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# 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, but only in firms with limited promotion opportunities.

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## 1 Introduction

Workers of every generation fear that their careers are being held back by their elders. 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 important not only 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 necessarily an oversight but rather a result of standard economic reasoning, as follows. 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 indeed come at the expense of his or her coworkers' success.

In this paper, we show that career spillovers do 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 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, known as the Fornero reform, created a reasonably close approximation of this ideal. This reform, swiftly

[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.

implemented in December 2011 to contain public expenditures, led to an overall increase in the minimum retirement-eligibility age. Grandfather clauses were limited, and the reform unexpectedly caused retirement delays among senior employees who had been slated to retire soon after December 2011. Moreover, the change in 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 in which to study the effect of retirement delays among senior workers on the careers of younger workers. Our main identification strategy compares changes in wage growth and internal promotions of younger employees across firms that have experienced 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. Specifically, we exploit the variation in treatment that reflects small differences in gender, age, and years of contribution to Social Security only among workers close to retirement before the reform, while controlling for broader firm-level differences in age and gender of the whole workforce. We choose this approach to avoid relying on cross-firm differences in broad demographic composition, because those could affect internal career trajectories through other channels.

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, nonagricultural 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 pension-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 do exist: longer retirement delays among older workers cause larger decreases in wage growth for younger workers. A one-year increase in the average retirement delay among close-to-retirement workers decreases the wage growth of their younger colleagues by 2.5 percent per year relative to their pre-reform wage growth. These effects persist throughout the four years of the treatment period. We also find that career spillovers are larger in firms with larger shares of workers who are close to retirement.

To better understand the underlying mechanism behind these career spillovers, we develop

[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.

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 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 a 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-reform employment growth and look at how the effects of the treatment differ by the growth rate of the firms. For every one-year increase in the average retirement delay among workers who were close to retirement before the reform, the decrease in annual wage growth is (i) 8 percent for younger workers in bottom-tertile firms, which are all shrinking in size, and (ii) approximately zero for younger 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-reform 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 may reflect the firms' use of seniority as one of the criteria for determining 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, workers will not leave for other firms, where they might have to enter at a lower rung on 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 some workers and hiring fewer new workers. Empirically, we find that a one-year increase in retirement delays leads firms to increase layoffs by 10 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-year increase in retirement delays amounts to a monetary loss of up to €718 over the course of four years. However, workers who were 55 years or older but not close to retirement in 2011 experienced monetary losses up to €2,951 over the course of four years. The magnitudes of these latter losses are comparable to 87 percent of the median wage gain associated with a promotion to a white-collar job and 29 percent of the median wage gain associated with a promotion to a managerial position.

We conclude the analysis by evaluating the extent to which our findings are consistent with other career-spillover channels. For example, firms that are already financially constrained when they face retirement delays may simply be unable to afford to promote their workers. Although we find evidence consistent with financial-constraint-driven career spillovers, financial constraints alone cannot easily 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] However, most leading models of internal labor markets in economics ignore 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] In contrast, our findings suggest the importance of incorporating firm-level factors, such as slot constraints, for better understanding workers' career dynamics.[8]

Our paper documents the impact on workers of career spillovers due to blocked promotion opportunities. Other papers in the literature establish a number of other channels through which career spillovers may occur. Hayes, Oyer, and Schaefer (2006) and Jäger and Heining (2019) show, for example, that career spillovers can arise because of team production. While several other papers emphasize the role of limited career opportunities, 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 Panova, 2008).[9] Our paper shows that scarce career opportunities can 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. This contrasts with several recent papers that 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).

[6] For early conceptual work, see Simon (1951) and White (1970). See Stewman and Konda (1983) and Stewman (1986) for earlier surveys, and see Bidwell and Keller (2014) for recent empirical evidence on the importance of available slots for a 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.

[8] 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).

[9] 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 a large influx 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.

We also contribute to the growing literature that shows how workers' careers are shaped by luck. Many studies have already documented 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. Specifically, we show that a worker's career progression can be affected at any time by whether or not senior workers leave their positions and thereby open up advancement opportunities for others.

Finally, we provide new evidence on the consequences of the Fornero reform, arguably the most important Italian reform of the last decade. Two other recent papers (Boeri, Garibaldi, and Moen (2017) and Carta, D'Amuri, and von Watcher (2020)) also study the effects of the Fornero reform, using firm-level variation in the exposure to the reform to study how pension reforms affect youth hiring and firm performance. There is some overlap between these papers and ours, both in terms of the type of data used and the nature of the main treatment variable. The primary way that our paper differs from those papers is one of focus: we are interested mainly in understanding how career spillovers arise in internal labor markets. Our model allows us to make specific predictions about when and where such spillovers arise, which guides our empirical analysis.

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, while Section 6 investigates heterogeneous effects and turnover. Section 7 discusses alternative mechanisms, and Section 8 concludes the paper.

# 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.[10] 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 analy-

[10]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.

sis. First, many workers experienced a substantial increase in their retirement-eligibility age (Table A1). 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 many 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 A2, panel A). In the case of seniority-based criteria, 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 A2, 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 applied only to workers who were eligible to claim a pension under the old rules by December 31, 2011, and to workers in a few other specific categories.[11] The paucity 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.[12] The reform was presented only 20 days after the appointment of a new technocratic government and started being enforced 26 days after its presentation.[13] Stock markets responded sharply on December 6, when the reform was officially presented, suggesting that at least some aspects of the reform had not been anticipated (Figure A3). 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 age 23 and contributed to Social Security without interruption. In spite of being born only one year earlier, the 1951

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

[12] Specifically, 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.

[13] 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.

cohort became eligible for a seniority-based pension in 2011 under the old rules, while the 1952 cohort faced a retirement delay of four years and seven months (Appendix B.3).

To summarize, the reform represents an unexpected and substantial shock to the minimum requirements for starting to receive one's public pension. Moreover, small demographic differences led to large differences in retirement delays for individuals. The reform, therefore, could have very different effects across firms whose workforces have similar demographic characteristics. 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 pension contribution but which 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 reform-induced retirement delays at the individual level.

The first dataset consists of matched employer–employee records for all private-sector, nonagricultural firms with at least one salaried employee. The dataset combines (i) 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 (ii) 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 to the pension system because they were relatively far from retiring at the time of the reform.[14] 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 (for example, maternity leave, sick leave, or disability leave) and bonuses. This greater stability is one reason we choose to use the contractual wage, rather than take-home pay, in measuring

[14] The results are robust to focusing on workers who were eligible in 2011 to retire within the following two, four, or five years (Section 5.1).

career progression. The second reason we do so is that the contractual wage closely reflects changes in job titles. Although job titles are not directly observable in the data, it happens that Italian law (Art. 2103 c.c.) requires that when an employee is assigned a new job title, his or her contractual wage must be modified to reflect the different responsibilities attached to the new position. And thus, a new job title gained by a worker will often be accompanied by a change in the contractual wage. In summary, our measure of monthly contractual wage growth captures the more enduring changes in job titles instead of reflecting transitory shocks to hours worked or one-time 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 pension-contribution histories of individuals who, between 2009 and 2015, worked in private-sector, nonagricultural firms that employed between 10 and 200 employees in the first quarter of 2009.[15] 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 who were close to retirement under pre-reform rules and precisely determining the firm-level shock to the retirement decisions of these 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 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.[16] Furthermore, in order to have a balanced sample, we consider only firms that operated every year between 2009 and 2015 and employed at least one full-time permanent worker in each year.

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

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

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.[17] 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 decreased after 2011, as did the wage growth and number of categorical promotions of younger workers (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, or voluntary or involuntary turnover), shows a similar percentage drop. Together with the decrease in turnover, we observe a decline both in average wage growth and in the number of promotions, whether from blue-collar to white-collar jobs or 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 by itself 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 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 into the Gibbons and Waldman (1999) framework the idea of limited career capacity, 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.

[17] In addition to the constraints just discussed, we limit the sample to firms that have nonmissing values for all measures of career progression. This step reduces the number of firms from 104,924 to 104,182.

## 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. Job 1 corresponds to a blue-collar job, and job 2 corresponds to a white-collar job. Workers' productivity depends on their effort, their innate ability, and the job to which 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 to: 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 $q_1 > q_2$. The blue-collar job 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 $\lambda$, 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 zero, 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 of 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 $t$; if it assigns $N_{j,t} \leq \overline{N}_{j,t}$ workers to job $j$ in period $t$ and they all exert effort, then it receives revenues $N_{j,t}Y_{j,t}$. Throughout, we also assume that in the first period the firm is endowed with $\overline{N}_{2,1}$ high-ability workers, all of whom 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 nonnegative 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 zero 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 $\theta$ 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 (i) a nonnegative wage $w_{j,t} \geq 0$ that the worker will receive if they are not caught shirking, and (ii) 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 zero. 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 $\theta$ 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 $(H_j)_j$ in order 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 incentive to exert effort in each period and to three additional sets of constraints. We explain 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 $j$ have a choice between (i) exerting effort, in which case they receive $w_{j,2} - c_j$, and (ii) shirking, 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 $\lambda$, will be assigned to job $k$ in the next period with probability $p_{j,k}(\theta_H)$, while a low-ability worker will be assigned to job $k$ in the next period with probability $p_{j,k}(\theta_L)$. Hence, a worker of unknown ability 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 (i) the number of new hires into job $j$, $H_j$, and (ii) 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 $j = 1, 2$, we have

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

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

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 the 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 a 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 wage growth 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 that workers in job 1 produce, $f_1$, is greater than $c_1 + R_1$, so workers in job 1 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, and 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 for other workers 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.[18]

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

[18] 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): If the nonnegativity constraint does not bind, a reduction in $d_2$ raises the wage that has to be paid to motivate workers in job 1 in the first period and therefore reduces within-job-1 wage growth.

the following corollary. We assume that $g + d_1 + d_2 < 1$ because that is the empirically relevant case.

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

(*i*) $\Delta w^*$ 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.

Our model is deliberately parsimonious, and additional elements could be incorporated, generating additional implications. First, if workers' abilities are only gradually revealed over time, those 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 had longer tenure in job 1. As a result, retirement delays will have a bigger impact on relatively more senior workers in job 1. Second, the model suggests that workers receive rents at their employer, and these rents are increasing over time. This result implies that even if promotion opportunities become more limited because older workers delay retirement, younger workers are not necessarily more likely to leave the firm voluntarily. Finally, 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 more workers to create more promotion opportunities when older workers delay retirement.

### 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 these 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 effect of retirement delays on wage growth is 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 7.

# 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, to whom we refer 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 and more experienced, have a longer tenure at the firm, and earn a higher wage (Table A3).

To identify CTR workers, we use data on gender, age, and years of pension contribution at the time of the reform, which are contained in the contribution histories provided by INPS. We use this information to compute the retirement-eligibility date under the pre-reform rules for each employee in the sample.[19] 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 $\psi$ represents the worker's group, which depends on their gender, age, and years of contribution as of 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 $D_\psi$ has a mean of 1.36 years and 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 main firm-level treatment, we weight the retirement delay for each worker

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

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 $\text{Delay}_i$ 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 $D_{\psi}$, 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 for those firms, we set $\text{Delay}_i = 0$. 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 with larger retirement delays are older and 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 and younger, and employ a younger workforce (Table A2, columns 2 and 3).

We now address in two ways the potential concern that these imbalances may confound our results. 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 $\text{Delay}_i$ 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 ages 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.

Our main treatment variable captures variation across firms in the average retirement delays among their CTR workers. Isolating this variation has advantages for our identification strategy, but it obscures another important source of variation: when $\text{Delay}_i$ is fixed, firms with a larger CTR share are more exposed to the reform. We use this additional source of variation in Section 5.2, where we estimate an alternative OLS specification to provide a broader picture of the career spillovers created by the reform.

## 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 $t \in \{2009, ..., 2015\}$.[20]

The dependent variable $y_{it}$ measures the average monthly contractual wage growth of non-CTR workers in firm $i$ and year $t$. The treatment $\text{Delay}_i$ is interacted with a time variable: either a post-reform dummy ($\text{Post 2011}_t$) to estimate the average treatment effect in the post-reform period, or a full set of year fixed effects ($\gamma_t$) 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 $\alpha_i$ and $\gamma_t$ represent 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 ($\gamma_t$) 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 $\leq 35$, share of workers with age between 36 and 55, and share of workers with age $> 55$ ($X_{ki}$).[21]

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

[20] As specified in Section 2.3, we use a balanced sample of firms that operated every year between 2009 and 2015. Moreover, using data on all firms in the INPS data, we show that the main treatment variable does not predict firm exit after 2011 (Table A4).

[21] Our results are robust to the use of alternative nonlinear trends and to the inclusion of several controls for the share and characteristics of CTR workers (Section 5.1).

pre-reform year, there was a categorical promotion in one out of twenty firms (Table 1). When 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 $\text{WC}_{it}$ 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 $\text{BC}_i$ is the average retirement delay among CTR blue-collar workers in firm $i$, and Delay $\text{WC}_i$ is the average retirement delay among CTR white-collar workers in firm $i$. Prediction (2) is that retirement delays among white-collar workers will have a larger effect on categorial promotions to white-collar jobs than 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 $\text{MNG}_{it}$ measures the number of blue- and white-collar workers promoted to managerial jobs in firm $i$ and year $t$. The variable Delay $\text{BWC}_i$ is the average retirement delay of CTR blue-collar and white-collar workers in firm $i$, while Delay $\text{MNG}_i$ 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 does not predict any changes in our career progression variables prior to the implementation of the reform (Table A5). This result holds even if we control for fewer confounding factors than those listed under equation (2); this shows that our identification strategy does not hinge upon the inclusion of a specific set of contemporaneous trends.

Specifically, we first regress contractual wage growth and the number of categorical promotions on the interaction between the treatment and a full set of year dummies while controlling for firm and year fixed effects only. The coefficients of the interaction terms are close to zero and not jointly statistically significant at five percent in both the full and the restricted sample (Table A5, panel A). Next, we add an increasing number of firm characteristics measured in 2009 and we interact them with year fixed effects (panels B and C). In these cases, the interactions between the treatment variables and the pre-reform year dummies become even smaller.

To provide further evidence of the lack of pre-reform effects, we estimate the changes in contractual wage growth and categorical promotions had the reform been implemented in either December 2009 or December 2010.[22] 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 A6).

# 5 Empirical Evidence of Career Spillovers

This section presents our main evidence on the existence of career spillovers on the contractual wage growth and categorical promotions of younger workers. Section 5.1 focuses on our primary source of variation: cross-firm differences in the average retirement delay among CTR workers. Section 5.2 incorporates cross-firm variation in the share of CTR workers.

## 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 (0.97 years; hereafter one-$\sigma$) increase in retirement delays (Table 3, column 1), which is 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 3, 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

[22]Specifically, we change the timing of the reform without changing its effects on workers.

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 $\text{Delay}_i$—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 reduce the rate of promotions to those positions. A one-$\sigma$ 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 3, 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 using 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-$\sigma$ increase in retirement delays among managers (0.3 years) decreases the number of non-CTR workers promoted to manager by 0.008 or 16 percent (Table 3, column 3). In contrast, retirement delays among lower-ranked workers 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 3, 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 A7). 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.[23] In addition, we control for interactions between time dummies

[23]We perform two separate tests. In the first, we divide firms into three mutually exclusive groups based

and each of the three sets of characteristics for CTR workers (that is, age, years of pension 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. We can also extend this test to control for much finer cross-firm differences in CTR workers. We divide the CTR workers into forty small groups based on their age, years of contribution in 2011, and gender (four bins for age, five for years of contribution in 2011, and two for gender). Then, we interact the firm-level share of CTR workers in each of these forty groups with year fixed effects. The results are robust to the inclusion of this large number of additional controls.

Finally, we show that the findings are robust to the inclusion of nonlinear trends for each province and two-digit NACE 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 A8). Second, we modify the definition of CTR workers in three ways, identifying them as those workers who were eligible in 2011 to retire in the following two, four, or five years (Table A9). 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 A10). The treatment effects on the share and number of categorical promotions are quantitatively similar.

## 5.2 The Effect of More Workers Facing Retirement Delays

Our analysis so far used cross-firm variation in the average retirement delays among CTR workers. In this section, we allow firms' treatment intensities to scale with the CTR share of their workforce, since a given average retirement delay will lead to a larger reduction in career capacity in firms with a larger CTR share.

Specifically, we can estimate an OLS specification that includes cross-firm variation in both average retirement delays among CTR workers and share of CTR workers. We modify our main treatment variable in Equation (1) by multiplying the variable $\text{Delay}_i$ by the share of CTR workers in firm $i$, as follows:

on their share of CTR workers: no CTR workers (only in the full sample), a below-median share of CTR workers conditional on having at least one CTR worker, and an above-median share of CTR workers conditional on having at least one CTR worker. We interact these dummies with year fixed effects. In the second test, we measure the actual share of CTR workers at baseline and we interact it with year fixed effects.

$$\text{Delay ALL}_i \quad = \text{Delay}_i \times \tfrac{\#\text{CTR workers}_i}{\#\text{ALL workers}_i} \quad = \frac{\sum_\psi \#\text{CTR workers}_{\psi,i} \times D_\psi}{\#\text{ALL workers}_i}. \tag{5}$$

The resulting variable divides the total retirement delays among a firm's CTR workers by the total size of the firm workforce at baseline, instead of by the number of CTR workers. Unlike our main treatment, this new specification takes into account the fact that firms in which the CTR workers are a smaller share of the workforce might be able to better absorb long retirement delays among those workers. Specifically, if firm A and B have the same average retirement delay per CTR worker, but CTR workers are a smaller share of the workforce in firm A, this alternative treatment variable will consider firm A as being less exposed to the pension reform in the short run.

Our results are robust to the adoption of this alternative treatment. The coefficient of Delay $\text{ALL}_i$ is negative and statistically significant at the 1 percent level, indicating that an increase in retirement delays per worker leads to lower wage growth among non-CTR employees. Specifically, a 0.97 year increase in the baseline treatment variable $\text{Delay}_i$ (one standard deviation) decreases the contractual wage growth by 0.016 percentage points (Table 3, column 1), while a 0.07 years per worker increase in this new alternative treatment variable Delay $\text{ALL}_i$ (one standard deviation) decreases the contractual wage growth by 0.017 percentage points (Table 4, column 1).

The results on categorical promotions follow a similar pattern. For example, in the case of promotions to white-collar jobs, a four-percentage point increase (one standard deviation) in Delay ALL $\text{WC}_i$, which divides the retirement delays among white-collar CTR workers by the size of the whole workforce, leads to 0.003 fewer categorical promotions to white-collar positions after 2011 (Table 4, column 2), which is a six-percent reduction in such promotions. In the case of promotions to managerial jobs, a 1.5-percentage point increase (one standard deviation) in Delay ALL $\text{MNG}_i$, which is the ratio between the retirement delays among CTR managers and the size of the workforce, leads to 0.005 fewer categorical promotions to managerial positions after 2011 (Table 4, column 3), which corresponds to a reduction in such promotions by 10 percent.

# 6 Further Evidence of Career Spillovers

## 6.1 Where Do Career Spillovers Arise?

In this section, we first test Prediction (3), which states that career spillovers are larger in slower-growing firms. In fast-growing firms, we would expect that retirement delays would be 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 three 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 three-year pre-reform period, and the minimum decline was 2.9 percent.

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

$$\begin{aligned} y_{it} = & \sum_t \beta_t^s \text{Delay}_i \times \text{time}_t \times \text{Slow}_i + \sum_t \beta_t^f \text{Delay}_i \times \text{time}_t \times \text{Fast}_i \\ & + \sum_t \kappa_t^s \text{time}_t \times \text{Slow}_i + \sum_t \kappa_t^f \text{time}_t \times \text{Fast}_i + \sum_t \kappa_t \text{Delay}_i \times \text{time}_t \\ & + \alpha_i + \gamma_t + \sum_k \sum_t \zeta_{kt} \cdot \gamma_t \cdot X_{ki} + \epsilon_{it}, \end{aligned} \tag{6}$$

where the dummy variable $\text{Slow}_i$ is equal to 1 for firms in the bottom tertile of pre-reform employment growth, while $\text{Fast}_i$ is equal to 1 for firms in the top tertile. The coefficients of interest, $\beta_t^s$ and $\beta_t^f$, 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-$\sigma$ increase (0.93 years) in average retirement delays (Table A11, column 1; and Figure 4, panel A). This triple interaction corresponds to a 6.6 percent larger decrease in wage growth. We can now move from triple interactions back to difference-in-differences estimates (Figure 4, panel B). In slow-growing firms, a one-$\sigma$ 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 3, column 1). In contrast, retirement delays 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), which states that career spillovers are concentrated among firms with larger spans, that is, firms in which a smaller share of jobs consist of high-level jobs. In such firms, there are likely to be fewer available jobs at the top, and retirement delays are thus 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 (that is, above- vs. below-median workforce size). We estimate a triple-difference specification analogous to Equation (6) in which we interact the baseline treatment variable with our indicator for firms with an above-median share of high-level jobs (Table A11, 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 0.021 additional percentage points for each one-$\sigma$ increase (0.96 years) 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.

## 6.2 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 those non-CTR workers who are relatively more senior. 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 non-CTR workers who are older and who have been with the firm longer.

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 Equation (2) (from Section 4.2) separately for each age group and find that the effects are concentrated among workers in the two older age bins, that is, 36 years old and older.[24]

In the full sample, contractual wage growth decreased by 0.02 percentage points after 2011 for each one-$\sigma$ increase (1.14 years) in average retirement delays among non-CTR workers aged 36 to 55 (Table A12, column 2; and Figure A5, panel B) and by 0.06 percentage points for each one-$\sigma$ increase (1.06 years) in average retirement delays among non-CTR workers older than 55 (Table A12, 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

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

effects are not statistically or economically significant for workers who are 35 years old or younger (Table A12, column 1; and Figure A5, panel A).

### 6.3 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 would 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)). Second, career spillovers have larger impacts on non-CTR workers who are relatively older and have been with the firm longer (Prediction (5) and Section 6.2). The combination of these two effects suggests that the non-CTR workers who are most affected by career spillovers have the most to lose from resigning. Retirement delays among CTR workers might therefore not be enough to push workers to leave the company. Ultimately, whether retirement delays lead to an increase in voluntary turnover is an empirical question whose answer depends on the extent to which firms 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 $t$.[25] Consistent with Prediction (6), retirement delays among CTR workers do not increase voluntary turnover for non-CTR workers (Table 5, 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.[26]

In addition to analyzing workers' responses, we study whether retirement delays had an effect on firms' choices in regard to 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-$\sigma$ increase (0.97 years) in retirement delays increased the number of layoffs by 0.049 non-CTR employees and decreased the number of new hires by 0.097 job candidates (Table 5, columns 2 and 5). These estimates mean that involuntary turnover increased by 10 percent per firm and

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

[26] We do not think that the recession Italy was going through at the time can fully explain these findings since 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).

year, while hiring decreased by 1.8 percent. If we include the layoffs of CTR workers in the computation, involuntary turnover increased by 12 percent (Table 5, column 3). Moreover, retirement delays *within* a given category (that is, blue-collar, white-collar, or managers) led to more layoffs and fewer hires within the same category than when compared to other categories (Table 5, columns 4 and 6). On net, a one-$\sigma$ increase (0.97 years) in retirement delays increased the total size of the workforce by 0.8 percent per firm and year (Table 5, column 7).

One might think that retirement delays slow the career progressions of younger workers only in labor markets with strict employment protection laws, such as Italy.[27] 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.[28] 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 relevant to policy making in other settings, given that many other OECD countries have similar, or even stricter, labor laws.[29]

Finally, the policy reform we study was both unanticipated and persistent. Its short-run effects may therefore differ from its long-run effects if most firms cannot quickly change the number of slots they have. Our main results as well as the results on turnover and hiring are consistent with the unanticipated nature of the shock. The persistent nature of the shock implies that its effects on future retirement decisions will be anticipated. For this reason, in the longer run, firms may make more comprehensive changes to their personnel policies over time in order to change the availability of slots, and workers may make different human-capital investments and labor-market choices.

## 6.4 Magnitudes

In this section, we discuss the magnitudes of the treatment effects we find (Table A14, columns 1 to 3). We start by converting the estimated decreases in monthly contractual wage growth to monetary annual losses. A one-$\sigma$ increase (0.97 years) in retirement delays decreases the annual contractual wage growth of non-CTR workers by €62. This estimate

[27] Appendix E provides more details about employment protection in Italy.

[28] Overall, the effects of the reform on firms' financial performance are either zero or slightly negative, depending on the specific measure of financial performance (Table A13), and there are no effects on firm exit (Table A4).

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

corresponds to a 2.5 percent decrease from an annual wage increase of €2,454.[30]

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.[31] 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 €1,589 and €1,928) and for non-CTR workers older than 55 (between €2,435 and €2,951).[32] In the latter group of workers, whose careers were penalized the most in the short run, monetary losses amounted to between 72 percent and 87 percent of the median wage gain associated with categorical promotions to white-collar jobs (€3,386) and to between 24 percent and 29 percent of the median wage gain associated with categorical promotions to managerial positions (€10,293).

## 7 Discussion of Alternative Career-Spillover Channels

In this section, we discuss the extent to which other wage-determination mechanisms can explain our findings. As already noted, our key finding is that the career progression of non-CTR workers is 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, since they treat workers' careers independently,[33] 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-spillover 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 constraints on firing workers, might increase the firm's future payroll costs and force financially constrained firms to postpone planned promotions.[34] 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

[30] 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.

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

[32] Repeating the analysis on the restricted sample leads to quantitatively similar findings (Table A14, columns 4 to 6).

[33] 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).

[34] There is evidence that financial constraints matter for firms' employment decisions along several dimensions. See, for example, Caggese and Cuñat (2013), Hut (2019), and Giupponi and Landais (2020).

depend only on the overall magnitude of the increase in future payroll costs and not directly on where in the organization payroll costs increase. In contrast, for the blocked-promotions channel, where retirement delays occur within the firm does matter. 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 firm's future payroll costs and thus reduce the firm's ability to afford promotions of 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-shock account of our findings.

Second, we examine the payroll-shock 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.[35] 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 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 A15, 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. Moreover, 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 A15, 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.

[35] 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.

Taken together, these patterns indicate that where payroll shocks occur within a firm does matter for the career progression of non-CTR workers. They also therefore conflict with a pure payroll-shock account of our main findings.

We carry out one additional exercise that focuses more on the financial-constraint side of the payroll-shock channel. In particular, we conduct an 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 (6) with an indicator for four-digit NACE 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 of retirement delays tend to be larger in sectors with a higher default risk and lower access to credit. We do, however, also find that even in firms operating in sectors with high access to credit, our earlier finding still holds: retirement delays decrease the contractual wage growth of non-CTR workers if those firms are slow-growing (Table A16, column 1).

The second potential 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 were 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. In fact, 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.

## 8 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 both 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 internalize the extent to which workers' careers become interdependent when their personnel policies promise careers to attract, retain and motivate employees. Hence, they 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 increasingly consequential (OECD, 2017).

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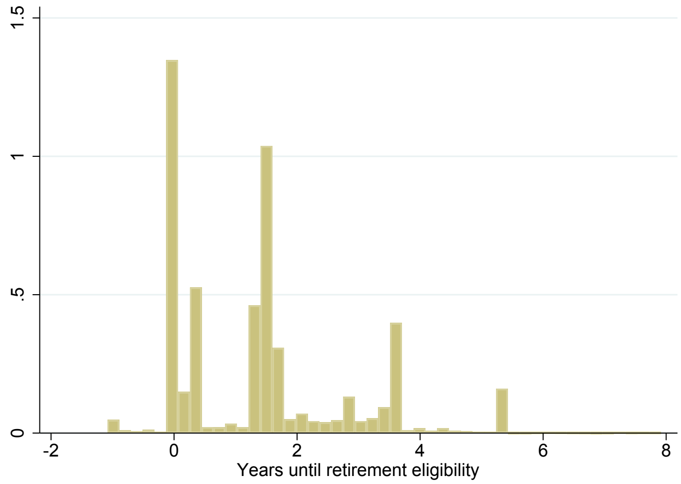
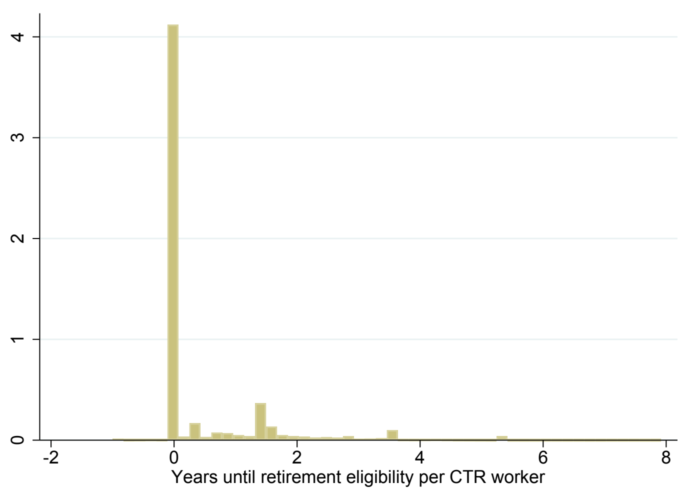
## Figures and Tables

**Figure 1:** Worker- and Firm-Level Treatment



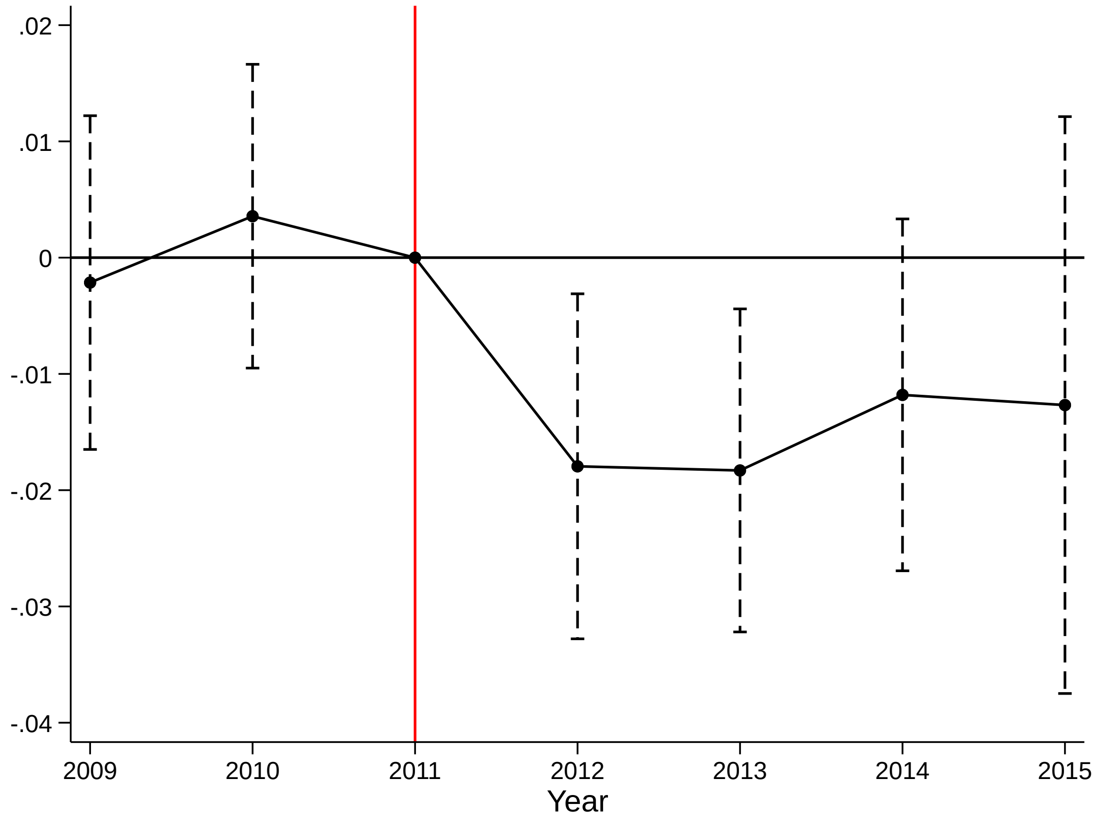

A. Distribution of worker-level retirement delays




B. Distribution of average firm-level retirement delays

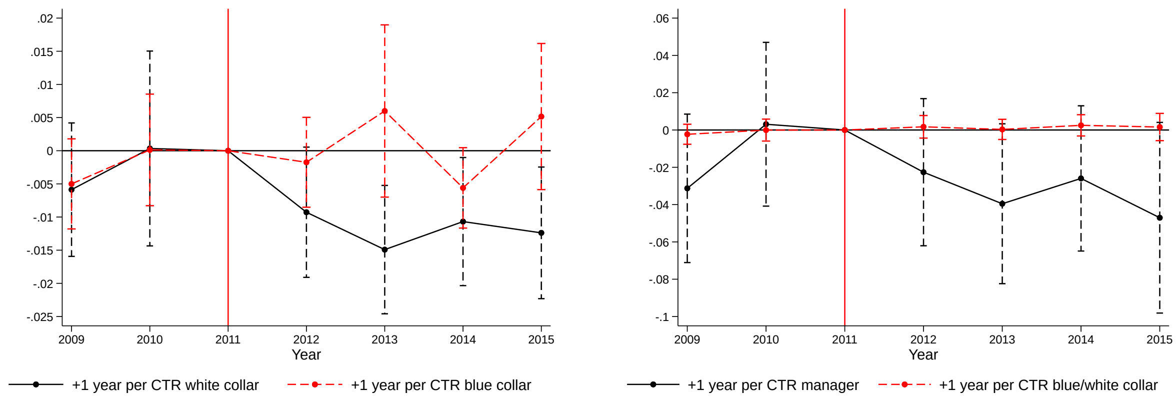
Notes: These graphs show the distribution of retirement delays among CTR (close-to-retirement) workers due to the reform. Workers are considered CTR if they were within three years of retirement in 2011. Panel A shows the distribution of retirement delays at the worker level among CTR workers. The average 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 $\leq 35$, 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



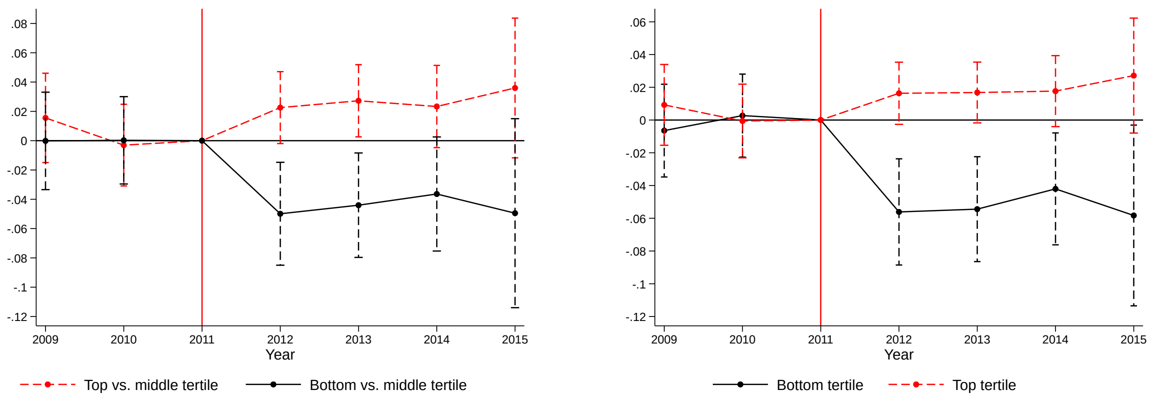

A. Promotion to white collar

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 $\leq 35$, 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 $\leq 35$, 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 | 2012-2015 |
|---|---|---|
| | (1) | (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 $\leq 35$ | -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 $p$ ($\gamma_{ip}$ 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:** 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 (WC) | 0.04 (MNG) | 1.36 | 0.53 (WC) | 0.14 (MNG) |
| Treatment std. dev. | 0.97 | 0.68 (WC) | 0.34 (MNG) | 1.28 | 1.12 (WC) | 0.59 (MNG) |
| 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 $\leq 35$, 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 4:** Alternative Definition of Retirement Delays: Averaged Over Entire Workforce

| | Wage growth (1) | Promotion to white (2) | Promotion to manager (3) | Wage growth (4) | Promotion to white (5) | Promotion to manager (6) |
|---|---|---|---|---|---|---|
| Delay ALL x Post 2011 | -0.248*** | | | -0.260** | | |
| | (0.085) | | | (0.105) | | |
| Delay ALL BC x Post 2011 | | 0.065* | | | 0.076* | |
| | | (0.037) | | | (0.046) | |
| Delay ALL WC x Post 2011 | | -0.077*** | | | -0.091*** | |
| | | (0.029) | | | (0.034) | |
| Delay ALL BWC x Post 2011 | | | 0.042*** | | | 0.051*** |
| | | | (0.011) | | | (0.017) |
| Delay ALL MNG x Post 2011 | | | -0.340* | | | -0.270 |
| | | | (0.189) | | | (0.178) |
| 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 | 0.03 | 0.01 (WC) | 0.002 (MNG) | 0.08 | 0.03 (WC) | 0.005 (MNG) |
| Treatment std. dev. | 0.07 | 0.04 (WC) | 0.015 (MNG) | 0.10 | 0.07 (WC) | 0.026 (MNG) |
| P-value WC<BC | | 0.001 | | | <0.001 | |
| P-value MNG<BWC | | | 0.022 | | | 0.036 |

Notes: This table uses an alternative definition of retirement delays. Specifically, it measures the firm-level exposure to the reform by dividing the retirement delays among a firm's CTR (close-to-retirement) workers by the total size of the workforce, instead of the number of CTR workers. Unlike our main treatment variable, this alternative specification assigns a lower exposure to firms with a larger share of non-CTR workers, keeping fixed the retirement delays among CTR workers. In other words, it takes into account that the effect of long retirement delays among a firm's CTR workers might be better absorbed if CTR workers do not represent a large share of the workforce. Workers are considered CTR if they were within three years of retirement in 2011. As in the main specification, retirement delays are computed 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 $\leq 35$, 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 5:** Effects of Increased Retirement Delays on Turnover and Hiring

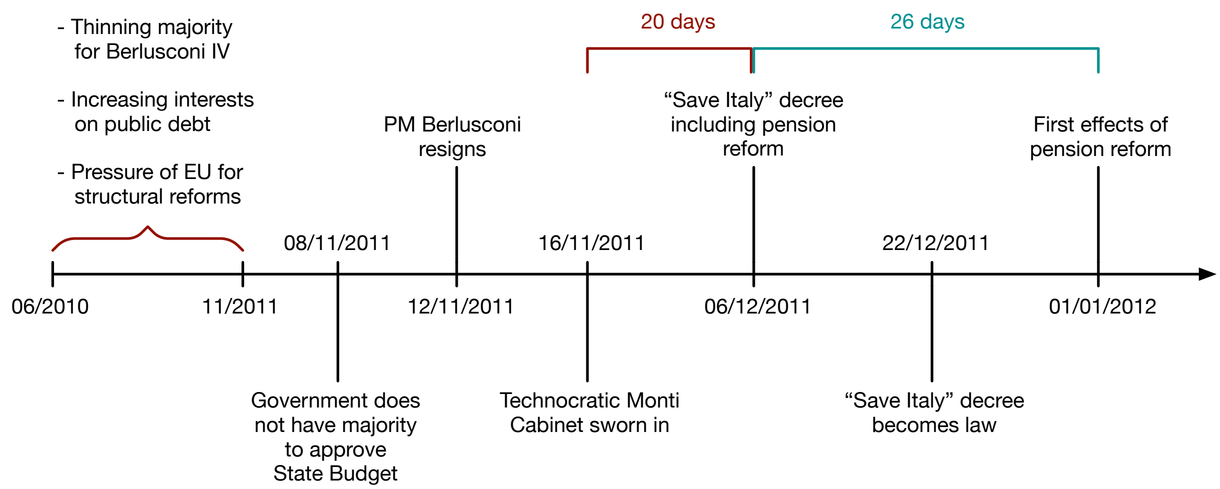
| | Voluntary quits | Layoffs | Layoffs | Layoffs | Hires | Hires | Total employment |
|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| | | | Panel A: Full sample | | | | |
| Delay x Post 2011 | -0.010** | 0.051*** | 0.059*** | | -0.100** | | 0.222** |
| | (0.005) | (0.006) | (0.006) | | (0.044) | | (0.093) |
| Delay Inside x Post 2011 | | | | 0.043*** | | -0.087* | |
| | | | | (0.006) | | (0.048) | |
| Delay Outside x Post 2011 | | | | 0.013*** | | -0.017 | |
| | | | | (0.003) | | (0.024) | |
| Observations | 729,274 | 729,274 | 729,274 | 1,133,265 | 729,274 | 1,133,265 | 729,274 |
| Mean outcome | 0.92 | 0.47 | 0.48 | 0.14 | 5.28 | 1.86 | 26.5 |
| Treatment mean | 0.44 | 0.44 | 0.44 | 0.15 | 0.44 | 0.15 | 0.44 |
| Treatment std. dev. | 0.97 | 0.97 | 0.97 | 0.59 | 0.97 | 0.59 | 0.97 |
| P-value Inside>Outside | | | | <0.001 | | | |
| P-value Inside<Outside | | | | | | 0.109 | |
| | | | Panel B: Restricted sample | | | | |
| Delay x Post 2011 | -0.001 | 0.035*** | 0.039*** | | -0.086* | | 0.120 |
| | (0.006) | (0.008) | (0.008) | | (0.045) | | (0.116) |
| Delay Inside x Post 2011 | | | | 0.035*** | | -0.114*** | |
| | | | | (0.007) | | (0.041) | |
| Delay Outside x Post 2011 | | | | 0.010*** | | -0.043 | |
| | | | | (0.004) | | (0.040) | |
| Observations | 237,272 | 237,272 | 237,272 | 536,760 | 237,272 | 536,760 | 237,272 |
| Mean outcome | 1.04 | 0.43 | 0.45 | 0.14 | 5.90 | 2.19 | 39.51 |
| Treatment mean | 1.36 | 1.36 | 1.36 | 0.32 | 1.36 | 0.32 | 1.36 |
| Treatment std. dev. | 1.28 | 1.28 | 1.28 | 0.82 | 1.28 | 0.82 | 1.28 |
| P-value Inside>Outside | | | | <0.001 | | | |
| P-value Inside<Outside | | | | | | 0.091 | |
| Workers | Non-CTR | Non-CTR | All | All | Non-CTR | Non-CTR | All |
| Unit of observation | Firm-year | Firm-year | Firm-year | Firm-job-year | Firm-year | Firm-job-year | Firm-year |

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. 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 (column 1), the number of layoffs of non-CTR workers (column 2), the number of layoffs of CTR and non-CTR workers (columns 3 and 4), the number of new hires of non-CTR workers (columns 5 and 6), and the total number of employees (column 7). The unit of observation is a firm-year pair in columns 1, 2, 3, 5, and 7, and a firm-job category (blue collar, white collar, or managers)-year combination is columns 4 and 6. 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 $\leq 35$, 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).

# Online Appendix - Not For Publication

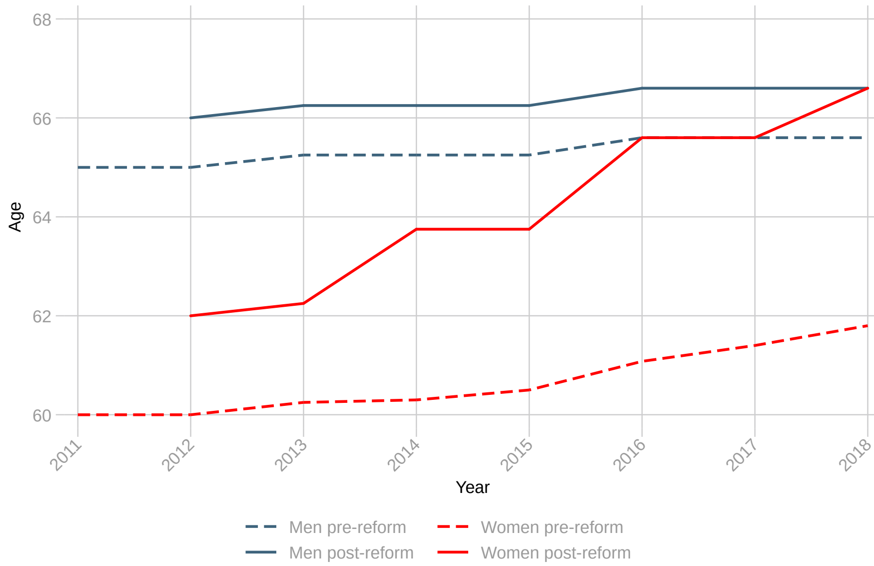
## A Additional Figures and Tables

**Figure A1:** Timeline of Fornero Pension Reform

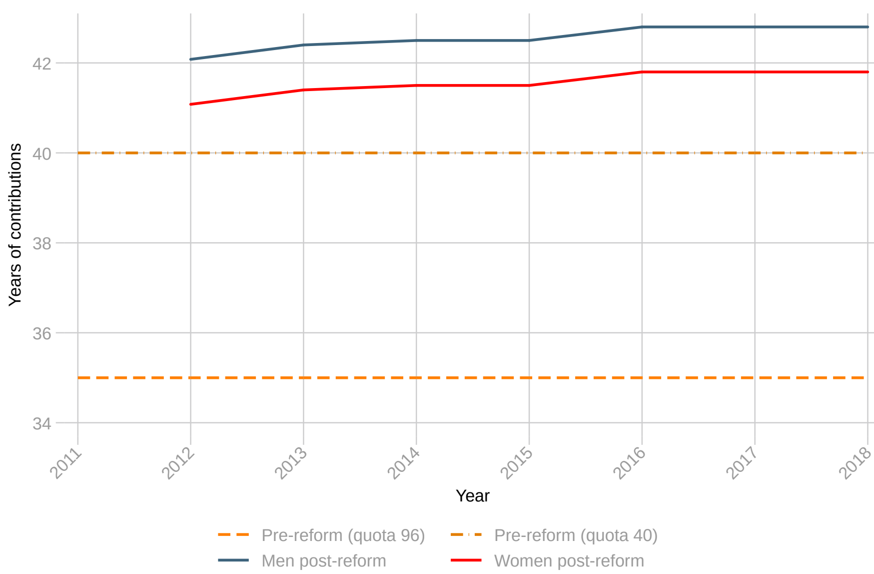



Notes: This figure describes the introduction of the 2011 pension reform in Italy.

**Figure A2:** 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 A3:** Selected Headlines

**12/06/2011 - "Stock markets up, spread down"**


LA CRISI
I MERCATI E L'ITALIA

Borse in alto, spread giù
Monti: l'Italia non fallirà

"Ma comprate i nostri prodotti". Berlusconi: fiducia, o la manovra salta

La giornata

574
Punti base con Silvio
Lo spread con il governo diretto da Silvio Berlusconi era arrivato a un record di 574 punti base il nove novembre

450
miliardi di euro
A tanto ammontano i titoli di Stato che scadono il prossimo anno e dovranno essere rifinanziati sui mercati

18
miliardi di euro
Un punto di rialzo dei tassi di interesse, consolidato per un anno, si tradurrà in 18 miliardi di euro di interessi da pagare sul debito pubblico in più ogni anno

**12/06/2011 - "The 1952 cohort in shock"**


La classe '52 sotto choc
"Lavoreremo fino a cinque anni in più"

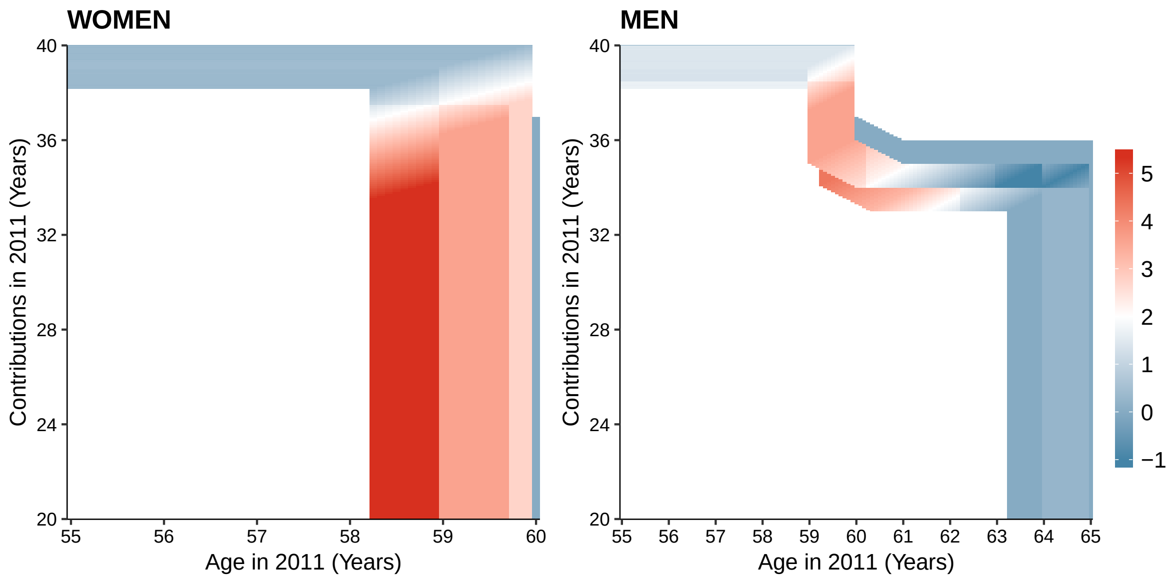
La protesta anche su Internet: "Molti di noi sono in mobilità, come pagheremo?"

Il calendario
Gli uomini andranno in pensione a 66 anni
Le donne a 62 e dal 2018 a 66

I FORTUNATI
Si salvano i nati nel '51 se lavorano dal '75 con 36 anni di contributi

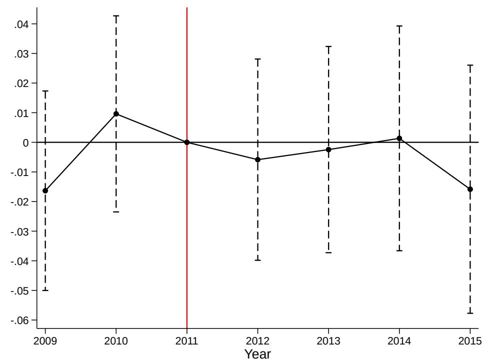
Notes: Headlines of the national newspaper *La Stampa*, http://archivio.lastampa.it/.

**Figure A4:** Effect of the Reform on Different Workers

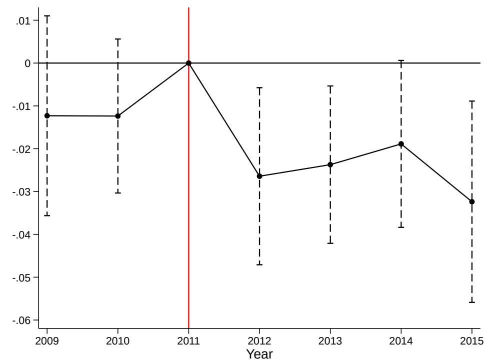



Notes: This heatmap shows how retirement delays differ by gender, age, and years of contribution for CTR workers.
Source: Authors' elaborations based on information from Istituto Nazionale della Previdenza Sociale (INPS).

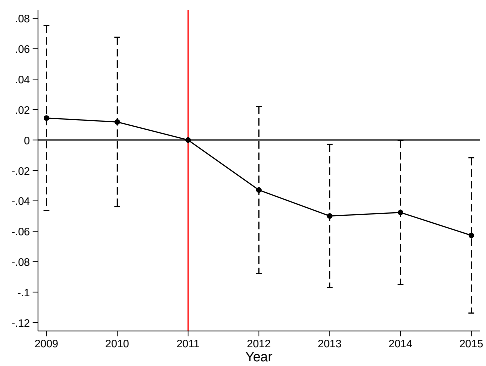
**Figure A5:** Effects of Increased Retirement Delays among CTR Workers on Different Age Groups




A. Age$\leq$35



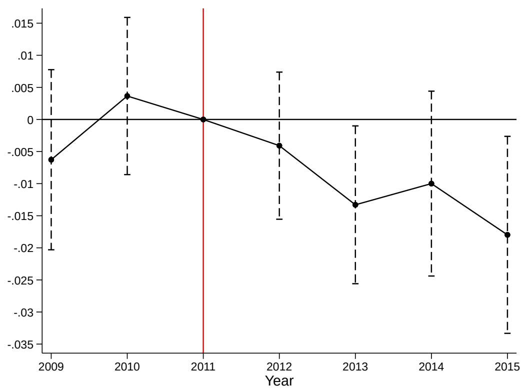

B. 35<Age$\leq$55



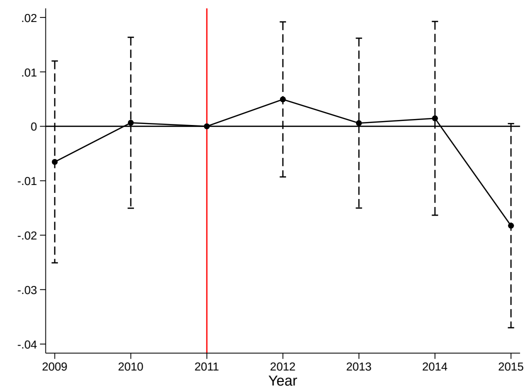

C. Age>55

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 $\leq$ 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 $\leq 35$, 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).

**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 | sd | mean | sd | mean | sd |
| | (1) | (2) | (3) | (4) | (5) | (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:** Correlation of Retirement Delays and Firm Exit

| | Firm exit | Firm exit |
|---|---|---|
| | (1) | (4) |
| Delay x Post 2011 | -0.00009 | -0.00016 |
| | (0.00013) | (0.00017) |
| Sample | Full | Restricted |
| Observations | 732,606 | 239,022 |
| Mean outcome (post 2011) | 0.007 | 0.007 |
| Treatment mean | 0.44 | 1.39 |
| Treatment std. dev. | 0.97 | 1.28 |

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 firm's probability of exiting the market. 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 variable is a dummy equal to one if a firm was not operating in a given year. 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 $\leq 35$, 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 A5:** 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: Controlling for firm and year fixed effects only | | | | | | |
| Delay x 2009 | -0.006 | -0.007 | -0.042* | 0.004 | -0.006 | -0.041* |
| | (0.007) | (0.005) | (0.022) | (0.009) | (0.006) | (0.022) |
| Delay x 2010 | -0.009 | 0.001 | -0.001 | 0.001 | 0.001 | -0.004 |
| | (0.006) | (0.008) | (0.024) | (0.008) | (0.008) | (0.024) |
| P-value for joint significance | 0.369 | 0.355 | 0.091 | 0.911 | 0.492 | 0.113 |
| Panel B: Adding some controls for firm baseline characteristics interacted with year fixed effects | | | | | | |
| Delay x 2009 | 0.001 | -0.006 | -0.032 | 0.003 | -0.004 | -0.023 |
| | (0.007) | (0.005) | (0.020) | (0.009) | (0.005) | (0.020) |
| Delay x 2010 | 0.001 | 0.001 | 0.003 | 0.001 | 0.001 | 0.001 |
| | (0.006) | (0.008) | (0.022) | (0.008) | (0.007) | (0.023) |
| P-value for joint significance | 0.999 | 0.408 | 0.186 | 0.938 | 0.668 | 0.433 |
| Panel C: Adding all controls for firm baseline characteristics interacted with year fixed effects | | | | | | |
| Delay x 2009 | -0.002 | -0.006 | -0.031 | 0.003 | -0.004 | -0.022 |
| | (0.007) | (0.005) | (0.020) | (0.009) | (0.005) | (0.020) |
| Delay x 2010 | 0.003 | 0.001 | 0.003 | 0.001 | 0.001 | 0.001 |
| | (0.007) | (0.007) | (0.022) | (0.008) | (0.007) | (0.023) |
| P-value for joint significance | 0.741 | 0.433 | 0.203 | 0.932 | 0.704 | 0.450 |
| 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 in columns 1 and 4, the average retirement delay among white-collar CTR in columns 2 and 5, and the average retirement delay among CTR managers in columns 3 and 6. 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. In panel A, the regressions include firm and year fixed effects. In panel B, they include the controls in panel A, as well as year dummies interacted with baseline firm characteristics measured in 2009 (sector dummies, dummy variables that identify firms above the median in terms of share of male workers, firm age, and average daily wage). In panel C, they include the controls in panel B, as well as year dummies interacted with more baseline firm characteristics measured in 2009 (dummy variables that identify firms above the median in terms of average worker age, number of employees, share of workers with age $\leq 35$, share of workers with age between 36 and 55, and 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: Database UNIEMENS and complete working histories, Istituto Nazionale della Previdenza Sociale (INPS).

**Table A6:** 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 $\leq 35$, 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:** 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) |
| High vs. low CTR | -0.013** | -0.011*** | -0.025* | -0.014** | -0.012*** | -0.022* |
| | (0.0065) | (0.0036) | (0.013) | (0.0065) | (0.0036) | (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) |
| Detailed CTR groups | -0.018** | -0.012*** | -0.024* | -0.017* | -0.014*** | -0.021* |
| | (0.0080) | (0.0037) | (0.013) | (0.010) | (0.0045) | (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: 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 $\leq 35$, 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 $\leq 35$, 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. **High vs. low CTR**: In addition to all the controls in the baseline, these regressions include a dummy for no CTR workers in 2011 (only in the full sample) and a dummy for a below-median share of CTR workers in 2011 (conditional on having at least one CTR worker), both interacted with year fixed effects. **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. **Detailed CTR groups**: In addition to all the controls in the baseline, these regressions include even more detailed variables describing the CTR workers in each firm. Specifically, we compute the share of CTR workers in forty groups defined using four bins for age ($< 54$, $54-59$, $60-65$, $> 65$), five bins for years of contribution in 2011 ($< 25$, $25-29$, $30-34$, $35-39$, $\geq 40$), and two bins for gender. These variables are also 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 A8:** 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 $\leq 35$, 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:** Alternative Definitions of CTR Workers

| | Wage growth | Promotion to white | Promotion to manager | Wage growth | Promotion to white | Promotion to manager |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (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 $\leq 35$, 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 A10:** 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 $\leq 35$, 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 A11:** 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 $\leq 35$, 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 A12:** 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 | $\leq 35$ | (35 , 55] | $> 55$ | $\leq 35$ | (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 $\leq 35$ (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 $\leq 35$, 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 A13:** Treatment Effects on Firm Performance

| | Revenues over assets | Value added over assets | EBIT over assets | ROE | ROI | ROA |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Delay x Post 2011 | 0.003 | -0.000 | -0.000 | -0.288 | -0.205** | -0.098*** |
| | (0.002) | (0.001) | (0.001) | (0.205) | (0.091) | (0.033) |
| Sample | Full | Full | Full | Full | Full | Full |
| Observations | 542,127 | 542,127 | 542,127 | 521,279 | 391,296 | 542,133 |
| $R^2$ | 0.64 | 0.51 | 0.21 | 0.34 | 0.40 | 0.55 |
| Mean outcome | 1.31 | 0.44 | 0.04 | 0.76 | 4.94 | 4.87 |
| Treatment—mean | 0.47 | 0.47 | 0.47 | 0.48 | 0.51 | 0.47 |
| Treatment—std. dev. | 0.97 | 0.97 | 0.97 | 0.98 | 0.98 | 0.97 |

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 different measures of firm performance. A worker is considered CTR if she was within three years of retirement in 2011. Delay measures the average retirement delay for all CTR workers. The dependent variables are: total revenues over total assets (column 1); value added over total assets (column 2); earnings before interests and taxes (EBIT) over total assets (column 3); return on equity (ROE), computed as net income divided by shareholders' equity (total assets minus debt; column 4); return on interest (ROI), computed as net income divided by the cost of investments (column 5); return on assets (ROA), computed as net income divided by total assets (total assets minus debt; column 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 $\leq 35$, 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 that were active every year between 2009 and 2015. Database UNIEMENS, Istituto Nazionale della Previdenza Sociale (INPS). Cerved data on firm performance: https://www.cerved-online.com/.

**Table A14:** 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 $\sigma$ 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 |
| 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-$\sigma$ increase in the treatment computes the wages lost over one year resulting from the reduced contractual wage growth from a one-$\sigma$ 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.
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 A15:** 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 $\leq 35$, 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 A16:** 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 (6) are further interacted with a variable that measures access to credit. Specifically, we interact the treatment variables in equation (6) 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 $\leq 35$, 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 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.[36] The dashed lines in Panel A of Figure A2 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 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 A2, 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

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

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.[37]

**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 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.

[37] 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.

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 $d_2$.■

**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 $d_2$.■

# E Additional Details on the Italian Labor Market

**A quick look at the size of the Italian labor market.** Italy has the third largest labor market among euro-area countries, totaling almost 22.7 million employees in 2019.[38] It has a record-high number of enterprises (close to 3.7 million in 2018), although most of them tend to be small. Firms with at least ten employees account for 57 percent of total employment, despite making up less than 6 percent of businesses.[39]

**Employment protection in the Italian labor market.** According to OECD Employment Protection Legislation (EPL) indicators, Italy can be classified as a country with high regulatory protection for workers. For example, the indicator that measures the overall strictness of regulations against individual and collective layoffs was equal to 2.86 in 2019, placing Italy among the ten countries with the most stringent regulations. In 2019, the same indicator was equal to 1.31 in the United States (the minimum in the OECD area), 1.90 in the United Kingdom, 2.33 in Germany, 2.68 in France, 2.71 in Belgium, and 2.88 in the Netherlands.[40] In short, employment protection in Italy is high, especially compared to the

[38]Source: Eurostat, European Union Labour Force Survey. 15-64 age bracket.

[39]Source: Eurostat, Structural Business Statistics. The data refer to total business economy and repairs of computer, personal, and household goods with the exception of financial and insurance activities.

[40]Source: OECD Employment Protection Legislation Database. The indicator on individual dismissals is the synthesis of 4 sub-indicators on: procedural requirements; notice and severance pay; regulatory framework

United States and other Anglo-Saxon countries. However, it is only marginally higher than employment protections in other major European counties, such as France.

**Recent trends in employment protection.** Another thing to consider is that the degree of protection in the Italian labor market has significantly declined in the 2013-2019 period after a few labor-market reforms. Specifically, the 2015 Jobs Act modified Article 18 of the Workers' Charter, which regulates layoffs in businesses with more than 15 employees. It narrowed the circumstances in which unfair individual dismissals are sanctioned with the reinstatement of the workers in their pre-layoff positions. Moreover, in addition to reducing the probability of reinstatements, it standardized the amount of monetary compensation to be paid by firms to unfairly dismissed workers. In fact, it made monetary compensations a deterministic function of workers' tenure, reducing the leeway that the courts enjoyed before the Jobs Act.[41] On the contrary, the 2018 labor-market reform increased the total amount of monetary compensations owed to unfairly dismissed workers. It is possible to track these legislative changes with the OECD EPL indicator on dismissals of regular workers. In Italy, this indicator declined from 3.1 in 2013 to 2.76 in 2018 and then it moved back up to 2.86 in 2019.

**Protections for temporary workers.** In 2019, almost 17 percent of salaried employees had temporary contracts. To discourage the overuse of temporary employment, some limitations are in place: for example, the obligation to provide a rationale for offering temporary rather than regular contracts, as well as caps on the contract duration and on the number of renewals. According to the OECD EPL indicators, these restrictions make Italy the OECD county with the third strictest regulations on hiring temporary workers. In 2018, the aforementioned labor-marker reform strengthened or reintroduced some restrictions that had been relaxed by previous reforms. According to the OECD, due to this reform, Italy witnessed the largest increase in the stringency of regulations on temporary contract between 2018 and 2019.

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for unfair dismissals; enforcement of unfair dismissal regulation. For a detailed description of the methodology, see: https://www.oecd.org/employment/emp/oecdindicatorsofemploymentprotection.htm. For details on Italian EPL, see: https://www.oecd.org/els/emp/Italy.pdf.

[41] These changes applied to workers hired on regular contracts after March 7, 2015. For a detailed review of the Jobs Act, see: https://ec.europa.eu/info/sites/info/files/economy-finance/dp072_en.pdf.