Later life human capital investment*

Evidence from the unintended effects of a pension reform

Simone Chinetti†

This draft: December 15, 2021

Abstract

This paper provides a novel empirical test of human capital theory by studying whether increases in residual working life induce additional training. By exploiting a sizable Italian pension reform, in a Difference-in-Differences setting, I find that a longer working horizon increases human capital investment. Additionally, I show that the response to the reform appears very heterogeneous and depends on gender, age profiles, education, marital status, sector of employment and firm size. However, my estimates suggest that these individual-level positive effects are not due 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.*

†University of Salerno, Department of Economics and Statistics, Via Giovanni Paolo II, 132 - 84084 Fisciano (SA) - Italy. E-mail: schinetti@unisa.it. I wish to thank for the insightful comments and suggestions Antonio Acconcia, Massimiliano Bratti, Elsa Fornero, Marco Pagano, Annalisa Tirozzi and seminar participants at the 36^{th} National Conference of Labour Economics, AIEL. The author gratefully acknowledges financial support from PRIN no. 2017KHR4MB-003.

1 Introduction

Many European countries face a sustained increase in the average age of their working population, raising the need to favour middle-aged workers' skills updating and lifelong learning (European Commission (2007))[1]. According to standard human capital theory, older workers are significantly less likely to be involved in on-the-job-training programs than relatively younger colleagues because the returns on such investments are disproportionately lower for older employees. Indeed, these returns crucially should be expected to depend on the time left before retirement (Ben-Porath (1967); Becker (1962))[2].

However, an open empirical question is whether pension policies that increase the minimum retirement age can stimulate middle-aged workers' training investments. Indeed, a tightening of the minimum age and contribution requirements crucially alter a given individual's retirement probability by directly increasing the length of his residual working horizon. As predicted by the theory, a (positive) variation in the distance to retirement affects training benefits, given that it widens the investment payback period (Blinder and Weiss (1976)).

In this paper, I exploit the *Fornero* pension reform introduced in Italy at the end of 2011 as a source of quasi-experimental variation to assess the (unintended) causal effect of a longer residual working life on middle-aged employees' human capital investments. 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' residual working life. Indeed, Italy and the *Fornero* reform represent an ideal framework to assess the impact of pension reforms increasing minimum retirement requirements on older workers training for several reasons. First, Italy has one of the oldest populations among advanced economies, well above the OECD and the EU averages, and low labour market participation at older ages. Second, the *Fornero* 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). Third, 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. Fourth, soon after its approval, a prolonged and inflamed public debate occurred, implying that the majority of the population understood (or at least were

[1] A skilled and educated workforce represents a critical factor for improving competitiveness, productivity, employments and economic development and growth (Acemoglu and Pischke (1999); Acemoglu and Pischke (1998); Evans et al. (1998); Mankiw et al. (1992); Lucas (1988); Romer (1987)).

[2] Also, early retirement institutions, human capital depreciation (Neuman and Weiss (1995)), and lower senior employees' learning ability and flexibility contribute to reducing their investments in training (Cunha and Heckman (2007); Heckman (2000)).

aware of the consequences brought by) the policy.

In order to provide causal evidence, I rely on a Difference-in-Differences approach where my treatment variable is given by a time-invariant measure of policy-induced shock. I construct a measure of exposure to the pension reform at the individual level by relying on the difference of the Minimum Retirement Age (MRA) in 2017, that is, the post-reform period, and 2011, the pre-reform period. Hence, the variation in MRAs provides the size of the reform-induced shock that mirrors the lengthen of the employees' residual working life, relative to the previous requirements in place before the *Fornero* 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*).

Individual-level data on labour market histories and human capital investments are drawn from the Italian Institute of Public Policy Analysis' (INAPP) Participation, Labour and Unemployment Survey (PLUS). I consider the survey's waves from 2007 to 2017 and a sample of individuals aged between 40 and 64 years with at least 10 and less than 40 years of accrued years of contribution, eligible to retire neither before nor after the 2011 pension reform. I develop this empirical test of human capital theory predictions by looking at three different outcomes. The first is the probability that individual i has attended seminars, conferences, training courses or professional refresher courses. Hence, I focus on activities that human capital theory defines as formal on-the-job training[3]. Then, I extend the empirical analysis by looking at other two outcomes that have not been investigated in the literature. Given that mandated positive variations in MRA translate in an increase in the payout period of the investment, older workers may find it profitable to increase their stock of knowledge by directly investing in it to bargain a higher wage. Hence, I test whether middle-aged workers' willingness to invest directly (and so pay for it) in human capital changed after the reform. Finally, although I cannot directly observe in the data whether firms directly finance training, I explore the role of firms in inducing their middle-aged workers to participate in training programs. I rely on a specific question asking whether the employer has strongly recommended the worker to attend or sponsored the training activity (without implying that the firm or the employer paid for it)[4]. Indeed, similar individual-level human capital

[3] In general, human capital refers to both formal training (formally organized activities such as apprenticeships, workshops, and courses) and informal training (learning-by-doing or work experience). While the Mincer (1962) definition of on-the-job training includes both types of activities, Arrow (1962), instead, highlights with more preponderance the importance of learning-by-doing. Furthermore, training can also be distinguished in general and specific training. The former represents skills that can be used at many other firms and are portable across companies as individuals change jobs, whereas the latter is by definition only valuable to the firm providing the training. However, 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.

[4] It has to be said that for the empirical test of the human capital effect, the information on whether the firms directly finance training is not required as I focus on the effect of the lengthening of working life on

effects can be also be found in a model of firms' investment: when the working life of employees increases, if workers are not perfectly mobile, overall firms' investment in human capital increases too (Acemoglu and Pischke (1999); Acemoglu and Pischke (1998); finding confirmed by empirical evidence: Berton et al. (2018); Berton et al. (2017); Quaranta and Ricci (2017)).

According to my estimates, I find that the causal effect of a longer working horizon due to the *Fornero* pension reform has a positive effect on human capital investment. For each additional year increase in MRA, an individual's probability of investing in human capital goes up to about 0.7 p.p. (1.7 percent when re-scaled in terms of the sample mean). However, the response to the reform is very heterogeneous and mainly driven by men (0.9 p.p. for each additional year or re-scaled in terms of sample average about 2.5 percent) and married women (1.3 percentage points). Furthermore, looking at the age profile of individuals, I find that increases in human capital investment occur only for those workers known as prime-aged (both men and women) and middle-aged (only men).

In terms of sector of employment and firms' economic sector of activity, I find that the positive effect on human capital investment comes from self-employed individuals (1.5 p.p. or to about a 4 percent increase when compared to the sample mean), from those who are employed in the service sector (0.8 p.p.) and from those employed in very small-sized firms (1.8 p.p.) operating in the service sector (2%).

In addition, my evidence supports the theoretical hypothesis that education and human capital investments are compliments: the higher the education level, the higher the training probability (particularly men and married women).

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 in the lengthening of 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 activities. However, only a few papers use individual-level data and assume an endogenous process of human capital investment[5]. These papers usually exploit (partial) pension reforms showing that an increase in the working life has a positive effect on human capital accumulation (Bauer and Eichenberger (2017); Fan et al. (2017); Battistin et al. (2012)). Similarly, Fouarge and Schils (2009) show that generous

training investment and not on the incidence of the human capital investment at the firm level.

[5]Fan et al. (2017) relying on a structural model show that curtailing pension benefits increases human capital accumulation, providing empirical evidence that data do not support the assumption of exogenous human capital process in many theoretical models.

early retirement options significantly reduced older worker human capital accumulation, or that, instead, pension reforms aimed at curtailing early retirement benefits can induce workers in increasing their stock of human capital (Montizaan et al. (2010)). However, several other papers reached opposite conclusions finding that training incidence decreases with age (Bassanini et al. (2005); De Grip and Van Loo (2002)).

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), channels not yet well understood. With regards to health-related outcomes, Bertoni et al. (2018) find that a reform-driven postponement of minimum retirement age has a positive effect on the (self-reported) health of affected individuals. Concerning labour market institutions, despite the limited empirical evidence, economic reasoning suggests that higher employment protection should favour firms-provided training. On this issue, Bratti et al. (2021) find that its reduction increases firm-provided training, whereas Messe and Rouland (2014) show that higher EPL does not affect older workers' training. With regards to productivity, Acemoglu and Restrepo (2017) find that an increase in the share of older workers relative to middle-aged ones is positively associated with the adoption of new technologies with ambiguous effects on overall labour productivity. On the other hand, Carta et al. (2019) find that a higher share of older workers does not harm younger workers' employment growth, leaving labour productivity and unit labour costs unchanged.

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.

Finally, this paper speaks to the strand of the literature that uses mortality rates change to assess variation in human capital accumulation (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 *Fornero* 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 many OECD countries, is characterized by a large first pillar, that is, public pension funds, and by almost marginal second and third pillars, that is compulsory and voluntary private pension funds[6]. Specifically, the main pillar of the Italian public pension system is a compulsory *pay-as-you-go*, meaning that the contributions that workers and companies pay to the Social Security Institute are used to pay the pensions of those who retired. 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

[6] 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.

the old-age or seniority pension requirements. Retiring early, however, comes at the cost of receiving sizably lower pension benefits. The cost of opting for it corresponds, on average, to a 35% reduction of the full pension benefit (Istituto Nazionale di Previdenza Sociale (2016)) given that 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.

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[7] introduced in the Italian pension system the notional defined-contribution (NCD) method, a pension benefits computation that links the lifetime paid contributions to total future pension benefits[8]. 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 pension reform was introduced[9]. The reform, known as

[7]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.

[8]The introduction of the NCD method was motivated by greater freedom of choice of workers regarding their retirement age. This principle of actuarial equity was not conceived in the DB pension benefits' computation. Indeed, the DB scheme pushed individuals to claim pension as soon as possible, as the pension's amount was not a function of the age of the worker at the start of the retirement period.

[9]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.

the *Fornero* reform (Law December 22, 2011, no. 201), 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[10]. The new rules applied to all workers who did not accrue the right to claim either pension by the end of 2011[11,12].

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[13] (see Table 1). The change in the

[10] According to Fondazione Itinerari Previdenziali (2020), after the implementation of the *Fornero* 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 3.

[11] 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.

[12] Finally, the *Fornero* 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.

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

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 *Fornero* 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, the take-up of *Opzione Donna* 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 human capital accumulation activities 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. It contains information on several aspects of the labour market with complete coverage of the Italian population and in particular of all employees.

Crucially for the empirical analysis, the survey provides a specific section that collects all the information regarding human capital investment activities attended by respondents, apart from those connected with formal education. Specifically, individuals are asked if they attended some activity to increase their knowledge and competencies during the last 12 months, such as seminars, conferences, training courses, or professional refresher courses; if they directly paid for attending them and if their employers (usually firms) sponsored the activity (that, however, do not necessarily imply that they paid on behalf of the worker)[14].

[14] There are also some other interesting questions regarding the type of course chosen and the number of hours spent per activity. However, these questions are not included in all of the survey's waves or these are asked in a format that is entirely different from the same question asked in the two years earlier survey. Furthermore, evidence suggests that it is more the incidence of a training spell than its duration that is

Hence, the availability of these data allows me to investigate the causal effect of an increase in the residual working life period on later life human capital investment. Furthermore, the data coupled with precise information on education levels allows me to check whether schooling education correlates with additional investment in human capital. Finally, the richness of these data allows me also to investigate the propensity of individuals in investing directly in (*i.e.* paying for) additional human capital and the role of firms in inducing middle-aged workers to increase or update their knowledge level.

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 *Fornero* pension reform.

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 have data also on their sector of employment and years of accrued contributions that represent a crucial piece of information to support the identifying assumptions and the soundness of the approach regarding the identification of the shock.

Even though the PLUS data has a longitudinal structure, where the panel follows a classic not rotated longitudinal design, the panel component across all the waves taken into account is very short (about less than 3,000 individuals), forcing me to conduct the empirical analysis using repeated cross-sections.

The working sample comprises individual-level data concerning individuals aged 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[15].

Identification of the shock. The reform generated different changes in years until re-

relevant (Pischke (2001)).

[15]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 *Fornero* 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.

tirement 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 becoming 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[16], 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[17]. 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 4 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,+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

[16] 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.

[17] 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.

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 $shock_q = MRA_{2017} - MRA_{2011}$[18]. 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).

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.

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 3 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 *Fornero* reform, her MRA increased, and the size of shock amounts to 6 years, that is, the increase in the residual working life.

[18]Other papers study the effects of the *Fornero* 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 *Fornero* reform, on firms' economic outcomes. Boeri et al. (2017) studies how the reform affected youth unemployment. This paper contributes to their findings by using the *Fornero* reform as a tool to study the human capital investment of middle-aged individuals.

Paola, instead, is a public sector worker with 26 years of paid contributions. Supposing she could have retired under the pre-*Fornero* 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 *Fornero* 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.

Empirical strategy. The *Fornero* 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 has participated to any activity involving human capital accumulation in the last 12 months in year t at the shock level q, then I also look for the propensity of individual i in paying for additional human capital investment and whether firms suggest employees to improve their knowledge; $shock_q$ 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 by starting with some graphical evidence, where I arranged individuals in two groups only for graphical and descriptive evidence purposes. In Figure 5, Panel 5a shows that the declared expected retirement age increases more around the reform (that is from 2011), and 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 with respect to the least affected group. Panel 5b, instead, shows that individuals more exposed to shock expect a lower pension income relative to job earnings, given that for these individuals, the pension benefits share computed according to the NDC method is higher. Overall, trends for both groups followed more or less the same patterns.

With regards to the primary outcome variable of interest, the participation in human capital activities, Figure 6 shows the age profiles of the average participation (Panel 6a) and by three different age classes (40-47, 48-56, 56-64; Panel 6b; that I also exploit in the empirical analysis) by the degree of exposure to the increase in the residual working life. Panel 6a shows that for individuals who experienced an increase in the MRA greater than 3 years, average participation in human capital investment is higher, primarily along with all the age profiles of the individuals included in the sample. This finding is also confirmed by looking at Panel 6b. Indeed, individuals whose shock is bigger than 3 years have higher participation relative to the least shocked ones: a difference of about 6 p.p. between the first age class. Concerning the middle and oldest age class, this difference reduces in size even though most shocked ones still display higher participation. Finally, Figures 7 and 8 shows the average trends in human capital investment according to exposure to the shock and also by gender. Figure 7 shows that individuals most shocked by the change in the minimum retirement age (shock greater than 3 years) display, on average, higher participation rate in activities involving human capital accumulation (seminars, conferences, training courses or professional refresher courses) in the aftermath of the *Fornero* pension reform, whereas in previous years their average participation was essentially the same as those least treated by the reform-induced shock. Looking at gender differences (Figure 8), most treated men after the late-2011 pension reform remarkably increased their human capital accumulation relative to the least treated group, especially in 2013 and 2015. On the other hand, women, independently of the size of the reform-induced shock, had more or less the same average participation rate.

Concerning the other two outcomes I consider in the empirical analysis, Figure 9a plots the probability of individual i, who has attended some human capital accumulation activity, in paying for it. Figure 9b shows the probability that human capital activities are directly sponsored (but not necessarily paid) by the firm[19]. Most affected individuals pay more often for participating in training activities, even though I cannot detect divergent patterns after the pension reform. Instead, for what concern human capital activities sponsored by firms, most shocked individuals, in the aftermath of the reform, appear less likely to be involved in training being suggested by their firm as if firms encouraged least affected individuals, that absent the reform would have retired, to invest in additional activities.

Finally, in Table 4 I present some descriptive statistics of the working sample. The first 3 columns regard all the survey waves taken into consideration, whereas the last 3 refer to the pre-reform waves. Furthermore, I differentiate each period by considering all the individuals in the sample and distinguishing between those most treated (*i.e.*, shock greater than 3 years) 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) and the shares of private-sector employee (considerably higher for least treated individuals) and self-employed individuals (greater for most exposed to changes in MRA).

[19]The sample includes individual who work or has worked for a firm and individuals who attended some training activities.

4 Results

Does an increase in the residual working life induce additional human capital investment?[20] As explained in Section 1, human capital theory predictions state that the value of human capital investment increases with the payout period of the investment. 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.

Table 5 reports the results obtained from estimating equation (1) on the main outcome of interest outlined in Section 3, that is, the probability that individual i has participated in any activity involving human capital accumulation in the last 12 months in year t at the shock level q. In addition to baseline results involving all the individuals included in my sample (column (1) of Table 5), I also conduct a sample-split analysis by gender (columns (2)-(3) of Table 5), both because men and women have different MRA shocks and because they tend to have heterogeneous labour market performances.

I find that the causal effect of an increase in the length of the residual working life, an increase in the minimum retirement age, has a positive effect on human capital investment. Concerning all the individuals included in the sample without distinguishing by gender (see column (1) of Table 5), the estimates suggest that if the length of the working life increases by one year, the probability of participating in activities aimed at improving human capital increases by 0.7 percentage points (statistically significant at the 1 percentage level). When evaluated at the sample mean of the dependent variable, the previous estimate translates into average training participation of about 1.7 percentage points. Instead, the gender-split analysis reveals that the effect is driven only by the response of men. For this group, an increase of 1 year in their residual working life implies a 0.9 p.p., or 2.5 percent in terms of the sample mean, increase in human capital activities participation. For what concerns women, despite a positive coefficient, it is not statistically different from zero. 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.

As a first heterogeneity exercise, I consider different age classes by looking at the response of human capital investment of individuals that more or less find themselves in the later part

[20]In Appendix A, I present additional results not discussed in this Section, based on an alternative definition of the treatment variable. Despite the coefficients measuring the causal effect of interest changing their interpretation, these additional results align with the presented evidence. However, the overall effects, that is, the coefficients re-scaled in terms of sample averages, are 3-4 times larger than those obtained using the variation in the MRA as treatment variable.

of their working life. In Table 6 I report the results for this exercise where columns (1), (2) and (3) report results for individuals aged 40-47, 48-56 and 57-64, respectively. The upper panel of the Table refers to all the individuals, whereas the last two to men and women, respectively. The first striking result is that, independently of the gender, the oldest individuals, those included in the age class 57-64, do not display any evidence of increased human capital investment due to the reform. Secondly, again, women of all age classes do not attend further activities connected with human capital investment. Instead, I find a positive and statistically significant effect for age classes 40-47 and 48-56 (first panel of Table 6). For the former class, an increase of 1 year in the residual working life increases the probability of additional human capital investment of about 1.3 p.p.; instead, the latter class increases about 0.7 percentage points. In terms of the sample mean, the previous estimates correspond to an average increase for each additional year of 3.6 and 1.9 p.p., respectively. Again, the gender-split exercise reveals that the whole variation is driven by men from 40-47 and 48-56 age classes. Youngest men, expecting at least one year increase in their working life, increase their participation in human capital activities of about less than 1.5 p.p. (3.9 percent in terms of the sample average for men), whereas those included in the age class 48-56 of about 1.1 percentage points, that corresponds to an average increase of 3%.

Furthermore, in Table 7 I look for the causal effect of an increase in the residual working life on human capital investment by splitting the sample according to the sectors in which the individual works, that is, public, private or whether the individual is self-employed. This splitting is motivated by the fact that these 3 different broad employment sectors may require their workers to update their knowledge and competencies to a different degree and extent. Usually, investment in additional human capital may be lower in the public sectors, given that the procedures that public employees accomplish are often standardized and may change very little over time. On the other hand, private-sector workers and self-employed tend to be exposed to working environments that are more constantly and rapidly changing. Columns (1), (2) and (3) of Table 7 refer to public workers, private sector employees and self-employed individuals, respectively. The only statistically significant effect comes from individuals working as self-employed for whom an increase of 1 year in their residual working life implies an increase in the probability of human capital investment of about 1.5 percentage points (statistically significant at the 5 percent level), or, in other words, to about a 4 percent increase when compared to the sample mean. The coefficients of interest are positive but not statistically at the conventional confidence level for public and private sector employees.

Finally, I perform another heterogeneity split-sample analysis by considering only private-sector and self-employed workers and distinguishing them according to the NACE code of the firms where they are employed. Specifically, I define two broad firm sectors based on the economic sectors' statistical code: the manufacturing and service sectors. The results, available in Table 8, show that, despite a positive coefficient for both groups of workers, only workers whose firms belong to the service sector increased (see column (2) of Table 8), at the conventional statistical level, their probability of training. In particular, for each additional 1-year increase in residual working life, service sector employees increase their probability of participating in human capital activities by about 0.8 percentage points (2.1 p.p. in terms of the sample subgroup mean).

Does human capital investment correlate with education level? Human capital theory suggests that the education level is likely to affect the worker's training probability (Griliches (1997)). Theory argues that workers with higher human capital levels tend to accumulate more skills and knowledge than individuals with lower education endowments, advocating that formal education and human capital investments are compliments. Henceforth, theory suggests a positive correlation between education and training participation. To check whether data support this theoretical prediction, I re-estimate equation (1) separately for three different education level groups, that is low (middle schools or lower), medium (high school) and high (bachelor or higher). Table 9 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 whole 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 (see the first panel, column (3)). For them, a 1-year increase in the residual working life due to the pension reform implies an increase in the probability of human capital accumulation of 1.4 p.p. or to a 2.3 p.p. average sample increase, suggesting that the higher the education level, the higher the propensity of training activities as predicted by theory. However, the complementarity between education level and human capital emerges when looking at the sample of men. Indeed, for this group, the coefficient measuring the causal effect of interest is positive for all the education levels considered, and it is also increasing in magnitude the higher the education endowment of the individuals, although with different statistical significance. While for low-educated affected men, the coefficient of interest is positive (about 0.6 percentage points) but not statistically significant at the conventional significance level, statistical significance, instead, is found for medium-educated (0.7 p.p. for each additional year at the 10 percent level) and high-educated (1.7 p.p. for each additional year at the 1 percent level) individuals. In terms of the sub-sample means, these estimates imply an average increase in the training participation rate of 2 and 2.3 percentage points for medium and high educated individuals, respectively. On the other hand, the positive correlation predicted by theory seems less clear-cut for the sample of women, even though those with higher education have for each additional year increase in their residual life an increase in the probability of attending human capital activities by about 1.5 percentage points (2.5 p.p. increase in terms of the sample mean).

Further women heterogeneity? As discussed so far, women do not modify their probabilities of attending training activities differently from men in the aftermath of the 2011 pension reform. However, this result is in striking contrast with the theoretical predictions about the lengthening of the residual working life horizon. Indeed, women expect to stay longer in the labour market, given that they are the most affected from the *Fornero* reform. In this short section, I focus on a factor that may influence women decision in investing in human capital activities. To carry out this further heterogeneity exercise, I split the sample of women into married and not married. In other words, I distinguish between female individuals that, in principle and according to solid empirical evidence, may be defined as more "family focused" (those who are married) and as more "career oriented" (those, instead, who are not married)[21]. Indeed, according to the gender and family economics literature (see, among many others, Goodpaster (2010); Leigh (2010); Munasinghe et al. (2008)), married women experience higher opportunity costs in terms of work and investments due to the household chores burden they shoulder. Therefore, they may be less willing or time-constrained in investing in additional human capital. However, an extension of the period they have to stay active in the labour market may provide married women higher incentives to invest in human capital as opposed to more "career focused" women.

To check this issue, I re-estimate the previous heterogeneity sample-spit exercises as well as the baseline specification (that in Table 5, column (3)), taking into account that the response of married women may be different from that of women that can be defined as more "career focused". Table 10 reports the results of re-estimating column (3) of Table 5 by distinguishing between married (columns (1), (2)) and not married women (columns (3), (4)). According to these estimates, there exists a different response to the pension reform given the marital status of the women. As reported by column (1) of Table 10, for a 1-year increase in the residual working life, married women increase their human capital investment probability by about 1.3 percentage points, translating into an average increase of about 3.6 p.p. if compared to the sample mean. In addition, the magnitude of the effect is the same when I also control for the number of kids and household size to take into account for family chores. On the contrary, for those women whose marital status is different from being married, the causal effect is negative, very close to 0 and not statistically significant.

Then, I re-estimated the results of Table 6 following the same reasoning above. For

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

what concerns women, Table 6 shows that independently of the age class taken into account, the estimated causal effects were not statistically different from 0. In Table 11 I show that indeed, again, married women in their 40s (up to 47 years, those that in the labour economics literature are known as prime-aged individuals) increased their probability of investing in human capital. For each additional year increase, the probability goes up by about 2.5 p.p. (that is a 6.8 percent increase with respect to the sample average for this sub-sample of women); for the 48-56 and 57-64 age classes, the coefficients of interest are positive, decreasing in magnitude, but not statistically significant at the conventional confidence levels. For what concerns not married women, all the estimated coefficients are not statistically different from zero and show a magnitude that decreases as age increases.

The last heterogeneity exercises involving women and their marital status concerns the relationship between education endowment and investment in human capital. In Table 9 it is shown that the positive correlation relationship between education, gender and investment in human capital was less clear-cut and supported by data for women rather than for men. This finding is again confirmed by looking at not married women (the second panel) in Table 12. Regarding married women (the first panel), despite none of the estimated coefficients being statistically significant, the positive relationship emerges: the higher the education level, the higher the probability of attending training activities.

Propensity to spend in additional human capital investment and the role of firms. Firms usually invest in their workers' human capital to enhance employees productivity and their growth prospects. However, after a careful cost-benefit analysis, they provide training only if productivity improvements outweigh the costs. Furthermore, provided that productivity returns from training are increasing in training more rapidly than wage returns (as usually happens in imperfect labour markets), firms will be willing to sustain the costs. Although I do not observe whether training is financed and provided directly by firms, I can gauge some evidence by looking at indirect proxies for firms involvement in middle-age workers training participation. I start exploring the role of firms in inducing their workers in investing in human capital by looking at the probability that individual i, affected by the 2011 pension reform, participated in human capital activities by the size of the firm at which she is employed. The results of this further heterogeneity check are available in Table 13, 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 an average response, in mean terms, of about 7 percentage points. As a further check, I also distinguish individuals by the firm's size and for two broad firm economic sectors: the manufacturing and service sectors. The results are available in Table 14 where the first panel is devoted to the manufacturing sector and the second one to firms operating in the service sector. For what concerns the manufacturing sector, individuals working in medium-sized firms (26-49 employees) saw a sizable increase in the probability of attending training activities, about to 4.8 p.p. 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 increased their probability of investing in human capital for a 1-year increase in the residual working life of about 2 percentage points, or about 8.7 p.p. when evaluated at the sample average (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, that is, the probability the firm sponsored the human capital activity and whether the individual directly financed her training. These results are available in Table 15, where the first 3 columns are devoted to the willingness of the affected individual in paying for his human capital investment, whereas the last column to the firm-sponsorship of the activity. Concerning the willingness to pay, I cannot find a statistically significant effect, even if I distinguish individuals according to the yearly median earnings of the sample, as a proxy for the individual budget constraint. For this outcome, the estimated coefficients are positive but not statistically significant at the conventional levels. On the other hand, for what concerns the probability that the employer sponsor the worker the training activity, I find that for one year increase in the residual life, this probability goes down by about 0.8 percentage points (-1.6 p.p. when evaluated at the sample average).

Parallel trend assumption. As standard for the estimation of Difference-in-Difference models, I need to show that the trends in human capital investment 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 participation in human capital activities of individuals 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:

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

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). Figure 10 reports the point estimates for φ_τ in equation (2) and 95% confidence intervals regarding the main outcome of interest referred to all the individuals included in the sample (that is this is the dynamic version of the estimate reported in column (1) of Table 5). As shown by the Figure, the coefficients relative to the pre-reform period are all close to 0 and not statistically significant, suggesting that individuals were on a parallel trend. In contrast, those relative from the post-reform period are positive and turn out to be statistically different from 0 from 2015 onward. Figure 11, instead, replicates Figure 10 splitting the sample according to gender. In this case, while for men the event-study confirms the common trend assumption during the pre-reform years and a strong and significant effect on the probability of human capital investment in the aftermath of the reform, women seem not to be perfectly on parallel trends before the *Fornero* reform. Figure 12 reports the event-study estimates relative to the probability of investing in human capital activities according to age classes (panel a), sector of employment (panel b), education (panel c) and economic sector of the firm where the individual is employed (panel d). The visual inspection of each sub-sample coefficients $\{\gamma_\tau\}_{2007}^{2011}$ shows that they were substantially on a parallel trend, excepts one relative to individuals employed in firms operating in the manufacturing sector. 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).

In Figure 13 are plotted the coefficient relative to the test of the parallel trend assumption considering only the sample of women and distinguishing them according to their marital status (married or not married, panel a) and by age classes (panel b, c) and education (panels d, e). 3 out of 5 figures clearly show that the considered sub-sample of women was on a parallel trend before implementing the reform. In contrast, in panels b and d, some of the estimates coefficients were statistically significant (at the 10%), suggesting that the parallel trend assumption holds weaker than the previous cases. Furthermore, most of the post-reform coefficients show very flat dynamics over time apart from the fact that they are never statistically significant at the conventional level.

Figure 14 shows the event-study estimates of the last heterogeneity exercise regarding firm size (panel a) and the economic sector, that is, the manufacturing (panel b) and service sector (panel c). The visual inspection of the coefficients does not evidence statistical significance during the pre-reform years, despite some of them being above 0. In the aftermath of the reform emerges the statistical significance of the coefficients associated with firms whose size is between 1-9 employees (panel 14a) and operating in the service sector (panel 14c).

Finally, Figures 15 and 16 graph the estimates relative to equation (2) and use as outcomes the probability in paying for human capital activities and the probability that the firm, where the worker is employed, sponsor the training activity, respectively. Regarding the pre-reform years, there is evidence of a parallel trend in both cases, given that the estimated coefficients are never statistically significant, despite being different from zero. Concerning Figure 15, in the aftermath of the pension reform, the probability that the individual pays for human capital investment has been very close to zero up to 2015 and slightly increasing in 2017. However, the post-reform coefficients are never statistically significant at the conventional confidence level. With regards to Figure 16, instead, during the post-*Fornero* reform years, there has been a decrease in the probability that the employer sponsors the training activity to the middle-aged workers, even though statistically significant only in 2015.

5 Conclusions

In this paper, I provide causal evidence for the theory of human capital accumulation. The standard prediction from human capital theory is that older workers are less likely to be involved in training activities than younger colleagues since senior workers and their employers have only a limited time to recoup the investment in skill before retirement occurs.

However, an open empirical question is whether pension policies that exogenously change the working life horizon by increasing the payout period for the human capital investment can stimulate additional training activities.

Specifically, I exploit a sizable 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 is given by an individual-level time-invariant measure of policy-induced shock, that is, the variation in *pre* and *post* MRA, that mirrors the lengthen of the employees' residual working life.

According to my estimates, I find that the causal effect of an increase in the length of the residual working life due to the *Fornero* pension reform has a positive effect on human capital investment. For each additional year increase in MRA, an individual's probability of investing in human capital goes up by about 0.7 percentage points. However, the response to the reform is very heterogeneous and mainly driven by men and married women. Furthermore, looking at the age profile of individuals, I find that increases in human capital investment occur only for those workers known as prime-aged (both men and women, that is, individuals aged 40 to 47) and middle-aged (only men, those aged between 48 and 56). In terms of sector of employment and firms' economic sector of activity, I find that the positive effect on human capital investment comes from self-employed individuals and those employed in (small-sized) firms operating in the service sector. Furthermore, my estimates provide evidence in support of the hypothesis of complementarity between education attainment and investment in human capital, given that individuals with higher education have a higher probability of investing in human capital. Finally, my estimates suggest to rule out that the positive variations in human capital investment in the aftermath of the reform are directly sponsored by employers.

Apart from being a novel test of human capital theory, this evidence may enrich the policy debates about pension policies, which usually do not consider human capital dynamics. My results suggest 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 J.-S. Pischke. Why do firms train? theory and evidence. *The Quarterly Journal of Economics*, 113(1):79–119, 1998.

D. Acemoglu and J.-S. Pischke. Beyond Becker: Training in Imperfect Labour Markets. *The Economic Journal*, 109(453):112–142, 1999.

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.

K. Arrow. The Economic Implications of Learning by Doing. *Review of Economic Studies*, 29(3):155–173, 1962.

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.

F. Berton, D. Guarascio, and A. Ricci. Increasing retirement age, workplace training and labor market outcomes. Evidence from italian firms. *INAPP WP (mimeo)*, 2018.

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.

F. Cunha and J. Heckman. The Technology of Skill Formation. *American Economic Review*, 97(2):31–47, 2007.

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.

European Commission. Strengthening Continuing Vocational Training at the Initiative of the Enterprise. *Employment in Europe 2007, Office for official publications of the European Communities, Luxembourg.*, 2007.

G. W Evans, S. Honkapohja, and P. Romer. Growth Cycles. *American Economic Review*, 88(3):495–515, 1998.

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.

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.

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.

J. J. Heckman. Policies to foster human capital. *Research in Economics*, 54(1):3 – 56, 2000.

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.

R. E. Lucas. On the mechanics of economic development. *Journal of Monetary Economics*, 22(1):3 – 42, 1988.

N. G. Mankiw, D. Romer, and D. N. Weil. A Contribution to the Empirics of Economic Growth. *The Quarterly Journal of Economics*, 107(2):407–437, 1992.

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.

S. Neuman and A. Weiss. On the effects of schooling vintage on experience-earnings profiles: Theory and evidence. *European Economic Review*, 39(5):943–955, 1995.

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.

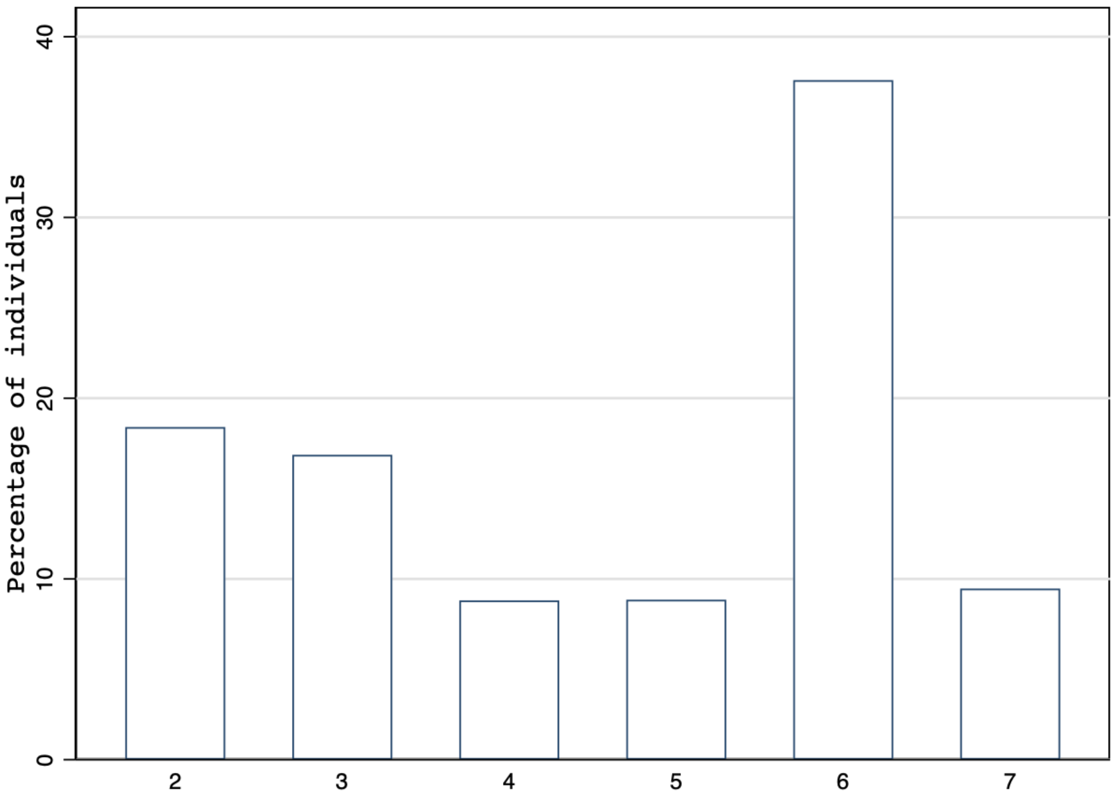
R. Quaranta and A. Ricci. Politiche di assunzione e struttura dell'occupazione: il ruolo della riforma delle pensioni. *Economia & Lavoro*, (3):173–188, 2017.

P. M. Romer. Growth Based on Increasing Returns Due to Specialization. *American Economic Review*, 77(2):56–62, 1987.

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

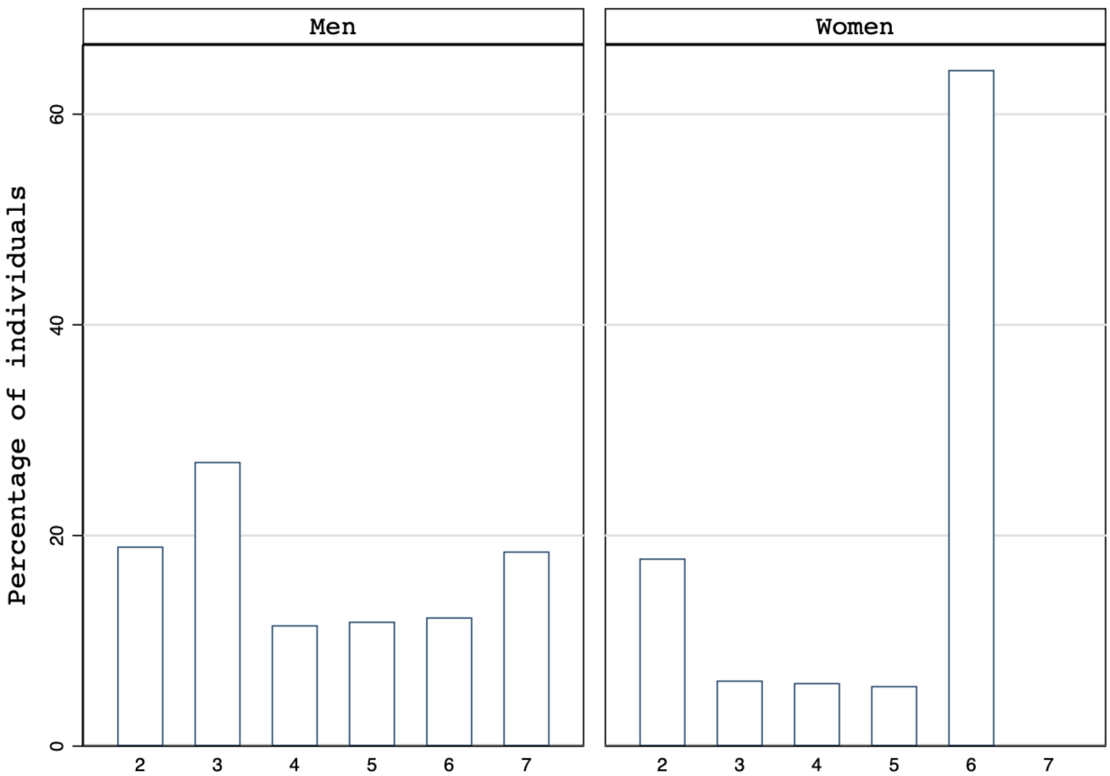
Figures and Tables

Figure 1: Shock distribution in the length of the "residual" working horizon (variation in pension rules between 2017 and 2011)



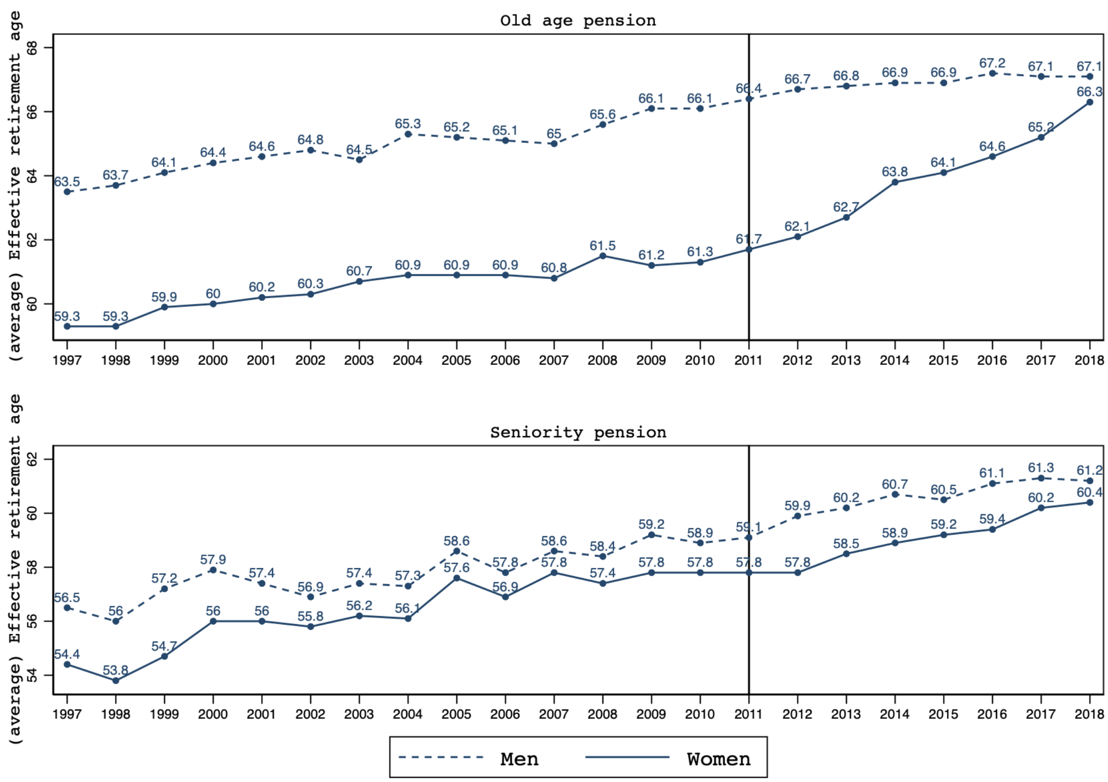
Source: PLUS (INAPP) 2007-2017.
Notes: The figure displays the distribution of the reform-induced shock to the "residual" working horizon. It shows the distribution of the difference between the minimum retirement age (MRA, the age at which individuals can claim their first pension benefit, either old age or seniority) under the post reform pension rules (2017) and the MRA under the pre-reform rules (2011). 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. Data are at the individual level, the y-axis reports the percentage of individuals for any given value of shock.

Figure 2: Shock distribution in the length of the "residual" working horizon by gender (variation in pension rules between 2017 and 2011)



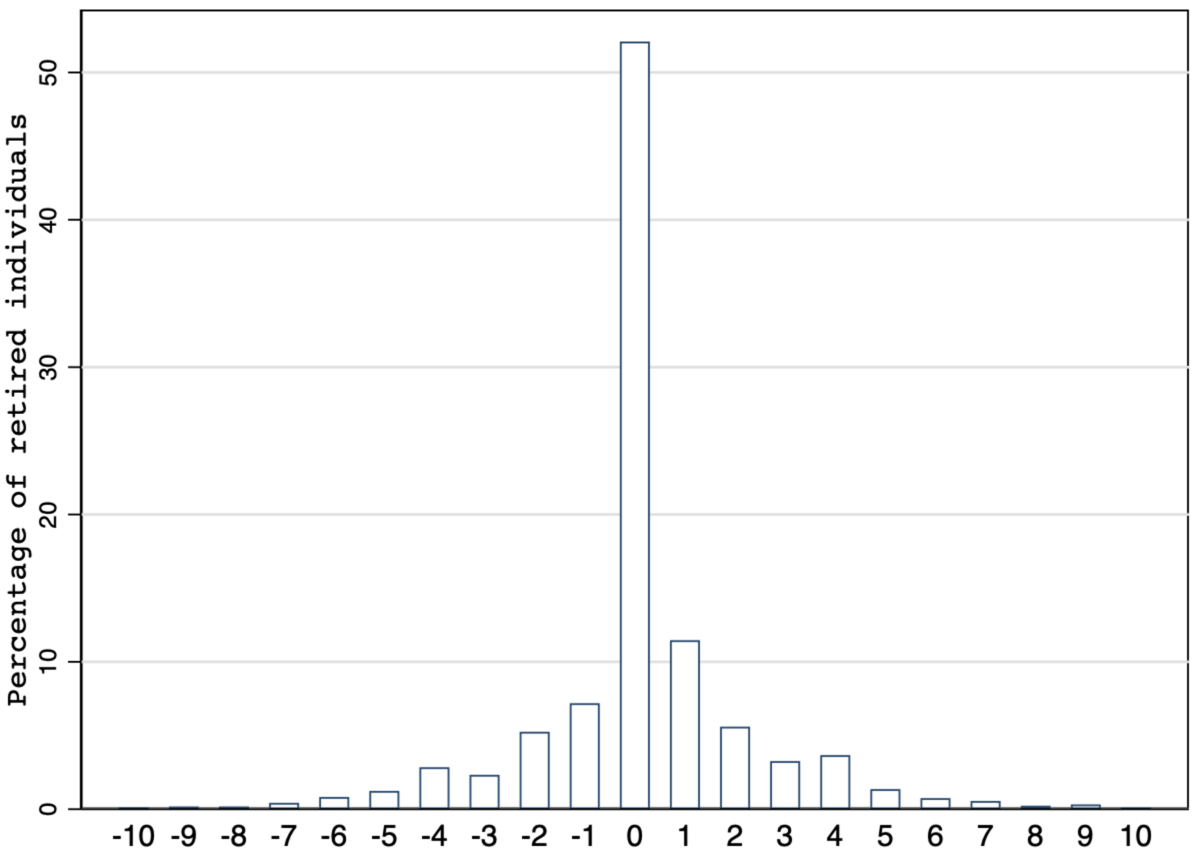
Source: PLUS (INAPP) 2007-2017.
Notes: The figure displays the distribution of the reform-induced shock to the "residual" working horizon, as in Figure 1, distinguishing by gender.

Figure 3: Effective (average) retirement age by gender and pension regime



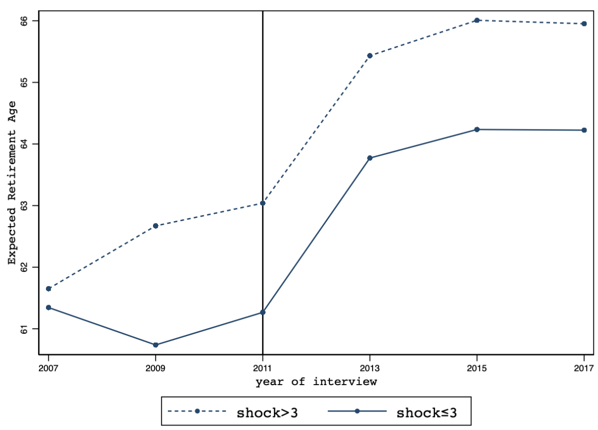
Source: Fondazione Itinerari Previdenziali (2020) based on social security records.

Figure 4: Percentage of individuals retired as function of distance to retirement (MRA_q - Retirement age)

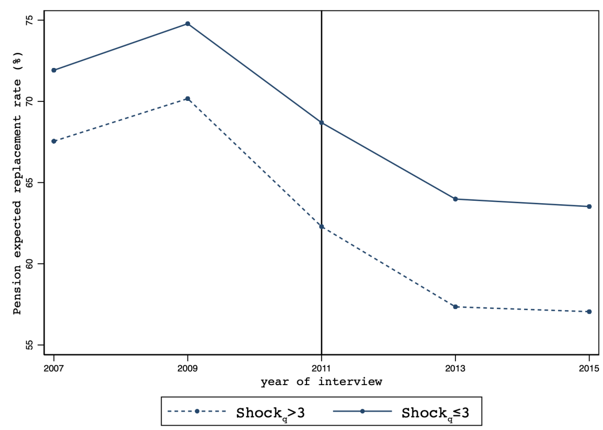


Source: Plus (INAPP) 2005-2015.
Notes: The figure plots the percentage of individuals who declare themselves as retired as a function of the distance to the minimum retirement age (MRA, the age at which individuals can claim their first pension benefit, either old age or seniority according to their gender and sector of employment). The sample of retired individuals is composed solely of those who entered in retirement between 2005 and 2015. Distance to MRA is the difference between the minimum retirement age according to the rules in place at the year of retirement and the individual's age at retirement. The Figure shows that individuals actually retire when they reach their MRA, *i.e.* when their distance to retirement approaches 0.

Figure 5: Declared expected retirement age and replacement pension income rate by exposure to the policy shock



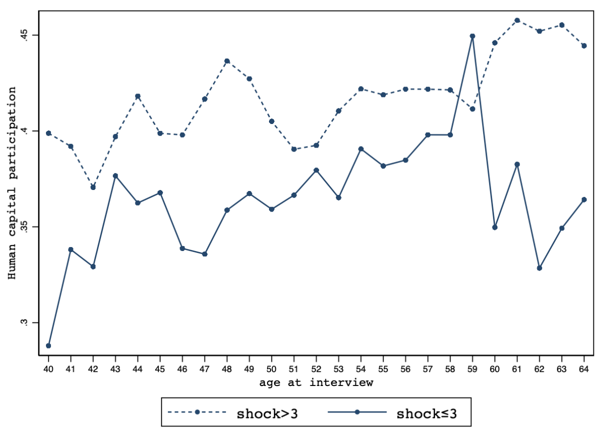
(a) Expected retirement age



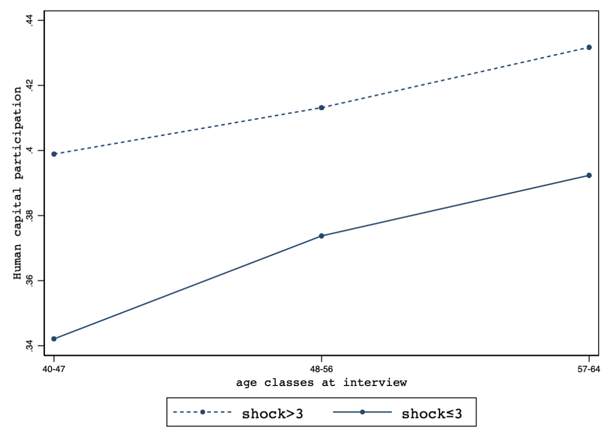
(b) Expected replacement rate

Source: Plus (INAPP) 2007-2017.
Notes: Panel **(a)** shows that the declared expected retirement age increases more around after the reform (that is from 2011) for individuals more exposed to the change in the minimum retirement age (most treated; *i.e.* $Shock_q > 3$). Panel **(b)**, instead, shows that individuals more exposed to shock expect a lower of pension income relative to job earnings. 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. The question on expected retirement age and expected replacement rate (not available in the 2017 wave) are asked only to individuals who have been employed at least once during their life.

Figure 6: Age profiles of later life human capital accumulation by most and least treated



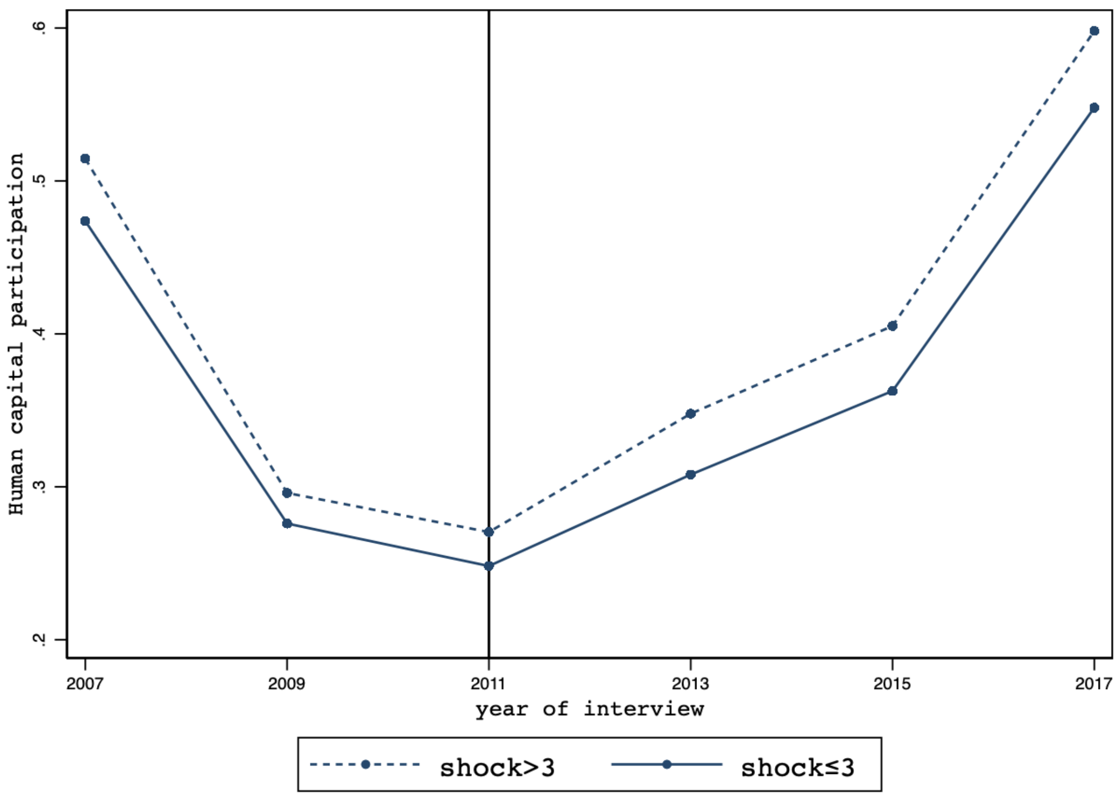
(a) Age profiles



(b) Age classes

Source: Plus (INAPP) 2007-2017.
Notes: The figures show human capital participation activity rate across ages (Panel **(a)**) included in the sample and across three age classes (Panel **(b)**). Individuals most shocked by the change in the minimum retirement age display, on average, higher participation rate in activities involving human capital accumulation, such as: seminars, conferences, training courses or professional refresher courses.

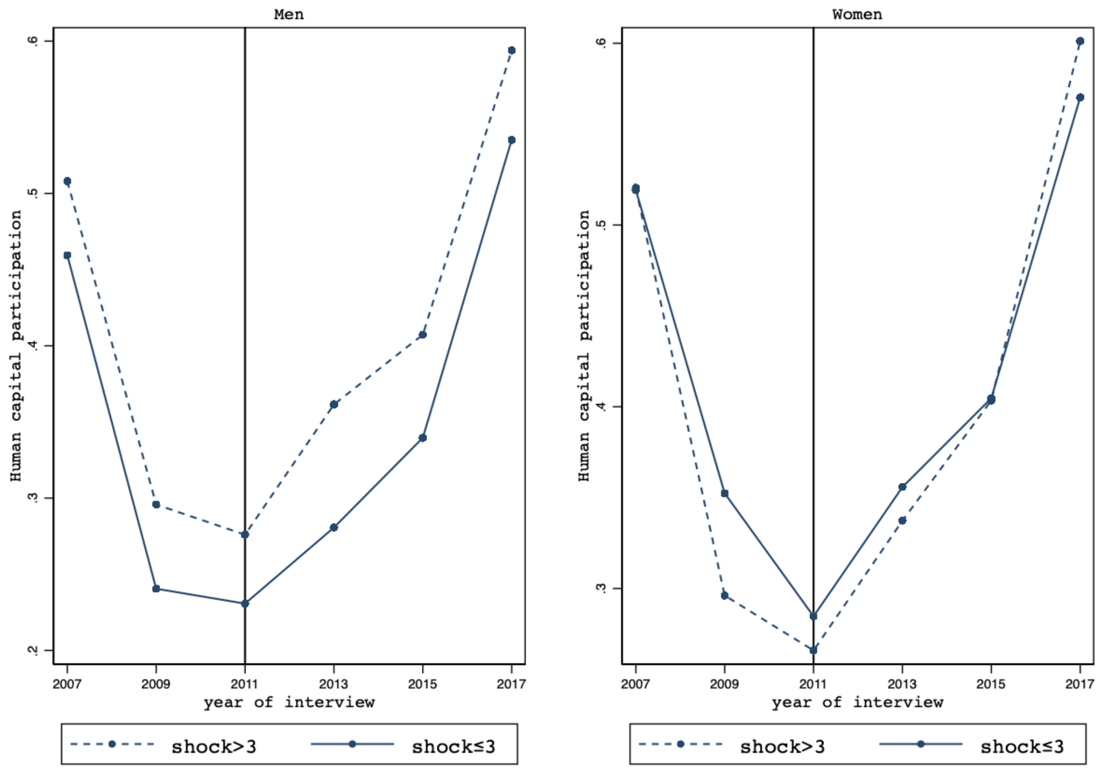
Figure 7: Human capital accumulation by most and least treated



Source: Plus (INAPP) 2007-2017.

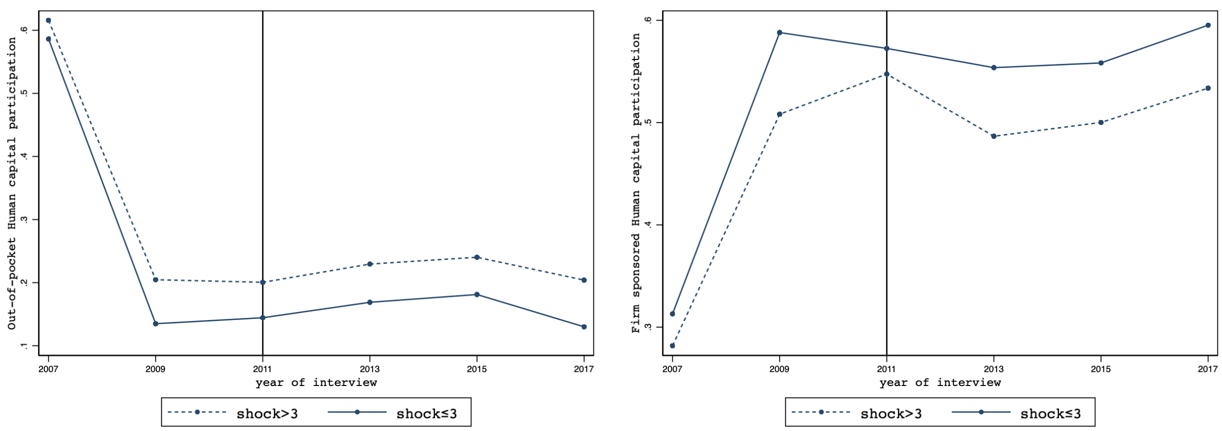
Notes: The figure shows that individuals most shocked by the change in the minimum retirement age ($Shock_q > 3$) display, on average, higher participation rate in activities involving human capital accumulation (seminars, conferences, training courses or professional refresher courses) in the aftermath of the *Fornero* pension reform, whereas in previous years their average participation was essentially the same as those least treated by the reform-induced shock.

Figure 8: Human capital accumulation by gender and most and least treated



Source: Plus (INAPP) 2007-2017.
Notes: The figure replicates Figure 7 distinguishing by gender. Women independently of the size of the reform-induce shock display, more or less, the same average participation rate. On the other hand, most treated men after the late-2011 pension reform remarkably increased their human capital accumulation relative to the least treated group.

Figure 9: Paid and firm-sponsored later life human capital accumulation by most and least treated

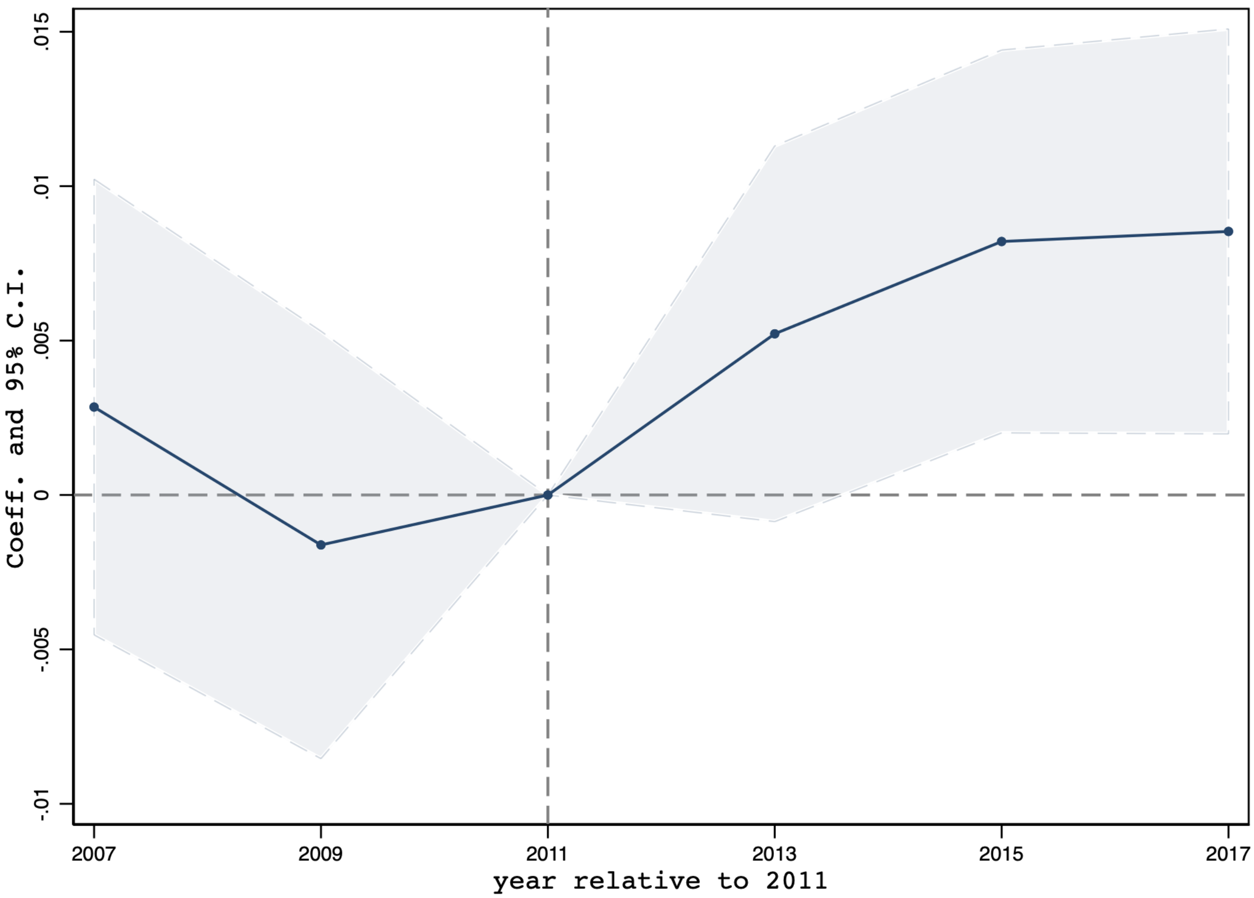


(a) Out-of-pocket HCA

(b) Firm sponsored HCA

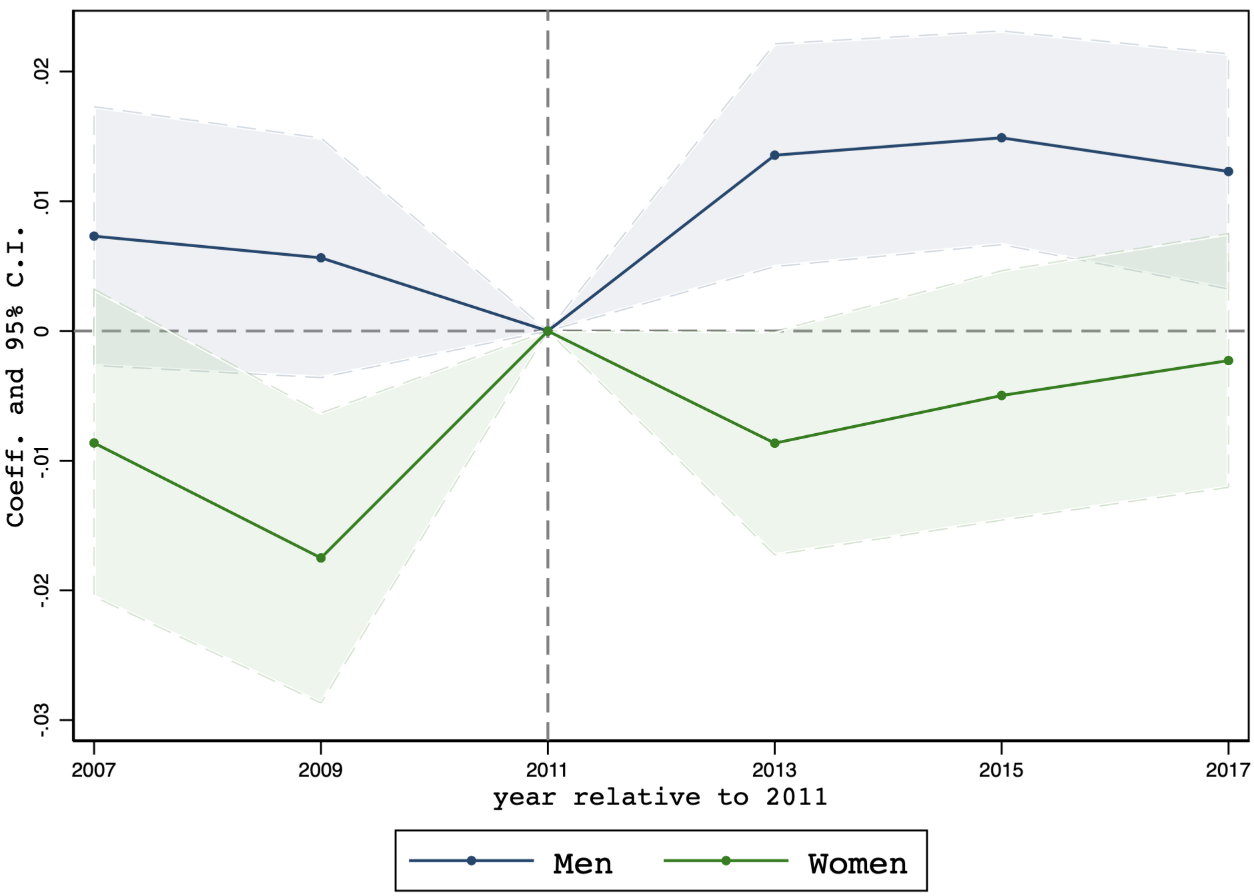
Source: Plus (INAPP) 2007-2017.
Notes: Panel **(a)** plot the probability of individual i, who has attended some kind of human capital accumulation activity, in paying for it. Panel **(b)**, instead, show the probability that human capital activities are directly sponsored (but not necessarily paid) by the firm; and the sample, apart from being composed of individuals who attended some training activities, includes individuals who work for a firm. Most affected individuals pay more often for taking part in training activities and are less sponsored by firms.

Figure 10: 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 human capital activities in the last 12 months.

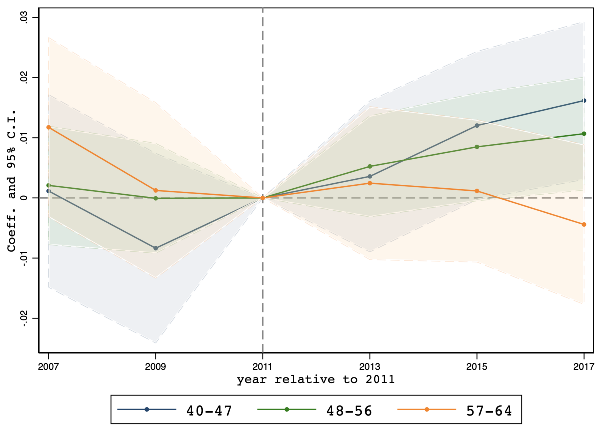
Figure 11: Event-study estimates by gender



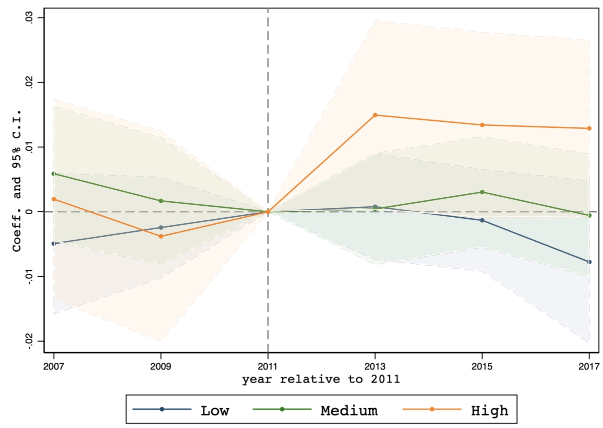
Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2) distinguishing the sample by gender. The dependent variable is a dummy variable that takes value of 1 if individual i has attended human capital activities in the last 12 months.

Figure 12: Event-study estimates by:

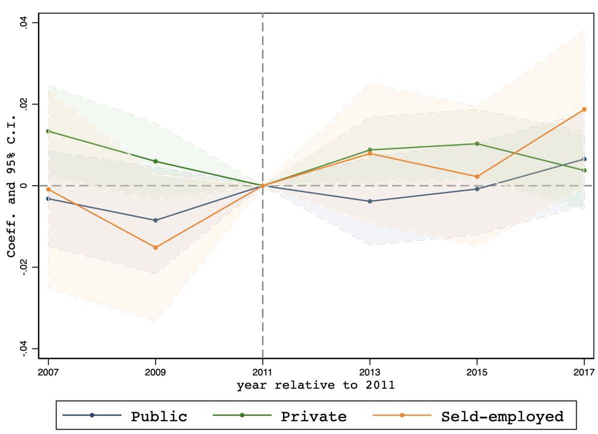
(a) Age classes



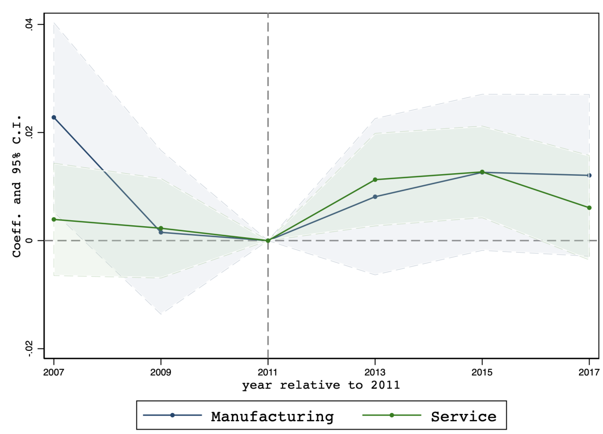
(c) Education



(b) Sector of employment



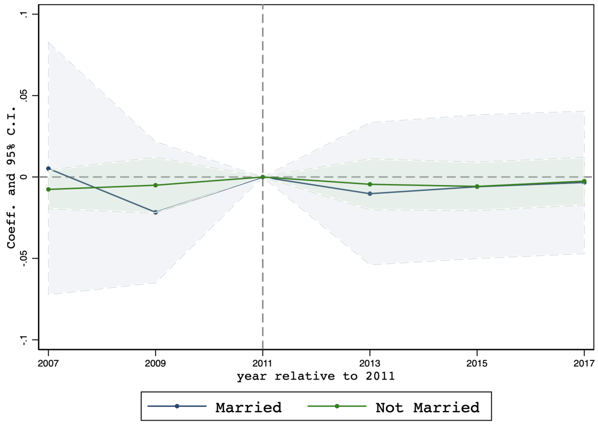
(d) Economic sector of the firm



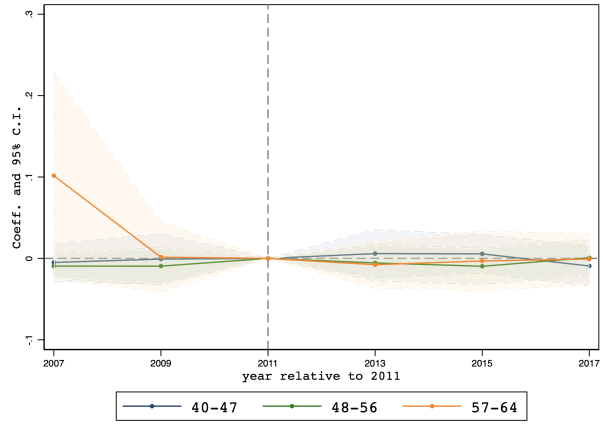
Source: PLUS (INAPP) 2007-2017.
Notes: Estimates based on equation (2) distinguishing by each sub-sample. The dependent variable is a dummy variable that takes value of 1 if individual i has attended human capital activities in the last 12 months.

Figure 13: Event-study estimates, women only, by:

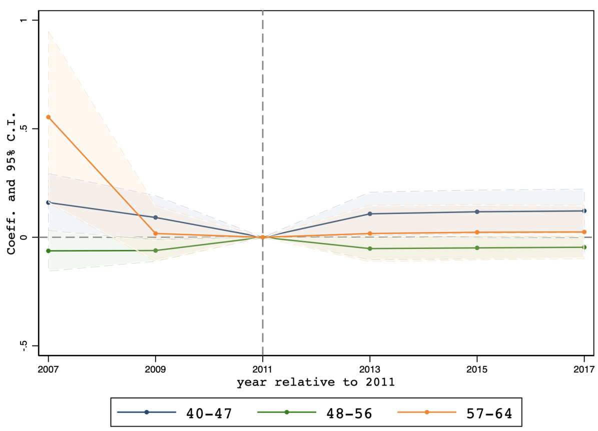
(a) Martial status



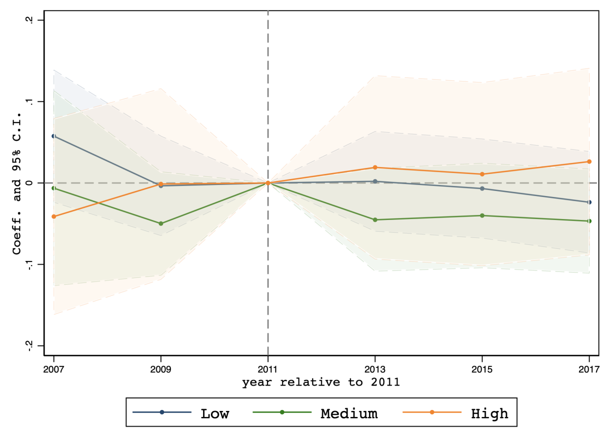
(c) Not married and age classes



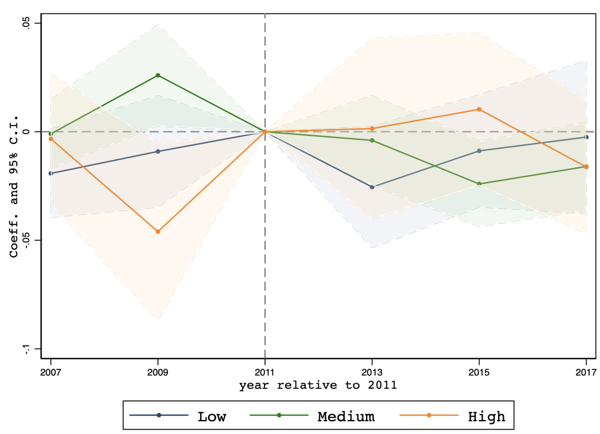
(b) Married and age classes



(d) Married and education



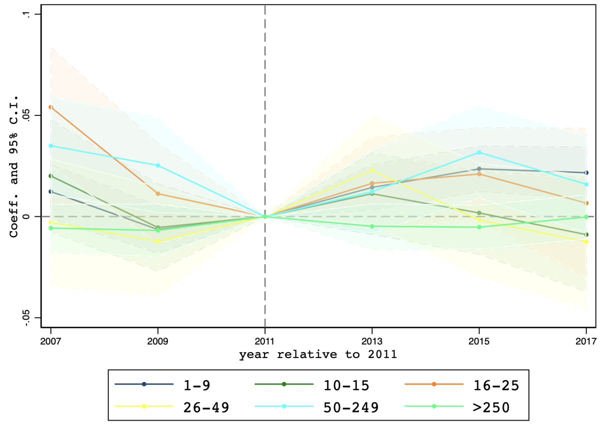
(e) Not married and education



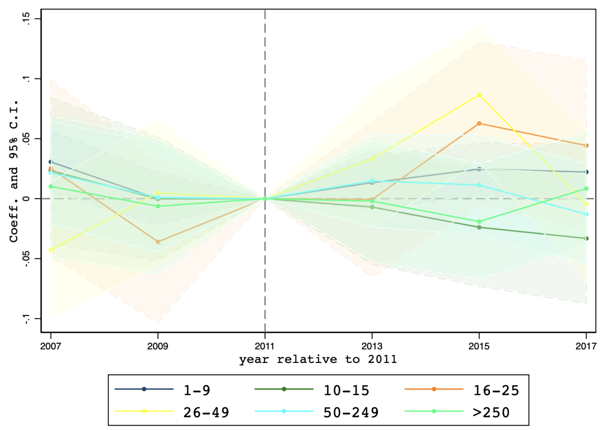
Source: Plus (INAPP) 2007-2017.
Notes: Estimates based on equation (2) considering only the sample of women and distinguishing them according to their martial status (married or not married) and by age classes and education level. The dependent variable is a dummy variable that takes value of 1 if individual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

Figure 14: Event-study estimates by:

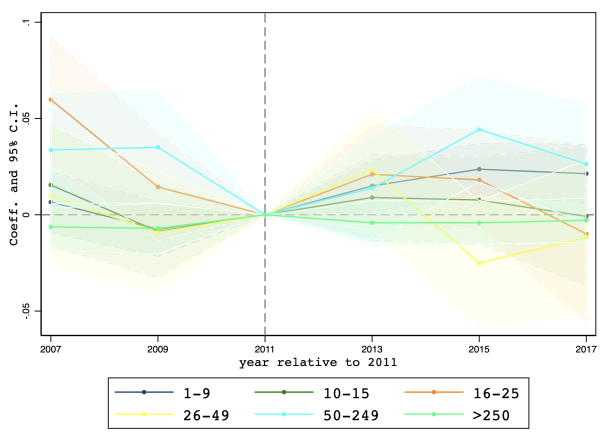
(a) Firm size



(b) Firm size, manufacturing sector

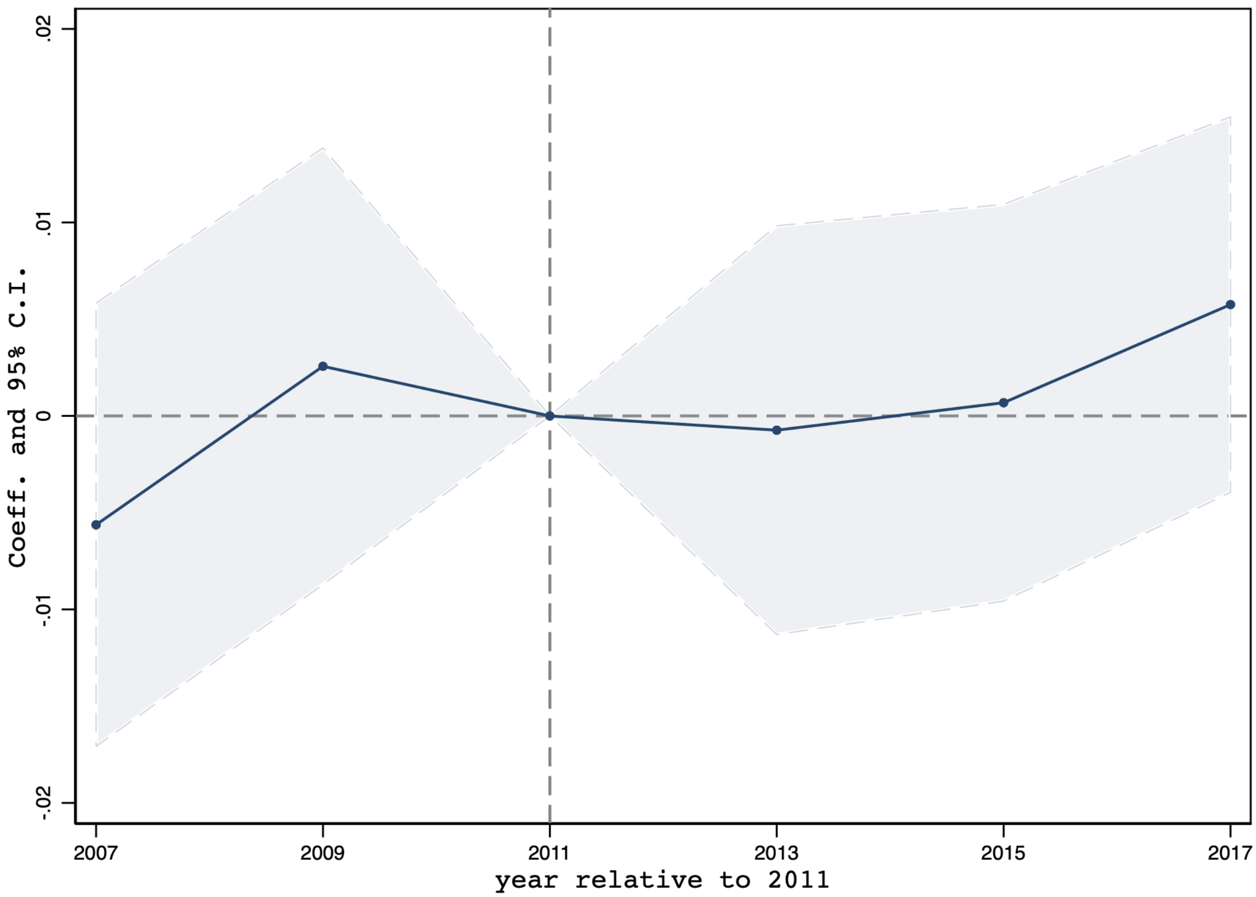


(c) Firm size, service sector



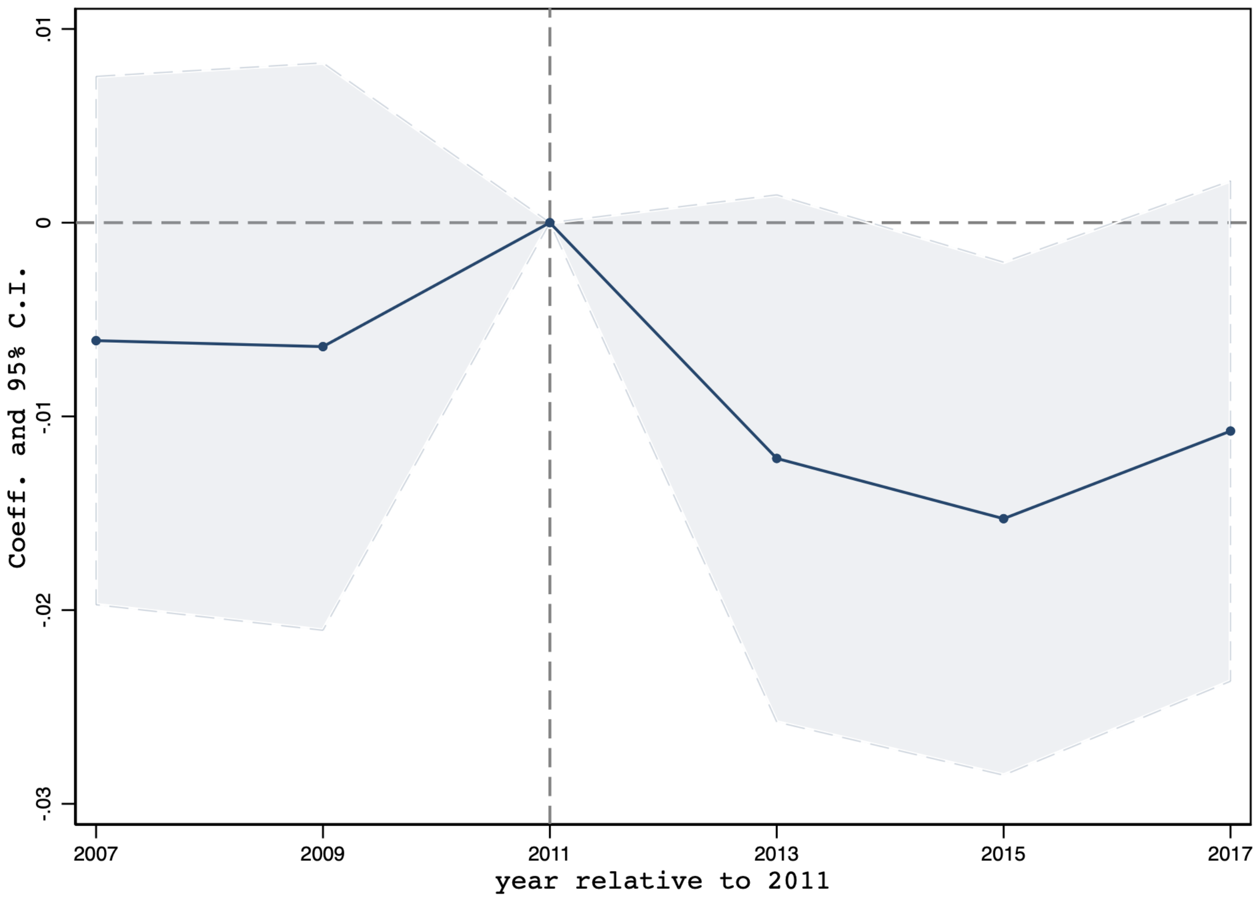
Source: Plus (INAPP) 2007-2017.
Notes: Estimates based on equation (2) according to size of the firm where the worker os employed and its economic sector. The dependent variable is a dummy variable that takes value of 1 if individual i has attended human capital activities in the last 12 months.

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 individual i, conditional on having invested in human capital in the last 12 months, has paid for it.

Figure 16: 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 the human capital activity has been sponsored by the firm/employer.

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: 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 *Fornero* 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.

Table 4: Descriptive statistics

	2007-2015:			2007-2011 (pre-reform period):		
	All	$Shock_q > 3$ (most treated)	$Shock_q \leq 3$ (least treated)	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 5: *Forward-looking* effect on human capital participation activities

	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
R^2	0.1314	0.1043	0.1750
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 6: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.1373	0.1353	0.1293
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
R^2	0.1181	0.1070	0.1109
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
R^2	0.1728	0.1831	0.1738
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 7: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.0793	0.0729	0.0876
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 8: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.0766	0.0861
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 9: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.0726	0.1083	0.0769
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
R^2	0.0722	0.0840	0.0772
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
R^2	0.1011	0.1548	0.0916
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 10: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.1633	0.1640	0.1990	0.1993
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 11: *Forward-looking* effect on human capital participation activities, women only, by

	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
R^2	0.1644	0.1645	0.1693	0.1703	0.1682	0.1708
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
R^2	0.2080	0.2088	0.2112	0.2114	0.1977	0.1992
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 12: *Forward-looking* effect on human capital participation activities, women only, by:

	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
R^2	0.1007	0.1011	0.1492	0.1508	0.0772	0.0772
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
R^2	0.1320	0.1332	0.1777	0.1778	0.1391	0.1391
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 13: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.1012	0.0906	0.0944	0.1130	0.0767	0.0828
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 14: *Forward-looking* effect on human capital participation activities by:

	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
R^2	0.1218	0.1465	0.2300	0.2216	0.1207	0.1015
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
R^2	0.1039	0.1077	0.1134	0.1428	0.0909	0.0938
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 15: *Forward-looking* effect on human capital participation activities by:

	Paid			Firm-sponsored
	All:	Wage above median:	Wage below 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
R^2	0.2081	0.2175	0.2036	0.0938
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: $^{+}$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

A Alternative treatment definition - ONLINE APPENDIX

In this Section, I re-estimate the findings discussed in Section 4, based on equation (1), changing the way I define the treatment variable. Specifically, the previous results were based on a time-invariant measure of exposure to the pension reform mirroring the difference between the MRA in 2017 and 2011, the number of years of increase in the residual working life. Now, I change the treatment definition using as treatment variable a binary indicator that takes value of 1 if affected individual i has experienced more than 3 years of increase in her MRA (that is $shock > 3$), and 0 otherwise—consequently, the interpretation of the Difference-in-Differences coefficient changes. Indeed, the β coefficient in equation (1), given by the interaction of the policy-induced shock measure and the post-reform dummy, measured the average human capital investment for each additional year increase in the MRA, exclusively depending on their degree of exposure to the policy, around its implementation. Now it measures the average difference in human capital investment, in the aftermath of the reform, between those more exposed to the increase in MRA (those with shock greater than 3 years) relative to the control group, composed of individuals whose shock in the "residual" working life is lower or equal to three years.

Table 16: *Forward-looking* effect on human capital participation activities

	All	Men	Women
	(1)	(2)	(3)
$S_i \times post2011$	0.0205*	0.0273*	0.0114
	(0.0086)	(0.0116)	(0.0132)
μ D.V. $S_i = 1$, $post2011 = 0$	0.3591	0.3591	
Coeff. rescaled	+5.7%	+7.6%	
Obs.	53,977	28,478	25,499
R^2	0.1313	0.1042	0.1750
Adj. R^2	0.1299	0.1014	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 17: *Forward-looking* effect on human capital participation activities by:

	Age class:		
	40-47	48-56	57-64
	(1)	(2)	(3)
		All:	
$S_i \times post2011$	0.0319^{+}	0.0282^{*}	-0.0317
	(0.0171)	(0.0115)	(0.0194)
μ D.V. $S_i = 1$, $post2011 = 0$	0.3368	0.3711	
Coeff. rescaled	+9.5%	+7.6%	
Obs.	13,600	27,289	13,088
R^2	0.1370	0.1353	0.1294
Adj. R^2	0.1330	0.1330	0.1246
		Men:	
$S_i \times post2011$	0.0290	0.0420^{**}	-0.0291
	(0.0246)	(0.0157)	(0.0249)
μ D.V. $S_i = 1$, $post2011 = 0$		0.3681	
Coeff. rescaled		+11.4%	
Obs.	6,103	14,703	7,672
R^2	0.1177	0.1071	0.1111
Adj. R^2	0.1084	0.1026	0.1027
		Women:	
$S_i \times post2011$	0.0243	0.0102	-0.0139
	(0.0259)	(0.0174)	(0.0328)
Obs.	7,497	12,586	5,416
R^2	0.1727	0.1831	0.1738
Adj. R^2	0.1657	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 18: *Forward-looking* effect on human capital participation activities by:

	Sector of employment:		
	Public	Private	Self-employed
	(1)	(2)	(3)
$S_i \times post2011$	0.0113	0.0104	0.0598*
	(0.0152)	(0.0112)	(0.0257)
μ D.V. $S_i = 1$, $post2011 = 0$			0.2989
Coeff. rescaled			+20%
Obs.	21,113	24,831	8,033
R^2	0.0792	0.0729	0.0875
Adj. R^2	0.0754	0.0696	0.0775

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 19: *Forward-looking* effect on human capital participation activities by:

	Firm's economic sector:	
	Manufacturing	Service
	(1)	(2)
$S_i \times post2011$	0.0246	0.0247*
	(0.0197)	(0.0116)
μ D.V. $S_i = 1$, $post2011 = 0$		0.38
Coeff. rescaled		+6.5%
Obs.	8,059	24,805
R^2	0.0767	0.0860
Adj. R^2	0.0665	0.0828

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 20: *Forward-looking* effect on human capital participation activities by:

	Education level:		
	Low	Medium	High
	(1)	(2)	(3)
		All:	
$S_i \times post2011$	0.0035	-0.0079	0.0452*
	(0.0129)	(0.0119)	(0.0185)
μ D.V. $S_i = 1$, $post2011 = 0$			0.5678
Coeff. rescaled			+8%
Obs.	11,645	27,057	15,275
R^2	0.0726	0.1083	0.0767
Adj. R^2	0.0655	0.1054	0.0713
		Men:	
$S_i \times post2011$	0.0255	0.0172	0.0556*
	(0.0174)	(0.0159)	(0.0243)
μ D.V. $S_i = 1$, $post2011 = 0$			0.5518
Coeff. rescaled			+10%
Obs.	6,694	14,319	7,465
R^2	0.0723	0.0839	0.0768
Adj. R^2	0.0598	0.0782	0.0656
		Women:	
$S_i \times post2011$	-0.0178	-0.0358*	0.0540
	(0.0222)	(0.0179)	(0.0331)
μ D.V. $S_i = 1$, $post2011 = 0$		0.358	
Coeff. rescaled		−10%	
Obs.	4,951	12,738	7,810
R^2	0.1012	0.1547	0.0916
Adj. R^2	0.0849	0.1488	0.0812

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 21: *Forward-looking* effect on human capital participation activities, women only

	Married		Not married	
	(1)	(2)	(3)	(4)
$S_i \times post2011$	0.0510*	0.0508*	-0.0047	-0.0044
	(0.0245)	(0.0244)	(0.0220)	(0.0219)
μ D.V. $S_i = 1$, $post2011 = 0$	0.3074			
Coeff. rescaled	+16.3%	+16.5%		
Obs.	14,991	14,991	10,508	10,508
R^2	0.1633	0.1640	0.1990	0.1993
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) add 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 22: *Forward-looking* effect on human capital participation activities, women only, by:

	Age class:					
	40-47		48-56		57-64	
	(1)	(2)	(3)	(4)	(5)	(6)
	Married:					
$S_i \times post2011$	0.0821$^+$	0.0825$^+$	0.0374	0.0376	0.0248	0.0246
	(0.0496)	(0.0495)	(0.0306)	(0.0304)	(0.0618)	(0.0614)
μ D.V. $S_i = 1$, $post2011 = 0$	0.2771					
Coeff. rescaled	+29%	+29%				
Obs.	4,610	4,610	7,127	7,127	3,254	3,254
R^2	0.1642	0.1643	0.1694	0.1703	0.1682	0.1708
Adj. R^2	0.1534	0.1531	0.1614	0.1621	0.1507	0.1528
	Not married:					
$S_i \times post2011$	0.0086	0.0090	-0.0051	-0.0053	-0.0234	-0.0275
	(0.0344)	(0.0344)	(0.0330)	(0.0330)	(0.0444)	(0.0443)
Obs.	2,887	2,887	5,459	5,459	2,162	2,162
R^2	0.2080	0.2088	0.2112	0.2114	0.1977	0.1992
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) add 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 23: *Forward-looking* effect on human capital participation activities, women only, by:

	Education level:					
	Low		Medium		High	
	(1)	(2)	(3)	(4)	(5)	(6)
	Married:					
$S_i \times post2011$	-0.0217	-0.0207	0.0121	0.0121	0.0944^{+}	0.0944^{+}
	(0.0331)	(0.0332)	(0.0325)	(0.0324)	(0.0505)	(0.0505)
μ D.V. $S_i = 1$, $post2011 = 0$					0.5268	
Coeff. rescaled					+17.9%	
Obs.	2731	2,731	7,585	7,585	4,639	4,639
R^2	0.1008	0.1011	0.1492	0.1508	0.0774	0.0774
Adj. R^2	0.0722	0.0719	0.1397	0.1411	0.0601	0.0601
	Not married:					
$S_i \times post2011$	-0.0212	-0.0219	-0.0501^{+}	-0.0500^{+}	0.0064	0.0064
	(0.0414)	(0.0414)	(0.0284)	(0.0284)	(0.0529)	(0.0529)
Mean $S_i = 1$, $post2011 = 0$			0.3803			
Coeff. rescaled			-13%			
Obs.	2,220	2,220	5,153	5,153	2,453	2,453
R^2	0.1320	0.1332	0.1774	0.1775	0.1390	0.1390
Adj. R^2	0.0966	0.0969	0.1633	0.1630	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) add 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 24: *Forward-looking* effect on human capital participation activities by:

	Firm size:					
	1-9	10-15	16-25	26-49	50-249	>250
	(1)	(2)	(3)	(4)	(5)	(6)
$S_i \times post2011$	0.0562^{***}	0.0144	-0.0158	0.0050	0.0206	-0.0209
	(0.0164)	(0.0304)	(0.0330)	(0.0389)	(0.0293)	(0.0237)
μ D.V. $S_i = 1$, $post2011 = 0$	0.2170					
Coeff. rescaled	+26%					
Obs.	11,975	2,827	2,113	1,864	3,909	8,536
R^2	0.1006	0.0907	0.0944	0.1126	0.0768	0.0828
Adj. R^2	0.0939	0.0614	0.0550	0.0686	0.0556	0.0732

Notes: Controls include: year, shock, gender, age, marital status, region, sector of employment and years of contribution fixed effects. 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. 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 25: *Forward-looking* effect on human capital participation activities by:

	Firm size:					
	1-9	10-15	16-25	26-49	50-249	>250
	(1)	(2)	(3)	(4)	(5)	(6)
	Manufacturing sector:					
$S_i \times post2011$	0.0266	-0.0414	0.1001	0.1259	0.0112	-0.0139
	(0.0321)	(0.0615)	(0.0886)	(0.0779)	(0.0513)	(0.0570)
Obs.	3,063	739	439	560	1,339	1,623
R^2	0.1215	0.1439	0.2294	0.2196	0.1208	0.1015
Adj. R^2	0.0955	0.0280	0.0384	0.0757	0.0589	0.0500
	Service sector:					
$S_i \times post2011$	0.0645***	0.0257	-0.0263	-0.0449	0.0359	-0.0321
	(0.0193)	(0.0352)	(0.0374)	(0.0463)	(0.0364)	(0.0270)
μ D.V. $S_i = 1$, $post2011 = 0$	0.2185					
Coeff. rescaled	+29.5%					
Obs.	8,912	2,088	1,674	1,304	2,570	6,913
R^2	0.1032	0.1078	0.1133	0.1435	0.0912	0.0937
Adj. R^2	0.0943	0.0685	0.0641	0.0814	0.0590	0.0820

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 26: *Forward-looking* effect on:

		Paid		Firm-sponsored
		Wage above	Wage below	
	All:	median:	median:	
	(1)	(2)	(3)	(4)
$S_i \times post2011$	0.0085	0.0125	0.0009	-0.0171
	(0.0124)	(0.0159)	(0.0194)	(0.0149)
Obs.	21,289	13,033	8,256	20,308
R^2	0.2081	0.2174	0.2036	0.0937
Adj. R^2	0.2048	0.2120	0.1949	0.0897

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: $^+$ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$