while the standard deviation is 0.97 years. Two-thirds of the firms in the sample did not employ a single CTR employee and therefore did not experience any retirement delays according to our measure. Among firms with at least one CTR worker, the average retirement delay is 1.36 years, and the standard deviation is 1.28 years.

We perform a series of balance tests to estimate the correlation between the treatment variable and a rich set of firm characteristics observed in 2009. Firms that experience higher retirement delays of senior workers are older, larger, and employ an older workforce (Table 2, column 1). These findings are not surprising, as the sample includes firms that did not have any CTR workers in 2011 and therefore have no retirement delays. Such firms tend to be smaller, younger, and employ a younger workforce (Table A2, columns 2 and 3).

We address the potential concern that these imbalances may confound our results in two ways. First, our main specifications include controls for nonlinear trends that differ based on firm characteristics. Second, we also perform our analysis on the subset of firms that had at least one CTR worker. In this restricted sample, the correlations between the treatment variable and firm characteristics are much weaker (Table 2, column 3). Relative to the full sample, these correlations are smaller because the treatment variable Delay_i does not depend on the presence of CTR workers, which is itself related to firm size and workforce age.

It is also important to note that the treatment variable does not predict large cross-firm differences in the gender composition of the workforce in either the full or restricted samples (Table 2, columns 1 and 3). As shown in Section 2, the reform led to different increases in retirement eligibility for men and women. This could in principle raise the concern that the treatment variable was capturing differences in firms' gender compositions, which could be correlated with other features of their internal labor markets. In addition to showing that this correlation is weak, we also explicitly control for nonlinear trends in career progression that are correlated with the share of male workers employed at baseline.

To summarize, the treatment variable and firm characteristics are correlated because the full sample includes firms without any CTR workers in 2011. In the rest of the analysis, we will show our results for both the full and the restricted samples. The results are similar across samples, although the estimates are noisier in the restricted sample due to the reduction in the number of observations.

4.2 Specifications

Our analysis compares the contractual wage growth and the number of categorical promotions of non-CTR workers across firms that experienced different retirement delays among CTR workers, both before and after the implementation of the pension reform. For our analysis of contractual wage growth, the baseline difference-in-differences specification is:

$$y_{it} = \sum_t \beta_t \cdot \text{Delay}_i \cdot \text{time}_t + \alpha_i + \gamma_t + \sum_k \sum_t \zeta_{kt} \cdot \gamma_t \cdot X_{ki} + \epsilon_{it}, \tag{2}$$

where the unit of observation is a firm i in year $t \in \{2009, ..., 2015\}$.

The dependent variable y_{it} measures the average monthly contractual wage growth of non-CTR workers in firm i and year t. The treatment Delay_i is interacted with a time variable: either a post-reform dummy (Post 2011_t) to estimate the average treatment effect in the post-reform period or a full set of year fixed effects (γ_t) to evaluate how the treatment effect changes over time. Prediction (1) from our model is that non-CTR workers will experience lower contractual wage growth in firms with greater retirement delays. This corresponds to negative post-reform coefficients.

The coefficients α_i and γ_t are firm and year fixed effects, respectively. In all specifications, we control for nonlinear trends interacted with several firm characteristics that were not balanced in the full sample before the reform (Section 4.1). We do so by including year dummies (γ_t) interacted with firm characteristics measured in 2009: sector fixed effects and multiple dummy variables that identify firms above the median in terms of average worker age, share of workers who are male, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, and share of workers with age > 55 (X_{ki}).[19]

We also study the effect of retirement delays on non-CTR workers' categorical promotions. These outcomes identify relatively rare career-changing promotions. In the average pre-reform year, there was a categorical promotion in one out of twenty firms (Table 1). When

[19] Our results are robust to the use of alternative nonlinear trends (Tables A5).

we analyze categorical promotions to white-collar jobs, we estimate the following difference-in-differences specification:

$$\begin{aligned} \text{Promotion WC}_{it} &= \sum_t \beta_t^{BC} \cdot \text{Delay BC}_i \cdot \text{time}_t + \sum_t \beta_t^{WC} \cdot \text{Delay WC}_i \cdot \text{time}_t \\ &\quad + \alpha_i + \gamma_t + \sum_k \sum_t \zeta_{kt} \cdot \gamma_t \cdot X_{ki} + \epsilon_{it}. \end{aligned} \tag{3}$$

The dependent variable Promotion WC_{it} measures the number of blue-collar workers promoted to white-collar jobs in firm i and year t. This regression includes two sets of treatment variables: Delay BC_i is the average retirement delay among CTR blue-collar workers in firm i, and Delay WC_i is the average retirement delay among CTR white-collar workers in firm i. Prediction (2) is that the effect on categorical promotions will be larger for retirement delays among white-collar workers than for retirement delays among blue-collar workers. This corresponds to $\beta_t^{WC} < \beta_t^{BC} \leq 0$ for $t > 2011$.

Similarly, we can estimate the following difference-in-differences specifications to analyze changes in the number of categorical promotions to managerial positions:

$$\begin{aligned} \text{Promotion MNG}_{it} &= \sum_t \beta_t^{BWC} \cdot \text{Delay BWC}_i \cdot \text{time}_t \\ &\quad + \sum_t \beta_t^{MNG} \cdot \text{Delay MNG}_i \cdot \text{time}_t \\ &\quad + \alpha_i + \gamma_t + \sum_k \sum_t \zeta_{kt} \cdot \gamma_t \cdot X_{ki} + \epsilon_{it}. \end{aligned} \tag{4}$$

The dependent variable Promotion MNG_{it} measures the number of blue- and white-collar workers promoted to managerial jobs in firm i and year t. The variable Delay BWC_i is the average retirement delay of CTR blue-collar and white-collar workers in firm i, while Delay MNG_i is the average retirement delay of CTR managers in firm i. Again, Prediction (2) is that $\beta_t^{MNG} < \beta_t^{BWC} \leq 0$ for $t > 2011$.

4.3 Pre-Reform Trends in Wage Growth and Promotions

The identifying assumption in our main specifications is that the career progression of younger workers in firms with differential exposure to the reform would have followed the same trends absent the reform. Although this assumption is inherently untestable, we can show that contractual wage growth and categorical promotions followed similar pre-reform trends across firms with different retirement delays among CTR workers. The data indicate that the treatment variable is not predictive of any changes in our career progression variables prior to the implementation of the reform (Table 3). This result does not change if we use different types of linear and nonlinear trends.

Specifically, we first estimate linear trends by regressing contractual wage growth and the number of categorical promotions on the interaction between the treatment and a linear trend while controlling for firm and year fixed effects. The coefficients of the interaction terms are close to zero and not statistically significant in both the full and the restricted sample (panel A, column 1). The findings are similar if we interact the treatment variable with a second-degree polynomial of time or with a full set of year dummies (panels B and C, respectively).

To provide further evidence of the lack of pre-reform effects, we estimate changes in contractual wage growth and categorical promotions had the reform been implemented in either December 2009 or December 2010. If anticipatory responses are present, we should be able to detect significant effects in 2011. As suggested by the hasty implementation of the reform, the placebo treatment effects are all small and not statistically significant at the 5 percent level (Table A4).

5 Empirical Evidence of Career Spillovers

5.1 Do Career Spillovers Exist?

Effects on Contractual Wage Growth We first estimate Equation (2) to analyze the effects of retirement delays on the monthly contractual wage growth of non-CTR workers. We find that contractual wage growth decreases by 0.016 percentage points after 2011 for each one-standard-deviation (hereafter one-σ) increase in retirement delays (Table 4, column 1), consistent with Prediction (1). Compared with a baseline mean of 0.64 percent, these estimates indicate that contractual wage growth falls by 2.5 percent every year after 2011. The results are quantitatively similar if we limit the sample to firms with at least one CTR worker (Table 4, column 4).

Year-specific difference-in-differences estimates allow us to evaluate how the effect changes over time (Figure 2). The coefficients are small and not statistically significant in 2009 and 2010. The treatment effects are negative and statistically significant between 2012 and 2013 and are slightly closer to zero in 2014 and 2015. This U-shaped pattern is consistent with the design of our empirical strategy. Once CTR workers started retiring under the new rules, the cross-firm differences in the short-term exposure to the reform—measured by Delay_i—became less relevant.

Effects on Categorical Promotions We then estimate Equation (3) to analyze changes in the number of categorical promotions to white-collar jobs. In this regression, we include two sets of treatment variables: the retirement delay among blue-collar workers and the retirement delay among white-collar workers.

The results are consistent with Prediction (2). Only retirement delays among those in higher-level positions reduces the rate of promotions to those positions. A one-σ increase in retirement delays among white-collar workers (0.7 years) leads to 0.007 fewer categorical promotions to white-collar positions after 2011 (Table 4, column 2), which corresponds to a reduction in such promotions by 14 percent. Moreover, retirement delays among blue-collar workers, do not have any effect on the number of categorical promotions to white collar positions.

We repeat this analysis with the number of categorical promotions to managerial positions as the dependent variable. In this specification, we include two treatment variables at the firm level: retirement delays among CTR blue- and white-collar workers and retirement delays among CTR managers. Again, consistent with Prediction (2), only retirement delays among managers affect the number of categorical promotions to managerial positions. A one-σ increase in retirement delays among managers decreases the number of non-CTR workers promoted to manager by 0.008 or 16 percent (Table 4, column 3). Retirement delays among lower-ranked workers, in contrast, have a small and statistically insignificant effect.

Year-specific coefficients show a pattern similar to the one we observe for contractual wage growth. The main difference is that the estimates remain negative until 2015 (Figure 3). These results also hold if we limit the sample to firms with at least one CTR worker (Table 4, columns 5 and 6).

Robustness Checks The main results are robust to several modifications to the baseline regressions. For example, instead of including indicators for firms with above-median characteristics, we can interact year dummies with indicators for different tertiles, quartiles, or quintiles of the distributions of firm characteristics observed in 2009 (Table A5). The treatment effects are unchanged across these specifications..

We also control for the share of CTR workers interacted with time dummies, and the effects remain the same. In addition, we control for interactions between time dummies and each of the three sets of characteristics for CTR workers (age, years of contribution, and gender) that determine their retirement-eligibility dates. While our main empirical strategy leverages simultaneous cross-firm variation in age, years of contribution, and gender of CTR workers, the results are the same when we exclude variation in only a single characteristic.

Finally, we show that the findings are robust to the inclusion of nonlinear trends for each province and two-digit sector.

In addition to including more controls, we can show that the results are robust to slight changes to the sample. First, we repeat the main analysis including all non-CTR workers instead of limiting the sample to full-time permanent employees (Table A6). Second, we modify the definition of CTR workers, identifying them as those workers who were eligible in 2011 to retire in the following two, four, or five years (Table A7). In all cases, the main findings are robust.

Finally, the results on categorical promotions are robust to modifications to the dependent variables. Specifically, we can define promotions as the share of categorical promotions per 10 employees rather than using their level (Table A8). The treatment effects on the share and number of categorical promotions are quantitatively similar.

5.2 Where Do Career Spillovers Arise?

In this section, we first test Prediction (3), that career spillovers are larger in slower-growing firms. In fast-growing firms, we would expect that retirement delays are less likely to limit the advancement opportunities for younger workers. The treatment effect that we estimated in Section 5.1 should therefore be most prominent in firms that were not growing before the reform.

To test this prediction, we compute the average yearly employment growth for every firm in the sample between 2009 and 2011. We categorize firms as fast growing if they are in the top tertile of the distribution. On average, employment in these firms increased by 13 percent in the years leading up to the reform. All these firms were growing: The minimum growth rate in the top tertile was 2.9 percent. Similarly, we categorize firms as slow growing if they are in the bottom tertile. In this group, the average firm shrank by 10 percent in the pre-reform period, and the minimum decline was 2.9 percent.

We then compare the differences in the effects of retirement delays on the contractual wage growth of non-CTR workers between fast-growing and slow-growing firms. We estimate the following triple-difference specification:

$$\begin{aligned} y_{it} = & \sum_t \beta_t^s \text{Delay}_i \times \text{time}_t \times \text{Slow}_i + \sum_t \beta_t^f \text{Delay}_i \times \text{time}_t \times \text{Fast}_i \\ & + \sum_t \kappa_t^s \text{time}_t \times \text{Slow}_i + \sum_t \kappa_t^f \text{time}_t \times \text{Fast}_i + \sum_t \kappa_t \text{Delay}_i \times \text{time}_t \\ & + \alpha_i + \gamma_t + \sum_k \sum_t \zeta_{kt} \cdot \gamma_t \cdot X_{ki} + \epsilon_{it}, \end{aligned} \tag{5}$$

where the dummy variable Slow_i is equal to 1 for firms in the bottom tertile of pre-reform employment growth, while Fast_i is equal to 1 for firms in the top tertile. The coefficients of interest, β_t^s and β_t^f, indicate whether retirement delays impacted the contractual wage growth of non-CTR workers differently in slow- and fast-growing firms, as compared with firms in the middle tertile of employment growth.

Consistent with Prediction (3), the overall effect of retirement delays on contractual wage growth is concentrated among slow-growing firms. Compared with firms in the middle tertile, the contractual wage growth in slow-growing firms decreased by 0.042 additional percentage points after 2011 for each one-σ increase in average retirement delays (Table A9, column 1; and Figure 4, panel A). This triple interaction corresponds to a 6.6 percent larger decrease in wage growth. We can move from triple interactions back to difference-in-differences estimates (Figure 4, panel B). In slow-growing firms, a one-σ increase in retirement delays decreases contractual wage growth among non-CTR workers by up to 0.051 percentage points. This effect is more than three times larger than the estimate for the average firm (Table 4, column 1). Retirement delays, in contrast, did not affect contractual wage growth in fast-growing firms. In fact, the estimate of $\beta_t^f + \kappa_t$ is positive, although it is small in magnitude (Figure 4, panel B).

Next, we test Prediction (4), that career spillovers are concentrated among firms with larger spans, that is, firms in which a smaller share of jobs are high-level jobs. In such firms, there are likely to be fewer available jobs at the top, and retirement delays are more likely to slow the careers of younger workers.

We measure the firm-level availability of high-level jobs with an indicator that is equal to 1 for firms with an above-median share of top earners. We define top earners as all workers with an above-median wage, relative to a wage distribution calculated within a province, two-digit sector, and firm-size category (above- vs. below-median workforce size). We estimate a triple-difference specification analogous to Equation (5) in which we interact the baseline treatment variable with our indicator for firms with an above-median share of high-level jobs (Table A9, column 2).

Consistent with Prediction (4), retirement delays decrease the contractual wage growth of non-CTR workers only among firms with a below-median share of high-level jobs. In these firms, the contractual wage growth decreased by 0.022 additional percentage points for each one-σ increase in average retirement delays. In contrast, the treatment effect is a precisely estimated zero in firms with an above-median share of high-level jobs.

5.3 Which Workers are Most Affected by Career Spillovers?

We now explore the patterns of career spillovers across different types of workers. Specifically, we look at whether the careers of different sets of non-CTR workers are differentially impacted by these career spillovers.

Our prediction regarding the heterogeneity of career spillovers across different non-CTR workers is Prediction (5), which states that the effects of retirement delays on contractual wage growth are larger for relatively more senior non-CTR workers. To test this prediction, we first divide employees into three age bins: workers who are 35 years or younger, workers who are between 36 and 55 years old, and workers who are above 55 years old. We then estimate (2) separately for each age group and find that the effects are concentrated among middle-aged and older workers.[20]

In the full sample, a one-σ increase in retirement delays decreased contractual wage growth by 0.02 percentage points after 2011 among non-CTR workers aged 36 to 55 (Table A10, column 2; and Figure A5, panel B) and by 0.06 percentage points among non-CTR workers older than 55 (Table A10, column 3; and Figure A5, panel C). These estimates correspond to decreases in contractual wage growth of 3.8 percent and 10.3 percent, respectively. The effects are not statistically or economically significant for workers who are 35 years old or younger (Table A10, column 1; and Figure A5, panel A).

These findings are consistent with Prediction (5), which states that the effect of retirement delays is larger on workers with longer tenure. In practice, if firms use seniority as one of the criteria to assign promotions, retirement delays are more likely to immediately stall the career progressions of older non-CTR workers who have been with the firm longer. In the data, this pattern is reflected by our finding that the treatment effect is larger in magnitude for older non-CTR workers.

5.4 Turnover and Hiring

In this section, we study whether retirement delays have extensive-margin consequences on turnover and hiring.

We start by looking at voluntary turnover. One might expect that having to wait longer to be promoted might lead some non-CTR workers to search for opportunities elsewhere. The model, however, suggests two reasons this may not be the case. First, leaving a firm erases firm-specific progress that has been made toward a promotion if firms use seniority as one of the criteria to promote internally (Prediction (6)). Moreover, career spillovers have larger impacts on the relatively older non-CTR workers who have been with the firm longer (Prediction (5) and Section 5.3). The combination of these two effects suggest that non-CTR workers who are most affected by career spillovers have longer tenure and are closer to earning a promotion, and therefore they have more to lose from resigning. Retirement delays among CTR workers might thus not be enough to push the marginal older non-CTR

[20] In this exercise, we use age as a proxy for tenure because the tenure variable in the dataset is heavily right censored.

worker to leave the company. Ultimately, this question is empirical in nature. Its answer depends on how heavily firms in the sample rely on seniority to promote internal candidates.

We address this question by using voluntary turnover as the dependent variable in Equation (2). Specifically, the dependent variable is the number of non-CTR workers who voluntarily leave firm i in year t.[21] Consistent with Prediction (6), retirement delays among CTR workers do not increase voluntary turnover for non-CTR workers (Table A11, column 1; Figure A6). If anything, the treatment effects are negative after 2011, although the estimates are small in magnitude and not precisely estimated. The same result holds in the restricted sample. Overall, the treatment effects correspond to changes in voluntary turnover between -0.9 percent and 0 percent.[22]

In addition to analyzing workers' responses, we study whether retirement delays had an effect on firms' choices, such as layoffs and hiring. As predicted by the model, firms value promotion opportunities and may respond to retirement delays by increasing involuntary turnover (Prediction (7)) and decreasing hiring (Prediction (8)). A one-σ increase in retirement delays increased the number of layoffs by 0.057 employees and decreased the number of new hires by 0.097 job candidates (Table A11, columns 2 and 4). These estimates suggest that involuntary turnover increased by 12 percent per firm and year, while hiring decreased by 1.8 percent. Moreover, retirement delays within a given category (blue-collar, white-collar, or managers) led to more layoffs and fewer hires of non-CTR workers within the same category than in other categories (Table A11, columns 3 and 5).

One might think that retirement delays only slow the career progressions of younger workers in labor markets with strict employment protection laws, such as Italy. However, our results show that Italian firms were able to respond by laying off part of their workforce despite stringent employment protection. These responses did not fully offset the consequences of the Fornero reform for the remaining non-CTR workers, but they potentially allowed Italian firms to partially ease the consequences of having limited career capacity. This fact suggests that our results are not driven exclusively by the inability of firms to fire unneeded employees, although one might expect the treatment effects to be smaller in more flexible labor markets. Moreover, the magnitudes of the effects estimated in the Italian setting are policy relevant in other settings, given that many other OECD countries have similar, or even stricter, labor laws.[23]

[21] The INPS data include the reason behind any firm separation, allowing us to distinguish voluntary from involuntary turnover.

[22] We do not think that the recession Italy was going through at the time can fully explain these findings as the recession did not push turnover to zero. The average number of vacancies, net of retirees, per firm and year after 2011 was 1.12 positions or 4 percent of the workforce (Table 1).

[23] https://www.oecd.org/els/emp/oecdindicatorsofemploymentprotection.htm.

5.5 Magnitudes

In this section, we discuss the magnitudes of the treatment effects we find. We start by converting the estimated decreases in monthly contractual wage growth to monetary annual losses. A one-σ increase in retirement delays decreases the annual contractual wage growth of non-CTR workers by €62. This estimate corresponds to a 2.5 percent decrease from an annual wage increase of €2,454.[24]

It is also possible to compute the overall effect of the reform over the four post-reform years in our sample. For the average non-CTR worker, the undiscounted four-year loss is equal to €718. Discounting future periods reduces this effect to a loss of between €592 and €676, depending on the discount rate.[25] In other words, the reform led to total wage losses for non-CTR workers of between 24 percent and 28 percent of a year's wage growth. These losses are much larger for non-CTR workers in slow-growing firms (between €1,598 and €1,928) and for older non-CTR workers (between €2,435 and €2,951).[26]

Another way to view the magnitudes of our results is to benchmark them against other drivers of wage growth.[27] For example, Wheeler (2006) finds a positive relationship between workers' wage growth and the population of the location in which they live. Our estimated 2.5 percent decrease in annual wage growth for a one-σ increase in retirement delays is approximately the same effect size as that generated by living in a location with $794,000$ fewer residents. This would be like moving from Boston, the 10th most populous metro area in the US, to Detroit, the 14th, in terms of its effects on wage growth.

As another example, Barron, Black, and Loewenstein (2002) study the relationship between employer size and wage growth. They find that new hires at larger employers receive a higher salary but experience lower wage growth. We estimate that a one-σ increase in retirement delays has the same size effect on wage growth as moving from a firm with 26 employees (the average in our sample) to a firm with 179 employees. As a final example, the magnitude of our baseline result is similar to the effect of receiving 1.6 days of on-the-job full-time training (Bartel, 2002), which is close to the amount of on-the-job training that the average worker receives over the course of six months.

[24] We compute the average yearly wage in the sample as the average daily gross wage (102.83; Table A3, column 3) multiplied by 300, the average number of working days in the Italian labor market.

[25] The discount rates are 10 percent and 3 percent, respectively.

[26] Repeating the analysis on the restricted sample leads to quantitatively similar findings (Table A12, columns 4 to 6).

[27] As a disclaimer, these coefficients measure correlations from fixed-effect specifications, not necessarily causal effects from randomized controlled trials or natural experiments.

5.6 The Effect of More Workers Facing Retirement Delays

Our main analysis uses cross-firm variation in the average retirement delays among CTR workers but does not take advantage of cross-firm differences in the share of the firm's workforce that is close to retirement and therefore directly impacted by the reform. In this section, we examine the effect of delaying the retirement of an additional CTR worker on the career progression of non-CTR workers.

To this end, we build an alternative treatment, which we call Share of treated workers$_i$, that measures the share of full-time permanent workers who are CTR and face retirement delays of at least one year due to the reform. The main concern with using this variable as our treatment is that it may be correlated with the total number of CTR workers employed by the firm. Using Share of treated workers$_i$ as the main treatment variable in Equation (2), therefore, could lead to biased estimates, as it may conflate the effect of the treatment with the fact that some firms have older workforces, and such firms may differ systematically in how they manage their workers' career progressions. To avoid this concern, we estimate instrumental variable specifications in which we instrument for Share of treated workers$_i$ with Delay$_i$. The first identifying assumption underlying this approach is that Delay$_i$ is strongly correlated with the endogenous variable Share of treated workers$_i$. A one-σ increase in retirement delays among a firm's CTR workers increases the share of treated workers by 2 percentage points or 200 percent, and the Kleibergen-Paap F-statistic exceeds 1,000 in all of our specifications(Table A13, panel A). The second identifying assumption is that Delay$_i$ is not correlated with unobservable factors determining the career progression of non-CTR workers, an assumption we discussed in Section 4.

The IV estimates show that a higher share of positions blocked by CTR workers leads to a larger decrease in the contractual wage growth of non-CTR coworkers. A one-σ increase in the share of CTR workers (3 percent) decreases the wage growth of non-CTR workers by 0.024 percentage points after 2011 (Table A13, panel A, column 2). These findings also hold if we compute the new treatment variable as the share of CTR workers whose retirement was delayed by at least two years (Table A13, panel B, column 2) or three years (Table A13, panel C, column 2).[28]

6 Discussion of Alternative Career-Spillover Channels

In this section, we discuss the extent to which other wage-determination mechanisms can explain our findings. Our key finding is that the career progression of non-CTR workers is

[28] We also estimate these IV effects on the probability of non-CTR workers being promoted to a white-collar or managerial position (Table A13, columns 2 and 3; Figure A8, and Figure A9) and on the restricted sample (Table A13, columns 5 to 6).

slowed when their senior colleagues face retirement delays, especially in slow-growing firms. Many workhorse models of wage determination, in their most basic forms, cannot capture the wage and promotion dynamics that arise from these career spillovers, as they treat workers' careers independently,[29] and so any explanation of our findings must involve career spillovers. Aside from the blocked-promotions channel for career spillovers that we describe in Section 3, there are at least three other potential career-spillovers channels that have been identified in the literature and that we will discuss: spillovers arising from firm-level financial difficulties, team production spillovers, and informational spillovers.

The first alternative channel through which career spillovers can arise is payroll shocks. Unexpected retirement delays, combined with limitations in firing workers, might increase the firm's future payroll costs and force financially constrained firms to postpone planned promotions. A key distinction between this payroll-shock channel and our blocked-promotions channel lies in the pattern of spillovers they imply. The effects of a payroll shock should depend only on the overall magnitude of the increase in future payroll costs and not directly on where in the organization payroll costs increase. For the blocked-promotions channel, where retirement delays occur within the firm matters. To explore this distinction, we first look at where retirement delays occur and ask whether it affects promotion opportunities differently. Second, we construct a measure of payroll shocks. We look at where payroll shocks occur and whether that affects wage growth and categorical promotions.

First, recall our findings in Section 5.1, where we find that retirement delays among blue-collar workers do not affect the probability of non-CTR blue-collar workers being promoted to white-collar jobs. Retirement delays among blue-collar workers, however, should increase the future payroll costs nevertheless and reduce the firm's ability to afford to promote its non-CTR workers. Similarly, retirement delays among blue- and white-collar workers do not decrease the number of internal promotions to managerial jobs, even though they too are an unexpected financial burden for firms. These patterns conflict with a pure payroll-shocks account of our findings.

Second, we examine the payroll-shocks channel directly by measuring the effects of payroll shocks on career progression. For this purpose, we create a new treatment variable, blocked wages, that measures the predicted additional wages that each firm was expected to pay to its average CTR worker as a result of retirement delays.[30] If payroll shocks are the sole driver of slower career progression, one additional dollar of blocked wages for CTR workers in any job category or anywhere in the firm's wage distribution should have the same effect

[29] See, for example, Lazear (1979), Jovanovic (1979), Harris and Holmström (1982), Prendergast (1993), Farber and Gibbons (1996), Gibbons and Waldman (1999), and Bose and Lang (2017).

[30] For each worker, we multiply their retirement delay by their wage (divided by $10,000). Then, we compute the average blocked wages at the firm level for different subgroups of workers.

on non-CTR workers' careers.

To examine this hypothesis, we first regress the average contractual wage growth of non-CTR workers on the average blocked wages of CTR workers in the top, middle, and bottom tertiles of the firm's wage distribution (Table A14, column 1). The effects differ depending on where in the wage distribution the blocked wages occur: Blocking $1 of wages in the middle tertile has a more negative effect on the wage growth of non-CTR workers than blocking $1 of wages in the top or bottom tertile. Similarly, blocking $1 of wages at the top has a more negative effect than blocking $1 of wages at the bottom, but these two coefficients are not statistically different from each other.

Next, we repeat this analysis using the number of categorical promotions as a dependent variable (Table A14, columns 2 and 3). In this case, blocking $1 of wages among CTR white-collar workers has a negative effect on the number of non-CTR blue-collar workers being promoted to white-collar jobs, but blocking $1 of wages among CTR blue-collar workers does not have any effect. The findings are similar, albeit less precise, for categorical promotions to managerial jobs.

Taken together, these patterns indicate that where payroll shocks occur within a firm matters for the career progression of non-CTR workers. They also therefore conflict with a pure payroll-shocks account of our main findings. We carry out one additional exercise that focuses more on the financial-constraint side of the payroll-shocks channel. In particular, we conduct industry heterogeneity analysis, using industry-level differences in firms' access to capital (Hut, 2019). To do so, we estimate a quadruple-difference specification in which we interact the treatment variables in Equation (5) with an indicator for four-digit sectors with an above-median share of firms at high risk of default. This variable, provided by one of the main credit rating agencies in Italy (Cerved), measures the sector-level incidence of firms with serious problems in meeting short-term financial commitments. It is commonly used by banks to make lending decisions. Consistent with the presence of financial constraints, the effects tend to be larger in sectors with higher default risk and lower access to credit. We do, however, also find that retirement delays decrease the contractual wage growth of non-CTR workers in slow-growing firms that operate in sectors with high access to credit (Table A15, column 1).

The second alternative source of career spillovers is team production (Hayes, Oyer, and Schaefer, 2006). For example, a worker's wages might increase if a coworker with complementary skills stays longer at the firm, and they might decrease if that coworker has substitute skills (Jäger and Heining, 2019). Such a team-production explanation does not, however, explain our finding that career spillovers arise only in shrinking firms. Our findings are also inconsistent with the view that higher-ranked workers are complements for younger workers. In our setting, CTR workers are older and tend to have higher wages than their coworkers (Table A3). If higher-ranked workers are complements for younger workers, then we would expect to find that the wage growth of non-CTR workers would increase when CTR workers face longer retirement delays. We find the opposite.

Finally, career spillovers can arise through informational channels (Gibbons and Katz, 1991; Acemoglu and Pischke, 1998; Li, 2013). For example, if the departure of a worker systematically affects the labor market's perception of the skill distribution of the remaining workers, it may affect their outside opportunities and hence their wages. In our setting, the pension reform led to a sudden decrease in the departure rate of older workers, which is plausibly exogenous to the skill of the individual workers affected.

7 Conclusions

This paper investigates whether and how career spillovers arise in internal labor markets. If firms use promotion-based personnel policies but are limited in their ability to promote qualified workers, then one worker's career success can come at the expense of the career progressions of their coworkers. We propose a theoretical framework that generates several testable implications regarding the patterns of these career spillovers in internal labor markets. We test these implications using the 2011 Italian pension reform that abruptly and substantially delayed impending retirements.

We report four main findings. First, retirement delays among older workers in a firm decrease the contractual wage growth and the number of categorical promotions of their younger coworkers. Second, the effects on categorical promotions occur in response to retirement delays among hierarchical superiors but not in response to retirement delays among hierarchical equals. Third, the career spillovers we identify are concentrated among shrinking firms and firms with larger spans—firms that were more likely constrained in their ability to create additional advancement opportunities. Finally, consistent with the use of seniority as an important criterion for allocating promotion slots, the career advancement of relatively more senior workers was relatively more affected.

Taken together, our results suggest that career spillovers play an important role in individual workers' career advancement, especially in firms with limited promotion opportunities. These results have implications for our understanding of internal labor markets. Firms should treat workers' careers as interdependent when developing personnel policies, and firms that attract, retain, and motivate their workers by promising them careers must ensure that they can deliver on those promises. Our results also have implications for the design of public policies. Policies that change eligibility requirements for public pensions might have significant consequences on the career trajectories of younger workers and not just on the older workers who are close to retirement. The gradual aging of the workforce in many OECD countries is projected to make these issues more pressing (OECD, 2017).

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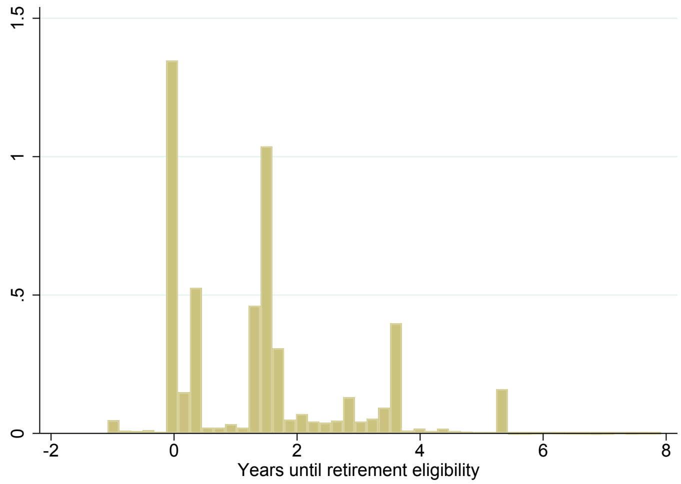
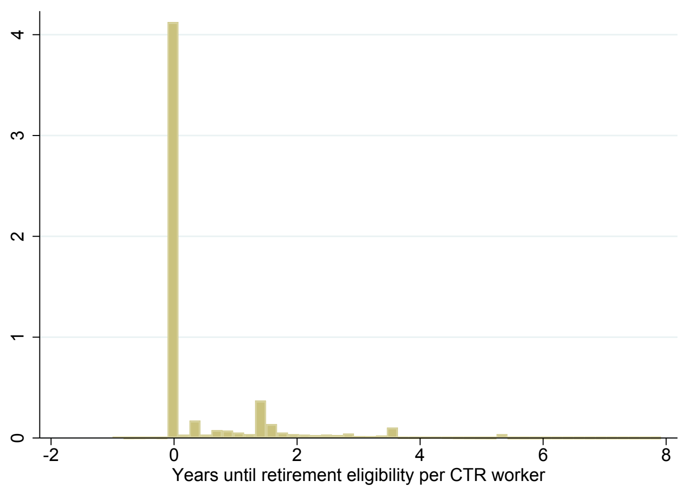
Figures and Tables

Figure 1: Worker- and Firm-Level Treatment



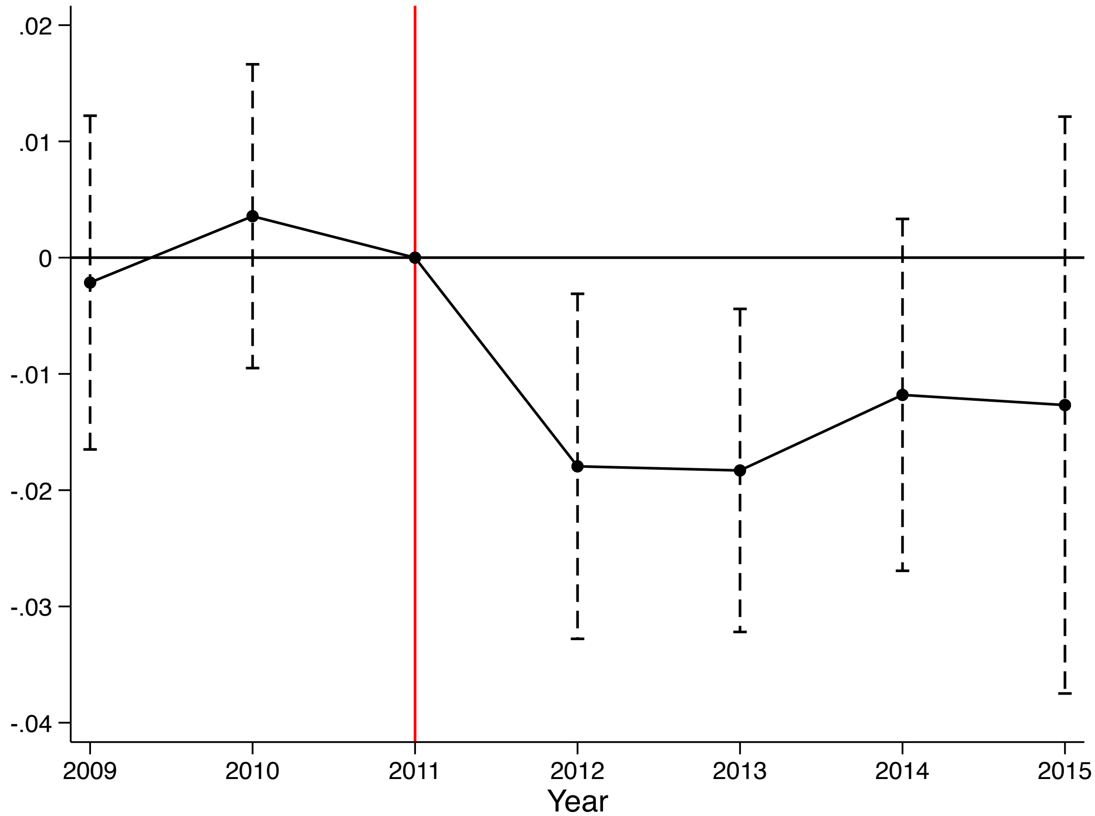
A. Distribution of worker-level retirement delays



B. Distribution of average firm-level retirement delays

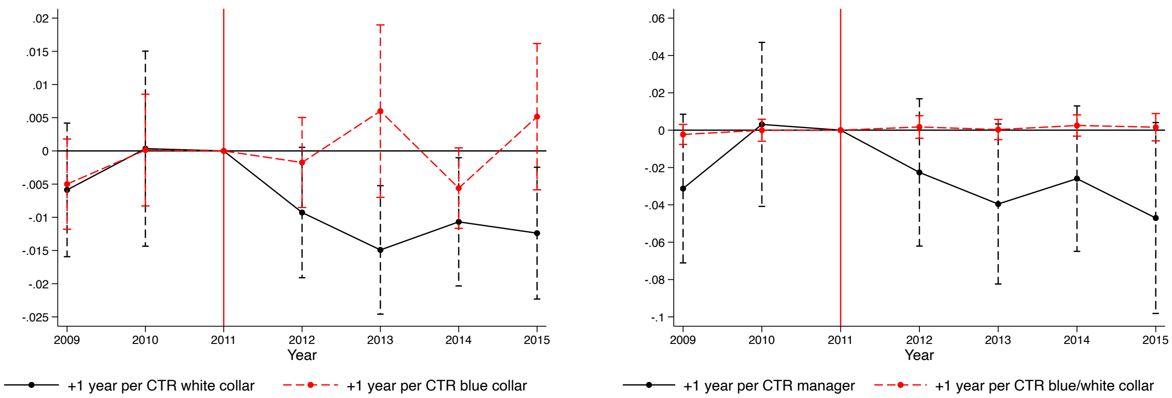
Notes: These graphs show the distribution of retirement delays among CTR (close-to-retirement) workers due to the reform. Workers are considered CTR if they were within three years of retirement in 2011. Panel A shows the distribution of retirement delays at the worker level among CTR workers. The average CTR worker experienced a retirement delay of 1.36 years with a standard deviation of 1.42 years. Panel B shows the distribution of average retirement delays at the firm level. The mean firm-level average retirement delay is 0.44 years, and the standard deviation is 0.97 years.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Figure 2: Effect of Increased Retirement Delays on Contractual Wage Growth of non-CTR Workers



Notes: This graph shows the effect of a one-year increase in the average retirement delay among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. The dependent variable is the average monthly contractual wage growth of workers who were not within three years of retirement in 2011. The treatment variable measures the average retirement delay of CTR workers within each firm. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Standard errors are clustered at the firm level. Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. Mean wage growth in the pre-reform period: 0.64. Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

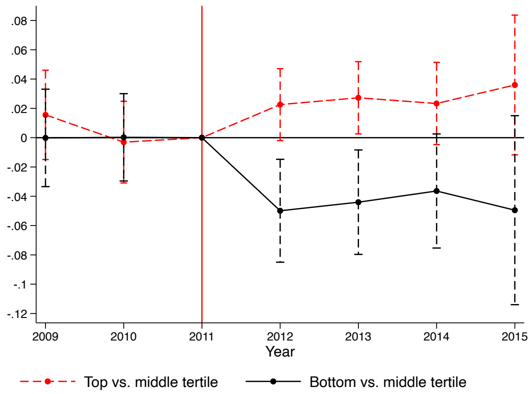
Figure 3: Effect of Increased Retirement Delays on Categorical Promotions of non-CTR Workers



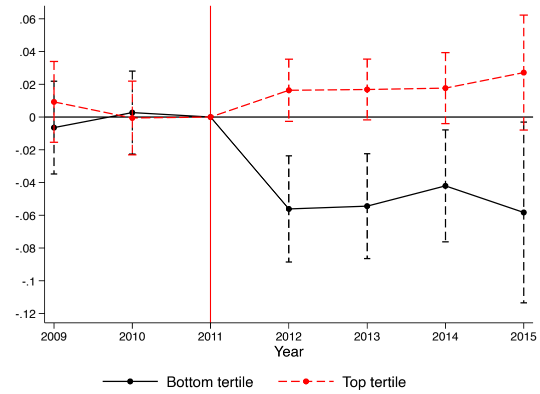
A. Promotion to white collar B. Promotion to manager

Notes: This graph shows the effect of a one-year increase in the average retirement delay among a firm's CTR (close-to-retirement) workers on the number of categorical promotions of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. In panel A, the dependent variable is the number of categorical promotions from blue-collar to white-collar positions per firm and year. This regression estimates the effects of retirement delays among blue-collar (red dashed line) and white-collar CTR workers (black solid line). In panel B, the dependent variable is the number of categorical promotions to managerial jobs per firm and year. This regression estimates the effects of retirement delays among blue- and white-collar CTR workers (red dashed line) and CTR managers (black solid line). The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Standard errors are clustered at the firm level. Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. Mean outcomes in the pre-reform period: 0.05 categorical promotions per firm and year for both panels.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Figure 4: Differential Effects of Increased Retirement Delays by Pre-Reform Employment Growth



A. Triple differences



B. Double differences

Notes: These graphs show the effects of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth of the firm's non-CTR workers, distinguishing between firms with different employment growth between 2009 and 2011. Workers are considered CTR if they were within three years of retirement in 2011. The treatment variable measures the average retirement delay of CTR workers in each firm. These regressions include the interaction between the treatment variable, year fixed effects, and two dummy variables that identify firms in the top and bottom tertile of employment growth before 2011. Panel A shows the estimates of these triple interactions. Panel B shows the difference-in-differences effect of the treatment on wage growth separately for firms in the bottom and top tertile of employment growth before 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Standard errors are clustered at the firm level. Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. The average monthly contractual wage growth in the pre-reform period is 0.64 percent.

Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table 1: Summary Statistics

	2009-2011 (1)	2012-2015 (2)
Turnover (all employees)		
Retirees	0.067	0.055
	(0.384)	(0.437)
Vacancies	1.370	1.114
	(2.805)	(2.874)
Contractual wage growth and categorical promotions (only workers not within three years of retirement eligibility in 2011)		
Contractual wage growth (percentage points)	0.641	0.484
	(2.233)	(2.848)
Promotions blue to white collar	0.048	0.039
	(0.4723)	(0.475)
Promotions blue/white collar to manager	0.052	0.045
	(0.499)	(0.574)
Observations	312,546	416,728

Notes: This table shows averages per firm and year before and after the December 2011 reform. Standard deviations in parentheses. Retirees measures the number of workers retiring per firm and year. Vacancies measure the total number of positions left available by employees leaving the firm (voluntarily or involuntarily): retirements, deaths, layoffs, and quits. Contractual wage growth and categorical promotions are measured only for workers who were not within three years of retirement in 2011. The contractual wage is the monthly wage that each employee should receive based on her labor contract. Unlike take-home pay, it is not affected by transitory shocks such as leaves of absence (maternity, injury, sick) and bonuses. It is, instead, closely related to job titles. Assigning a new job title to an employee, in fact, often requires by law a modification of the contractual wage to reflect the different responsibilities attached to the new position (Art. 2103 c.c.). Firms in the sample: 104,182.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table 2: Relationships between Pre-Reform Characteristics and the Treatment

	Full sample (1)	Mean outcome (2)	Restricted sample (3)	Mean outcome (4)
Firm age	1.604***	18.12	-0.107	21.66
	(0.052)		(0.065)	
Number of employees	6.596***	26.23	1.046***	38.99
	(0.195)		(0.155)	
Average worker age	1.141***	39.24	-0.001	41.66
	(0.018)		(0.019)	
Average daily wage	2.907***	90.69	0.139	96.71
	(0.697)		(0.949)	
Share of male workers	0.000	0.640	-0.022***	0.670
	(0.002)		(0.002)	
Share of full-time workers	0.014***	0.870	-0.005***	0.910
	(0.001)		(0.001)	
Share of blue-collar workers	0.001	0.590	-0.009***	0.600
	(0.002)		(0.002)	
Share of white-collar workers	0.005***	0.330	0.009***	0.330
	(0.002)		(0.002)	
Share of managers	0.004***	0.020	0.001*	0.030
	(0.000)		(0.000)	
Share of workers with age ≤ 35	-0.041***	0.380	-0.000	0.300
	(0.001)		(0.001)	
Share of workers with age between 36 and 55	0.022***	0.550	-0.005***	0.600
	(0.001)		(0.001)	
Share of workers with age > 55	0.019***	0.070	0.005***	0.100
	(0.000)		(0.000)	
Average worker experience	0.900***	14.18	-0.214***	16.35
	(0.023)		(0.025)	
Observations	104,182		33,896	
Treatment mean	0.44		1.36	
Treatment std. dev.	0.97		1.28	

Notes: Each row shows the estimated coefficient $\hat{\beta}_1$ from a different regression: Pre-2011 characteristic$_{fip} = \beta_0 + \beta_1 \cdot \text{Delay}_f + \gamma_{ip} + \varepsilon_{fip}$ in year 2009 for firm f in sector i and province p (γ_{ip} denotes sector–province fixed effects). The variable Delay$_f$ measures the average retirement delay among firm f's CTR (close-to-retirement) workers. A worker is considered close to retirement if she was within three years of retirement in 2011. The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors clustered by province and sector are displayed in parentheses, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table 3: Pre-Reform Trends in Contractual Wage Growth and Categorical Promotions

	Wage growth	Promotion to white	Promotion to manager	Wage growth	Promotion to white	Promotion to manager
	(1)	(2)	(3)	(4)	(5)	(6)
			Panel A: Linear trend			
Delay x Trend	0.0011			-0.0017		
	(0.0037)			(0.0048)		
Delay WC x Trend		0.0029			0.0018	
		(0.0026)			(0.0027)	
Delay MNG x Trend			0.016			0.011
			(0.010)			(0.0098)
			Panel B: Quadratic trend			
Delay x Trend	0.020			-0.0065		
	(0.025)			(0.033)		
Delay x Trend^2	-0.0046			0.0012		
	(0.0062)			(0.0080)		
Delay WC x Trend		0.016			0.014	
		(0.027)			(0.027)	
Delay WC x Trend^2		-0.0033			-0.0030	
		(0.0068)			(0.0066)	
Delay MNG x Trend			0.091			0.060
			(0.082)			(0.084)
Delay MNG x Trend^2			-0.019			-0.012
			(0.020)			(0.021)
			Panel C: Year dummies			
Delay x 2009	-0.0021			0.0035		
	(0.0073)			(0.0096)		
Delay x 2010	0.0036			0.00056		
	(0.0067)			(0.0085)		
Delay WC x 2009		-0.0059			-0.0036	
		(0.0051)			(0.0053)	
Delay WC x 2010		0.00034			0.0011	
		(0.0075)			(0.0073)	
Delay MNG x 2009			-0.031			-0.022
			(0.020)			(0.020)
Delay MNG x 2010			0.0031			0.00098
			(0.022)			(0.023)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	312,546	312,546	312,546	101,688	101,688	101,688
Mean outcome	0.64	0.05	0.05	0.53	0.07	0.09
Treatment mean	0.44	0.17	0.04	1.36	0.53	0.14
Treatment std. dev.	0.97	0.68	0.34	1.28	1.12	0.59

Notes: Delay measures the average retirement delay among all CTR workers, Delay WC measures it for white-collar CTR, and Delay MNG measures it for CTR managers. The dependent variables are the average monthly contractual wage growth (columns 1 and 4), the number of categorical promotions from blue to white collar (columns 2 and 5), and the number of categorical promotions from blue/white collar to manager (columns 3 and 6). They are computed on workers who were not within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009. The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Source: Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table 4: Effects of Increased Retirement Delays on Career Progression of non-CTR Workers

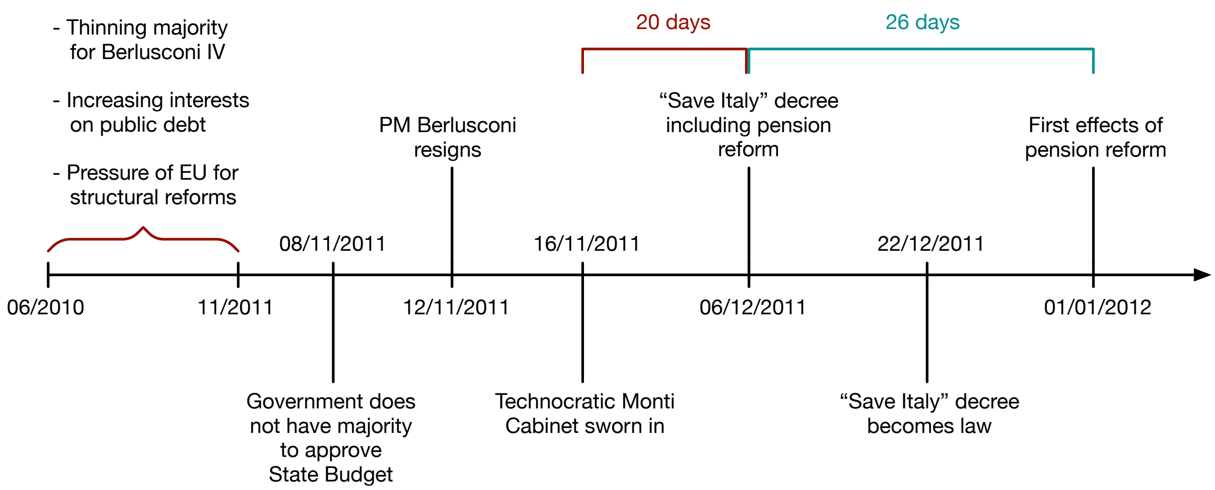
	Wage growth (1)	Promotion to white (2)	Promotion to manager (3)	Wage growth (4)	Promotion to white (5)	Promotion to manager (6)
Delay x Post 2011	-0.016***			-0.014**		
	(0.005)			(0.007)		
Delay BC x Post 2011		0.003			0.002	
		(0.003)			(0.003)	
Delay WC x Post 2011		-0.010***			-0.012***	
		(0.003)			(0.004)	
Delay BWC x Post 2011			0.002			0.002
			(0.001)			(0.002)
Delay MNG x Post 2011			-0.024*			-0.022*
			(0.013)			(0.012)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	729,274	729,274	729,274	237,272	237,272	237,272
R^2	0.26	0.19	0.26	0.29	0.19	0.28
Mean outcome	0.64	0.05	0.05	0.53	0.07	0.09
Treatment mean	0.44	0.17	0.04	1.36	0.53	0.14
Treatment std. dev.	0.97	0.68	0.34	1.28	1.12	0.59
P-value WC<BC		<0.001			<0.001	
P-value MNG<BWC			0.019			0.027

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth and number of categorical promotions of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. Delay measures the average retirement delay for all CTR workers, for white-collar CTR workers (WC), for blue-collar CTR workers (BC), for blue- and white-collar CTR workers (BWC), or for CTR managers (MNG). The dependent variables are the average monthly contractual wage growth (columns 1 and 4), the number of categorical promotions from blue to white collar (columns 2 and 5), and the number of categorical promotions from blue/white collar to manager (columns 3 and 6). They are computed on workers who were not within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Online Appendix - Not For Publication

A Additional Figures and Tables

Figure A1: Timeline of Fornero Pension Reform

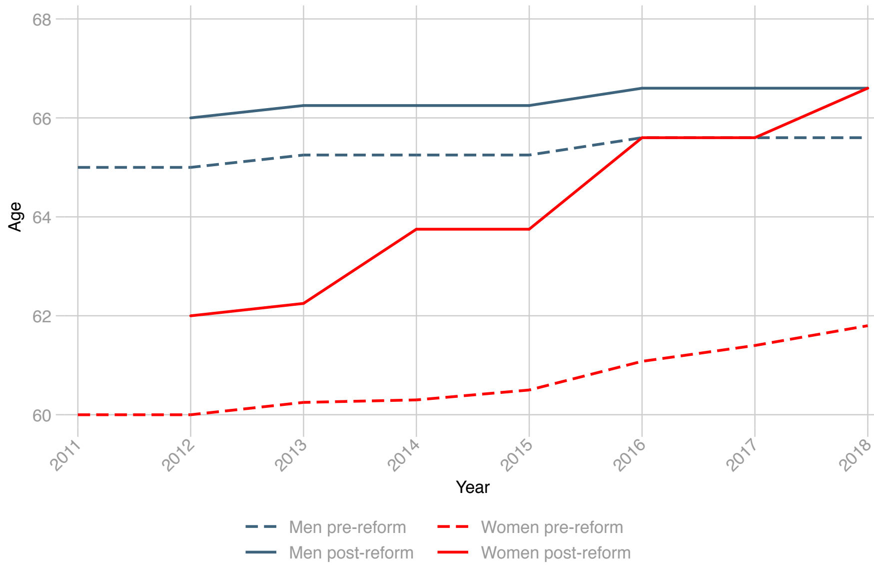


Notes: This figure describes the introduction of the 2011 pension reform in Italy.

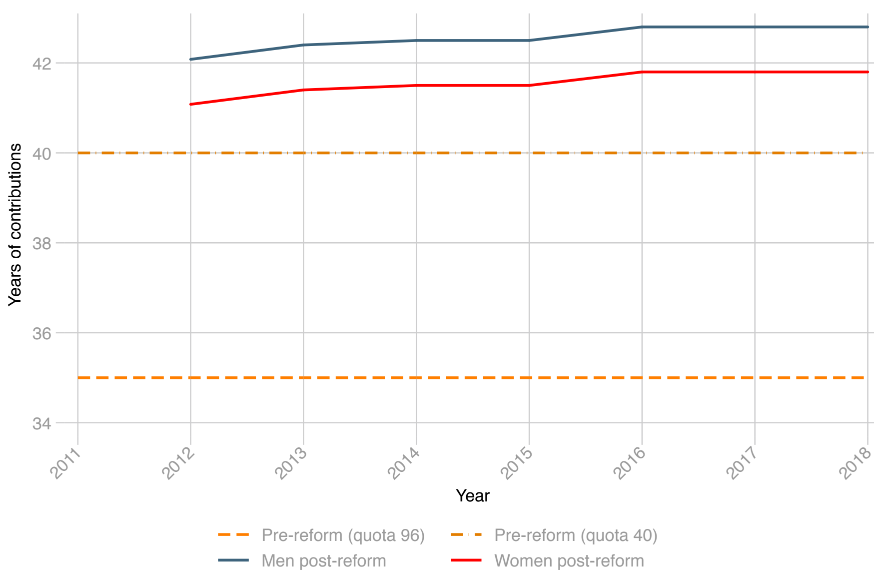
Figure A2: Selected Headlines

Notes: Headlines of the national newspaper *La Stampa*, http://archivio.lastampa.it/.

Figure A3: Fornero Reform Changes to the Pension-Eligibility Criteria



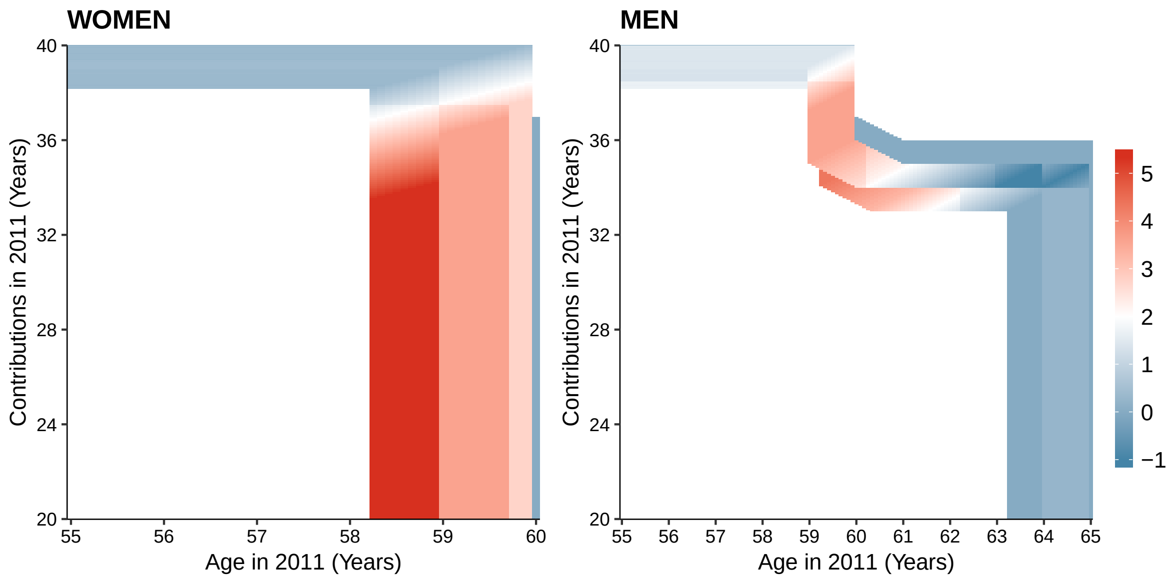
A. Age-based critera



B. Seniority-based criteria

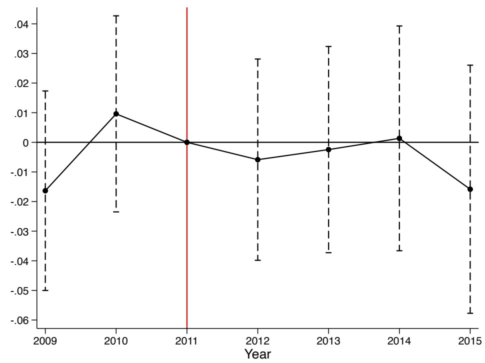
Notes: This graph shows how requirements for claiming an age-based (Panel A) or a seniority-based (Panel B) pension changed after 2011. Panel A shows the change in age requirements for age-based pensions by gender. The requirement on years of retirement contributions (20 years) is constant before and after the reform and across genders. Panel B shows the change in contribution requirements for seniority-based pensions. Before the reform was implemented, man and women had the same requirements. Quota 40 had no additional requirement on age, while quota 96 required more than 60 years of age. After the reform, there is no requirement on age.
Source: Authors' elaborations based on information from Istituto Nazionale della Previdenza Sociale (INPS).

Figure A4: Effect of the Reform on Different Workers

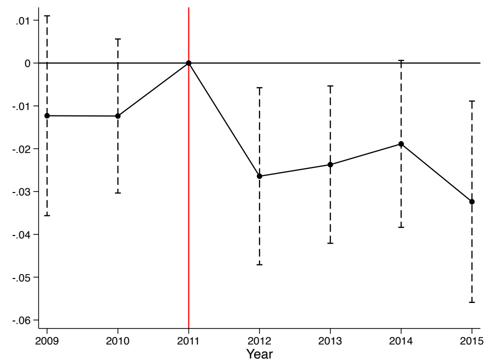


Notes: This heatmap shows how retirement delays differ by gender, age, and years of contribution for CTR workers.
Source: Authors' elaborations based on information from Istituto Nazionale della Previdenza Sociale (INPS).

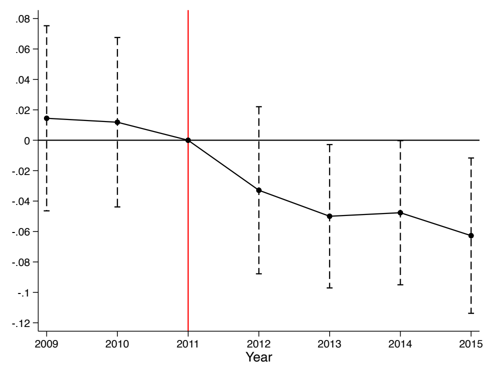
Figure A5: Effects of Increased Retirement Delays among CTR Workers on Different Age Groups



A. Age≤35



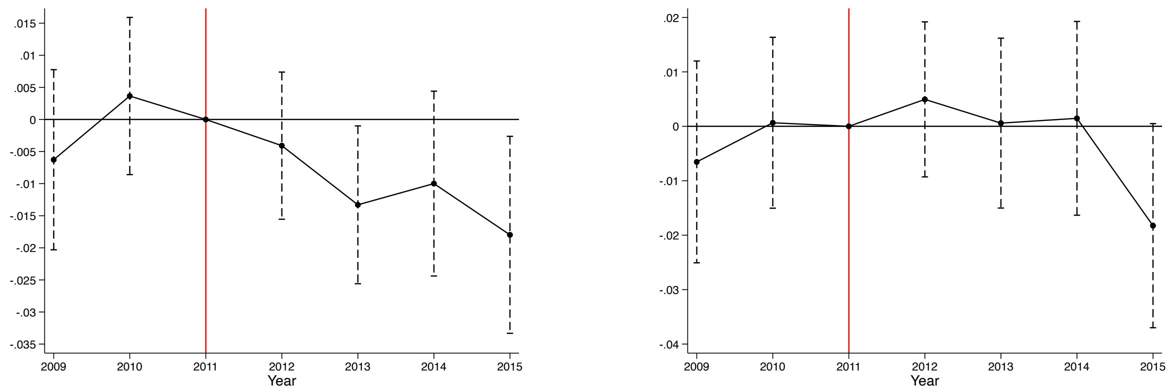
B. 35<Age≤55



C. Age>55

Notes: These graphs show the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth of the firm's non-CTR workers, distinguishing by age group. Workers are considered CTR if they were within three years of retirement in 2011. The treatment variable measures the average retirement delay of CTR workers. The dependent variable is the average contractual wage growth of non-CTR workers in different age groups. Panel A focuses on workers who are 35 years old or younger, panel B on workers aged between 36 and 55 years old, and panel C on workers who are older than 55. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Standard errors are clustered at the firm level. Number of observations: 401,630 (panel A), 402,722 (panel B), or 351,343 (panel C) firm-year pairs. The average monthly wage growth in the pre-reform period is 0.81 (panel A), 0.52 (panel B), or 0.58 (panel C) percent.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Figure A6: Effect of Increased Retirement Delays among CTR Workers on Voluntary Turnover

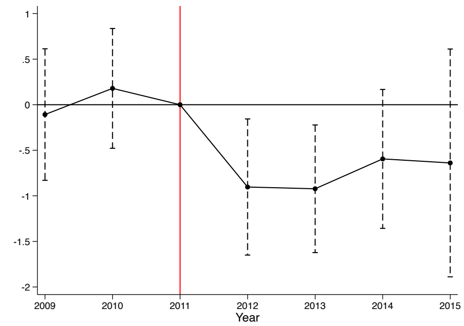


A. Full sample

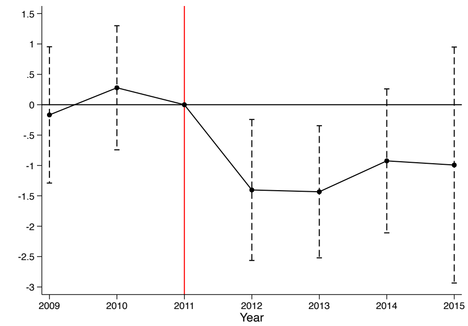
B. Restricted sample

Notes: These graphs show the effects of a one-year increase in retirement delays among a firm's CTR (close-to-retirement) workers on the number of the firm's non-CTR workers who voluntarily leave in each year. Workers are considered CTR if they were within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Panel B limits the sample to firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Number of observations: 729,274 (panel A) and 237,272 (panel B) firm-year pairs. The average number of voluntary separations in the pre-reform period is 0.92 in panel A and 1.04 in panel B. Standard errors are clustered at the firm level.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

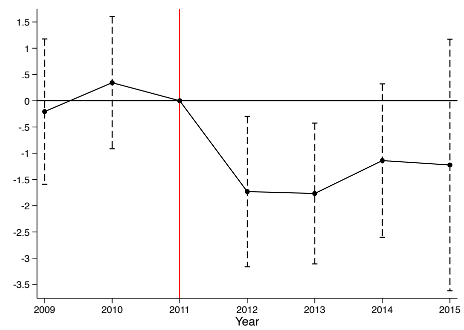
Figure A7: IV, Effect of Blocking More CTR Workers on Wage Growth of Other Employees



A. Effect of 100% increase in share of workers with increase in retirement age ≥ one year



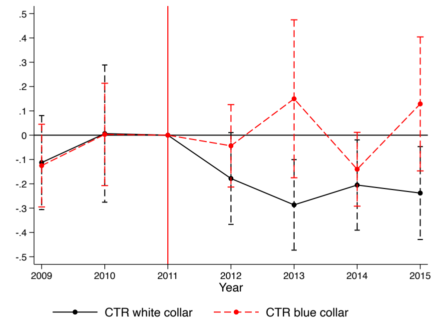
B. Effect of 100% increase in share of workers with increase in retirement age ≥ two years



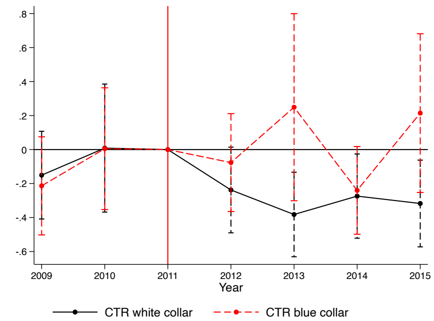
C. Effect of 100% increase in share of workers with increase in retirement age ≥ three years

Notes: This graph shows the effect of a 100 percent increase in the share of CTR (close-to-retirement) workers whose retirement age increased by at least one year (panel A), two years (panel B), or three years (panel C) on the wage growth of other full-time permanent employees at the firm. The regressions instrument the share of CTR workers (for example, number of CTR workers whose retirement age increased by at least one year over total workforce) with the average retirement delay of CTR workers (sum of all the additional years that CTR workers will have to spend at the firm over number of CTR workers). The dependent variable is the average wage growth of workers who were not within three years of retirement in 2011. The regressions also include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector fixed effects, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤35, share of workers with age between 36 and 55, share of workers with age >55). Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. The average monthly wage growth in the pre-reform period is 0.64 percent. Standard errors are clustered at the firm level.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

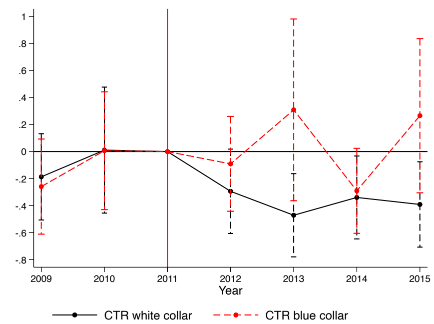
Figure A8: IV, Effect of Blocking More CTR Workers on Promotion to White-Collar Jobs



A. Effect of 100% increase in share of workers with increase in retirement age ≥ one year



B. Effect of 100% increase in share of workers with increase in retirement age ≥ two years



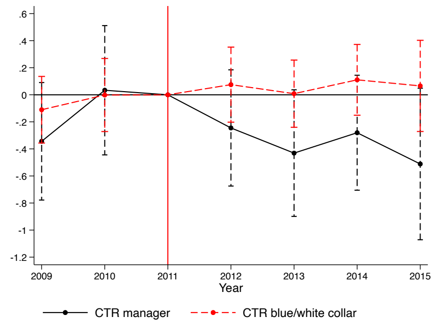
C. Effect of 100% increase in share of workers with increase in retirement age ≥ three years

Notes: This graph shows the effect of a 100 percent increase in the share of CTR (close-to-retirement) workers whose retirement age increased by at least one year (panel A), two years (panel B), or three years (panel C) on the number of promotions to white-collar jobs of other full-time permanent employees at the firm. The regressions instrument the share of CTR workers (for example, number of CTR workers whose retirement age increased by at least one year over total workforce) in blue-collar and white-collar jobs with the average retirement delay of CTR workers in the same positions (in the case of white-collar workers, for example, sum of all the additional years that CTR white-collar workers will have to spend at the firm over number of CTR white-collar workers). The dependent variable is the number of promotions from blue collar to white collar positions per firm and year. The regressions also include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector fixed effects, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤35, share of workers with age between 36 and 55, share of workers with age >55). Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. Mean outcomes in the pre-reform period: 0.05 promotions per firm and year. Standard errors are clustered at the firm level.

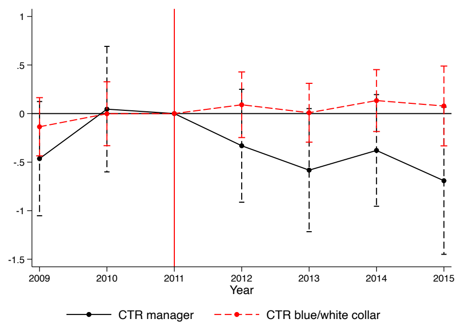
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Figure A9: IV, Effect of Blocking More CTR Workers on Promotions to Managerial Jobs

A. Effect of 100% increase in share of workers with increase in retirement age $\geq$ one year



B. Effect of 100% increase in share of workers with increase in retirement age $\geq$ two years



C. Effect of 100% increase in share of workers with increase in retirement age $\geq$ three years

Notes: This graph shows the effect of a 100 percent increase in the share of CTR (close-to-retirement) workers whose retirement age increased by at least one year (panel A), two years (panel B), or three years (panel C) on the number of promotions to managerial jobs of other full-time permanent employees at the firm. The regressions instrument the share of CTR workers (for example, number of CTR workers whose retirement age increased by at least 1 year over total workforce) in blue-/white-collar positions and in managerial jobs with the average retirement delay of CTR workers in the same positions (in the case of managers, for example, sum of all the additional years that CTR managers will have to spend at the firm over number of CTR managers). The dependent variable is the number of promotions from blue-/white-collar to managerial positions per firm and year. The regressions also include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector fixed effects, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Number of observations: 729,274 firm-year pairs. Firms in the sample: 104,182. Mean outcomes in the pre-reform period: 0.05 promotions per firm and year. Standard errors are clustered at the firm level.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A1: Eligibility Rules for Pensions

Panel A: Age-based criteria

	Men		Women	
	Pre-reform	Post-reform	Pre-reform	Post-reform
		Age requirement		
2011	65 y.o.	Not in place	60 y.o.	Not in place
2012	65 y.o.	66 y.o.	60 y.o.	62 y.o.
2013	65.25 y.o.	66.25 y.o.	60.25 y.o.	62.25 y.o.
2014	65.25 y.o.	66.25 y.o.	60.33 y.o.	63.75 y.o.
2015	65.25 y.o.	66.25 y.o.	60.50 y.o.	63.75 y.o.
2016	65.58 y.o.	66.58 y.o.	61.08 y.o.	65.58 y.o.
2017	65.58 y.o.	66.58 y.o.	61.42 y.o.	65.58 y.o.
2018	65.58 y.o.	66.58 y.o.	61.83 y.o.	66.58 y.o.
		Contribution requirement		
	20 y.c.	20 y.c.	20 y.c.	20 y.c.
		Waiting window		
	12 months	None	12 months	None

Panel B: Seniority-based criteria

	Pre-reform	Post-reform	
	Men and Women	Men	Women
2011	Quota 96 (min 60 y.o. and 35 y.c.)	Not in place	
2012	Quota 96 (min 60 y.o. and 35 y.c.)	42.08 y.c.	41.08 y.c.
2013	Quota 97.3 (min 61.25 y.o. and 35 y.c.)	42.42 y.c.	41.42 y.c.
2014	Quota 97.3 (min 61.25 y.o. and 35 y.c.)	42.50 y.c.	41.50 y.c.
2015	Quota 97.3 (min 61.25 y.o. and 35 y.c.)	42.50 y.c.	41.50 y.c.
2016	Quota 97.6 (min 61.58 y.o. and 35 y.c.)	42.83 y.c.	41.83 y.c.
2017	Quota 97.6 (min 61.58 y.o. and 35 y.c.)	42.83 y.c.	41.83 y.c.
2018	Quota 97.6 (min 61.58 y.o. and 35 y.c.)	42.83 y.c.	41.83 y.c.
	Waiting window		
	12 months	None	

Notes: This table shows the age-based and seniority-based eligibility criteria under the old and new rules. Age-based eligibility also requires at least 20 years of contribution to social security, both under the old and new rules. In the table, "y.o." stands for "years old," "y.c." stands for "years of contribution." Under the old rules, workers also became eligible under the seniority-based criteria after 40 years of contribution, regardless of age. The waiting window is the number of months between retirement eligibility and actual pension disbursement.
Source: Authors' elaborations based on information from Istituto Nazionale della Previdenza Sociale (INPS).

Table A2: Summary Statistics of Firms in Sample

	All Firms		Firms with $\geq$ 1 CTR worker		Firms with no CTR workers	
	mean (1)	sd (2)	mean (3)	sd (4)	mean (5)	sd (6)
Firm size	26.23	26.62	38.99	36.56	20.08	17.00
Firm age	18.12	12.03	21.66	12.86	16.41	11.22
Share in manufacturing	0.45	0.50	0.52	0.50	0.41	0.49
Share in services	0.54	0.50	0.46	0.50	0.58	0.49
Share male	0.64	0.29	0.67	0.27	0.63	0.30
Avg. workforce age	39.24	4.63	41.66	3.79	38.06	4.54
Share aged $\leq$ 35	0.38	0.20	0.30	0.16	0.42	0.21
Share aged (35-55]	0.55	0.18	0.60	0.15	0.53	0.19
Share aged $>$ 55	0.06	0.08	0.10	0.08	0.05	0.07
Avg. workforce tenure	7.07	4.21	8.61	4.42	6.32	3.89
Avg. workforce experience	14.66	4.30	16.74	3.68	13.65	4.21
Share blue collar	0.59	0.32	0.60	0.30	0.58	0.33
Share white collar	0.33	0.30	0.33	0.27	0.33	0.31
Share manager	0.02	0.07	0.03	0.07	0.02	0.07
Share full-time contracts	0.87	0.18	0.91	0.13	0.86	0.19
Share temporary contracts	0.09	0.14	0.07	0.11	0.10	0.15
Avg. real daily wage	90.69	133.24	96.71	178.10	87.79	104.85
Share CTR workers	0.02	0.04	0.06	0.05	0.00	0.00
Observations	104,182		33,896		70,286	

Notes: This table shows summary statistics in 2009 for firms in the sample. Tenure and experience are censored before 1983. Columns 1 and 2 show means and standard deviations for all 104,182 firms in the full sample, which includes all private-sector non-agricultural firms that (1) employed between 10 and 200 employees in 2009, (2) were active every year between 2009 and 2015, and (3) employed at least one full-time permanent worker in every year. The remaining columns divide these firms into two subgroups: (1) firms with at least one CTR (close-to-retirement) worker in 2011 and (2) firms with no CTR workers in 2011. Columns 3 and 4 show means and standard deviations for all 33,896 firms that employed at least 1 CTR worker in 2011. Columns 5 and 6 show means and standard deviations for all 70,286 firms that employed no CTR workers in 2011.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A3: Summary Statistics of Workers in Sample

	CTR workers		Non-CTR workers	
	mean	sd	mean	sd
	(1)	(2)	(3)	(4)
Male	0.71	0.45	0.71	0.45
Age	57.68	2.80	40.29	9.64
Tenure	15.09	9.22	8.76	6.92
Experience in private sector	24.55	7.93	15.01	9.37
Years in labor market	39.72	10.40	19.72	15.36
Blue collar	0.64	0.48	0.56	0.50
White collar	0.29	0.46	0.36	0.48
Manager	0.06	0.24	0.04	0.20
Daily gross real wage	113.71	113.72	102.83	114.40
Observations	87,354		2,736,586	

Notes: This table shows summary statistics in 2009 for workers in the sample, that is, workers who work for private-sector non-agricultural firms that (1) employed between 10 and 200 employees in 2009, (2) were active every year between 2009 and 2015, and (3) employed at least one full-time permanent worker every year. Tenure and experience are censored before 1983. Columns 1 and 2 show means and standard deviations for CTR (close-to-retirement) workers. Columns 3 and 4 show means and standard deviations for non-CTR workers.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A4: Placebo Reforms

	Wage growth	Promotion to white	Promotion to manager	Wage growth	Promotion to white	Promotion to manager
	(1)	(2)	(3)	(4)	(5)	(6)
			Panel A: Placebo reform in December 2009			
Delay x Post 2009	0.004	0.004	0.004*	-0.003	0.003	0.002
	(0.007)	(0.002)	(0.002)	(0.009)	(0.003)	(0.003)
			Panel B: Placebo reform in December 2010			
Delay x Post 2010	-0.001	0.001	0.001	-0.002	0.000	0.002
	(0.006)	(0.003)	(0.002)	(0.008)	(0.003)	(0.003)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	312,546	312,546	312,546	101,688	101,688	101,688
Mean outcome	0.64	0.05	0.05	0.53	0.07	0.09
Treatment mean	0.44	0.17	0.04	1.36	0.53	0.14
Treatment std. dev.	0.97	0.68	0.34	1.28	1.12	0.59

Notes: These regressions estimate the effect of a placebo reform that would have been implemented in December, 2009 (panel A) or December, 2010 (panel B), instead of December, 2011. We include only data from 2009, 2010, and 2011. Delay measures the average retirement delay among CTR (close-to-retirement) workers. The dependent variables are the average monthly contractual wage growth (columns 1 and 4), the number of categorical promotions from blue to white collar (columns 2 and 5), and the number of categorical promotions from blue/white collar to manager (columns 3 and 6) for non-CTR workers. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A5: Additional Controls to Baseline Specifications

	Wage growth (1)	Promotion to white (2)	Promotion to manager (3)	Wage growth (4)	Promotion to white (5)	Promotion to manager (6)
Baseline	-0.016***	-0.010***	-0.024*	-0.014**	-0.012***	-0.022*
	(0.005)	(0.003)	(0.013)	(0.007)	(0.004)	(0.012)
Tertiles	-0.016***	-0.0093***	-0.023*	-0.014**	-0.011***	-0.022*
	(0.005)	(0.003)	(0.013)	(0.007)	(0.003)	(0.012)
Quartiles	-0.016***	-0.0091***	-0.023*	-0.014**	-0.011***	-0.022*
	(0.005)	(0.003)	(0.013)	(0.007)	(0.003)	(0.012)
Quintiles	-0.016***	-0.0092***	-0.022*	-0.014**	-0.011***	-0.021*
	(0.005)	(0.003)	(0.013)	(0.007)	(0.003)	(0.012)
Share CTR	-0.025***	-0.011***	-0.025*	-0.014**	-0.012***	-0.022*
	(0.007)	(0.003)	(0.013)	(0.007)	(0.004)	(0.012)
CTR features	-0.018***	-0.011***	-0.025**	-0.012*	-0.012***	-0.023*
	(0.007)	(0.003)	(0.013)	(0.007)	(0.004)	(0.012)
Province-sector-year	-0.017***	-0.0089***	-0.017	-0.015**	-0.012***	-0.010
	(0.006)	(0.003)	(0.011)	(0.006)	(0.003)	(0.011)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	729,274	729,274	729,274	237,272	237,272	237,272
Mean outcome	0.64	0.05	0.05	0.53	0.07	0.09
Treatment mean	0.44	0.17	0.04	1.36	0.53	0.14
Treatment std. dev.	0.97	0.68	0.34	1.28	1.12	0.59

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth and number of categorical promotions of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. Each cell contains the estimated coefficient of Delay x Post 2011 (columns 1 and 4), Delay WC x Post 2011 (columns 2 and 5), or Delay MNG x Post 2011 (columns 3 and 6) from separate regressions. **Baseline**: The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). **Tertiles**: Instead of dummies for firms above the median, these regressions include dummy variables that identify different *tertiles* of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, and share of workers with age > 55. **Quartiles**: These regressions include dummy variables that identify different *quartiles* of the distributions of the same variables. **Quintiles**: These regressions include dummy variables that identify different *quintiles* of the distributions of the same variables. **Share CTR**: In addition to all the controls in the baseline, these regressions include the share of CTR workers in 2011 interacted with year fixed effects. **CTR features**: In addition to all the controls in the baseline, these regressions include three variables describing the CTR workers in each firm (mean age, mean years of contribution in 2011, and male share) interacted with year fixed effects. **Province-sector-year**: In addition to all the controls in the baseline, these regressions include province-sector (two-digit NACE Rev. 2)-year fixed effects. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, INPS.

Table A6: All non-CTR Workers

	Wage growth	Promotion to white	Promotion to manager	Wage growth	Promotion to white	Promotion to manager
	(1)	(2)	(3)	(4)	(5)	(6)
Delay x Post 2011	-0.016***			-0.023***		
	(0.006)			(0.007)		
Delay BC x Post 2011		0.004			0.003	
		(0.003)			(0.004)	
Delay WC x Post 2011		-0.013***			-0.014***	
		(0.004)			(0.004)	
Delay BWC x Post 2011			0.002			0.002
			(0.002)			(0.002)
Delay MNG x Post 2011			-0.024*			-0.022*
			(0.013)			(0.013)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	729,274	729,274	729,274	237,272	237,272	237,272
R^2	0.44	0.20	0.26	0.43	0.20	0.28
Mean outcome	1.10	0.07	0.06	0.87	0.10	0.09
Treatment mean	0.44	0.17	0.04	1.36	0.53	0.14
Treatment std. dev.	0.97	0.68	0.34	1.28	1.12	0.59
P-value WC<BC		<0.001			<0.001	
P-value MNG<BWC			0.023			.032

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR workers on the average monthly contractual wage growth and number of categorical promotions of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. Delay measures the average retirement delay for all CTR workers, for white-collar CTR workers (WC), for blue-collar CTR workers (BC), for blue- and white-collar CTR workers (BWC), or for CTR managers (MNG). The dependent variables are the average monthly contractual wage growth (columns 1 and 4), the number of categorical promotions from blue to white collar (columns 2 and 5), and the number of categorical promotions from blue/white collar to manager (columns 3 and 6) for non-CTR workers. They are computed on the same baseline sample of firms, but using data on **all non-CTR workers, including those who were not employed full-time or did not have permanent contracts**. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A7: Alternative Definitions of CTR Workers

	Wage growth (1)	Promotion to white (2)	Promotion to manager (3)	Wage growth (4)	Promotion to white (5)	Promotion to manager (6)
	Panel A: CTR workers within two years of retirement in 2011					
Delay x Post 2011	-0.017***			-0.013*		
	(0.006)			(0.007)		
Delay WC x Post 2011		-0.008**			-0.008**	
		(0.004)			(0.004)	
Delay MNG x Post 2011			-0.033**			-0.003**
			(0.015)			(0.015)
Observations	729,274	729,274	729,274	184,702	184,702	184,702
P-value WC<BC		0.030			0.025	
P-value MNG<BWC			0.026			0.048
	Panel B: CTR workers within four years of retirement in 2011					
Delay x Post 2011	-0.018***			-0.015**		
	(0.005)			(0.006)		
Delay WC x Post 2011		-0.008***			-0.009***	
		(0.002)			(0.002)	
Delay MNG x Post 2011			-0.027***			-0.029***
			(0.009)			(0.009)
Observations	729,274	729,274	729,274	288,869	288,869	288,869
P-value WC<BC		<0.001			<0.001	
P-value MNG<BWC			<0.001			<0.001
	Panel C: CTR workers within five years of retirement in 2011					
Delay x Post 2011	-0.013***			-0.009		
	(0.005)			(0.006)		
Delay WC x Post 2011		-0.008***			-0.009***	
		(0.002)			(0.002)	
Delay MNG x Post 2011			-0.025***			-0.026***
			(0.006)			(0.007)
Observations	729,274	729,274	729,274	331,716	331,716	331,716
P-value WC<BC		<0.001			<0.001	
P-value MNG<BWC			<0.001			<0.001
Sample	Full	Full	Full	Restricted	Restricted	Restricted

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR workers on the average monthly contractual wage growth and number of categorical promotions of the firm's non-CTR workers. The definition of CTR workers changes across panels. In the baseline specification, workers are considered CTR if they were within *three* years of retirement in 2011. **Panel A**: Workers are considered CTR if they were within *two* years of retirement in 2011. **Panel B**: Workers are considered CTR if they were within *four* years of retirement in 2011. **Panel C**: Workers are considered CTR if they were within *five* years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A8: Categorical Promotions as Number of Promotions per 10 Employees

	Promotions to white per 10 employees	Promotions to mng per 10 employees	Promotions to white per 10 employees	Promotions to mng per 10 employees
	(1)	(2)	(3)	(4)
Delay BC x Post 2011	0.002*		0.002	
	(0.001)		(0.001)	
Delay WC x Post 2011	-0.002***		-0.002***	
	(0.001)		(0.001)	
Delay BWC x Post 2011		0.001***		0.001*
		(0.000)		(0.000)
Delay MNG x Post 2011		-0.006*		-0.007**
		(0.003)		(0.003)
Sample	Full	Full	Restricted	Restricted
Observations	729,274	729,274	237,272	237,272
R^2	0.17	0.23	0.17	0.26
Mean outcome	0.02	0.01	0.02	0.02
Treatment mean	0.17	0.04	0.53	0.14
Treatment std. dev.	0.68	0.34	1.12	0.59
P-value WC<BC	0.001		0.001	
P-value MNG<BWC		0.033		0.015

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the number of categorical promotions of the firm's non-CTR workers scaled by total employment divided by 10. Workers are considered CTR if they were within three years of retirement in 2011. Delay measures the average retirement delay for white-collar CTR workers (WC), for blue-collar CTR workers (BC), for blue- and white-collar CTR workers (BWC), or for CTR managers (MNG). The dependent variables are the number of categorical promotions from blue to white collar per 10 employees (columns 1 and 3) and the number of categorical promotions from blue/white collar to manager per 10 employees (columns 2 and 4). They are computed on workers who were not within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A9: Treatment Effects by Employment Growth and Span of Control

	Wage growth (1)	Wage growth (2)	Wage growth (3)	Wage growth (4)
Delay x Post 2011	-0.006	-0.022***	-0.013	-0.022**
	(0.007)	(0.007)	(0.009)	(0.009)
Delay x Post 2011 x Fast	0.023**		0.013	
	(0.010)		(0.013)	
Delay x Post 2011 x Slow	-0.045***		-0.024	
	(0.013)		(0.017)	
Delay x Post 2011 x Share top earners		0.020**		0.019
		(0.010)		(0.012)
Sample	Full	Full	Restricted	Restricted
Observations	729,274	724,451	237,272	236,817
R^2	0.27	0.27	0.30	0.29
Mean outcome	0.64	0.62	0.53	0.52
Treatment effect—Fast growing	0.017**		0.002	
	(0.008)		(0.009)	
Treatment effect—Slow growing	-0.051***		-0.035**	
	(0.011)		(0.141)	
P-value Slow<Fast	<0.001		0.016	
Treatment effect—Share top earners		-0.002		-0.003
		(0.007)		(0.009)

Notes: We estimate triple differences in which the treatment variable is further interacted with two sets of variables that measure the ability of firms to add positions to their organizations. In columns 1 and 3, the treatment variable is interacted with two dummy variables measuring employment growth in the years leading to the 2011 reform: "Fast" is 1 for firms in the top tertile of employment growth between 2009 and 2011; "Slow" is 1 for firms in the bottom tertile of employment growth between 2009 and 2011. In columns 2 and 4, the treatment variable is interacted with a measure of span of control. "Share top earners" is a dummy equal to 1 for firms with an above-median share of top earners in their workforce in 2011. Top earners are defined as workers with above-median wage relative to wage distributions computed within a province, two-digit sector (NACE Rev. 2), and category of firm size (above vs. below median workforce). The dependent variable is the average monthly contractual wage growth for workers who were not within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A10: Treatment Effects by Age Group

	Wage growth (1)	Wage growth (2)	Wage growth (3)	Wage growth (4)	Wage growth (5)	Wage growth (6)
Delay x Post 2011	-0.004	-0.017***	-0.056***	-0.024*	-0.022***	-0.057***
	(0.011)	(0.006)	(0.016)	(0.013)	(0.007)	(0.020)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Age group	≤ 35	(35 , 55]	> 55	≤ 35	(35 , 55]	> 55
Observations	401,630	402,722	351,343	199,563	200,009	157,462
R^2	0.29	0.35	0.31	0.29	0.36	0.30
Mean outcome	0.81	0.52	0.58	0.70	0.46	0.46
Treatment mean	0.69	0.69	0.61	1.39	1.39	1.36
Treatment std. dev.	1.14	1.14	1.06	1.28	1.28	1.23

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on the average monthly contractual wage growth of the firm's non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. Delay measures the average retirement delay of CTR workers. The dependent variables are the average monthly contractual wage growth of non-CTR employees with age ≤ 35 (columns 1 and 4), with age between 36 and 55 (columns 2 and 5), and with age > 55 (columns 3 and 6). The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A11: Effects of Increased Retirement Delays on Turnover and Hiring

	Voluntary quits	Layoffs	Layoffs	Hires	Hires	Voluntary quits	Layoffs	Layoffs	Hires	Hires
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Delay x Post 2011	-0.010**	0.059***		-0.100**		-0.001	0.039***		-0.086*	
	(0.005)	(0.006)		(0.044)		(0.006)	(0.008)		(0.045)	
Delay Inside x Post 2011			0.043***		-0.087*			0.035***		-0.114***
			(0.006)		(0.048)			(0.007)		(0.041)
Delay Outside x Post 2011			0.013***		-0.017			0.010***		-0.043
			(0.003)		(0.024)			(0.004)		(0.040)
Sample	Full	Full	Full	Full	Full	Restricted	Restricted	Restricted	Restricted	Restricted
Unit of observation	Firm-year	Firm-year	Firm-job-year	Firm-year	Firm-job-year	Firm-year	Firm-year	Firm-job-year	Firm-year	Firm-job-year
Observations	729,274	729,274	1,133,265	729,274	1,133,265	237,272	237,272	536,760	237,272	536,760
R^2	0.56	0.34	0.31	0.66	0.69	0.68	0.34	0.32	0.75	0.64
Mean outcome	0.92	0.48	0.14	5.28	1.86	1.04	0.45	0.14	5.90	2.19
Treatment mean	0.44	0.44	0.15	0.44	0.15	1.36	1.36	0.32	1.36	0.32
Treatment std. dev.	0.97	0.97	0.59	0.97	0.59	1.28	1.28	0.82	1.28	0.82
P-value Inside>Outside			<0.001					<0.001		
P-value Inside<Outside					0.109					0.091

Notes: This table shows the effect of a one-year increase in average retirement delays among a firm's CTR (close-to-retirement) workers on hiring and turnover of non-CTR workers. Workers are considered CTR if they were within three years of retirement in 2011. The dependent variables measure the number of non-CTR workers voluntarily leaving a firm in each year (columns 1 and 6), the number of layoffs of non-CTR workers (columns 2, 3, 7, and 8), and the number of new hires (columns 4, 5, 9, and 10). The unit of observation is a firm-year pair in columns 1, 2, 4, 6, 7, and 9, and a firm-job category (blue collar, white collar, or managers)-year combination is columns 3, 5, 8, and 10. Delay measures the average retirement delay for all CTR workers, Delay Inside measures the average retirement delay for CTR workers within a job category, Delay Outside measures the average retirement delay for CTR workers outside a job category. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The specifications at the level of firms, job categories, and years also include firm-job category fixed effects. The restricted sample includes only firms with at least one CTR worker in 2011. Firms in the full sample: 104,182. Firms in the restricted sample: 33,896. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A12: Magnitudes

	(1)	(2)	(3)	(4)	(5)	(6)
Average wage	€30,849	€30,849	€34,113	€30,849	€30,849	€34,113
Average one-year wage increase	€2,454	€2,454	€2,451	€2,020	€2,020	€1,931
Annual effect of one σ increased treatment	−€62	−€166	−€258	−€70	−€118	−€301
Effect over 4 years						
A. Not discounted	−€718	−€1,928	−€2,951	−€797	−€1,342	−€3,334
B. Discounted at 3 percent	−€676	−€1,814	−€2,777	−€750	−€1,264	−€3,139
C. Discounted at 5 percent	−€650	−€1,745	−€2,671	−€722	−€1,215	−€3,020
D. Discounted at 10 percent	−€592	−€1,589	−€2,435	−€658	−€1,108	−€2,754
Benchmark annual effect against effect of population (Wheeler, 2006)—Moving from Boston metro area (4.8mil; 10th) to						
Population	4,043,026	2,985,192	2,279,900	3,772,918	3,181,622	1,590,611
Closest location	Detroit	Denver	Las Vegas	Seattle	Tampa	Milwaukee
Aggregation	Metro area	Metro area	Metro area	Metro area	Metro area	Metro area
Population in closest location	4,313,002	2,888,227	2,204,079	3,867,046	3,091,399	1,576,236
Rank	14	19	28	16	18	39
Benchmark against effect of employer size (Barron, Black, and Loewenstein, 2002)—Moving from mean firm (26 workers) to						
Firm size	179	4,697	85,585	377	2,364	4,132,258
Benchmark annual effect against effect of training (Bartel, 2002)						
Days of training	1.57	4.22	6.58	2.17	3.66	9.73
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Age group	All	All	> 55	All	All	> 55
Firms	All	Slow growth	All	All	Slow growth	All

Notes: This table computes the reform-induced wage loss for the average non-CTR worker. The annual effect of a one-σ increase in the treatment computes the wages lost over one year resulting from the reduced contractual wage growth from a one-σ increase in retirement delays among a firm's CTR workers. We also compute this annual effect over the four years of our post-reform period. We provide this computation for different values of the discount rate. This table also benchmarks the annual effect of a one-σ increase in the treatment against other estimates in the literature that look at the effects of other factors on wage growth. Specifically, Wheeler (2006) studies the effect of population size on the annual wage growth of residents. The paper finds that one log point increase in population size increases average yearly wage growth by 14 percent. We use this coefficient to estimate what change in population would lead to an effect of a similar magnitude as the annual effect of our treatment. We take Boston to be the reference metro area. It had 4,836,531 residents in 2017 and was the 10th-most-populous metro area in the United States. We show both the population level that corresponds to the annual effect we measure ("Population") and the metro area with the closest population. Population data is accessible at https://www2.census.gov/programs-surveys/popest/datasets/2010-2017/metro/totals/cbsa-est2017-alldata.csv. We follow a similar procedure to benchmark the magnitude of our effects against the effects of employer size (Barron, Black, and Loewenstein, 2002) and on-the-job training (Bartel, 2002).
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

Table A13: IV, Effects of Blocking More CTR Workers on Coworkers' Careers

	First stage (1)	Wage growth (2)	Promotion to white (3)	Promotion to manager (4)	First stage (5)	Wage growth (6)	Promotion to white (7)	Promotion to manager (8)
				Panel A: Share of CTR workers $\geq$ 1 years				
Delay x Post 2011	0.020***				0.016***			
	(0.002)				(0.001)			
Share of CTR workers x Post 2011		-0.789***				-0.924**		
		(0.266)				(0.418)		
Share of BC CTR workers x Post 2011			0.064				0.048	
			(0.067)				(0.098)	
Share of WC CTR workers x Post 2011			-0.192***				-0.218***	
			(0.061)				(0.071)	
Share of BWC CTR workers x Post 2011				0.061				0.065
				(0.040)				(0.071)
Share of MNG CTR workers x Post 2011				-0.183*				-0.164*
				(0.097)				(0.092)
Treatment std. dev.	0.97	0.03	0.06	0.06	1.28	0.04	0.10	0.11
KP F-Stat		15,643	1,673	3,435		9,103	2,384	1,778
P-value WC<BC			<0.001				0.002	
P-value MNG<BWC				0.010				0.024
				Panel B: Share of CTR workers $\geq$ 2 years				
Delay x Post 2011	0.013***				0.014***			
	(0.001)				(0.001)			
Share of CTR workers x Post 2011		-1.225***				-1.039**		
		(0.413)				(0.470)		
Share of BC CTR workers x Post 2011			0.106				0.091	
			(0.114)				(0.132)	
Share of WC CTR workers x Post 2011			-0.255***				-0.284***	
			(0.081)				(0.086)	
Share of BWC CTR workers x Post 2011				0.102				0.090
				(0.068)				(0.085)
Share of MNG CTR workers x Post 2011				-0.263*				-0.236*
				(0.139)				(0.130)
Treatment std. dev.	0.97	0.02	0.04	0.05	1.28	0.03	0.08	0.08
KP F-Stat		12,450	1,432	1,740		11,311	1,712	1,602
P-value WC<BC			0.002				0.003	
P-value MNG<BWC				0.010				0.023
				Panel C: Share of CTR workers $\geq$ 3 years				
Delay x Post 2011	0.010***				0.012***			
	(0.001)				(0.001)			
Share of CTR workers x Post 2011		-1.511***				-1.218**		
		(0.510)				(0.551)		
Share of BC CTR workers x Post 2011			0.132				0.121	
			(0.139)				(0.156)	
Share of WC CTR workers x Post 2011			-0.315***				-0.347***	
			(0.101)				(0.105)	
Share of BWC CTR workers x Post 2011				0.124				0.107
				(0.082)				(0.100)
Share of MNG CTR workers x Post 2011				-0.357*				-0.318*
				(0.188)				(0.176)
Treatment std. dev.	0.97	0.01	0.04	0.04	1.28	0.02	0.07	0.07
KP F-Stat		9,315	1,045	1,312		9,265	1,340	1,253
P-value WC<BC			0.002				0.002	
P-value MNG<BWC				0.010				0.023
Sample	Full	Full	Full	Full	Restricted	Restricted	Restricted	Restricted
Observations	729,274	729,274	729,274	729,274	237,272	237,272	237,272	237,272
Mean outcome	0.01	0.64	0.05	0.05	0.04	0.53	0.07	0.09

Notes: This table shows the effect of a 100 percent increase in the share of CTR workers on the average monthly wage growth and number of promotions of the firm's non-CTR workers. The treatment variable (Share of CTR workers) measures the number of CTR workers whose retirement age increased by at least one year (panel A), two years (panel B), or three years (panel C) divided by total employment. This variable is instrumented by the average retirement delay of CTR workers. In columns 3, 4, 7, and 8, both the treatment and instrument are created for different subgroups of CTR workers. Columns 1 and 5 show the first stage. The regressions also include firm FEs, year FEs, and year dummies interacted with baseline firm characteristics measured in 2009 (sector fixed effects, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤35, share of workers with age between 36 and 55, share of workers with age >55). "KP F-Stat" is the F statistic from the Kleibergen-Paap test for weak instruments. Standard errors are clustered at the firm level, *** p<0.01, ** p<0.05, * p<0.1. Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, INPS.

Table A14: Effects of CTR Wage Bill on Career Progression of non-CTR Workers

	Wage growth	Promotion to white	Promotion to manager	Wage growth	Promotion to white	Promotion to manager
	(1)	(2)	(3)	(4)	(5)	(6)
Blocked wages TOP x Post 2011	-0.001			-0.002		
	(0.001)			(0.001)		
Blocked wages MID x Post 2011	-0.008**			-0.009**		
	(0.004)			(0.004)		
Blocked wages BOT x Post 2011	0.002			0.002		
	(0.004)			(0.005)		
Blocked wages BC x Post 2011		0.002			0.002	
		(0.001)			(0.002)	
Blocked wages WC x Post 2011		-0.003***			-0.004***	
		(0.001)			(0.001)	
Blocked wages BWC x Post 2011			0.002***			0.003***
			(0.001)			(0.001)
Blocked wages MNG x Post 2011			-0.001			-0.000
			(0.003)			(0.002)
Sample	Full	Full	Full	Restricted	Restricted	Restricted
Observations	662,908	729,274	729,274	226,882	237,272	237,272
R^2	0.34	0.18	0.25	0.33	0.19	0.28
Mean outcome	0.60	0.05	0.05	0.51	0.07	0.09
Treatment std. dev.—TOP	2.33			3.66		
Treatment std. dev.—MID	0.96			1.54		
Treatment std. dev.—BOT	0.65			1.07		
Treatment std. dev.—BC		1.23	1.23		1.88	1.88
Treatment std. dev.—WC		1.49	1.49		2.48	2.48
Treatment std. dev.—BWC		1.85	1.85		2.69	2.69
Treatment std. dev.—MNG		1.61	1.61		2.78	2.78
P-value TOP≠MID	0.076			0.090		
P-value TOP≠BOT	0.531			0.475		
P-value MID≠BOT	0.097			0.090		
P-value WC<BC		0.001			0.001	
P-value MNG<BWC			0.136			0.144

Notes: This table shows the effect of a $10,000 increase in the total wages of the average CTR (close-to-retirement) worker blocked by the reform (yearly wage x retirement delay) on the average monthly contractual wage growth and number of categorical promotions of the firm's non-CTR workers. For each worker, we multiply her retirement delay by her wage (divided by $10,000). Then, Blocked wages is the average "blocked wages" at the firm level for different subgroups of workers. We first distinguishing between CTR workers in the top tertile of the firm's wage distribution (TOP), in the second tertile (MID), and in the bottom tertile (BOT). For categorical promotions, we distinguish between white-collar CTR workers (WC), blue-collar CTR workers (BC), blue- and white-collar CTR workers (BWC), or CTR managers (MNG). The dependent variables are the average monthly contractual wage growth (columns 1 and 4), the number of categorical promotions from blue to white collar (columns 2 and 5), and the number of categorical promotions from blue/white collar to manager (columns 3 and 6). They are computed on workers who were not within three years of retirement in 2011. The regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. The sample is smaller in columns 1 and 4 because we consider only firms with at least one worker in each tertile of the wage distribution. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, INPS.

Table A15: Treatment Effects by Access to Credit

	Wage growth (1)	Wage growth (2)
Delay x Post 2011 for slow-growing high-risk firms	-0.07***	-0.05***
	(0.02)	(0.02)
Delay x Post 2011 for fast-growing high-risk firms	0.02**	0.02
	(0.01)	(0.01)
Delay x Post 2011 for slow-growing low-risk firms	-0.04**	-0.02
	(0.02)	(0.02)
Delay x Post 2011 for fast-growing low-risk firms	0.01	-0.01
	(0.01)	(0.02)
Sample	Full	Restricted
Observations	713,391	230,006
R^2	0.26	0.29
Mean outcome	0.64	0.52
Treatment std. dev.—Slow-growing high-risk firms	0.88	1.21
Treatment std. dev.—Fast-growing high-risk firms	0.95	1.37
Treatment std. dev.—Slow-growing low-risk firms	0.96	1.29
Treatment std. dev.—Fast-growing low-risk firms	0.99	1.35
P-value Slow<Fast for high-risk firms	<0.001	0.002
P-value Slow<Fast for low-risk firms	0.007	0.34

Notes: We estimate quadruple differences in which the treatment variables in equation (5) are further interacted with a variable that measures access to credit. Specifically, we interact the treatment variables in equation (5) with a variable that is equal to 1 if the four-digit sector of a firm has an above-median share of firms at high risk of default. This measure of default risk for each four-digit sector is provided by Cerved, one of the main credit rating agencies in Italy. High-risk firms have a credit rating in the bottom three categories, out of 13 total (https://ratingagency.cerved.com/sites/ratingagency.cerved.dev/files/CRA_MetodologiaRating_0.pdf; page 4). These are firms that present serious or extremely serious problems that jeopardize their ability to meet commitments. Firms in these categories are unlikely to be able to receive loans from banks. In addition to all the necessary triple and double interactions, the regressions include firm fixed effects, year fixed effects, and year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, multiple dummy variables that identify firms above the median in terms of average worker age, share of male workers, firm age, number of employees, average daily wage, share of workers with age ≤ 35, share of workers with age between 36 and 55, share of workers with age > 55). The restricted sample includes only firms with at least one CTR worker in 2011. Standard errors are clustered at the firm level, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Universe of workers employed by firms with 10 to 200 employees in Q1 2009 that were active every year between 2009 and 2015. Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

B Additional Details on the Italian Pension System and the 2011 Pension Reform

B.1 The Italian Pension System

The Italian social security tax rate is 33 percent, a third of which is paid by the employee and the remaining fraction by the employer.

Two methods for computing social security benefits coexist. The first is an earning-based method, whereby entitlements are a function of the average salary that the worker earned in the final stages of the career. The second is a notional contribution-based method: social security contributions are credited into a notional account, earn a return that depends on the performance of the Italian economy, and are then converted into a stream of benefits. The conversion factor is more favorable the longer the workers delay claiming benefits.

The 2011 Fornero pension reform expanded the adoption of the contribution-based method. Before the reform, it was used only for workers who had less than 18 years of contribution in 1995. Moreover, it only applied to their pension contributions from 1996. After the reform, the contribution-based method started being used to compute the benefits also for the more experienced workers with more than 18 years of contribution in 1995. However, it only applied to the pension contributions paid from 2012 onward.

Retirement in the private sector is not mandatory and working past retirement is allowed, although workers rarely choose to do so.

B.2 Additional Details on the 2011 Pension Reform

Why was the pension system changed? The government specifically targeted the pension system because it was one of the main drivers of the increase in the national debt. In 2011, public pension spending amounted to 14 percent of the GDP, twice as much as the OECD average of 7 percent (OECD, 2011). This discrepancy between Italy and other OECD countries was due to a combination of more generous pension benefits and a more rapidly aging population. In 2011, 33 percent of the Italian population was over age 65, compared with only 23.6 percent among other OECD countries. Moreover, it was normal for retired workers to rely exclusively on public pensions. In 2009, only 12.5 percent of the working age population (16-64 years old) invested in private pension funds (OECD, 2011).

Main changes. The right to claim full benefits is based either on age (age-based pension) or on years of contribution (seniority-based pension).

In regard to age-based pensions, the reform gradually increased the age requirement, while maintaining a 20-year contribution prerequisite.[31] The dashed lines in Panel A of Figure A3 plot how the pre-reform age requirement would have increased over the period 2012-2018, absent the reform. Men (women) could have claimed an age-based pension upon turning 65 (60) years old in 2012, and the minimum retirement age would have gradually reached 65.6 (61.8) in 2018. The continuous lines plot the evolution of the age requirement under post-reform rules. The minimum retirement age increased slightly for men, from 66 in

[31] The comparison between pre- and post-reform rules is also displayed in Table A1.

2012 to 66.6 in 2018. The change was far larger for women, as the age requirement jumped to 62 in 2012 and then quickly rose to 66.6 by 2018.

In regard to seniority-based pensions, the reform simultaneously raised and simplified eligibility rules (Figure A3, Panel B). Before the reform, seniority-based pensions could be claimed either upon totaling 40 years of contribution or as soon as the sum of age and years of contribution reached a certain threshold (the so-called *quota* system). Absent the reform, in 2012 the *quota* would have been set at 96, conditional on being at least 60 years old and having 35 years of contribution. The reform abolished the *quota* system, based eligibility exclusively on years of contribution, and raised the contribution requirement differentially for men and women. In 2012, it was set to 42.08 for men and 41.08 for women.[32]

Early retirement. There is only one option to claim pension benefits before meeting the eligibility criteria for either an age-based or a seniority-based pension. It is called *opzione donna* and is available only to women.

Until 2011, it allowed female employees to claim benefits three years before they became eligible for an age-based pension (i.e., at age 57). The take-up, however, was very low, due to the fact that *opzione donna* reduced benefits by a sizable amount. This reduction is driven by two main factors. First, early retirement leads to fewer years of contribution and, all else equal, lower pension wealth. Second, choosing *opzione donna* implies that the contribution-based formula applies to all contribution years. For workers who retire relatively young, the adoption of the contribution-based method on their whole contribution history usually translates into lower entitlements; the average cut is estimated to be 35 percent of the seniority-based pension (INPS, 2016).

After 2011, the number of women choosing early retirement through *opzione donna* increased because the Fornero reform significantly raised eligibility requirements for a public pension for women. However, the take-up of *opzione donna* remained low, reaching at most 20 percent in 2015 within our sample period. Furthermore, job-to-retirement transitions accounted for only 80 percent of the cases, while the rest were cases in which the early retirees were unemployed or not part of the labor force at the time of retirement (INPS, 2016).

In summary, most workers could not or chose not to retire early even after the Fornero reform. As a consequence, they could not undo the effects of the reform.

Grandfather clauses. The reform did not apply to workers who could have claimed an age-based or a seniority-based pension by December 31, 2011 under pre-reform rules. Moreover, a limited group of workers on short-time work or redundancy schemes were grandfathered in. All other workers, the vast majority of the Italian workforce in 2011, were subject to the new retirement rules.

B.3 The Example in Section 2

Here, we provide more details on the example described in Section 2. Consider two male workers born in 1951 and 1952. If these individuals started working at 23 and contributed

[32] Men and women who would have qualified for *quota* 96 under pre-reform rules could exceptionally retire at 64.25 in 2013-2015 and at 64.6 from 2016 onward.

to social security without interruptions, they would each have accumulated 37 years of contribution upon turning 60. Even if these workers were born only one year apart, they would have faced drastically different consequences after 2011. The 1951 worker was grandfathered, while the 1952 worker faced a 4-year and 7-month delay in retirement.

The explanation for these calculations follows. The 1951 cohort could claim a seniority pension in 2011 and was therefore grandfathered. Workers born in 1952 could have claimed a seniority pension at age 60 in 2012 under pre-reform rules. Under post-reform rules, however, they had to be at least 64 years and 7 months old to retire with an age-based pension. Their retirement delay is therefore equal to 4 years and 7 months.

As a second example, consider two women born in August 1951. Due to different interruptions in their careers, they accrue 20 years of contribution—the minimum contribution requirement for an age-based pension—in December 2013 and January 2014, respectively. At this time, pre-reform rules would have allowed them both to claim an age-based pension. Under the new rules, the former worker faces no changes: she satisfies the higher age requirement prevailing in 2013 by turning 62 years and 4 months old in December 2013. The latter worker, however, has her pension eligibility delayed by 1 year and 4 months. She, in fact, can no longer claim an age-based pension in January 2014, but has to wait until May 2015. This delay is due to the fact that the minimum age requirement was further raised in 2014 and 2015 to 63 years and 9 months.

C The Computation of Predicted Retirement Dates

To predict retirement dates under pre- and post-reform rules, we rely on information about gender, age and years of contribution in 2011. For each worker, we start from the contribution history up to 2011. Moreover, we make two assumptions on the behavior of the worker after the reform:

i) There are no gaps in the post-reform contribution history, from January, 2012 to the retirement date;

ii) Employees retire as soon as they can claim either an age-based or a seniority-based pension.

The first assumption requires that individuals continue making monthly contributions to social security until they retire without any gap. This assumption is supported by the data. The median annual contributions for workers aged 60 or above is 52 weeks and the average is 45 weeks. The second assumption requires that most workers do not further delay retirement after becoming eligible for a public pension. Again, this assumption is consistent with the available evidence. In the data, 88 percent of workers retire as soon as they can. When computing the predicted retirement date under pre-reform rules, we take into account the existence of the "waiting window" (the so-called *finestra mobile*): abolished by the Fornero reform, it made it possible to claim the first pension benefit only 12 months after becoming eligible for either type of public pension.

D Proofs of Propositions

Lemma 1. Suppose $N_{2,2} > (1-d_2)N_{2,1}$. Then, in an optimal personnel policy, the following are true:
(i.) $H_2^* = 0$, so no period-2 hires are assigned to job 2;
(ii.) $w_{2,t}^* > w_{1,t}^*$, so job 2 pays more than job 1, and $w_{1,2}^* > w_{1,1}^*$, so wages increase with tenure;
(iii.) $p_{1,2}^*(\theta_H) > 0$, so high-ability workers assigned to job 1 in the first period may be promoted;
(iv.) $p_{1,2}^*(\theta_L) = 0$, so low-ability workers assigned to job 1 in the first period will not be promoted;
(v.) $p_{2,1}^* = 0$, so workers assigned to job 2 in the first period will not be demoted.

Proof of Lemma 1. For part $(i.)$, if the firm hires a new worker and assigns her to job 2, it receives $(1-\lambda)(f_2 + h_2\theta_L) + \lambda(f_2 + h_2\theta_H)$, which we assumed to be negative, so the firm will never assign second-period new hires to job 2. Part $(iv.)$ holds for the same reason.

For parts $(ii.)$, $(iii.)$, and $(v.)$, we will take $\{N_{i,t}\}$ as given. Optimal personnel policies therefore minimize the rents that are paid to new hires. Below, we first establish a lower bound on the total rents the firm pays, and then we construct a personnel policy satisfying $(i.) - (v.)$ attains this lower bound.

To establish this lower bound, notice that the total rents paid to new hires consist of three parts. First, $N_{2,1}R_2$ is the lower bound for the rents paid to new hires into job 2 in period 1. Next, the firm will hire at least $N_{1,2} + N_{2,2} - (1-d_1)(N_{1,1} + N_{2,1})$ new workers into job 1 in the second period, and these workers will get at least R_1. Finally, workers hired into the bottom job in the first period will not be promoted if they are low ability, and they will be promoted with probability $(N_{2,2} - (1-d_2)N_{2,1})/((1-d_1)N_{1,1})$ if they are high ability. If they are paid a wage of 0 in their first period of employment, they will therefore receive a rent of at least $max\{R_1, \tilde{R}_1\}$, where

$$\tilde{R}_1 = -c_1 + \delta(1-d_1)[(1-\lambda)R_1 + ((N_{2,2} - (1-d_2)N_{2,1})/((1-d_1)N_{1,1}))R_2].$$

Taken together, these results establish that a lower bound on the total rents paid to new hires is

$$N_{2,1}R_2 + N_{1,1}max\left\{R_1, \tilde{R}_1\right\} + \delta(N_{1,2} + N_{2,2} - (1-d_1)(N_{1,1} + N_{2,1}))R_1.$$

For the last part of the proof, consider a personnel policy with $p_{1,2}^*(\theta_H) = (N_{2,2} - (1 - d_2)N_{2,1})/((1-d_1)N_{1,1})$ and $p_{2,1}^* = 0$, and let $w_{j,2}^* = c_j + R_j$, $w_{2,1}^* = c_2 + (1 - \delta(1-d_2))R_2$, and

$$w_{1,1}^* = max\left\{c_1 + R_1 - \delta(1-d_1)[(1-\lambda)R_1 + \frac{N_{2,2} - (1-d_2)N_{2,1}}{(1-d_1)N_{1,1}}R_2], 0\right\}.$$

The expression for $w_{1,1}^*$ reflects the idea that if $p_{1,2}^*(\theta_H)$ is sufficiently high, the limited-liability constraint will bind in the first period for period-1 new hires into job 1, and they will be paid $w_{1,1}^* = 0$. And if $p_{1,2}^*(\theta_H)$ is sufficiently low, then they will be paid the wage $w_{1,1}^*$

at which their first period rents are equal to R_1. Such a personnel policy satisfies the firm's flow constraints and each worker's incentive constraints. Moreover, it satisfies $w^*_{2,t} > w^*_{1,t}$ for $t = 1, 2$. This establishes the proposition. The result described in footnote 17 follows from the fact that $w^*_{1,1}$ is decreasing in d_2, while $w^*_{1,2}$, $w^*_{2,1}$, and $w^*_{2,2}$ are independent of d_2.$\blacksquare$

Proposition 1. Suppose $f_1 > c_1 + R_1$. A worker assigned to job 1 in period 1 will receive an expected wage increase of

$$\Delta w^* = w^*_{1,2} - w^*_{1,1} + \lambda p^*_{1,2}(w^*_{2,2} - w^*_{1,2}),$$

where

$$p^*_{1,2} = min\left\{\frac{g + d_2}{(1 - d_1)\lambda s}, 1\right\}.$$

Moreover, the number of new hires in the second period satisfies $H^*_1 = N^*_{1,2} + N^*_{2.2} - (1 - d_1)N^*_{1,1} - (1 - d_2)N^*_{2,1}$ and is increasing in d_1 and d_2.

Proof of Proposition 1. Given the optimal personnel policies described in Lemma 1, we only need to show that high-ability workers are optimally promoted with probability $p^*_{1,2}$. Doing so requires that we establish that the firm optimally operates at capacity in the first period, that is, $N^*_{1,1} = \overline{N}_{1,1}$ and $N^*_{2,1} = \overline{N}_{2,1}$. The result that the firm fills all its job-2 slots in the first period follows directly from the assumption that the firm has $\overline{N}_{2,1}$ high-ability legacy workers, and high-ability workers generate strictly positive profits for the firm when assigned to job 2.

Next, since $f_1 > c_1 + R_1$, job-1 workers in the second period produce strictly positive profits for the firm, so the firm will optimally choose $N^*_{1,2} = \overline{N}_{1,2}$. In the first period, the firm could hire new workers into job 1 and not rehire them in the second period. This would require paying them $w_{1,1} = c_1 + R_1$ in the first period, which by the argument above, the firm is willing to do. The firm could of course do better by retaining these workers and paying them less in the first period, but in any case, it will choose $N^*_{1,1} = \overline{N}_{1,1}$.

The expression for the number of new hires in the second period follows directly from result (*i.*) in Lemma 1. It remains to show the comparative-static result. There are two cases. First, if the firm has limited career capacity, so $N^*_{2,2} = \overline{N}_{2,2}$, we have $H^*_1 = \overline{N}_{1,2} + \overline{N}_{2,2} - (1 - d_1)\overline{N}_{1,1} - (1 - d_2)\overline{N}_{2,1}$, which is increasing in d_1 and d_2. Second, if the firm has abundant career capacity, then $N^*_{2,2} = \overline{N}_{2,1}(1 - d_2) + \overline{N}_{1,1}(1 - d_1)\lambda$. In this case, we have $H^*_1 = \overline{N}_{1,2} - (1 - d_1)(1 - \lambda)\overline{N}_{1,1}$, which is increasing in d_1 and (weakly) in d_2.$\blacksquare$