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LATE-IN-LIFE INVESTMENTS IN HUMAN CAPITAL. EVIDENCE FROM THE (UNINTENDED) EFFECTS OF A PENSION REFORM

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Late-in-life investments in human capital*

Evidence from the (unintended) effects of a pension reform

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Abstract

This paper provides a novel empirical test of human capital formation by studying whether forced increases in the residual working life, determined by a restrictive pension reform, induce additional training activities. By exploiting a sizable Italian pension reform, in a Difference-in-Differences setting, I find that a lengthening of the working horizon increases, through training, workers' human capital. Additionally, I show that the response to the reform appears very heterogeneous and depends on gender, age, education, marital status, sector of employment and firm size. My estimates suggest, furthermore, that these individual positive effects are not attributable to employers' sponsorship.

JEL codes: J24; J26

Keywords: human capital, pension reform, longer working horizon, middle-aged workers

* *The views expressed in the article are those of the author only and do not involve the responsibility of the Italian Institute of Public Policy Analysis.*

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1 Introduction

Pension reforms that tighten eligibility requirements, usually motivated by public finance motives, end up prolonging workers' stay in the labour market. Recent research suggests that a longer working horizon influences employees' choice by increasing employment (Hairault et al. (2010)), (middle-age) female labour force participation (Carta et al. (2019)), and affecting health behaviours (Bertoni et al. (2018)).

However, an open empirical question is whether an exogenous widening of the working life affects middle-aged employees' training investments because of pension rule variations. Human capital theory, starting from Ben-Porath (1967) and Becker (1962), predicts that the value of human capital investment increases with its payout period. An unanticipated postponement of pension eligibility extends the time left before retirement and, therefore, human capital investment returns (Blinder and Weiss (1976)).

In this paper, I leverage a 2011 Italian pension reform as a source of quasi-experimental variation to study whether a longer working horizon causally (and unintendedly) affects middle-aged employees' human capital investment decisions. Previous evidence has shown that an increase in mandated retirement age for certain workers has sizable, positive and statistically significant effects on human capital. Diversely from these individual-level studies, I focus on a major pension reform that abruptly extended almost all employees' career.

The 2011 pension reform is particularly well suited to provide causal evidence on the effects of working life extension on training participation. First, the 2011 reform has represented for almost all older Italian workers a sudden tightening of the minimum requirements for claiming a public pension with a considerable increase in residual working life (up to 6-7 years). Second, the pension reform was rapidly implemented, with very limited grandfathered clauses, avoiding, crucially for the empirical analysis, any anticipation effects from both employees and employers. Third, soon after its approval, a prolonged and inflamed public debate occurred, implying that the majority of the population understood (or at least were aware of the consequences brought by) the policy change.

I estimate the effect of the increase in minimum retirement age by relying on a Difference-in-Differences approach where my treatment variable is given by a time-invariant measure of policy-induced shock. I construct an individual level measure of exposure to the pension reform exploiting the difference between the Minimum Retirement Age (MRA) in 2017, that is, the post-reform period, and 2011, the pre-reform period. Accordingly, the variation in MRAs provides the size of the reform-induced shock that mirrors the lengthen of the residual career, relative to the previous requirements in place before the 2011 pension reform. Hence, I exploit increases in the distance to retirement, known in the literature also as the *horizon* *effect* (alternatively as *forward looking* or *perspective effect*; Hairault et al. (2010)).

Individual-level data on labour market histories and post-schooling human capital investments are drawn from the 2007-2017 Participation, Labour and Unemployment Survey (PLUS), managed by the Italian Institute of Public Policy Analysis (INAPP). I consider a sample of individuals between 40 and 64 years who have at least 10 but less than 40 years of cumulative contribution years and are not eligible to retire before or after the 2011 pension change. In addition to the effect on formal on-the-job training (Mincer (1962))[1], I estimate the implications of the pension reform on individuals' propensity to pay for their training participation and the role of employers in inducing their middle-aged workers to attend training programs[2].

My main finding is that an increase in the working life causally increases human capital investment: for each year increase in MRA, I show that training - measured as participation in training in the last 12 months - increases by about 0.7 percentage points. The point estimate corresponds to a relative increase of 1.7%, suggesting that an exogenous career extension considerably affects training. I investigate heterogeneity and find that the positive reform effect is driven by men (+2.5%), married women (+1.3%), prime-aged (men and women) and middle-aged (men) workers, and individuals with higher education. Exploring further heterogeneity, I find evidence that this positive response also comes from self-employed individuals (+4 percent), those working in the service sector (0.8 p.p.) and those employed in very small-sized service firms (+2 percentage points).

Finally, I do not detect any change in individuals' propensity to pay for human capital activities in the aftermath of the reform. In contrast, for each additional year increase in the residual working life, affected individuals experienced a decrease of about 0.8 percentage points in the probability that the employer sponsored the training activity.

This paper is related to several strands of the literature. Most importantly, I contribute to the empirical studies related to the human capital theory estimating the effect of pension requirements variations on training programs. However, only a few papers use individual-level data and assume an endogenous process of human capital investment. These papers usually exploit (limited) pension reforms showing that an increase in the working life or a reduction in future pension benefits have a positive effect on human capital accumulation (Bauer and Eichenberger (2017); Fan et al. (2017); Battistin et al. (2012); Montizaan et al. (2010)). However, several other papers reached opposite conclusions finding that training

[1]The focus in this paper is on formal on-the-job training, but the data I exploit does not allow me to discern between general or specific training investment.

[2]For the sake of clarity, the information on whether the firms directly finance training is not required as I focus on the effect of a longer working horizon on training investment and not on the incidence of the human capital investment at the firm level.

incidence decreases with age (Bassanini et al. (2005); De Grip and Van Loo (2002)) or with early retirement options (Fouarge and Schils (2009)).

Instead, a closer strand of the literature, mainly at the firm level, analyzes how investments in human capital benefit overall firm performance (Martins (2020); Dostie (2018); Almeida and Carneiro (2009)). This study is also indirectly related to the literature that analyzes the consequences of increases in retirement age, workforce ageing, and firms' productivity, overall performance and interactions with labour market institutions (see Brunello and Wruuck (2020) for an extensive survey; Bratti et al. (2021); Carta et al. (2019); Acemoglu and Restrepo (2017); Messe and Rouland (2014)), channels not yet well understood[3].

Finally, this paper speaks to the strand of the literature that studies variation in human capital accumulation using mortality rates changes (for an extensive survey see Bloom et al. (2019)) which, however, provides mixed findings (Hansen and Strulik (2017); Oster et al. (2013); Lorentzen et al. (2008); Jayachandran and Lleras-Muney (2009); Acemoglu and Johnson (2007); Kalemli-Ozcan et al. (2000)). Nonetheless, these studies suffer from at least two criticisms. First, as discussed by Cervellati and Sunde (2013) and Hazan (2009) what matters the most for investment in human capital are the survival rates during adult life rather than the change in the life *per-se*. Second, variation in life expectancy is rarely random or unexpected, complicating causal estimation and results interpretation.

The rest of the paper is organized as follows. Section 2 provides details on the Italian pension system and describes the 2011 pension reform that I exploit as source of quasi-experimental variation in the empirical analysis. Section 3 introduces a description of the data and explains the identification and the empirical strategy adopted to recover the (unintended) causal effect of interest. In Section 4, I report the results of the empirical analysis. Section 5 concludes.

2 The Italian pension system and the 2011 reform

The Italian pension system, as well as that of OECD countries, is characterized by a large first pillar, that is, public pension schemes, and by almost marginal second and third pillars, that is compulsory and voluntary private pension funds[4]. Specifically, the main pillar of the Italian pension system is a compulsory *pay-as-you-go*, meaning that all workers are enrolled

[3] A further connection of this paper is with the literature studying how the characteristics of social security systems, specifically legal retirement age (Manoli and Weber (2016); Lalive and Staubli (2015); Staubli and Zweimüller (2013); Mastrobuoni (2009)) and pension benefit rules (Liebman et al. (2009); Krueger and Pischke (1992)), affect agents' behaviours.

[4] For a brief overview of private pension funds reform please refer to Appendix A.

and their contributions are used to pay the pension of current retirees. Furthermore, the system offers two schemes claiming full retirement: the old-age and the seniority pension schemes. They both feature requirements on age and years of contributions. Under the old-age pensions scheme, individuals retire after having achieved a certain minimum age, whereas, under the seniority pensions scheme, individuals retire after having accrued a given number of years of contribution. Pension benefits are computed using a combination of defined-benefits (DB) and notional defined-contributions (NDC) methods. Specifically, under the DB regime, benefits are computed according to the following earning-based formula: $b = \rho N w_r$ where ρ is the accrual rate, N are years of contributions, and w_r is the average salary earned during the last r years of a worker's career. Under the NDC scheme, instead, social security contributions accrue into a notional account which is capitalized using a five-year moving average of the nominal GDP growth rate. They are then transformed into annual benefits through a transformation coefficient that depends on the age at retirement and life expectancy.

Apart from the old-age and seniority schemes, there exists only one early retirement option called *Opzione Donna* introduced in 2004 on an experimental basis (and still in place), that, however, is only available for women. It allows claiming benefits before meeting the old-age or seniority pension requirements. Retiring early, however, comes at the cost of receiving less generous pension benefits with respect to those computed according to the DB or mixed method (a 35% reduction, on average; Istituto Nazionale di Previdenza Sociale (2016)). Indeed, pension entitlements under this option are computed applying the NDC regime to contributions accrued both before and after 1996.

The private and public-sector social security tax rate is 33 percent: one-third is paid by the employee and two-thirds by the employer. For self-employed that pay contributions to the Social Security Institute the social security tax rate ranges between 24 and 34 percent. Retirement is not mandatory, and working past retirement is allowed although the employer when retirement age is reached ask the worker to retire.

The Italian pension system was dramatically revised through a long reform process to improve its financial sustainability during the last three decades. Indeed, the progressive increase in Italian population ageing has meant that pensions have to be paid for a more extended period implying that the flow of Social Security Institute's income (represented by contributions) was not in balance with the expenses (the pensions paid). In addition, the slowdown in economic growth has further decelerated contribution income. Therefore, a series of reforms have been implemented to bring pension expenditure under control. In 1995, the *Dini* reform[5] introduced in the Italian pension system the notional defined-contribution

[5]Please see Appendix B for earlier intervention to curb pension expenditures.

(NCD) method, a pension benefits computation that links the lifetime paid contributions to total future pension benefits. However, the transition from a defined-benefit (DB) to a notional defined-contribution (NDC) basis was gradual, involving only those who had less than 18 years of paid contribution before January 1, 1996. Several legislative interventions from 1996 onward, motivated by public finance reasons, increased the requirements for claiming a pension.

At the end of December 2011, the new technocratic government approved an emergency package of measures, the *Salva Italia* decree, in response to the pressure of the financial markets on the Italian sovereign debt that reached unprecedented levels. Among the emergency measures approved, a substantial and restrictive pension reform was introduced[6]. The reform, known as the *Fornero* reform (Law December 22, 2011, no. 201)[7], entered into force on January 1, 2012 (ten days after its approval) and raised age and contribution requirements to claim old-age and seniority pensions[8]. The new rules applied to all workers who did not accrue the right to claim either pension by the end of 2011[9].

The technocratic government specifically targeted the pension system because it was one of the main drivers of the national debt increase. 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 ageing 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 is customary 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)).

The reform raised the age requirement for old-age pensions, leaving the contribution requirement (20 years) unchanged. The statutory retirement age was 60 (61) for women (women employed in the public sector) and 65 for men (irrespective of their sector of employment) in 2011. Absent the reform, it would have risen to reach 61 years and 10 months for women and 65 years and 7 months for men and women employed in the public sector in 2018. Per effect of the reform, the old-age statutory retirement age has gradually increased to reach 66 years and 7 months for both genders in 2018 (in compliance with the ruling of the

[6]Although the pension reform was the central component of the decree, other measures were legislated to increase taxation on real estate, cars, and consumption. The whole text of the law can be accessed at Decreto Salva Italia, Gazzetta Ufficiale.

[7]Fornero (2015) provides an exhaustive account of the Italian situation during Autumn 2011 and a detailed technical description of the 2011 pension reform.

[8]See Appendix C for the evolution of the effective retirement age after the 2011 pension reform.

[9]For a brief discussion of other grandfathering clauses and other provisions of the 2011 pension reform, please see Appendices D and E.

European Court of Justice on the equalization of retirement age between men and women)[10] (see Table 1). The change in the age requirement was thus considerably more significant for women than for men.

In addition, the reform modified the rules for claiming seniority pensions. A "Quota system" was in place until 2011. Workers could retire as soon as their age and years of contributions summed to a certain "Quota", conditional on both surpassing a certain threshold. In 2011 the quota was set to 96, conditional on being at least 60 years old and having at least 35 years of contributions. Alternatively, workers could retire upon totalling 40 years of contributions, regardless of their age. The 2011 reform abolished the "Quota system", and it legislated that a seniority pension can be claimed upon totalling at least 41 years of contribution for women and 42 for men (irrespective of their age; see Table 2). Thus, workers planning to retire under the "Quota system" faced a large increase in years until pension eligibility, up to 6-7 years.

However, the reform did not change the early retirement rules. The take-up of early retirement was very low before the reform because of the cut in benefits. After the reform, which heavily raised requirements for women, the take-up of *Opzione Donna* increased. As a result, its take-up remains limited, involving only less than 65,000 women over the period 2008-2016 (representing around 20% of women who could have exercised the early retirement option; Istituto Nazionale di Previdenza Sociale (2016)).

3 Data and empirical strategy

Data. In this analysis, I draw information on on-the-job training participation and labour market histories from the Participation, Labor and Unemployment Survey (PLUS), a biannual survey administered by the Italian Institute of Public Policy Analysis (INAPP) to a sample of Italian individuals, about 55,000 respondents per wave.

The survey provides a specific section on training activities attended by respondents, apart from schooling education. Specifically, individuals are asked if they attended training activities in the last 12 months, such as seminars, conferences, training courses, or professional refresher courses; if they directly paid for attending them and whether their employers (usually firms) sponsored the activity (that, however, do not necessarily imply that they paid on behalf of the worker)[11]. Hence, the availability of these data allows me to investigate the

[10] The reform allowed all individuals to retire at 70, as long as they have accrued at least 5 years of paid contribution.

[11] Unfortunately, data on type of course chosen and the number of hours spent per activity are not always available. However, evidence suggests that it is more the incidence of a training spell than its duration that is relevant (Pischke (2001)).

causal effect of an increase in the working life on human capital accumulation. Furthermore, the richness of these data allows me also to investigate the propensity of individuals in investing directly in (*i.e.* paying for) training participation and to gauge some evidence on the role of firms in providing training to their workers.

The empirical analysis builds on the most recent waves of the survey, that is, from 2007 up to 2017, that include the years around the 2011 pension reform. Even though the PLUS data has a longitudinal structure, the panel component is very short (about less than 3,000 individuals), forcing me to conduct the empirical analysis using repeated cross-sections.

The working sample comprises individuals between 40 and 64 years, with at least 10 and less than 40 years of paid contributions, eligible to retire neither before nor after the 2011 pension reform[12].

Crucially for the empirical analysis, the PLUS data allows me to construct pension eligibility criteria because it includes information on age, gender, sector and type of employment and, importantly, on accrued years of contribution; this allows me to build for each individual the Minimum Retirement Age (MRA) based on the eligibility rules in place each year.

Moreover, it collects information on workers' expected retirement age and pensioners' effective retirement age. In particular, for retired individuals, I also have data on their employment sector and years of accrued contributions, allowing me to check the soundness of the shock variable's approach and identifying assumptions.

Identification of the shock. The reform generated different changes in years until retirement eligibility among otherwise similar older workers, given that slight demographic differences led to significant differences in retirement delays for individuals. The different mandated retirement age by gender, age, sector and, mainly, by previously accrued years of contributions implies that individuals have been differently affected by the reform in terms of how much the length of the residual working period before retirement did increase.

In order to estimate the increasing shift in the residual working life, I predict the minimum retirement dates under pre- and post-reform rules by drawing on information about individuals' gender, age, sector and years of contribution. I use as a starting point the contribution declared by the worker in each wave of the survey, and I make two assumptions on their working histories: *i*) workers accrue full contributions (52 weeks per year) until retirement; *ii*) the predicted retirement date is the earliest date the worker can collect the first pension instalment by claiming either an old-age or a seniority pension.

The first assumption requires that individuals work year-long spells and full-time. The second one, instead, requests that most workers do not further delay retirement after becom-

[12] For other information on the sample selection, please refer to Appendix F.

ing eligible for a public pension. While the former assumption may appear more problematic to check and can imply an underestimation of the expected shock to the MRA[13], the latter assumption can be more easily checked by looking at the behaviour of individuals who retired in the past. In particular, to show that a significant share of individuals retire when they reach their minimum retirement age (MRA), I use the sample of individuals who declare themselves as retired in the PLUS data. By exploiting information on their effective retirement age (ERA), years of contribution and sector of employment for all pensioners between 2005 and 2015, I compute the minimum retirement age for each individual retired in year t, with $t \in [2005, 2015]$, that I compare with their effective retirement age[14]. In this way, I define the distance to retirement, the difference between the MRA and ERA. If the distance to retirement is zero, individuals indeed retire when reaching their minimum eligibility requirement. In Figure 3 I plot the percentage of individuals retired, considering only the sample of pensioners, as a function of distance to retirement. The figure clearly shows that when the distance equals zero, MRA equals ERA, more than one out of two individuals retire. If, instead, I take into account the interval ± 1, given that I am exploiting survey and not administrative data and there may be small errors in reporting ERA and years of paid contributions, this percentage increases up to 70 percent. Overall, it seems that the second assumption provides sound evidence in support for the identification of the shock.

Hence, to compute the individual level shock in the increase of the expected residual working life, which can also be interpreted as the degree of exposure to the pension reform, I construct a time-invariant measure of exposure to the policy change by taking the difference between the expected MRA under the post-reform (at 2017) and under the pre-reform rules (at 2011), that is $shock_q = MRA_{2017} - MRA_{2011}$[15]. This measure of cross-sectional variation in the exposure to the pension reform is based on the full interaction of all the characteristics necessary to determine the MRA, that is, age, gender, years of contribution and sector of employment (whether it is private, public or if the individual is self-employed).

[13] Bianchi et al. (2019) exploiting contribution histories from the Social Security Institute show that for several types of workers (in 2012), the median annual contribution is 52 weeks, and the average is 45 weeks.

[14] I also take into account that the reform abolished the "waiting window", a rule whereby the first pension instalment could be collected only 12 months after becoming eligible for either type of pension. However, I do not consider the sample of retired individuals in the 2017 wave, given that for these individuals, information on accrued years of contribution is not available.

[15] Other papers study the effects of the 2011 pension reform using as identification of the policy-induced shock similar versions to that one I am exploiting in this paper. Bovini and Paradisi (2019) examines how firms adjusted their hiring and firing decisions in response to the reform, Bianchi et al. (2019) the effects on internal labour markets. Carta and De Phillipis (2019) the effect of the pension reform on the labour force participation of middle-aged individuals and their partners. Carta et al. (2019) study the increase in retirement ages, due to the 2011 policy change, on firms' economic outcomes. Boeri et al. (2017) studies how the reform affected youth unemployment. This paper contributes to their findings by using the 2011 pension reform as a tool to study the human capital investment of middle-aged individuals.

In Figure 1, I plot the percentage of individuals according to the values of the reform-induced $shock_q$, ranging between 2 and 7 years of expected increase in the residual working life (with an average value of 4 years and 7 months). According to the figure, individuals whose expected residual working life increased more than 3 years are about slightly less than 64% in the sample. Figure 2, instead, plots the reform-induced shock distribution in the length of the residual working life by gender. Regarding men, about 55 percent experiences an increase in the residual working life greater than 3 years, and this is coherent with the fact that Italian working men have more stable career trajectories and start working earlier than women. On the other hand, about 75 percent of women in the sample experienced increases in their expected residual working horizon greater than 3 years.

Empirical strategy. The 2011 pension reform has at least two characteristics that are important for the empirical analysis. First, many workers experienced a substantial increase in their retirement-eligibility age, meaning that the reform represents an unexpected and substantial shock to the minimum requirements for pension eligibility. Second, as highlighted in Section 2, the decision and implementation lags of the reform were both very short, implying that anticipatory effects were likely negligible. Hence, the changes introduced by the reform provide a clean empirical setting to study how changes in the expected residual working life would affect workers' human capital investment.

The identification of the shock described above aims at evaluating the magnitude of the perspective effect (or the *forward looking effect*). Therefore, it studies the human capital investment of individuals who would not have been eligible to retire even under the pre-reform rules but whose MRA increased due to the 2011 pension reform. Hence, using the variation in distance to retirement exclusively induced by the pension reform and given by the cross-sectional time-invariant measure of exposure to the policy, I estimate the following empirical model:

$$y_{iqt} = \beta shock_q \times post2011 + \delta_q + \alpha_t + \mathbf{X}_{it} + \eta_{iqt} \tag{1}$$

where: y_{iqt} is an outcome of interest at the individual level i in year t at the shock level q. My main outcome of interest is a dummy variable that indicates whether individual i attended on-the-job training in the last 12 months in year t at the shock level q, then I also investigate the propensity of individual i in paying for training participation and whether firms sponsored training activities; $shock_q$ is the change in the residual working life induced by the reform (as described above), that is a time invariant measure of exposure to the policy; $post2011$ is a dummy that indicates the post-reform period, that is years 2013, 2015 and 2017; α_t are year fixed effects, absorbing long term or cyclical developments that affect all individuals in the same way; δ_q are fixed effects at the shock level absorbing all pre-reform permanent differences in distance to MRA; $\mathbf{X}_{it}$ is a vector of fixed effects at the individual level (marital status, region of residence, sector of employment, gender, age, years of contribution) absorbing cross sectional time-invariant heterogeneity among individuals. Finally, η_{iqt} is the error term. Standard errors are clustered at the age-sector of employment-gender-years of contribution level.

As usual in any Difference-in-Differences model, the coefficient of interest is β, that is, the interaction between the treatment variable and the post-reform variable, which estimates the average human capital investment effect among individuals that experienced a larger or a minor increase in MRA, exclusively depending on their degree of exposure to the policy, around its implementation.

Descriptive statistics. Before discussing the Difference-in-Differences estimates, I briefly provide descriptive statistics, where I arranged individuals in two groups only for graphical and descriptive evidence purposes. Individuals more exposed to the change in the minimum retirement age (most treated; *i.e.* $shock_q > 3$) expect to stay active in the labour market two more years (Figure 4a) and a lower pension income relative to job earnings (Figure 4b) with respect to the least affected group. Overall, trends for both groups followed more or less the same patterns. Figures 5 and 6 show the average trends in training participation by exposure to the shock and also by gender. Most shocked individuals display, on average, higher participation rate in training in the aftermath of the pension reform, and this is driven by most affected men (see Figure 6).

Most affected individuals, in addition, pay more often for participating in training activities, even though I cannot detect divergent patterns after the pension reform (see Figure 7a). Instead, for what concern on-the-job training sponsored by firms, most shocked individuals, in the aftermath of the reform, appear less likely to be involved in training sponsored by their employer as opposed to least shocked ones (Figure 7b).

Table 3 presents some descriptive statistics of the working sample. The first 3 columns regard the 2007-2017 survey waves, whereas the last 3 refer to the pre-reform waves. Furthermore, I present statistics for the entire sample and distinguishing between those most treated and least shocked. Overall, no remarkable differences there exist between least and most treated groups, either in the entire sample or in the pre-reform waves, with the only exceptions regarding gender composition of the groups (men are over-represented in the least treated group), the shares of private-sector employee (considerably higher for least treated individuals) and self-employed individuals (greater for most exposed to changes in MRA).

4 Results

Longer working horizon and training participation. As explained in Section 1, human capital theory predicts that the value of human capital investment increases with its payout period. Therefore, the 2011 pension reform represents an unanticipated and exogenous shock that induces a sizable increase in the working life (*i.e.*, an increase of the period to recoup human capital investment benefits), affecting a large share of the middle-aged working population.

To measure causally the effect of an increase in the working horizon on the investment into human capital, I rely on the Difference-in-Differences approach described in Section 3. In Table 4 I present the estimation results relative to the entire sample (Column (1)) and the gender sample-split analysis (Columns (2) and (3)). All specifications include the vector of controls. The results of these three specifications, although with differences in estimates' magnitude and statistical significance, suggest that an exogenous and unanticipated extension in the working horizon causally and positively affects human capital accumulation. Column (1) shows that for each additional year increase in MRA, participation in training increase by 0.7 percentage points, or in relative terms of about 1.7 percent given the pre-reform average of the dependent variable. However, this considerable positive effect is only driven by men (Column (2)), whose training participation increases of 0.9 percent (+2.5%) in the aftermath of the 2011 reform. These results are broadly in line with Montizaan et al. (2010), who find that public sector workers affected by a pension reform, lowering their pension rights, increased training participation of about 2.7-3.2 percentage points.

Furthermore, I extend the main analysis with a series of heterogeneity checks. Table 5 reports results for individuals aged 40-47, 48-56 and 57-64 (Columns (1), (2) and (3), respectively). The upper panel of the Table refers to the entire sample, whereas the last two to men and women, respectively. Fortunately, all the estimates relative to the oldest cohort (Column (3)) are never statistically significant. Indeed, an overall effect driven only by oldest individuals would have contradicted the theoretical result about the length of the payout period. The first panel of Table 5 reports a positive and statistically significant effect for age classes 40-47 and 48-56, that is prime-aged and middle-aged workers. The point estimates show a training participation increase of 1.3% and 0.7%, respectively, that translate into a relative increase of 3.6 and 1.9 percent. Again, the gender-split exercise reveals that the whole variation is driven by prime-aged (3.9 percent average increase) and middle-aged (3 percent relative increase) men.

Table 6 reports the results when I split the sample according to the sectors in which the individual works, that is, public, private or whether the individual is self-employed (Columns

(1), (2) and (3), respectively). The only statistically significant effect comes from individuals working as self-employed for whom an increase of a 1-year in MRA implies an increase in training participation of about 1.5 percentage points, or, in other words, to about a 4 percent relative increase when compared to the sample mean. The last heterogeneity considers only private-sector and self-employed workers. Specifically, I define two broad firm sectors based on the economic sectors' statistical code of the firm where they are employed: the manufacturing and service sectors. The results, available in Table 7, show that, despite a positive coefficient for both groups of workers, only workers whose firms belong to the service sector increased (Column (2)), at the conventional statistical level, their probability of training of about 0.8 percent (or 2.1 percentage points).

Training participation and schooling education. Human capital theory suggests that the education level is likely to affect the worker's training probability (Griliches (1997)). Theory advocates that schooling education and human capital investments are complements. Henceforth, theory suggests a positive correlation between education and training participation. In this analysis, I test if the reform effect also varies by schooling educational levels (low: middle schools or lower, medium: high school, and high: bachelor or higher). Table 8 reports the results for this heterogeneity check. Columns (1), (2) and (3) refer to low, medium and high education, respectively, whereas the first panel to the entire sample and the last two panels to men and women separately, respectively. Overall, I find that individuals with higher education have a higher probability of investing in human capital. For them, a 1-year increase in the working life due to the pension reform implies an increase in the probability of human capital accumulation of 1.4 p.p. (entire sample), 1.7 percent for men and 1.5 percent for women (although only statistically significant the 10 percent level). These estimates translates into a relative increase of 2.3% (both entire sample and men) and of 2.5 percent for high educated women.

Further women heterogeneity? So far, my evidence shows that women, differently from men, do not increase training participation in the aftermath of the 2011 pension reform. However, this finding is in striking contrast with the direct effect of the reform: women experience the largest increase in minimum retirement age. The gender and family economics literature (see, among many others, Goodpaster (2010); Leigh (2010); Munasinghe et al. (2008)), indeed, suggests that married women experience higher opportunity costs in terms of work and investments due to the household chores burden. Therefore, they may be less willing or time-constrained in investing in additional human capital. Nevertheless, an increase in pension eligibility requirements may provide married women higher incentives to invest in human capital as opposed to more “career focused” women.

Therefore, I focus on the sample of women distinguishing them according to their marital status: married and not married[16]. I replicate the previous sample-split exercises and baseline specification. The results, available in Tables 9, 10 and 11, suggest marital status is an important determinant of training participation after the pension rule changes. Overall, I find that married women increase training participation of about 1.3% (see Columns (1) and (2) of Table 9), corresponding to a relative increase of 3.6 percentage points. The positive effect of the reform is driven by prime-aged women (+6.8 p.p. in relative terms, see Table 10) and high educated married women. For what concerns not married women all the estimated coefficients are not statistically different from zero and have a magnitude very close to zero.

Propensity to spend in human capital accumulation and the role of firms. As previously highlighted, I do not observe whether training is financed and provided directly by firms. However, I can gauge some evidence by looking at indirect proxies for firms involvement in middle-age workers training participation. I examine if the training effect differs by the size of the company. The results of this further heterogeneity check are available in Table 12, where columns (1)-(6) refer to firms whose size is 1-9 employees, 10-15, 16-25, 26-49, 50-249 and > 250 workers, respectively. According to these estimates, only employees working in very small-sized firms, those with at least 1 and maximum 9 employees, increased their training probability. Indeed, for each additional year of residual working life, this probability increases by about 1.8 percentage points, statistically significant at 1 percent level, translating into a relative increase of about 7 percentage points. As a further check, I also distinguish individuals by firm size and two broad firm economic sectors: the manufacturing and service sectors. The results are available in Table 13 where the first panel is devoted to the manufacturing sector and the second one to individuals working for firms operating in the service sector. For what concerns the manufacturing sector, individuals working in medium-sized firms (26-49 employees) experience a sizable increase in training participation, about 4.8 percent for each additional year of delay in pension eligibility in the aftermath of the reform. With regards to the service sector, individuals working in small-sized firms increases human capital accumulation, through more training, of about 2 percentage points, or about 8.7% in relative terms (as found by Berton et al. (2017) that, instead, use firm-level data).

Finally, I conclude the empirical analysis by looking at the other two outcomes I outlined in Section 3. These results are available in Table 14, where the first 3 columns are devoted to the willingness of the affected individual in paying for training participation, whereas the

[16]I consider those who declare themselves as single, divorced or widows as not married women.

last columns to the firm-sponsorship of on-the-job training. Concerning the willingness to pay, I cannot find a statistically significant effect despite being positive, even if I distinguish individuals according to the yearly median earnings of the sample, as a proxy for the individual budget constraint. On the other hand, for what concerns the probability that the training is employer-sponsored, I find that for each additional year increase MRA, this probability goes down by about 0.8 percentage points (-1.6 p.p. in relative terms).

Parallel trend assumption. As standard for the estimation of Difference-in-Difference models, I need to show that the trends in training participation would have been parallel for individuals with different exposure to the shock, absent the change in the pension rules. In order to test this assumption, I show that the difference in the individuals' on-the-job training participation more or less exposed to the shock was constant before 2011 and started changing exactly after the introduction of the new pension rules, from 2012 onward. Specifically, I estimate Eq. (1) by interacting the coefficient of the reform-induced shock with year-dummies (from 2007 to 2017) while omitting the year 2011 as the reference category. That is, I estimate the following equation, which consists of an event-study that estimates the baseline regression with different treatment years:

$$y_{iqt} = \sum_{\tau=2007}^{2017} \varphi_\tau shock_q \times \mathbf{1}(t=\tau) + \delta_q + \alpha_t + \mathbf{X}_{it} + \eta_{iqt} \qquad (2)$$

Equation (2) includes interactions between the shock variable and year dummies for every year excluded 2011. Under the assumption of parallel trends $\varphi_\tau \approx 0$ for τ<2011 (or at least not statistically significant at the conventional level of confidence). Figures 9, 10, 11, 12, 13, 14, 15 report the point estimates for φ_τ in equation (2) and 95% confidence intervals replicating all the specifications previously discussed. The visual inspection of each sub-sample coefficients $\{\gamma_\tau\}_{2007}^{2011}$ shows that individuals were substantially on a parallel trend, excepts some instances relative to women (Panels b and d of Figure 12) and individuals employed in manufacturing firms (Panel d of Figure 13). In the post-reform period, essentially, the dynamic estimates go in favour of the coefficients obtained by estimating its compact version counterpart, that is, equation (1).

5 Conclusions

In this paper, I provide causal evidence for human capital accumulation theory, whose key prediction is that middle-aged human capital investment returns crucially depend on the time left before retirement. An unanticipated and exogenous change in the working horizon affects the payout period and these returns.

Specifically, I leverage a restrictive pension reform affecting all Italian workers from 2011 that abruptly increased minimum retirement age (MRA) requirements. The analysis is based on a sample of individuals eligible to retire neither before nor after the 2011 pension reform. It exploits a Difference-in-Differences approach where the treatment variable measures the variation in *pre* and *post* reform MRAs, that is the exogenous increase in employees' residual working life.

My evidence shows that an increase in working life induced by the 2011 pension reform causally affects human capital investment through more training. In more detail, my results reveal that for each year increase in MRA, that training increases by about 0.7 percentage points, a relative increase of 1.7%. Additionally, I show that the response to the reform appears very heterogeneous and depends on gender, age, education, women's marital status, sector of employment and firm size. Unfortunately, in the aftermath of the reform, individuals' propensity to finance their training does not change. In contrast, I find that these individual positive effects are not attributable to employers' sponsorship.

Overall, my results supports the key prediction of human capital theory and have the potential to enrich the policy debates about pension policies, which usually do not consider human capital dynamics. My evidence suggests that policies aimed at increasing MRAs, mainly due to public finance motives, may have positive unintended consequences that may pay off also in terms of higher training, possibly because they may have contributed to extending relatively short working horizons and increasing the perceived benefits from additional training.

References

D. Acemoglu and S. Johnson. Disease and Development: The Effect of Life Expectancy on Economic Growth. *Journal of Political Economy*, 115(6):925–985, 2007.

D. Acemoglu and P. Restrepo. Demographics and Robots: Theory and Evidence. *mimeo*, 2017.

R. Almeida and P. Carneiro. The return to firm investments in human capital. *Labour Economics*, 16(1):97 – 106, 2009.

A. Bassanini, A. Booth, G. Brunello, M. De Paola, and E. Leuven. Workplace Training in Europe. IZA Discussion Papers 1640, Institute of Labor Economics (IZA), 2005.

E. Battistin, G. Brunello, S. Comi, and D. Sonedda. Train More, Work Longer The Effects of Pension Reforms and Training Incentives on the Retirement and Training Decisions of Older Workers. 2012.

A. B. Bauer and R. Eichenberger. Endogenous aging: How statutory retirement age drives human and social capital. *mimeo*, 2017.

G. S. Becker. Investment in Human Capital: A Theoretical Analysis. *Journal of Political Economy*, 70(5):9–49, 1962.

Y. Ben-Porath. The Production of Human Capital and the Life Cycle of Earnings. *Journal of Political Economy*, 75(4):352–365, 1967.

F. Berton, D. Guarascio, and A. Ricci. Ageing and firm-sponsored training: the role of pension reforms in Italy. *INAPP WP (mimeo)*, 2017.

M. Bertoni, G. Brunello, and G. Mazzarella. Does postponing minimum retirement age improve healthy behaviors before retirement? Evidence from middle-aged Italian workers. *Journal of Health Economics*, 58:215–227, 2018.

N. Bianchi, G. Bovini, J. Li, M. Paradisi, and M. Powell. Career Spillovers in Internal Labor Markets. *mimeo*, 2019.

A. Blinder and Y. Weiss. Human Capital and Labor Supply: A Synthesis. *Journal of Political Economy*, 84(3):449–72, 1976.

D. E. Bloom, M. Kuhn, and K. Prettner. Health and Economic Growth. *mimeo*, 2019.

T. Boeri, P. Garibaldi, and E. Moen. Closing the Retirement Door and the Lump of Labor. *mimeo*, 2017.

G. Bovini and M. Paradisi. Labor substitutability and the impact of raising the retirement age. *mimeo*, 2019.

M. Bratti, M. Conti, and G. Sulis. Employment protection and firm-provided training in dual labour markets. *Labour Economics*, 69, 2021.

G. Brunello and P. Wruuck. Employer Provided Training in Europe: Determinants and Obstacles. *mimeo*, 2020.

F. Carta and M. De Phillipis. Working Horizon and Labour Supply:the Effect of Raising Minimum Retirement Age on Middle-aged Individuals. *mimeo*, 2019.

F. Carta, F. D'Amuri, and T. von Wachter. Aging worforce, pension reforms, and firms' dynamics. *mimeo*, 2019.

M. Cervellati and U. Sunde. Life Expectancy, Schooling, and Lifetime Labor Supply: Theory and Evidence Revisited. *Econometrica*, 81(5):2055–2086, 2013.

Commissione di Vigilanza sui Fondi Pensione. Relazione annuale per l'anno 2018. 2019.

A. De Grip and J. Van Loo. The economics of skills obsolescence: A review. *Research in Labor Economics*, 21:1–26, 2002.

B. Dostie. The Impact of Training on Innovation. *ILR Review*, 71(1):64–87, 2018.

X. Fan, A. Seshadri, and C. Taber. Understanding Earnings, Labor Supply, and Retirement Decisions. *mimeo*, 2017.

Fondazione Itinerari Previdenziali. Settimo Rapporto sul bilancio del sistema previdenziale Italiano. 2020.

E. M. Fornero. "Reform, Inform, Educate": A new paradigm for pension systems. In B. Marin, editor, *The Future of Welfare in a Global Europe*, pages 297–324. Ashgate, 2015.

D. Fouarge and T. Schils. The Effect of Early Retirement Incentives on the Training Participation of Older Workers. *LABOUR*, 23(s1):85–109, 2009.

N. K. Goodpaster. Leaves and leaving: the family and medical leave act and the decline in maternal labor force participation. *The BE Journal of Economic Analysis & Policy*, 10 (1), 2010.

Z. Griliches. Education, Human Capital, and Growth: A Personal Perspective. *Journal of Labor Economics*, 15(1):S330–44, 1997.

J.-O. Hairault, T. Sopraseuth, and F. Langot. Distance to retirement and older workers' employment: The case for delaying the retirement age. *Journal of the European Economic Association*, 8(5):1034–1076, 2010.

C. Hansen and H. Strulik. Life expectancy and education: evidence from the cardiovascular revolution. *Journal of Economic Growth*, 22(4):421–450, 2017.

M. Hazan. Longevity and Lifetime Labor Supply: Evidence and Implications. *Econometrica*, 77(6):1829–1863, 2009.

Istituto Nazionale di Previdenza Sociale. XV Rapporto Annuale. 2016.

S. Jayachandran and A. Lleras-Muney. Life Expectancy and Human Capital Investments: Evidence from Maternal Mortality Declines. *The Quarterly Journal of Economics*, 124(1): 349–397, 2009.

S. Kalemli-Ozcan, H. E. Ryder, and D. Weil. Mortality decline, human capital investment, and economic growth. *Journal of Development Economics*, 62(1):1–23, 2000.

A. B. Krueger and J.-S. Pischke. The Effect of Social Security on Labor Supply: A Cohort Analysis of the Notch Generation. *Journal of Labor Economics*, 10(4):412–437, 1992. doi: 10.1086/298294.

R. Lalive and S. Staubli. How does raising women's full retirement age affect labor supply, income, and mortality. *mimeo*, 2015.

A. Leigh. Informal care and labor market participation. *Labour Economics*, 17(1):140–149, 2010.

J. B. Liebman, E. F. P. Luttmer, and D. G. Seif. Labor supply responses to marginal Social Security benefits: Evidence from discontinuities. *Journal of Public Economics*, 93(11): 1208 – 1223, 2009. ISSN 0047-2727.

P. Lorentzen, J. McMillan, and R. Wacziarg. Death and development. *Journal of Economic Growth*, 13(2):81–124, 2008.

D. Manoli and S. Weber. Nonparametric Evidence on the Effects of Financial Incentives on Retirement Decisions. *American Economic Journal: Economic Policy*, 8(4):160–82, 2016.

P. Martins. Employee Training and Firm Performance: Quasi- experimental evidence from the European Social Fund. *mimeo*, 2020.

G. Mastrobuoni. Labor supply effects of the recent social security benefit cuts: Empirical estimates using cohort discontinuities. *Journal of Public Economics*, 93(11):1224 – 1233, 2009. ISSN 0047-2727.

P.-J. Messe and B. Rouland. Stricter employment protection and firms' incentives to sponsor training: The case of French older workers. *Labour Economics*, 31:14–26, 2014.

J. Mincer. On-the-Job Training: Costs, Returns, and Some Implications. *Journal of Political Economy*, 70(5):50–79, 1962.

R. Montizaan, F. Corvers, and A. De Grip. The effects of pension rights and retirement age on training participation: Evidence from a natural experiment. *Labour Economics*, 17(1): 240 – 247, 2010.

L. Munasinghe, T. Reif, and A. Henriques. Gender gap in wage returns to job tenure and experience. *Labour economics*, 15(6):1296–1316, 2008.

OECD. Pensions at a Glance 2011: Retirement-Income Systems in OECD and G20 Countries. 2011.

E. Oster, I. Shoulson, and E. R. Dorsey. Limited Life Expectancy, Human Capital and Health Investments. *American Economic Review*, 103(5):1977–2002, 2013.

J.-S. Pischke. Continuous training in Germany. *Journal of Population Economics*, 14(3): 523–548, 2001.

S. Staubli and J. Zweimüller. Does raising the early retirement age increase employment of older workers? *Journal of Public Economics*, 108:17 – 32, 2013.

Figures and Tables

Figure 1: Increase in working life (variation in pension rules between 2017 and 2011)

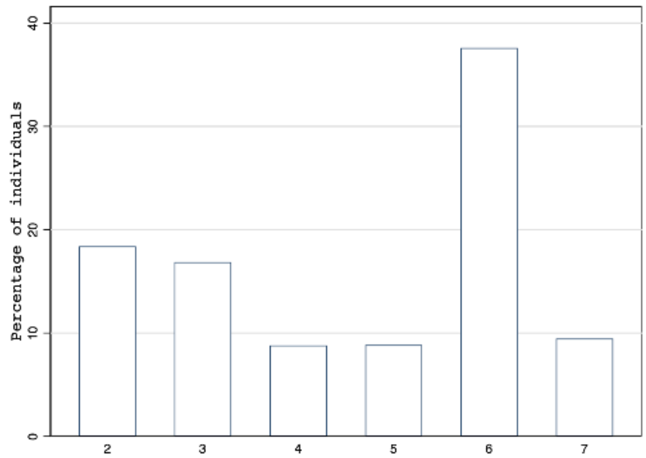


Figure 2: Increase in working life by gender (variation in pension rules between 2017 and 2011)

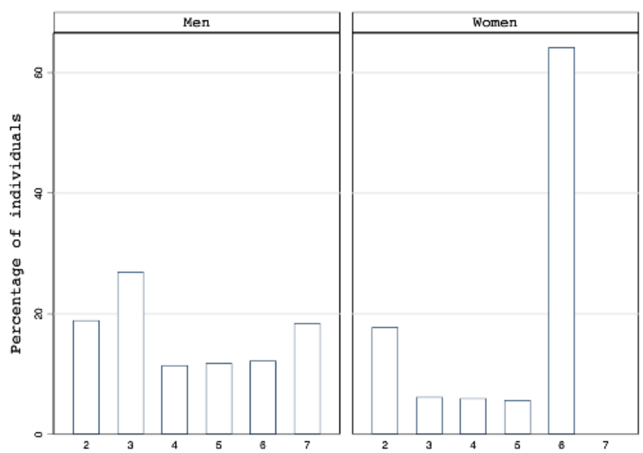


Figure 3: Percentage of individuals retired as function of distance to retirement (MRA_q - Retirement age)

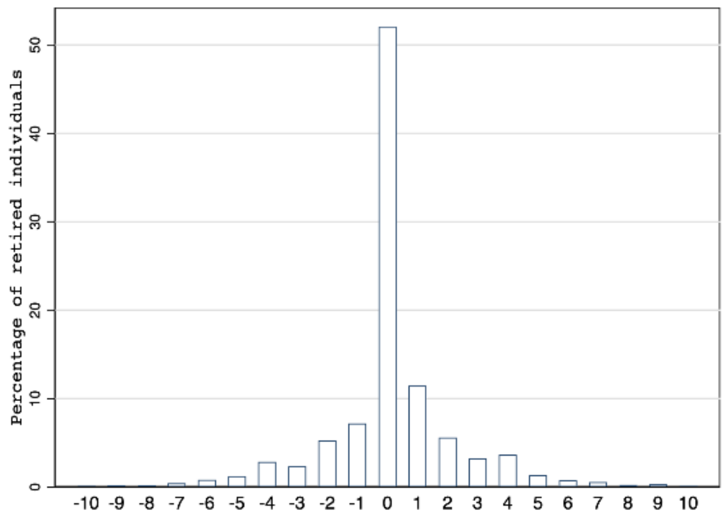
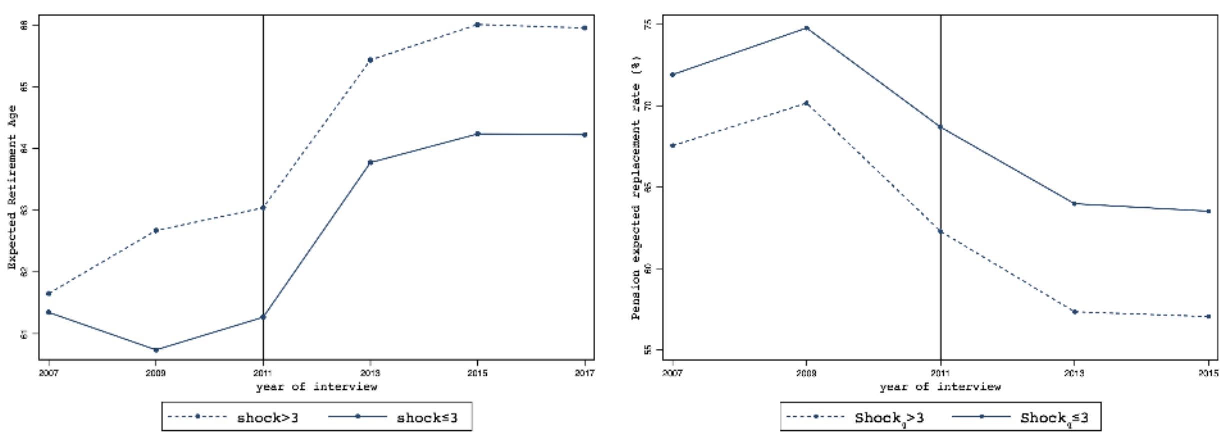


Figure 4: Declared expected retirement age and replacement pension income rate



(a) Expected retirement age **(b)** Expected replacement rate

Figure 5: Training participation by most and least treated

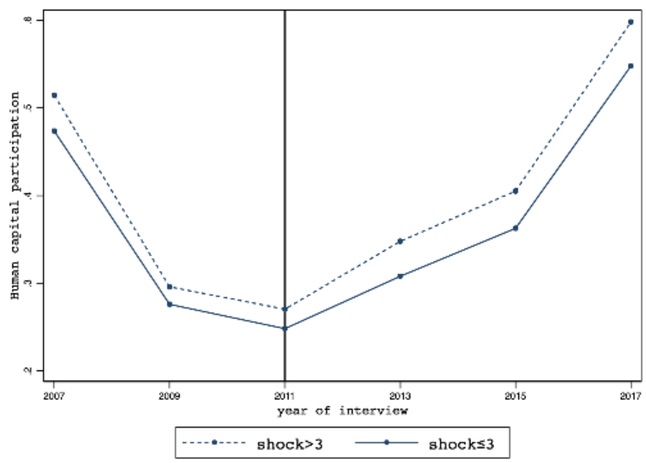


Figure 6: Training participation by gender and exposure to the 2011 reform

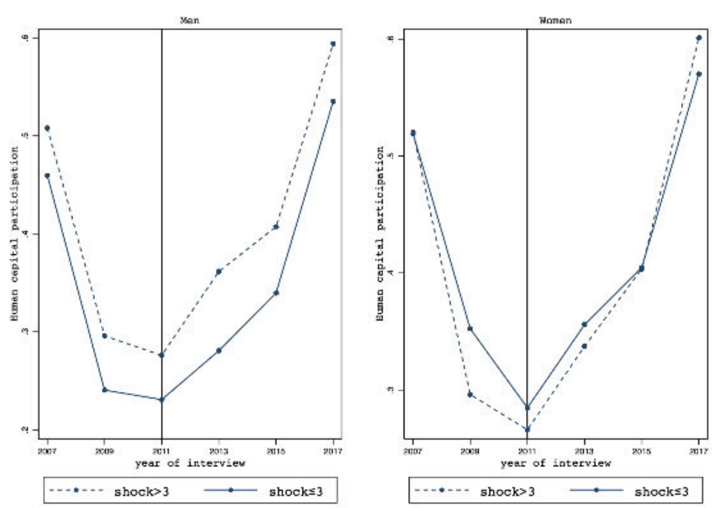
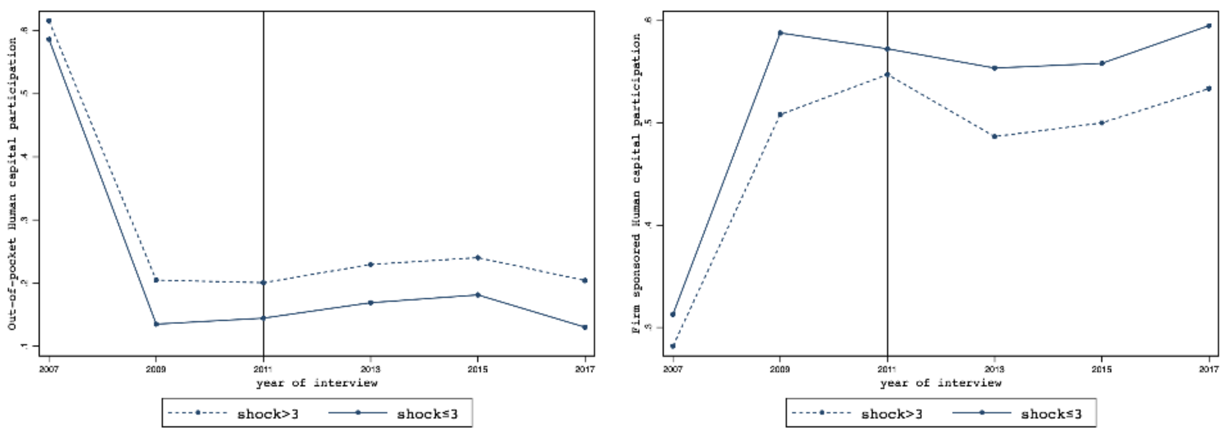


Figure 7: Paid and firm-sponsored training participation



(a) Paid training

(b) Firm sponsored training

Table 1: Old age pension eligibility rules

	Men			Women		
Year	Public	Private	Self-employed	Public	Private	Self-employed
Before *Fornero* reform:						
2007	65	65	65	60	60	60
2008	65	65	65	60	60	60
2009	65	65	65	60	60	60
2010	65	65	65	61	60	60
2011	65	65	65	61	60	60
After *Fornero* reform:						
2012	66	66	66	66	62	63
2013	66	66	66	66	62	64
2014	66	66	66	66	64	65
2015	66	66	66	66	64	65
2016	67	67	67	67	66	66
2017	67	67	67	67	66	66
2018	67	67	67	67	67	67

Notes: Old age pension eligibility requires the legal retirement age (reported above) and at least 20 accrued years of contribution.

Table 2: Seniority pension eligibility rules

	Men				Women			
	Public-Private		Self-employed		Public-Private		Self-employed	
Year	only C	Quota	only C	Quota	only C	Quota	only C	Quota
Before *Fornero* reform:								
2007	39yc.	57y.+ 35yc.	40yc.	58y. + 35yc.	39yc.	57y.+ 35yc.	40yc.	58y. + 35yc.
2008	40yc.	58y.+ 35yc.	40yc.	59y. + 35yc.	40yc.	58y.+ 35yc.	40yc.	59y. + 35yc.
2009	40yc.	59y.+ 35yc.	40yc.	60y. + 35yc.	40yc.	59y.+ 35yc.	40yc.	60y. + 35yc.
2010	40yc.	59y.+ 35yc.	40yc.	60y. + 35yc.	40yc.	59y.+ 35yc.	40yc.	60y. + 35yc.
2011	40yc.	60y.+ 35yc.	40yc.	61y. + 35yc.	40yc.	60y.+ 35yc.	40yc.	61y. + 35yc.
After *Fornero* reform:								
2012	42yc.	-	42yc.	-	41yc.	-	41yc.	-
2013	42yc.	-	42yc.	-	41yc.	-	41yc.	-
2014	42yc.	-	42yc.	-	41yc.	-	41yc.	-
2015	42yc.	-	42yc.	-	41yc.	-	41yc.	-
2016	43yc.	-	43yc.	-	42yc.	-	42yc.	-
2017	43yc.	-	43yc.	-	42yc.	-	42yc.	-
2018	43yc.	-	43yc.	-	42yc.	-	42yc.	-

Notes: Under the seniority pension regime individuals can be granted eligibility when the number of accrued years of contribution reached a minimum amount; that is 39 in 2007, 40 between 2008 and 2011 and so on. When individuals retire using the option that I labelled as “only C” there is no binding age requirement. The second option, instead, available up to 2011 was the *Quota* system according to which individuals can retire if they have at least 35 years of contribution and a minimum age requirement. In the pre-reform period, self-employed and private-public employees were subject to different seniority pension rules, both in terms of “only C” and the *Quota* system, but not with respect gender. In the post-reform period requirement have been levelled out between sectors of employment but not with respect gender.

Table 3: Descriptive statistics

	2007-2015: All	$Shock_q > 3$ (most treated)	$Shock_q \leq 3$ (least treated)	2007-2011 (pre-reform period): All	$Shock_q > 3$ (most treated)	$Shock_q \leq 3$ (least treated)
Men	0.528	0.449	0.666	0.562	0.473	0.703
	(0.499)	(0.497)	(0.472)	(0.496)	(0.499)	(0.457)
Age	51.861	52.053	51.525	51.790	51.883	51.643
	(5.978)	(6.170)	(5.608)	(5.548)	(5.622)	(5.424)
Years of contrib.	25.946	24.661	28.202	25.750	24.318	28.044
	(7.904)	(7.767)	(7.634)	(7.745)	(7.421)	(7.704)
High educ.	0.283	0.327	0.206	0.242	0.278	0.185
	(0.450)	(0.469)	(0.404)	(0.428)	(0.448)	(0.388)
Married	0.577	0.574	0.582	0.291	0.274	0.318
	(0.494)	(0.494)	(0.493)	(0.454)	(0.446)	(0.466)
Household size	3.167	3.154	3.189	3.176	3.161	3.201
	(1.157)	(1.166)	(1.140)	(1.153)	(1.163)	(1.138)
If children	0.800	0.804	0.793	0.821	0.825	0.814
	(0.400)	(0.397)	(0.405)	(0.383)	(0.380)	(0.389)
Annual earnings	28,138.844	28,000.584	28,380.898	28,377.006	28,652.243	27,944.502
	(28,374.396)	(29,097.370)	(27,061.327)	(28,428.983)	(30,558.296)	(24,711.057)
Public sector	0.391	0.400	0.376	0.407	0.414	0.396
	(0.488)	(0.490)	(0.484)	(0.491)	(0.493)	(0.489)
Private sector	0.460	0.403	0.561	0.451	0.392	0.547
	(0.498)	(0.490)	(0.496)	(0.498)	(0.488)	(0.498)
Self-employed	0.149	0.198	0.063	0.142	0.194	0.057
	(0.356)	(0.398)	(0.244)	(0.349)	(0.396)	(0.232)
HAC	0.398	0.415	0.370	0.346	0.359	0.324
	(0.490)	(0.493)	(0.483)	(0.476)	(0.480)	(0.468)
Paid HAC	0.258	0.279	0.218	0.373	0.398	0.328
	(0.438)	(0.449)	(0.413)	(0.484)	(0.489)	(0.470)
Firm-sponsored HAC	0.497	0.477	0.534	0.430	0.409	0.467
	(0.500)	(0.499)	(0.499)	(0.495)	(0.492)	(0.499)
Obs.	53,977	34,386	19,591	20,600	12,681	7,919

Notes: The sample is composed of individuals aged between 40 and 64 years, with at least 10 and less than 40 accrued years of contribution, eligible to retire neither before nor after the reform. HAC stands for human capital accumulation. Mean averages and standard deviation in parentheses.

Table 4: Training participation and the 2011 pension reform

	All	Men	Women
	(1)	(2)	(3)
$shock_q \times post2011$	0.0069**	0.0093**	0.0034
	(0.0024)	(0.0032)	(0.0035)
Obs.	53,977	28,478	25,499
Adj. R^2	0.1299	0.1015	0.1722

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 5: Heterogeneity by age

	Age class:		
	40-47	48-56	57-64
	(1)	(2)	(3)
		All:	
$shock_q \times post2011$	0.0131**	0.0074*	-0.0038
	(0.0048)	(0.0033)	(0.0049)
Obs.	13,600	27,289	13,088
Adj. R^2	0.1332	0.1330	0.1245
		Men:	
$shock_q \times post2011$	0.0140*	0.0113*	-0.0007
	(0.0068)	(0.0046)	(0.0062)
Obs.	6,103	14,703	7,672
Adj. R^2	0.1087	0.1026	0.1025
		Women:	
$shock_q \times post2011$	0.0088	0.0027	-0.0041
	(0.0069)	(0.0047)	(0.0082)
Obs.	7,497	12,586	5,416
Adj. R^2	0.1658	0.1784	0.1629

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 6: Heterogeneity by sector of employment

	Sector of employment:		
	Public	Private	Self-employed
	(1)	(2)	(3)
$shock_q \times post2011$	0.0042	0.0016	0.0154^{*}
	(0.0043)	(0.0033)	(0.0064)
Obs.	21,113	24,831	8,033
Adj. R^2	0.0754	0.0696	0.0776

Notes: Controls include: year, shock, gender, age, marital status, region and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 7: Heterogeneity by firms broad economic sectors:

	Firm's economic sector:	
	Manufacturing	Service
	(1)	(2)
$shock_q \times post2011$	0.0038	0.0083^{**}
	(0.0055)	(0.0032)
Obs.	8,059	24,805
Adj. R^2	0.0664	0.0829

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 8: Heterogeneity by education level:

	Education level:		
	Low	Medium	High
	(1)	(2)	(3)
		All:	
$shock_q \times post2011$	0.0005	-0.0013	0.0143^{**}
	(0.0035)	(0.0033)	(0.0047)
Obs.	11,645	27,057	15,275
Adj. R^2	0.0655	0.1054	0.0715
		Men:	
$shock_q \times post2011$	0.0067	0.0073^{+}	0.0173^{**}
	(0.0048)	(0.0043)	(0.0060)
Obs.	6,694	14,319	7,465
Adj. R^2	0.0597	0.0783	0.0661
		Women:	
$shock_q \times post2011$	-0.0039	-0.0114^{*}	0.0154^{+}
	(0.0058)	(0.0048)	(0.0088)
Obs.	4,951	12,738	7,810
Adj. R^2	0.0848	0.1489	0.0813

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 9: Heterogeneity by women marital status:

	Married		Not married	
	(1)	(2)	(3)	(4)
$shock_q \times post2011$	0.0136^{*}	0.0135^{*}	-0.0009	-0.0008
	(0.0066)	(0.0066)	(0.0059)	(0.0059)
Obs.	14,991	14,991	10,508	10,508
Adj. R^2	0.1586	0.1591	0.1924	0.1925

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Columns (2) and (4) include, also, no. of kids and household size as further controls. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 10: Heterogeneity by women marital status and age:

	Age class:					
	40-47		48-56		57-64	
	(1)	(2)	(3)	(4)	(5)	(6)
	Married:					
$shock_q \times post2011$	0.0245$^+$	0.0246$^+$	0.0089	0.0090	0.0042	0.0043
	(0.0139)	(0.0139)	(0.0083)	(0.0082)	(0.0152)	(0.0152)
Obs.	4,610	4,610	7,127	7,127	3,254	3,254
Adj. R^2	0.1536	0.1533	0.1613	0.1621	0.1507	0.1528
	Not married:					
$shock_q \times post2011$	0.0020	0.0022	0.0001	0.0000	-0.0051	-0.0063
	(0.0091)	(0.0091)	(0.0087)	(0.0087)	(0.0123)	(0.0122)
Obs.	2,887	2,887	5,459	5,459	2,162	2,162
Adj. R^2	0.1906	0.1908	0.2008	0.2007	0.1708	0.1716

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Columns (2), (4) and (6) include, also, no. of kids and household size as further controls. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 11: Heterogeneity by women marital status and education:

	Education level:					
	Low		Medium		High	
	(1)	(2)	(3)	(4)	(5)	(6)
	Married:					
$shock_q \times post2011$	-0.0051	-0.0049	0.0034	0.0036	0.0213	0.0213
	(0.0088)	(0.0088)	(0.0089)	(0.0089)	(0.0141)	(0.0141)
Obs.	2,731	2,731	7,585	7,585	4,675	4,675
Adj. R^2	0.0722	0.0718	0.1397	0.1411	0.0599	0.0599
	Not married:					
$shock_q \times post2011$	-0.0056	-0.0058	-0.0172*	-0.0172*	0.0073	0.0073
	(0.0108)	(0.0108)	(0.0078)	(0.0078)	(0.0135)	(0.0135)
Obs.	2,220	2,220	5,153	5,153	3,135	3,135
Adj. R^2	0.0966	0.0969	0.1635	0.1633	0.1139	0.1139

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Columns (2), (4) and (6) include, also, no. of kids and household size as further controls. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 12: Heterogeneity by firm size:

	Firm size:					
	1-9	10-15	16-25	26-49	50-249	>250
	(1)	(2)	(3)	(4)	(5)	(6)
$shock_q \times post2011$	0.0185***	-0.0003	-0.0040	0.0106	-0.0003	-0.0082
	(0.0042)	(0.0086)	(0.0091)	(0.0115)	(0.0086)	(0.0069)
Obs.	11,975	2,827	2,113	1,864	3,909	8,536
Adj. R^2	0.0945	0.0614	0.0550	0.0690	0.0554	0.0733

Notes: The estimates refer only to self-employed and private sector workers. Firm size refers to the number of employees, including the interviewed, working in the firm at the year of interview. Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 13: Heterogeneity by firms size and economic sectors:

	Firm size:					
	1-9	10-15	16-25	26-49	50-249	>250
	(1)	(2)	(3)	(4)	(5)	(6)
	Manufacturing sector:					
$shock_q \times post2011$	0.0109	-0.0282	0.0335	0.0480*	-0.0017	-0.0048
	(0.0083)	(0.0180)	(0.0264)	(0.0234)	(0.0171)	(0.0177)
Obs.	3,063	739	439	560	1,339	1,623
Adj. R^2	0.0958	0.0310	0.0391	0.0781	0.0588	0.0500
	Service sector:					
$shock_q \times post2011$	0.0208***	0.0057	-0.0081	-0.0026	0.0033	-0.0107
	(0.0049)	(0.0096)	(0.0101)	(0.0135)	(0.0105)	(0.0077)
Obs.	8,912	2,088	1,674	1,304	2,570	6,913
Adj. R^2	0.0949	0.0684	0.0642	0.0808	0.0586	0.0821

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 14: Willingness to pay and employer-sponsored training:

	Paid training			Firm-sponsored
		Wage above	Wage below	
	All:	median:	median:	
	(1)	(2)	(3)	(4)
$shock_q \times post2011$	0.0041	0.0056	0.0018	-0.0079^{+}
	(0.0036)	(0.0046)	(0.0055)	(0.0041)
Obs.	21,289	13,033	8,256	20,308
Adj. R^2	0.2048	0.2121	0.1949	0.0898

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. Robust standard errors, in parentheses, clustered at the age-sector of employment-gender-years of contribution level. Statistical significance denoted as follows: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Appendix - Additional info, figures and tables

A The 2007 severance pay reform

In 2007, the implementation of the severance pay (Trattamento di fine rapporto, TFR) reform had introduced an automatic enrolment mechanism for voluntary pension funds. According to the reform, the private sector workers' severance pay will be automatically paid into an occupational pension plan and no anymore retained in the firm if they do not opt-out. However, according to Commissione di Vigilanza sui Fondi Pensione (2019), only one-third of private-sector workers have a contract with a private pension fund, whose benefits are conditional on the eligibility for a public pension.

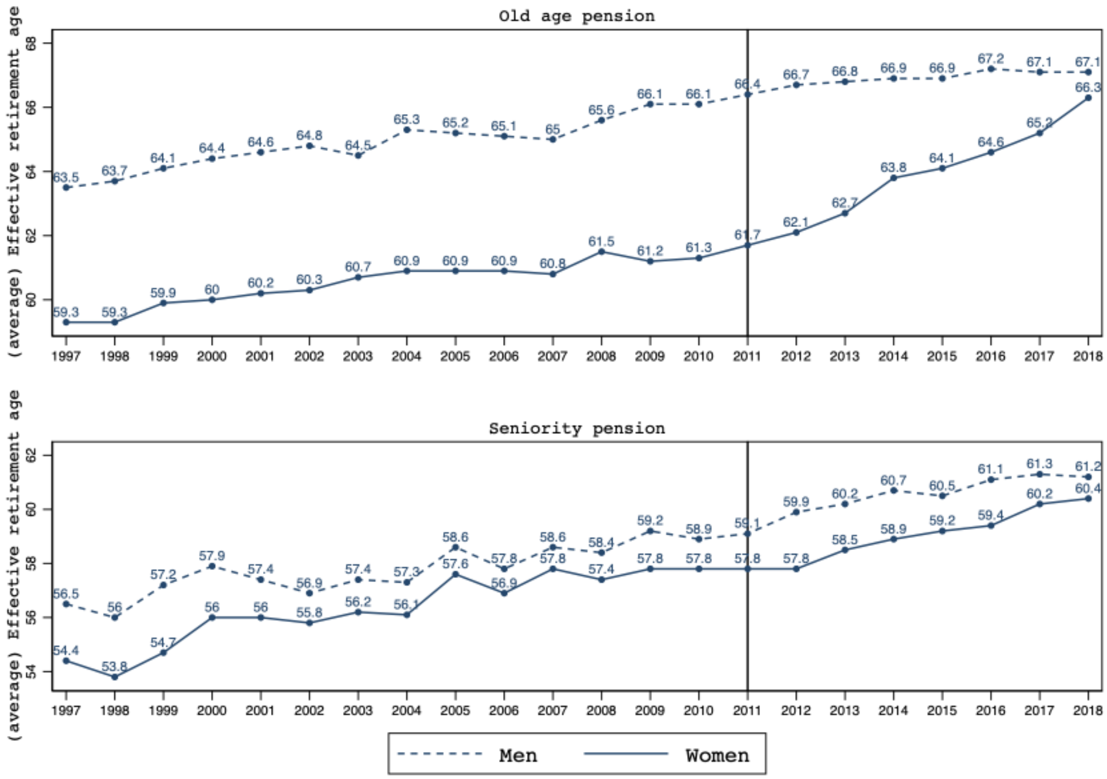
B *Amato* reform

Three years earlier than the *Dini* reform, another policy measure was legislated to curb pension expenditures. The *Amato* reform (legislative decree no. 503/1992) increased the requirements for claiming an old-age pension. According to the decree's directives, the retirement age for old-age pensions, managed by the Social Security Institute, was raised from 55 to 60 for women and from 60 to 65 for men, while the necessary contribution years became 20 (15 before the reform). In addition, once fulfilled the requirements, pension benefits were calculated based on the salary of the last 5 years according to the DB method.

C Effective retirement age evolution

According to Fondazione Itinerari Previdenziali (2020), after the implementation of the 2011 reform, the (average) effective retirement age has increased. However, the rise in the average age at which first pension instalments are claimed differently evolved. The highest increase, on average, has been experienced by women retiring under the old-age scheme (about 4 years and 6 months). For men, instead, the rise has been of about 7 months. Regarding the seniority scheme, the (average) effective retirement age evolved according to the increase in the required years of accrued contributions (43 and 42 for men and women, respectively; whereas up to 2011, the requirement was set to 40 years of paid contributions). Women retiring under this regime faced increases of about 2 years and 6 months, whereas men 2 years and 1 month. However, it should be reminded that retiring according to the seniority regime only implies requirements in terms of accrued years of paid contributions and not in age. For more details, see Figure 8.

Figure 8: Effective (average) retirement age by gender and pension regime



Source: Fondazione Itinerari Previdenziali (2020) based on social security records.

D Grandfathering clauses

An essential feature of the reform is that grandfather clauses were very limited. They only applied to workers eligible to claim a pension under the old rules by December 31, 2011, and to a couple of other specific categories: workers *collocati in mobilità* according to law 223/91 and based on collective agreements signed before 31/10/2011; workers who, as of October 31, 2011, were beneficiaries of *prestazioni straordinarie a carico dei fondi di solidarietà di settore*; workers who, as of October 31, 2011, had ceased to work but had been authorized to continue to pay contributions. The lack of grandfather clauses meant the reform had an immediate effect on the retirement decisions of most Italian workers.

E Other provisions of the 2011 pension reform

Finally, the 2011 pension reform, in addition to increasing the mandated retirement age, changed the pension benefit formula for those who were still covered by the defined-benefit method of calculation (individuals with at least 18 years of accrued contribution by January 1996), moving them to the notional defined-contribution method for working years after 2011.

F More on sample selection and *esodati*

For the sake of clarity, I drop from each survey's wave all those individuals eligible to retire under the old-age pension scheme according to the pension rules in place in that year. I do not have to check for seniority requirements since I consider only individuals with less than 40 years of accrued contributions, but I drop all of them that are eligible to retire under the "Quota" system up to 2011. Furthermore, I can drop from the sample all those individuals that after 2011 declare themselves as *esodato* (a question contained in the survey). An *esodato* is a worker who, when he comes close to retirement, has reached an agreement with his company to leave his job in exchange for economic coverage until he reaches the pension. According to Istituto Nazionale di Previdenza Sociale (2016), there have been 7 *salvaguardie* from 2011 (up to 2016) in order to ensure that these *esodati* would have been able to obtain pension installments even though they did not meet the post-reform eligibility rules. The total number of *esodati salvaguardati* amounts to more than 101,837 individuals for a total cost, borne by taxpayers, of more than 9 billion euros.

G An illustrative example

To better understand the source of cross-sectional variation in the exposure to the pension reform that I exploit in the empirical analysis, a simple example may be illustrative. Table 15 considers six different individuals: 3 women (the first panel) and 3 men (the second panel), all aged 59 years, however, with different years of paid contributions and sector of employment. For instance, consider Beatrice, a private-sector worker with 35 years of paid contributions. According to the pre-reform rules, she would have met eligibility criteria in access to the public pension at 64 years if she had chosen to retire under the seniority scheme or 60 years under the old-age or quota system. Hence, her minimum retirement age was 60 years. Under the post-reform rules, she can only choose to retire under the seniority or old-age regime. In both cases, her retirement age will be 66. Because of the reform, her MRA increased, and the size of shock amounts to 6 years, that is, the increase in the residual working life. Paola, instead, is a public sector worker with 26 years of paid contributions. Supposing she could have retired under the pre-reform rules, she would have retired at 61 years under the old-age requirements, which corresponds to her MRA. Following the rules in

2017, instead, now she would retire at 67 years, six years later than expected. Hence, women experienced the greatest and least heterogeneous increase in the residual working life.

Men, conversely, have been affected differently by the 2011 pension reform. Alessandro is a private sector employee with 35 years of contributions. If he could have retired under the 2011 rules, his MRA was 60 years, but because of the reform, his MRA was 67 in 2017. That is a 7 years shock. Alternatively, Leonardo has 26 years of paid contributions as a public sector worker. In 2011, his MRA was 65 years. Because of the reform, in 2017, his MRA equals 67 years, which is a two years shock. In this case, the shock's source of variation for men is larger for those who planned to retire under the quota system before the reform.

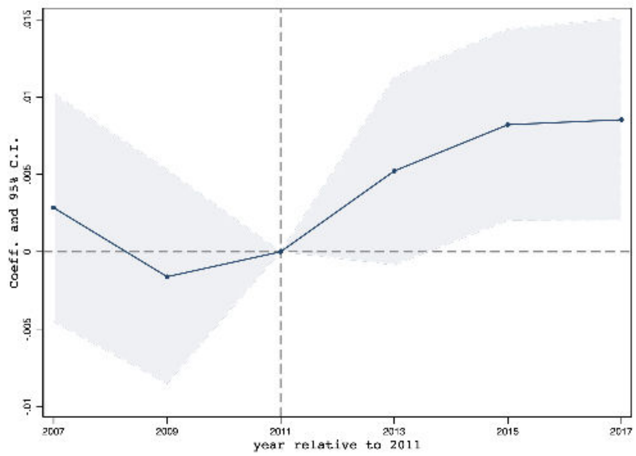
Table 15: Example of heterogeneity in gender, years of social security contributions, sector of employment, pension regimes and eligibility rules

	Pension rules in 2011:				Pension rules in 2017:			
	Seniority	Old age	Quota	MRA_{2011}	Seniority	Old age	MRA_{2017}	Shock:
Women, 59 years:								
Beatrice, C=35yc.; S=Priv.	64	60	60	**60**	66	66	**66**	6
Lucrezia, C=30yc.; S=Self-emp.	69	60	65	**60**	71	66	**66**	6
Paola, C=26yc.; S=Publ.	73	61	68	**61**	75	67	**67**	6
Men, 59 years:								
Alessandro, C=35yc.; S=Priv.	64	65	60	**60**	67	67	**67**	7
Francesco, C=30yc.; S=Self-empl.	69	65	64	**64**	72	67	**67**	3
Leonardo, C=26yc.; S=Publ.	73	65	68	**65**	76	67	**67**	2

Notes: This table reports an example of how individuals are differently affected by the increase in the MRA, given their accrued years of contribution, gender and sector of employment. The Table displays the age at which individuals can claim the old age or the seniority pension (including the *Quota* system in place before the 2011 reform). The minimum retirement age takes the first age of eligibility among the three pension regimes in the pre-reform period and among the two pension regimes in the post-reform period. C stands for the number of accrued years of contribution and S for the sector of employment (private, public or self-employed). Shock, the last column of the table, measures the difference between the minimum retirement age after and before the reform enacted at the end of 2011.

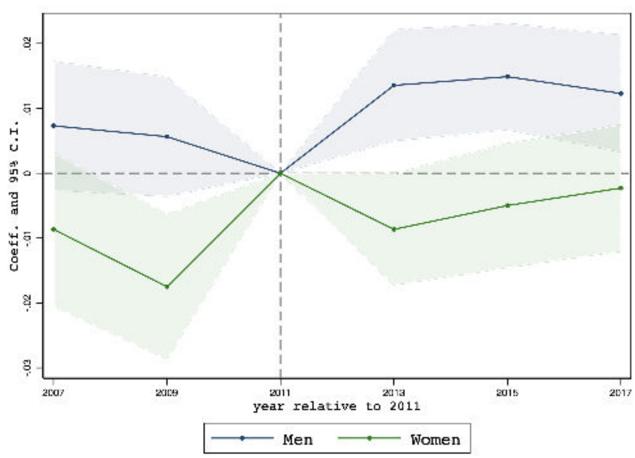
H Parallel trend assumption

Figure 9: Event-study estimates



Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2). The dependent variable is a dummy variable that takes value of 1 if individual i has attended on-the-job training in the last 12 months.

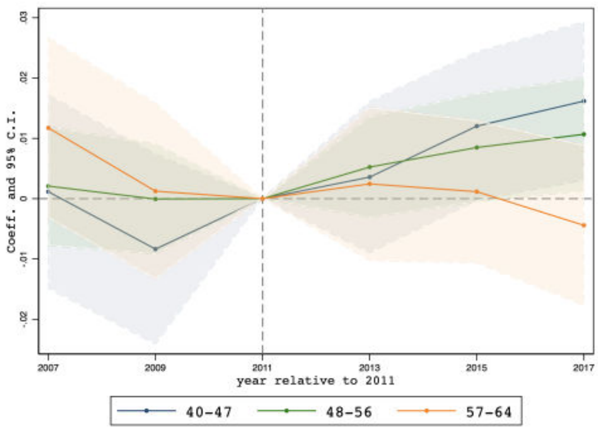
Figure 10: Event-study estimates by gender



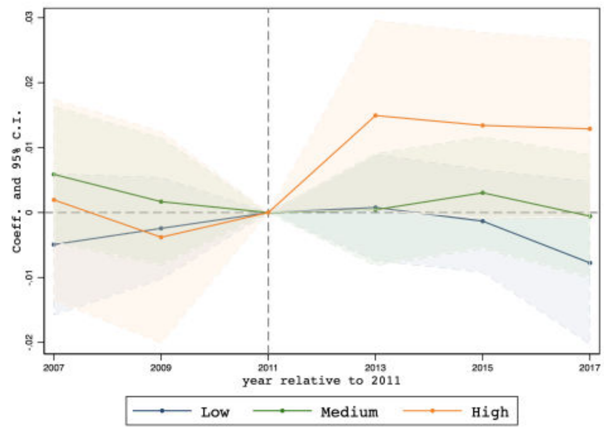
Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2) distinguishing the sample by gender. The dependent variable is a dummy variable that takes value of 1 if individual i has attended on-the-job training in the last 12 months.

Figure 11: Event-study estimates by:

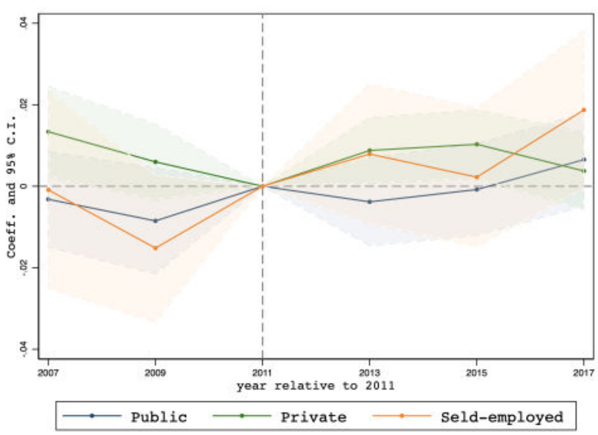
(a) Age classes



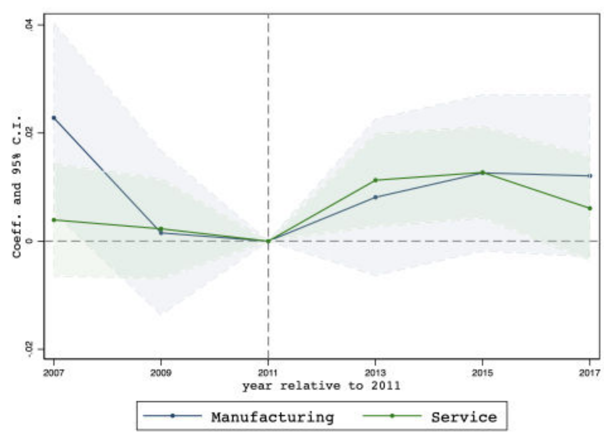
(c) Education



(b) Sector of employment



(d) Economic sector of the firm

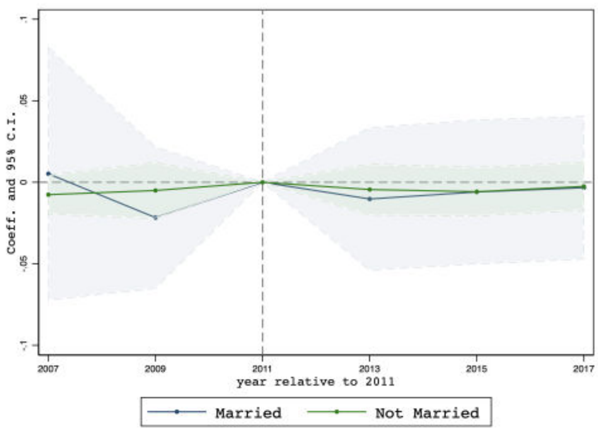


Source: PLUS (INAPP) 2007-2017.

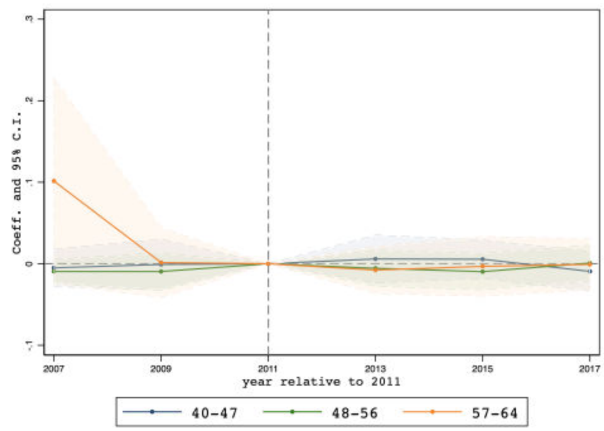
Notes: Estimates based on equation (2) distinguishing by each sub-sample. The dependent variable is a dummy variable that takes value of 1 if individual i has attended on-the-job training in the last 12 months.

Figure 12: Event-study estimates, women only, by:

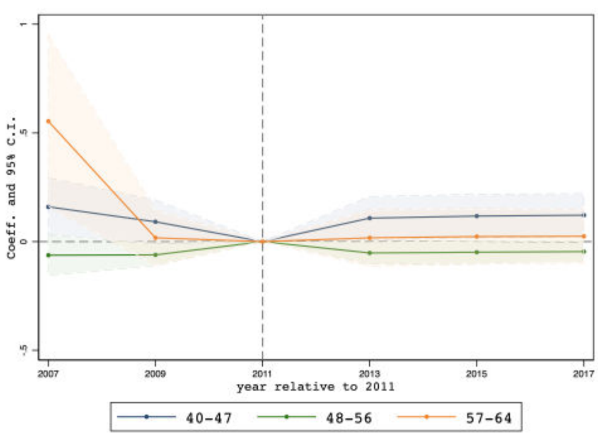
(a) Martial status



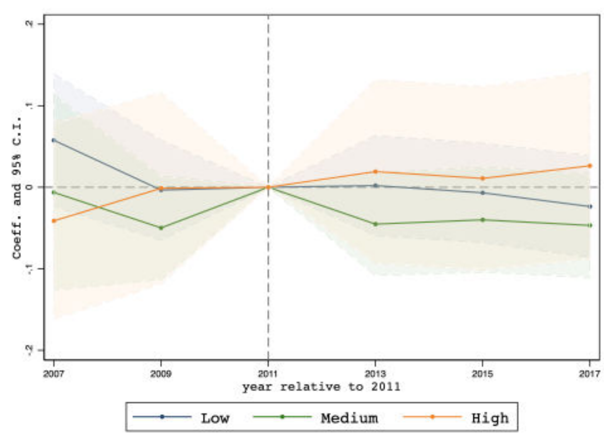
(c) Not married and age classes



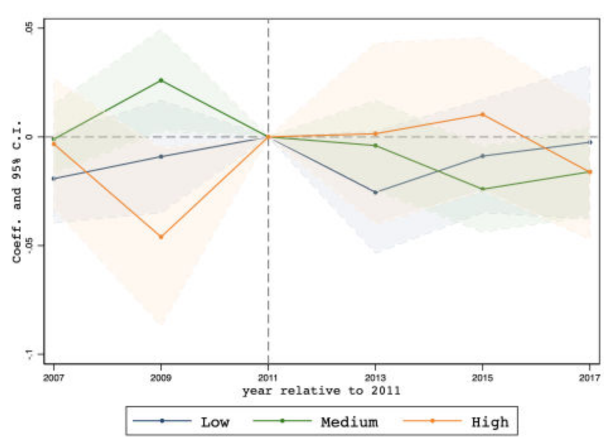
(b) Married and age classes



(d) Married and education



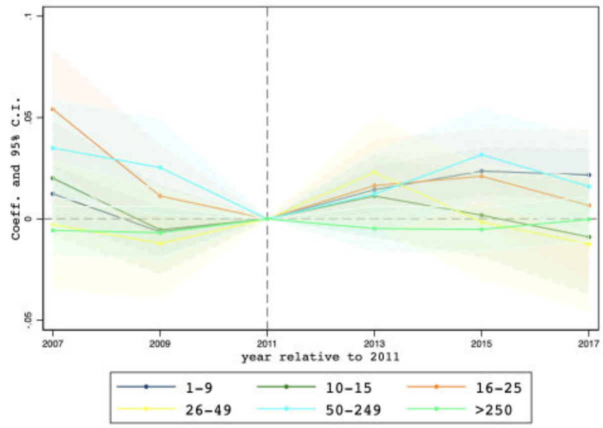
(e) Not married and education



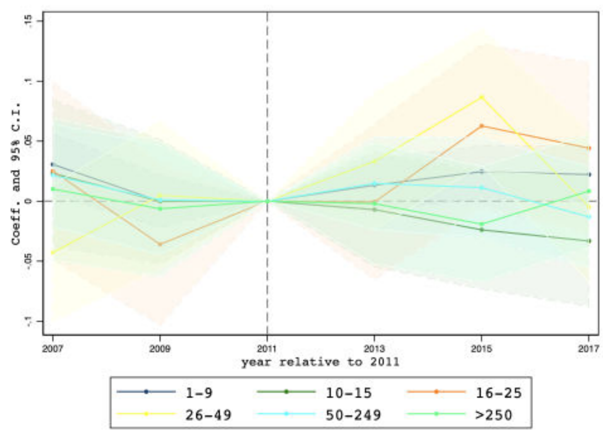
Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2) considering only the sample of women and distinguishing them according to their martial status (married or not married) and by age classes and education level. The dependent variable is a dummy variable that takes value of 1 if individual i has attended on-the-job training in the last 12 months.

Figure 13: Event-study estimates by:

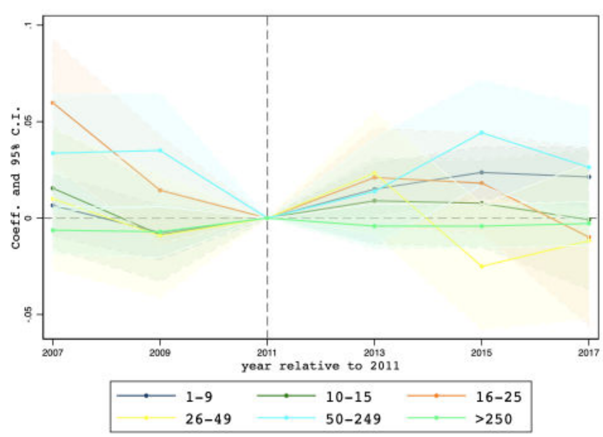
(a) Firm size



(b) Firm size, manufacturing sector

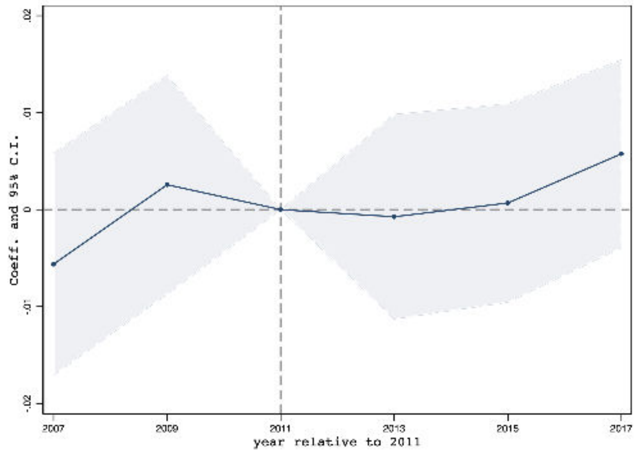


(c) Firm size, service sector



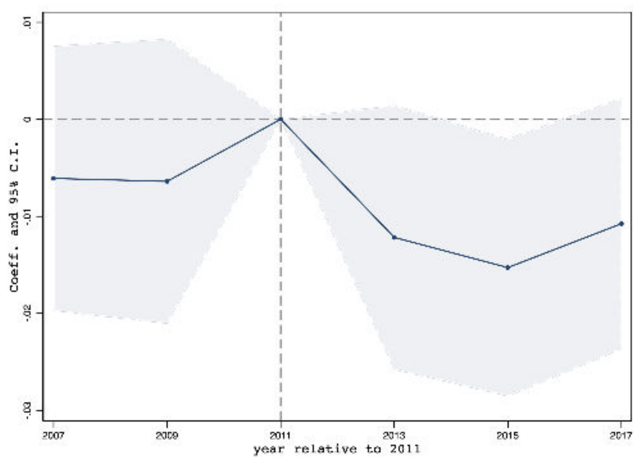
Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2) according to size of the firm where the worker os employed and its economic sector. The dependent variable is a dummy variable that takes value of 1 if individual i has attended on-the-job training in the last 12 months.

Figure 14: Event-study estimates



Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2). The dependent variable is a dummy variable that takes value of 1 if individual i, conditional on training participation in the last 12 months, has paid for it.

Figure 15: Event-study estimates



Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2). The dependent variable is a dummy variable that takes value of 1 if training activities are firm/employer-sponsored.

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