Pension Reforms, Longer Working Horizons and Absence from Work

by

Giorgio Brunello (Padua)

Maria De Paola (INPS)

Lorenzo Rocco (Padua)

Abstract

Using matched employer-employee for Italy and newly available data on sick leaves certificates, we study the effect of an exogenous increase in the length of the work horizon – triggered by a pension reform that increases minimum retirement age - on the probability that middle-aged employees take sick leaves. We find that this effect is positive for females and negative for males. After excluding health as a plausible mechanism, we argue that the intertemporal substitution of leisure prevailed on the fear of job loss for females, while the opposite happened for males. We show that the effect of a longer working horizon on firm productivity that operates by changing sick leaves is negligible.

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Introduction

To cope with the challenges posed by aging, many OECD countries have implemented reforms to increase the sustainability of their pension and social security systems, for instance by raising minimum retirement age. By delaying retirement, these reforms have affected not only individuals who would have retired in the absence of changes, but also middle aged individuals not too close to retirement, who experienced an increase in their expected residual working horizon.

In this paper, we study the effects of an exogenous increase in this horizon – triggered by an important pension reform that increased minimum retirement age - on sick leaves, a major component of absence from work that has been found to negatively affect firm productivity (Grinza et al, 2020).[1] In most countries, sick leaves require a medical certificate from a physician, and should therefore be motivated by poor health conditions. However, it is often difficult for the physician to refuse approval because symptoms are uncertain.[2] Asymmetric information and the coverage provided by sickness benefit programs may induce employees who are fit to work to behave opportunistically and consume leisure by calling in sick or by requesting benefits for longer periods than their health status calls for (Ferman et al, 2021; Moscarola et al, 2016).

[1] Sick leaves could also affect individual careers (Markussen, 2011; Henningsen and Hægeland, 2008)

[2] Examples of health conditions that are difficult to measure include bruise / contusion, headache, nausea, sprain/strain, illness, carpal tunnel disease and emotional/stress/mental disorder. See Butler et al, 2014.

From a theoretical point of view, the effect of an increase in the length of workers' residual working horizon on sick leaves is uncertain. On the one hand, workers may pretend to be sick and compensate for the expected reduction in future leisure with additional current leisure, or consume more leisure because their expected lifetime income increases. On the other hand, senior workers who have to stay with the firm longer because of the higher minimum retirement age may fear to be dismissed, as they are typically costlier than younger workers, and therefore reduce sick leaves to boost or maintain their productivity. Health behaviors may also improve, for instance because workers need to stay healthy to be able to work longer (see Bertoni et al, 2018), with negative effects on sick leaves.

Our empirical study focuses on the Monti-Fornero pension reform, that took place at the end of 2011 in Italy. We combine matched administrative employer-employee data with individual data on sick leave certificates, that are available from 2010. We focus on private sector employees aged 45 to 50, who at the time of the pension reform (end of 2011) had at least five years to go before reaching the minimum retirement age, and follow them for five years after the reform, from 2012 to 2016. Adopting a difference-in-differences strategy, we compare the change of sick leaves before and after the reform for males and females exposed to different increases in their working horizon, measured at the end of 2011.

We find that, following an exogenous increase in the working horizon, both the probability of taking sick leaves and the percent of total working days spent in sick leaves increase for females but decline for males. These effects are driven by relatively short episodes of sick leaves, suggesting that changes in health are unlikely to be important.

A longer working horizon that reduces expected future leisure is likely to increase the current demand for leisure. One way to consume additional leisure is to take opportunistic sick leaves, or leaves that are not motivated by poor health conditions. The cost of this behavior is the expected sanction if caught cheating on health conditions. Since sanctions may include the termination of the employment relationship, the cost is higher for those with a higher wage.

In support of this view, we find that the probability of taking sick leaves increases with the longer working horizon for females with lower than gender-specific median wages, but not for females with higher wages. This result could also be due to an implicit contract that exchanges lower wages with higher employer's tolerance of sick leaves.

Although the substitution of current for future leisure can explain our results for females, it does not account for the fact that males reduce their sick leaves when the working horizon increases. To explain this, we turn to the fear of job loss hypothesis. If wages increase with age faster than productivity in the latter part of a worker's life, a pension reform that raises retirement age above the optimal value reduces profits, forcing employers to either adjust wages, or rise productivity or finally terminate the employment relationship.

The pressure on employers is likely to be higher for males than for females, because males are typically better paid (see Casarico and Lattanzio, 2021) and more likely to be on open ended contracts than females. This higher pressure, and the associated higher fear of job loss, suggest that male workers reduced their sick leaves after the reform to increase their productivity and avoid dismissal. In support of this explanation, we find that a longer working horizon increases the hazard of dismissals and separations for males but not for females.

Overall, our findings can be explained if the intertemporal substitution of leisure – which raises sick leaves – prevails upon the fear of job loss – which reduces them - for females, while the opposite happens for males.

If sick leaves reduce firm productivity, pension reforms that increase the residual working horizon can affect the productivity of firms by increasing average sick leaves. We estimate that adding one year to the working horizon increases the average number of days of sick leave per worker by 0.01. Multiplying this increase by the effect of sick leaves per worker (per year) on firm productivity (-1.9 percent), we obtain that the effect on firm productivity is 0.019 percent, a negligible effect.

Our paper contributes to the literature that studies the effects of a longer working horizon on middle aged individuals, by considering job search behavior and employment (Hairault et al, 2010), labor supply (Carta and De Philippis, 2020; Lalive et al., 2017, Cribb et al. 2016), human capital investment (Montizaan et al, 2016, Brunello and Comi, 2015 and Chinetti, 2019), mortality (Bloemen et al, 2017) and informal care (Fischer and Muller, 2020).[3]

This literature includes the study by Moscarola et al, 2016, who also examine the impact of the Monti-Fornero reform on sick leaves in Italy. This earlier research considers only sick leaves lasting 7 days or longer,

[3] A related literature examines the spillover effects on younger workers of pension reforms affecting older workers (Bianchi et al, 2020; Boeri et al, 2021; Bertoni et al, 2021).

that in 2013 contributed to less than 35 percent of all sickness episodes in the private sector (see Figure 1), looks only at females and at the year immediately after the reform (2012). Compared to this study, we use data on all sickness episodes and consider the impact of the reform on both genders and over a five – years horizon. In addition, while Moscarola et al. focus on women who before the reform were close to retirement (born from 1947 to 1959), we focus on a younger group of individuals (born from 1959 to 1966).

We also contribute to the literatures that examine the impact of a longer working horizon on individual health-promoting behaviors. On the one hand, workers who face a longer working horizon may invest more on their health to keep fit and be able to work longer. Bertoni et al, 2018, use an earlier pension reform in Italy and show that middle-aged males have increased regular exercise, with positive effect on obesity and self-reported satisfaction with health. On the other hand, longer working years can cause additional stress and strain. De Grip et al., 2012, show that a Dutch reform postponing the minimum retirement age of public sector workers reduced their mental health (see also Carrino et al., 2020).

We also provide new evidence on the relationship between absences and productivity (Allen, 1983; Zhang et al, 2017 and Grinza et al, 2020).

Finally, this paper also speaks to the literature that investigates the determinants of absenteeism and emphiasizes the role of working conditions (Osborg Ose, 2005), sickness insurance (De Paola and Scoppa, 2014, Marie and Castillo, 2020), monitoring (D'Amuri, 2017; Boeri et. al, 2021) and employment protection legislation (Scoppa and Vuri, 2014). We add to this list pension reforms increasing working horizons.

The rest of the paper is organized as follows: in Section 1 we present the key features of the Monti-Fornero Reform of 2011, showing how minimum retirement age changed with respect to the pre-reform period, and briefly describe the sick leaves insurance system. Section 2 introduces the data and Section 3 presents the empirical approach. Our results are shown and discussed in Section 4. Section 5 and 6 are devoted to the investigation of the mechanisms and Section 7 studies the effects of sick leaves on firm-level outcomes, including productivity. Conclusions follow.

1. Institutional Background

1.1 The Monti Fornero Reform

Before 1992, the minimum age to access *old-age* pension for Italian males and females was 60 for employees in the private sector and the self-employed, and 65 for public sector employees – conditional on having paid social security contributions for at least 15 years. Earlier retirement with a *seniority* pension was possible at any age for workers who had paid social security contributions for at least 35 and 25 years in the private and public sector respectively (see Angelini et al, 2009). Starting from the early nineties, a sequence of reforms changed eligibility conditions for both old age and seniority pensions.

The Monti-Fornero reform was approved in December 2011 and produced its effects starting from January 2012. It abolished seniority pensions and increased the years of paid contributions required for retirement independently of age. Without seniority pensions, minimum age requirements became those for old age pensions.

Before the reform, the Sacconi Law was in place, which established that a male private sector employee in 2012 could retire either with 60 years of age and 36 years of paid contributions, or with 61 years of age and 35 years of contributions, or finally with 40 years of contributions at any age. In the year immediately after the Monti-Fornero reform, he could retire instead either at age 66 and with 20 years of contributions or with 42 years (plus one month) of contributions at any age. Five years after the reform, he could retire at 66 years and 7 months or with 42 years and 10 months of social security contributions. On the other hand, a female private sector worker, who could have retired before the reform at the same age established for males, in the year after the reform could only retire at 62, or with 41 years and one month of contributions at any age. Five years after the reform, she could retire at 65 and seven months, or with 41 years and 10 months of social security contributions (see Tables A1 and A2 in the Appendix).

1.2 Sick leave insurance system

In Italy, individuals absent from work because of illness need to inform both their employer and their physician, who issues a certificate establishing the type of illness and the period of absence from work. This certificate must be sent by the worker both to the employer and to the National Social Security Institute (INPS). Employees enjoy almost full insurance for the earnings lost due to sick leave. The employer pays the full wage during the first 3 days of absence, while public insurance starts from the fourth day of the sickness spell and covers from one half to two thirds of full pay depending upon absence duration (with a maximum duration cap of 180 days in the same calendar year). Despite the partial coverage offered by public insurance, workers typically end up obtaining close to 100 percent of their wage because almost all Italian labor contracts establish that the uncovered part is paid by the employer.

Employees on sick leave are subject to monitoring, which takes the form of random medical visits during day hours, when employees are required to stay at home. Home visits are inspections made by an INPS physician with the purpose of verifying whether the medical certificate truthfully reports the employee's health conditions. If the employee is found either in good health or not at home, the matter is reported to the employer, who can undertake disciplinary actions, including dismissal in extreme cases.

2. The data

Our analysis is based on a unique dataset provided by the Italian Social Security Institute (INPS), which links several administrative sources. We have access to all the sickness certificates issued to private sector employees from 2010, with information on the starting day and the duration of the sick leave.[4] We merge these data with social security accounts ("*estratti conto*"), which contain the complete working history of a 1/15 sample of all private sector workers, including paid social security contributions, a key ingredient for the computation of the treatment variable, as described below.

We further add annual information on workers' employment status, type of work, occupation, type of contract, voluntary and involuntary separations, and about the sector, size, location of the firm, by using a matched employee-employer dataset that includes all private-sector, non-agricultural firms with at least one salaried employee. Finally, we

[4] Unfortunately, the diagnosed sickness is not disclosed because of privacy restrictions.

merge these data with firms' balance sheets, which have information on value added, employment and fixed assets.

Since individuals can retire in the post-treatment period, changes in the composition of the population under study as time goes by will affect the impact of the reform on sick leaves, as "stayers" may react differently than "leavers". To avoid this, we restrict our working sample to individuals who, at the time of the reform, had an expected working horizon of at least five years and therefore could not retire before 2016 even in the absence of the reform. We also choose the years 2012 to 2016 as the post - treatment period. As certificates are only available from 2010, our pre-treatment period consists of the years 2010 and 2011.

The time invariant treatment variable T – the increase in the working horizon - is the difference between the expected number of months to retirement (MR) after the Monti-Fornero law was introduced and the MR before the reform, when the Sacconi law was in place.[5] To compute the treatment T for each individual, the rules about minimum retirement age established by both laws require information on age, gender and years of paid social security contributions before the reform was implemented, at the end of 2011. As shown in Tables A1 and A2 in the Appendix, both the Sacconi and the Monti-Fornero laws did not set a time invariant minimum retirement age but allowed this minimum to increase over the period 2012 to 2016 (and after) to track the increase of life expectancy.

[5] The change in the working horizon is conditional on continuing employment, an assumption regularly made in this literature. Carta and De Philippis, 2020, compute instead their treatment variable by considering 2010 as the pre-treatment year. By so doing, they evaluate the joint impact of both the Sacconi and the Monti-Fornero reforms.

Therefore, to each individual, the treatment T is time invariant but depends on the year he/she is expected to retire. To illustrate, suppose that individual A is expected to retire in 2012 according to the new law. For this individual, we compute T as the difference between MR in 2012 according to the Monti-Fornero rules, and MR in 2012 according to the Sacconi rules. On the other hand, if individual B is expected to retire in 2015, we compute T as the difference between MR in 2015 as established by the Monti-Fornero and the Sacconi rules.

We exclude from our working sample individuals younger than 45 in 2011, who were too far away from retirement, and individuals older than 50 in the same year, because the share of those with less than five years from retirement according to the Sacconi rules is relatively high (30 percent or higher for males, as shown in in Figures 2 and 3).

We further trim the sample by excluding outliers with less than 10 or more than 50 years of paid social security contributions and those who retired before 2016.[6] We end up with an unbalanced panel covering the period 2011-2016 composed by individuals who were employed in 2011 and spent at least one month per year in employment.

We also define a secondary sample by selecting individuals who were aged 45 to 52 in 2011, with at least 3 years to retirement, that we follow from 2010 to 2014 (three years after the reform). Table 1 shows that, in the main working sample, average years of paid social security contributions in 2011 were 25.94 for men and 24.53 for women; mean age was 47.24 for men and 47.27 for women; mean years to retirement before and after the reform were 11.79 and 14.59 for men and 11.20 and 12.82

[6] In very few cases this happens because of measurement error in the residual working horizon before the reform.

for women; and mean years of treatment were 3.05 for men and 3.15 for women, or 25.8 and 28.1 percent of pre-treatment mean years.

Figures 4 and 5 show, separately for men and women aged 45 to 50 in 2011, the distribution of the treatment rounded in years. We find that more than 60% of men have T=3. On the other hand, close to 50% of females have T=2 and close to 40% have T=5.

3. The empirical setup

We investigate empirically the effect of exogenous changes in the length of the working horizon on sick leaves by comparing individuals with different level of treatment T before and after the pension reform. We use the following difference – in – differences specification

$$Y_{iqt} = \beta_0 + \beta_1 X_{iqt} + \beta_2 T_{iq} * D2010_t + \sum_{y=2}^{6} \beta_{3y} T_{iq} D201y + \gamma_q + \gamma_t + \varepsilon_{iqt} \quad (1)$$

where Y is the outcome, the index i is for the individual, the index t for time and the index q is for the cell defined by age, gender and paid social security contributions in 2011 (rounded in years); D2010 is the pre-treatment dummy for 2010 and D201y is the post-treatment dummy for the year y; γ_q is a cell dummy, γ_t a year dummy and ε the residual error. As in Carta and De Philippis, 2020, the fixed effects γ_q absorb the cross-section variation in the outcome Y that depends on age, gender and years of labor market experience. Since the treatment is defined in months rather than in years, we absorb the residual cross-sectional variation by adding T to the controls in (1).

Finally, the vector X includes the treatment in months, labor market experience in 2011, the occupation in 2011, the type of contract (open ended or not) in 2011, regional, sectoral dummies, the interactions of linear time trends with years to retirement before the reform and the potential number of days worked in 2011. Estimates of pre- and post-treatment effects are relative to the baseline year 2011, and standard errors are clustered by cell q.

4. Results

We use two alternative definitions of the outcome Y: a) the dummy variable DY equal to one if individual *i* in year *t* has at least one day of sick leave, and to zero otherwise; b) the percent PY of potential days of work.

By so doing, we control for the fact that individuals who work more months are more likely to be absent because of sick leaves. Table 2 shows the summary statistics of the variables used in the estimates, by gender. On average, the percent of days lost because of sick leaves is 2.13 for males and 2.37 for females, and the probability of taking at least one day of sick leave in the year is 24.71 for males and 27.44 for females.

Our baseline estimates are reported in Table 3, separately for males and females, using both DY (columns (1) and (2)) and PY (columns (3) and (4)) as outcomes. Supporting the view that individuals with a different value of T were not on different trends before the Monti-Fornero reform was introduced, we always find that the interaction of T with the pre-treatment year 2010 is not statistically significant.

We also find that, while females with a longer reform-induced working horizon take more sick leaves, males tend to take fewer leaves. For females, the positive effect is persistent and statistically significant from 2013 onward. When we consider the last year in the sample, 2016, we estimate that adding one year to the expected working horizon raises both the probability of taking at least a day of sick leave and the percent of days of sick leaves by 1.5 percent.[7]

For men, we find that one additional year of treatment temporarily reduces the percent of days of sick leave by 1.1 percent in 2013 (-0.002 x 12 /2.13, statistically significant at the 5 percent level of confidence). This effect fades away in later years. One more year of T also reduces the probability of taking at least a day of sick leave by 1.1 percent in 2012 (statistically significant at the 5 percent level of confidence) and by 1.45 percent in 2013 (statistically significant at the 1 percent level of confidence). The negative effect persists between 2014 and 2016 but is smaller and not statistically significant.[8]

A potential concern with the estimates in Table 3 is that the pre-treatment period is too short for us to test the assumption of parallel trends. Since the data on private sector certificates are only available from 2010, we turn for further supporting evidence to an alternative dataset, the Italian Labor Force Survey, which asks individuals absent from work whether this absence was due to either sickness or injury. [9] Using the longer data from 2006 to 2016, we test whether the interactions of the treatment T with the year dummies from 2006 to 2010 are jointly different from zero, which would support the assumption of parallel

[7] These effects are computed as (0.034 x 12/27.4) and (0.003 x 12 /2.37) respectively. They are statistically significant at the 1 percent level of confidence.

[8] Table A3 in the Appendix replicates our estimates on the secondary sample and the post-treatment shorter period 2012-14. We confirm that a longer time horizon increases sick leaves for females and has negative but often imprecisely estimated effects on the sick leaves of males.

[9] These data do not have information on actual years of social security contributions. We therefore follow Bertoni et al, 2018, and use potential years of contributions, defined as age minus age at the end of education.

trends.[10] We find that the tests cannot reject the null, both for males (F-test: 0.73, p-value: 0.601) and for females (F-test:0.40, p-value:0.851).

5. Mechanisms

The key finding in Section 4 is that, when time to retirement rises because of a pension reform, the propensity to take sick leaves increases for females but decreases for males. In this sub-section, we discuss candidate explanations of these results.

5.1. Health

One reason why sick leaves vary with the treatment is that a longer working horizon affects individual health. This could happen, for instance, because individuals expecting to work longer try to stay healthy (see Bertoni et al, 2018). The empirical literature has pointed out that while opportunistic sick leaves, which occur without a poor health condition, are typically short-term, leaves driven by poor health have longer duration (see for instance Campolieti, 2004 and Ziebarth, 2013).[11]

If poorer health is driving our results, we should find that a longer working horizon increases long sick leave spells more than short spells. To verify this, we estimate the effect of a longer working horizon on the probability of having at least one short spell – lasting 1 to 9 days – and a long spell (lasting 10 days or more). Table 4 reports our estimates by

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[10] Since the sample size of the LFS is much smaller than in our working sample (131,708 observations for males and 83,939 observations for females), the absence of pre-treatment effects could be due to lack of power. We partly address this problem by testing whether *all* estimated effects between 2006 and 2010 are jointly zero. In the joint test, the null can be rejected even if all year-specific effects are insignificant, since the variables T_{iq}*D20yy for y=06…10 are correlated.

[11] A reason for this is that mild illness, which is usually associated with short leaves, is less verifiable, and therefore more exposed to opportunism, than serious health conditions.

gender. For females, a higher T has no effect on longer spells and positive and statistically significant effects on shorter spells. For males, both types of sick leaves decline with T. While the effect on shorter spells persists over time, the one on longer spells fades away quickly. We conclude from this that health conditions are unlikely to be the story driving our results.[12]

5.2 Demand for leisure

A longer working horizon may increase the demand for leisure as workers try to smooth its consumption over the life cycle. Additional leisure can be obtained opportunistically by taking sick leaves or, alternatively, by choosing work arrangements with fewer hours of work, for instance by switching from full to part-time contacts, from career to bridge jobs (Ruhm, 1990), or by reducing the supply of overtime.

Since sick leaves are paid, the cost of taking these leaves without being sick is the expected sanction if caught cheating on health conditions. Sanctions range from simple admonitions to outright firing and include a loss of reputation which negatively affect career prospects. Although in Italy the probability of being caught cheating on sick leaves is relatively small, because controls can be done only on a sample of all the workers on leave, we expect the penalty in the case of being found cheating to be larger for those with higher wages, who have more to lose in the event of dismissal and are therefore less likely to take sick leaves when the working horizon increases.

[12] Moscarola et al, 2016, consider the effect of the same reform on the weeks of sick leaves taken in 2012 by females with an average age close to 55. They find no statistically significant effect in the full sample, and a positive and statistically significant effect for low income grandmothers.

We also expect the effect of a longer working horizon on opportunistic sick leaves to be smaller for those with higher wages. We investigate whether this is the case by running separate regressions for males and females, distinguishing between workers who earn median or below median gender - specific wages and those who earn above median wages.

For males, we find that sick leaves decline with the treatment, especially for those with wages below the median, suggesting that the substitution of future with current leisure is unlikely to be a key factor at play (see columns (1) and (2) of Table 5). For females, we find instead that the effect of the treatment on the probability of taking sick leaves is positive and statistically significant for those with wages below the median, and small and not statistically different from zero for those with wages above the median, consistently with the lower expected cost of sick leaves (See Table 5, column (3) and (4)).

We would also expect a higher demand for leisure in those geographical characterized by more conservative gender roles where women might need time to support elderly persons in their family[13]. Even if in our analysis we consider women who had at least three years to retirement even before the reform, it could be that the increase in the working horizon has induced them to re-arrange their family organization. For instance, those women who before the reform were intensively relying on external services (or on the care provided by other members of the family) for support to their elderly (however planning to be soon able to take care of them personally), now being afraid of cost and implications

[13] As we are considering individuals aged 45-50 or 45-52, it is unlikely that they have young children or that they are grandfathers.

deriving from the prolonged time to retirement might have increased their demand for leisure. As the obligation to take care of family members might depend on social norms, we have elicited gender stereotypes from the World Value Survey at regional level and run separate regressions for individuals living in areas with more and less conservative gender attitudes. As shown in Table 6, the increase in women sick leave absences is concentrated among those living in areas with traditional gender roles.

A higher demand for leisure can be satisfied also by reducing working hours, for instance by switching from full to part time jobs. We estimate the effects of a longer working horizon on the probability that an individual is employed full-time or part-time and report our findings in Table 7. Although there is evidence that pre-treatment trends are not parallel, which suggests caution in the interpretation of results, the estimated effect of the treatment T on the probability of working part-time is positive for females and negative for males, in line with our results on sick leaves.[14]

5.3 Fear of job loss

Because of the Monti-Fornero reform, senior employees aged 45 to 50 need to stay longer in the labor market to attain the minimum retirement age. Earnings profiles typically increase with labor market experience and tend to overtake productivity profiles in the later part of a worker

[14] These results are somewhat in contrast with those by Carta and De Philippis, 2020, who use Bank of Italy survey data and find that the effect of the treatment on the probability of working part-time is negative for females and absent for males. An important difference with our study is that they do not condition on employment, as we do.

career to compensate for the productivity premium that characterizes the earlier part.

According to Lazear, 1979, from the employer's perspective the optimal retirement age is when the present value of wages is equal to the present value of productivity. If wages increase with age faster than productivity in the latter part of a worker's life, a higher T that raises retirement age above the optimal value reduces profits, forcing employers to either adjust wages, or rise productivity or finally terminate the employment relationship.

The pressure on employers is likely to be higher for males than for females, because the former group is typically better paid and more likely to be on open ended contracts than the latter. This higher pressure, and the higher fear of job loss, could have motivated male workers to increase their productivity, for instance by reducing sick leaves, as shown by the estimates in Table 3.[15]

We verify the "fear of job loss" hypothesis by estimating – separately by gender - the effect of the treatment T on the hazard of being dismissed for employees with an open-ended contract in 2011, whom we follow from 2012 to 2016. We use a Cox survival model, where employment corresponds to survival, and dismissal to failure. Table 8 presents our results, both for dismissals and for separations (dismissals plus quits).[16] Conditional on age and contributions paid until 2011, we find that a longer working horizon increases both dismissals and separations

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[15] Bovini and Paradisi, 2019, show that delayed retirement has increased dismissals in Italy.

[16] We do so because voluntary and involuntary separations are often not distinguishable empirically.

hazards for males but not for females, in line with the view that the threat of job loss is more salient for the former than the latter.[17]

6. Sick leaves and productivity

One route connecting pension reforms that increase the residual working horizon of employees to firm productivity is via the effect on sick leaves. Letting A and Y/L be the average number of days of sick leave per worker and labor productivity in a generic firm, this route is captured by $\frac{\partial A}{\partial T}\frac{\partial Y/L}{\partial A}$, where the first term is the effect of increasing the treatment T on the average days of sick leave per worker and the second term is the effect of the change in A on labor productivity.

To compute $\frac{\partial A}{\partial T}$, we first estimate a version of Eq. (1) which assumes that the post-treatment effect is the same between 2012 and 2016, using the number of days of sick leaves as the dependent variable. We report the results in Table A.4 in the Appendix. Next, we assume that employment consists of two groups, the young Y (aged 18 to 44) and the old O (aged 45 to 50). Let A_y and A_o be the average days of sick leave per private sector employee in each group; A_{fo} and A_{mo} the averages for senior females and males; s_y and s_{fo} the share of junior employees on total employment and the share of female senior employment on total senior employment.

The average days of sick leave per employee and per senior worker in a firm are given by $A = s_y A_y + (1 - s_y)A_o$ and $A_o = s_{fo}A_{fo} + (1 - s_{fo})A_{mo}$ respectively. We assume that

[17] We also control for experience in 2011, foreign status in 2011, days worked in 2011, occupation, region and sector dummies.

(i) $\partial A_y / \partial T$=0 (the young are unaffected by changes in the treatment)

(ii) $\partial s_{fo} / \partial T$=0 (the effect of a higher T on the composition by gender of senior workers is small and can be disregarded)

(iii) $\left(A_y - A_o\right) \partial s_y / \partial T \cong 0$

Under these assumptions, the marginal effect of a higher treatment T on average days of absence per employee is given by

$$\partial A / \partial T = (1 - s_y) \left[s_{fo} \, \partial A_{fo} / \partial T + (1 - s_{fo}) \, \partial A_{mo} / \partial T \right] \tag{2}$$

Based on Table A.4, we estimate that one year of treatment increase the number of days of sick leave for females by 0.168 (0.014 x 12) and reduce the number of days of sick leave for males by 0.012 (-0.001 x 12). Using data from the Italian Labor Force Survey of 2011, we obtain that $s_y = 0.78$ and s_{fo}=0.31. Placing these numbers in (2) we estimate that one year of treatment increases the average number of days of sick leave per employee in a firm by 0.0096.[18]

To evaluate the impact of average sick leaves on firm productivity, we merge our working sample with firm balance sheet data from the Cerved databank. We retain only firms with at least 5 employees and estimate for the period 2010-2016 the following dynamic panel

$$lnY_{it} = \theta_0 + \theta_1 lnY_{it-1} + \theta_2 lnL_{it} + \theta_3 lnK_{it} + \theta_4 X_{it} + \theta_5 A_{it} + \varepsilon_{it} \tag{3}$$

[18] If we extend the senior age group to include also those aged 51-64 and assume that the effect on the days of sick leaves is the same for the age groups 45-50 and 50-64, the estimated impact of the average number of days of sick leave per employee increases to 0.016.

where Y is real value added, L employment, K the capital stock, and X is a set of controls, which include the share of blue-collar employees, temporary workers and part timers, average age and the share of female employees.

We estimate (3) using the method proposed by Arellano and Bond, 1991, which consists of two steps: a) in the first step, the time invariant component of the error term is differenced away by taking first differences; b) in the second step, endogenous variables such as lagged value added, employment and average days of sick leave per employee, and predetermined variables, such as the capital stock, are instrumented with lags in levels. We select all available lags from the second for lagged value added, from the second to the fourth for employment and sick leaves, and from the first to the fourth lag for the capital stock.

Our estimates are reported in Table 9. When the idiosyncratic errors are independent and identically distributed, the first differenced errors are first order serially correlated. Serial correlation of order two in the first-differenced errors should not be present, however, for the moment conditions required by identification to be valid. We test this hypothesis and find that we cannot reject the null (p-value: 0.19). We find that adding one day a year of sick leave per employee reduces value added pro capita by -1.9 percent. We therefore conclude that adding one additional year to the working horizon reduces average firm productivity by 0.018 percent (-1.9 x 0.0096), a negligible amount.[19]

The effect of sick leaves on productivity is attenuated by the strategies that firms adopt to deal with the consequences of these leaves, which

[19] This effect is 0.03 percent (0.016 x 1.9) if we include also seniors aged 51 to 64 years.

may include employing additional labor, using overtime and re-organizing tasks. These strategies, however, are costly, and can cause temporary deviations of total employment, labor costs and operating profits from the profit maximizing path. The loss of operating profits with respect to the optimal path is due to the resources that firms must use to compensate for the damage that sick leaves inflict to normal operations.

Let g_{it} be the percent change of firm outcome W_{it} between year t and t+1, where W includes output per worker, operating profits per worker, total, temporary and part-time employment. We assume that the deviations of g_{it} from the firm-specific profit maximizing growth path $\bar{g}_i$ are correlated with the deviations of average days of sick leave per employee from the steady state days $\bar{A}_i$ and the deviations of other factors, captured by the vector Ω

$$g_{it} - \bar{g}_i = \beta_1(A_{it} - \bar{A}_i) + (\Omega_{it} - \bar{\Omega}_i) \quad (4)$$

where $g_{it} = \ln(W_{it+1}) - \ln(W_{it})$ and $\Omega_{it} = \gamma_1 lnL_{it} + \gamma_2 lnK_{it} + \gamma_3 \frac{K_{it}}{L_{it}} + \gamma_4 \frac{T_{it}}{L_{it}} + \gamma_5 \frac{PT_{it}}{L_{it}} + \omega_t + \varepsilon_{it}$.

The vector Ω_{it} includes firm size (captured by log employment and log physical capital), technology (captured by the capital labor ratio $\frac{K_{it}}{L_{it}}$), internal labor organization (captured by the share of employment with temporary ($\frac{T_{it}}{L_{it}}$) and part-time contracts ($\frac{PT_{it}}{L_{it}}$)), the macroeconomic cycle (captured by year fixed effects ω_t) and the random shocks ε_{it}.

We estimate (4) on our working sample and report the results in Table 10. We find that one additional day of sick leave per worker is associated with higher total employment (+0.57 percent), temporary employment

(+0.85 percent) and part-time employment (+0.42 percent), confirming that firms with more sick leaves use extra-labor to manage the consequences. We also find that both output and profits per worker decline when days of sick leave increase, by 0.22 and 0.24 percent respectively. The estimated decline of labor productivity is significantly lower than the one estimated in Table 9, thereby confirming that the route connecting changes in the working horizon due to pension reforms to labor productivity via sick leaves is not quantitatively relevant.

Conclusions

Pension reforms designed to enhance the financial sustainability of pension systems can produce several unintended side effects by altering the behavior of economic agents. In this paper we have investigated the effects of a recent Italian reform on sick leaves, by making use of a unique dataset which combines social security records with certified sick leaves for a large sample of private sector workers. We have found that delaying retirement affects the propensity to take sick leaves in opposite directions for middle aged males and females. Sick leaves increase for females but decrease for males.

We have argued that delayed retirement increases current leisure because of income and substitution effects. Additional leisure can be consumed by taking "opportunistic" sick leaves. The decision to take these leaves depends on the perceived cost, or the sanction applied by the employer if the employee is caught cheating. We expect this cost to be higher for males than for females, because male workers have more at stake, both in terms of labor income and of social recognition.

In addition, sick leaves taken by females may be more easily tolerated by employers because of the implicit contract exchanging lower pay for higher flexibility in the organization of work (see Casarico et al, 2021). Conversely, middle aged men who have to stay longer in the firm because of the reform and are paid more than females may need to perform well to avoid dismissal, for instance by reducing sick leaves. The relevance of the fear of job loss as motivator of the behavior of males is supported by the evidence that the probability of dismissal significantly increased after the reform for middle aged males but not for middle aged females.

Since we have shown that the average days of sick leave negatively affect the productivity of firms, a potential concern is that pension reforms, by raising the expected working horizon, increase sick leaves and therefore reduce productivity. The evidence presented in this paper for a specific Italian reform contributes to dispel this concern, mainly because the estimated effect of the reform on the average number of days lost because of sick leaves is small. Because of this, the impact of a longer working horizon on firm productivity that operates via the change in average sick leaves is negligible.

Tables and Figures.

Table 1. Years to retirement and treatment, by age group and gender

	$YR_s > 5,\ 45 \leq age \leq 50$, Post-treatment: 2012-16	$YR_s > 3,\ 45 \leq age \leq 52$, Post-treatment: 2012-14
Males		
Contributions paid (years)	25.94 (5.74)	26.82 (6.25)
Age	47.24 (1.66)	48.11 (2.22)
Years to retirement (pre-reform)	11.79 (3.40)	10.84 (3.82)
Years to retirement (post-reform)	14.59 (3.74)	13.58 (4.13)
Treatment (in months)	3.05 (1.02)	2.99 (1.03)
Females		
Contributions paid (years)	24.53 (6.23)	25.27 (6.70)
Age	47.27 (1.67)	48.09 (2.22)
Years to retirement (pre-reform)	11.20 (2.41)	10.27 (2.91)
Years to retirement (post-reform)	12.82 (2.65)	11.88 (3.14)
Treatment (in months)	3.15 (1.39)	3.15 (1.40)

Note: YR_s: years to retirement according to the Sacconi Law.

Table 2. Summary statistics. Individuals aged 45 to 50 in 2011 and with at least 5 years to retirement before the Monti-Fornero reform

	Males	Females
Days of sick leave per year	4.71 (15.67)	5.36 (16.96)
Percent days of sick leave (percent)	2.13 (8.00)	2.37 (7.95)
At least one day of sick leave per year (percent)	24.71	27.44
Treatment in months	36.55 (12.19)	37.79 (16.70)
Work experience (years)	19.86 (4.03)	20.05 (3.74)
Foreign citizen	0.04	0.03
Blue collar	0.60	0.38
White collar	0.29	0.56
Senior manager	0.03	0.01
Junior manager	0.07	0.04
Temporary contract	0.08	0.09
Days worked per year in 2011	229.0 (37.7)	229.3 (37.2)

Table 3. Effect of the treatment T on the probability of positive sick leaves and on the percent of potential days of work spent in sick leave. Males and females aged 45 to 50 with at least five years to retirement before the reform.

	Days sick leave – binary Males	Days sick leave – binary Females	Percent days sick leave Males	Percent days sick leave Females
Pre - treatment: 2010-2011				
2010 x T	0.006	-0.001	0.001	-0.000
	(0.010)	(0.009)	(0.001)	(0.001)
Post - treatment: 2012-2016				
2012 x T	-0.022**	0.014	-0.000	0.001
	(0.009)	(0.009)	(0.001)	(0.001)
2013 x T	-0.030***	0.028***	-0.002**	0.003**
	(0.012)	(0.010)	(0.001)	(0.001)
2014 x T	-0.002	0.036***	0.000	0.003***
	(0.011)	(0.010)	(0.001)	(0.001)
2015 x T	-0.016	0.027**	0.001	0.002**
	(0.012)	(0.011)	(0.001)	(0.001)
2016 x T	-0.008	0.034***	0.001	0.003***
	(0.012)	(0.012)	(0.001)	(0.001)
Mean dependent variable	24.71	27.44	2.13	2.37
Number of observations	827,104	525,285	827,104	525,285
R Squared	0.1194	0.1165	0.030	0.043

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, months of treatment, the interaction of time to retirement before the reform with a linear trend, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table 4. Effect of the treatment T on the probability of positive sick leaves lasting 1-9 days and 10 days or more. Males and females aged 45 to 50 with at least five years to retirement before the reform.

Countries	Probability of taking sick leaves lasting 1 to 9 days - Males	Probability of taking sick leaves lasting 10+ days - Males	Probability of taking sick leaves lasting 1 to 9 days - Females	Probability of taking sick leaves lasting 10+ days - Females
Pre - treatment: 2010-2011				
2010 x T	0.004	0.002	-0.002	0.002
	(0.010)	(0.005)	(0.009)	(0.006)
Post - treatment: 2012-2016				
2012 x T	-0.018**	-0.010*	0.019*	0.000
	(0.009)	(0.005)	(0.009)	(0.005)
2013 x T	-0.024**	-0.006	0.033***	-0.001
	(0.011)	(0.005)	(0.010)	(0.005)
2014 x T	-0.005	0.001	0.040***	0.002
	(0.011)	(0.006)	(0.010)	(0.006)
2015 x T	-0.021*	-0.001	0.032***	-0.001
	(0.012)	(0.006)	(0.011)	(0.006)
2016 x T	-0.009	-0.001	0.040***	0.000
	(0.012)	(0.006)	(0.012)	(0.006)
Mean dependent variable	23.01	3.403	25.67	3.661
Number of observations	827,104	827,104	525,285	525,285
R Squared	0.112	0.014	0.109	0.018

Table 5. Heterogeneous effects of the treatment T on the probability of positive sick leaves. Males and females aged 45 to 50 with at least five years to retirement before the reform. By earnings at/above and below gender-specific medians

	Males		Females	
	Below median	Above/at median	Below median	Above/at median
Pre - treatment: 2010-2011				
2010 x T	-0.001	0.003	0.015	-0.024
	(0.017)	(0.015)	(0.016)	(0.016)
Post - treatment: 2012-2016				
2012 x T	-0.032**	-0.014	0.039**	-0.005
	(0.015)	(0.014)	(0.016)	(0.015)
2013 x T	-0.055***	-0.026*	0.064***	0.013
	(0.017)	(0.014)	(0.016)	(0.014)
2014 x T	-0.038**	0.008	0.075***	0.010
	(0.019)	(0.017)	(0.018)	(0.014)
2015 x T	-0.033*	-0.022	0.057***	-0.004
	(0.020)	(0.017)	(0.020)	(0.015)
2016 x T	-0.048**	-0.004	0.059***	0.001
	(0.021)	(0.017)	(0.020)	(0.015)
Mean dependent variable	27.44	23.17	30.43	24.71
Number of observations	300,231	489,684	244,606	261,092
R Squared	0.055	0.184	0.077	0.181

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, months of treatment, the interaction of time to retirement before the reform with a linear trend, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table 6. Heterogeneous effects of the treatment T on the probability of positive sick leaves. Males and females aged 45 to 50 with at least five years to retirement before the reform: Geographical area with more and less traditional gender norms.

	More traditional gender norms		Less traditional gender norms	
	Women	Men	Women	Men
Pre - treatment: 2010-2011				
2010 x T	-0.0001	0.0006	-0.0018	0.0029
	(0.0015)	(0.0019)	(0.0028)	(0.0036)
Post - treatment: 2012-2016				
2012 x T	0.0014	-0.0006	0.0042	-0.0031
	(0.0015)	(0.0017)	(0.0033)	(0.0037)
2013 x T	0.0050***	-0.0037**	0.0031	-0.0004
	(0.0017)	(0.0017)	(0.0035)	(0.0038)
2014 x T	0.0052***	-0.0011	0.0062*	0.0029
	(0.0017)	(0.0021)	(0.0036)	(0.0041)
2015 x T	0.0041**	0.0002	0.0009	0.0018
	(0.0018)	(0.0023)	(0.0035)	(0.0057)
2016 x T	0.0056***	-0.0010	0.0033	0.0035
	(0.0018)	(0.0021)	(0.0038)	(0.0048)
Mean dependent variable	1.554	1.403	1.653	1.469
Number of observations	405,535	655,263	100,163	134,652
R Squared	0.0437	0.0308	0.0517	0.0411

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, the interaction of YRS with a linear trend, months of treatment, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table 7. Effect of the treatment T on the probability of working part-time. Males and females aged 45 to 50 with at least five years to retirement before the reform.

	Males	Females
Pre - treatment: 2010-2011		
2010 x T	-0.014***	0.013**
	(0.005)	(0.005)
Post - treatment: 2012-2016		
2012 x T	-0.014***	0.012**
	(0.005)	(0.006)
2013 x T	-0.015***	0.031***
	(0.006)	(0.007)
2014 x T	-0.024**	0.044***
	(0.007)	(0.008)
2015 x T	-0.016**	0.046***
	(0.007)	(0.009)
2016 x T	-0.014	0.061***
	(0.009)	(0.009)
Mean dependent variable	6.84	42.63
Number of observations	789,915	505,698
R Squared	0.090	0.116

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, months of treatment, the interaction of time to retirement before the reform with a linear trend, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table 8. Effect of the treatment T on dismissal and separation hazards - Males and females aged 45 to 50 with at least five years to retirement before the reform. Cox survival model.

	Dismissals hazard Males	Dismissals hazard Females	Separations hazard Males	Separations hazard Females
	(1)	(2)	(3)	(4)
Treatment in months	0.002***	-0.000	0.002***	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Contributions paid in 2011	-0.003***	-0.002***	-0.003***	-0.002***
	(0.000)	(0.000)	(0.000)	(0.000)
Age in 2011	0.054***	0.018***	0.028***	0.004
	(0.005)	(0.006)	(0.004)	(0.005)
Experience in 2011	-0.029***	-0.018***	-0.026***	-0.024***
	(0.003)	(0.003)	(0.002)	(0.003)
Foreigner in 2011	0.358***	0.252***	0.342***	0.0339***
	(0.036)	(0.055)	(0.031)	(0.040)
Days worked in 2011	-0.000**	0.0000	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Number of observations	112,239	71,485	108,036	69,191

Note: each regression includes a constant, qualification, region and sector dummies. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table 9. The effect of the average number of days of sick leave per employee on productivity. Arellano Bond estimates. Years 2010-2016. Dependent variable: log real value added. Firms with at least 5 employees

	(1)
Lagged log value added	0.205***
	(0.004)
Log real capital stock	0.042***
	(0.002)
Log employment	0.433***
	(0.020)
Days of sick leave per employee	-0.019***
	(0.002)
Share of blue-collar employees (lagged)	-0.236***

	(0.011)
Share of workers on temporary contracts (lagged)	-0.022***
	(0.007)
Share of part time workers (lagged)	-0.162***
	(0.012)
Average age	-0.000
	(0.000)
Percent females	-0.001
	(0.007)
p-value test second order autocorrelation	0.190
Number of observations	198,360

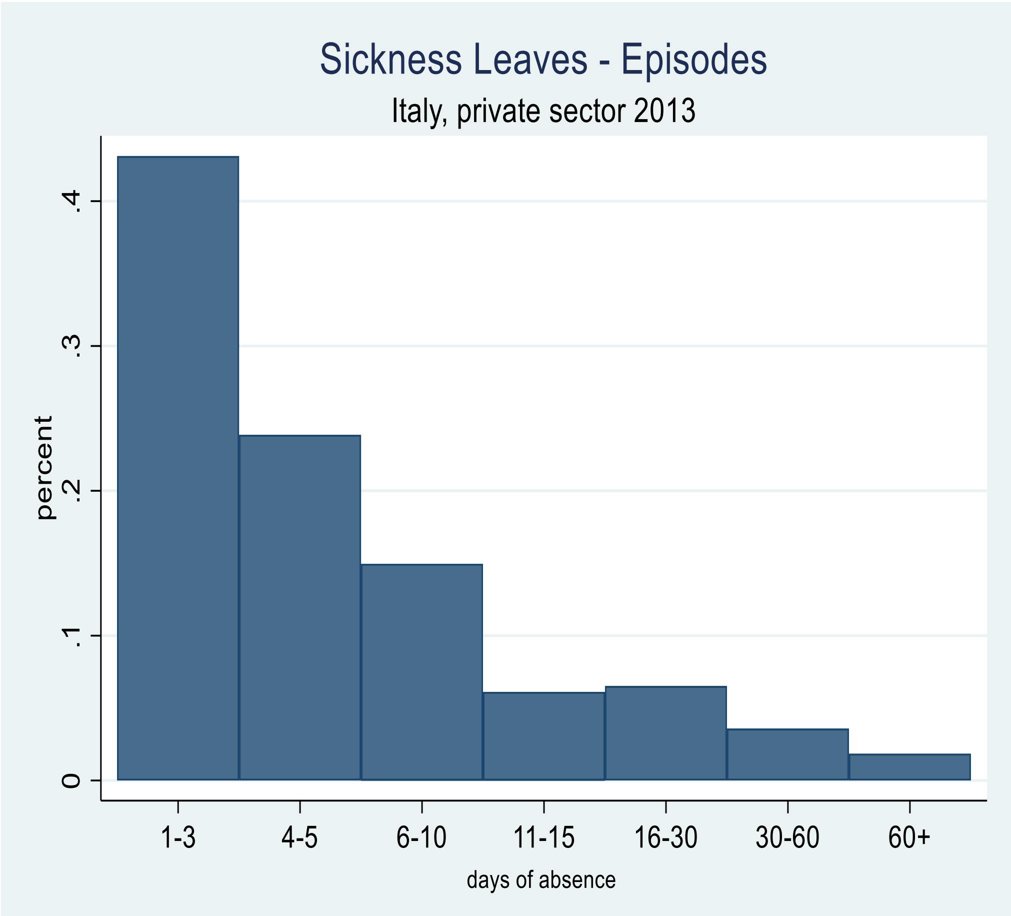
Note: the regression includes a constant, year and sector dummies. Standard errors are clustered by gender x age x years of contributions in 2011. Robust standard errors within parentheses. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence. The first differences of lagged value added, current employment and days of sick leave per employees are treated as endogenous and instrumented with all available lags of real value added in level, starting from lag 2; the first difference of employment and days of sick leaves are instrumented with lags 2 to 4 of employment and days of sick leave (in levels). The first difference of pre-determined capital is instrumented with the first to fourth lag of capital in levels.

Table 10. – The effect of days of sick leaves per worker on the growth rate of firm outcomes (percent points).

	(1) Employment	(2) Temporary Employment	(3) Part-time employment	(4) Output per worker	(5) Operating profits per worker
Days of sick leave per worker	0.569***	0.854***	0.420***	-0.222***	-0.243***
	(0.0315)	(0.0879)	(0.0651)	(0.0338)	(0.0904)
Observations	379,804	259,433	286,016	379,804	378,760
R-squared	0.447	0.515	0.578	0.505	0.405

Note: each the regression includes a constant and year dummies. Standard errors are clustered by firm.

Figure 1. The distribution of sickness episodes, by duration of episodes



Source: INPS

Figure 2. Years to retirement before the reform. Males

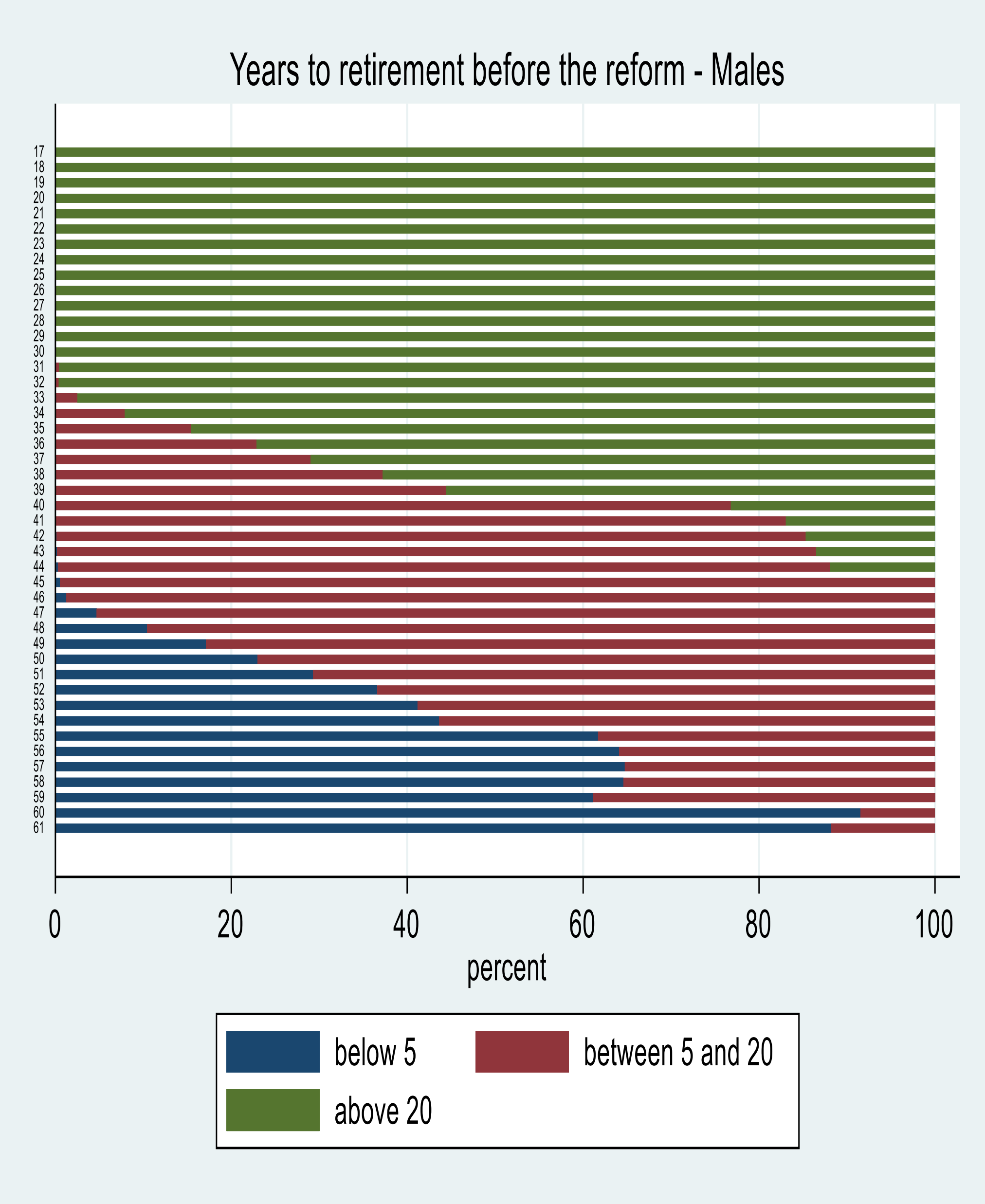


Figure 3. Years to retirement before the reform. Females

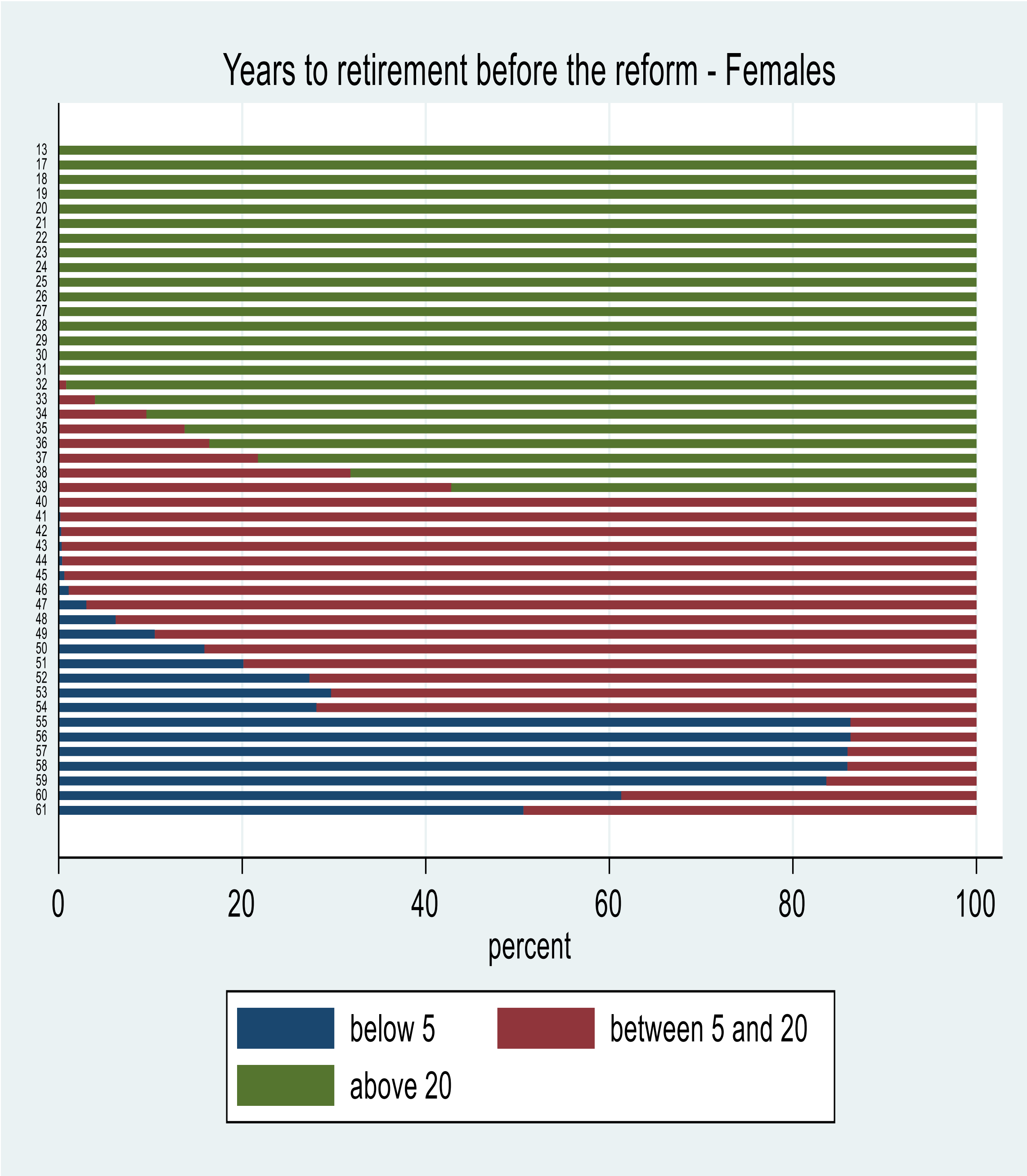


Figure 4. Distribution of the treatment T in years - Males

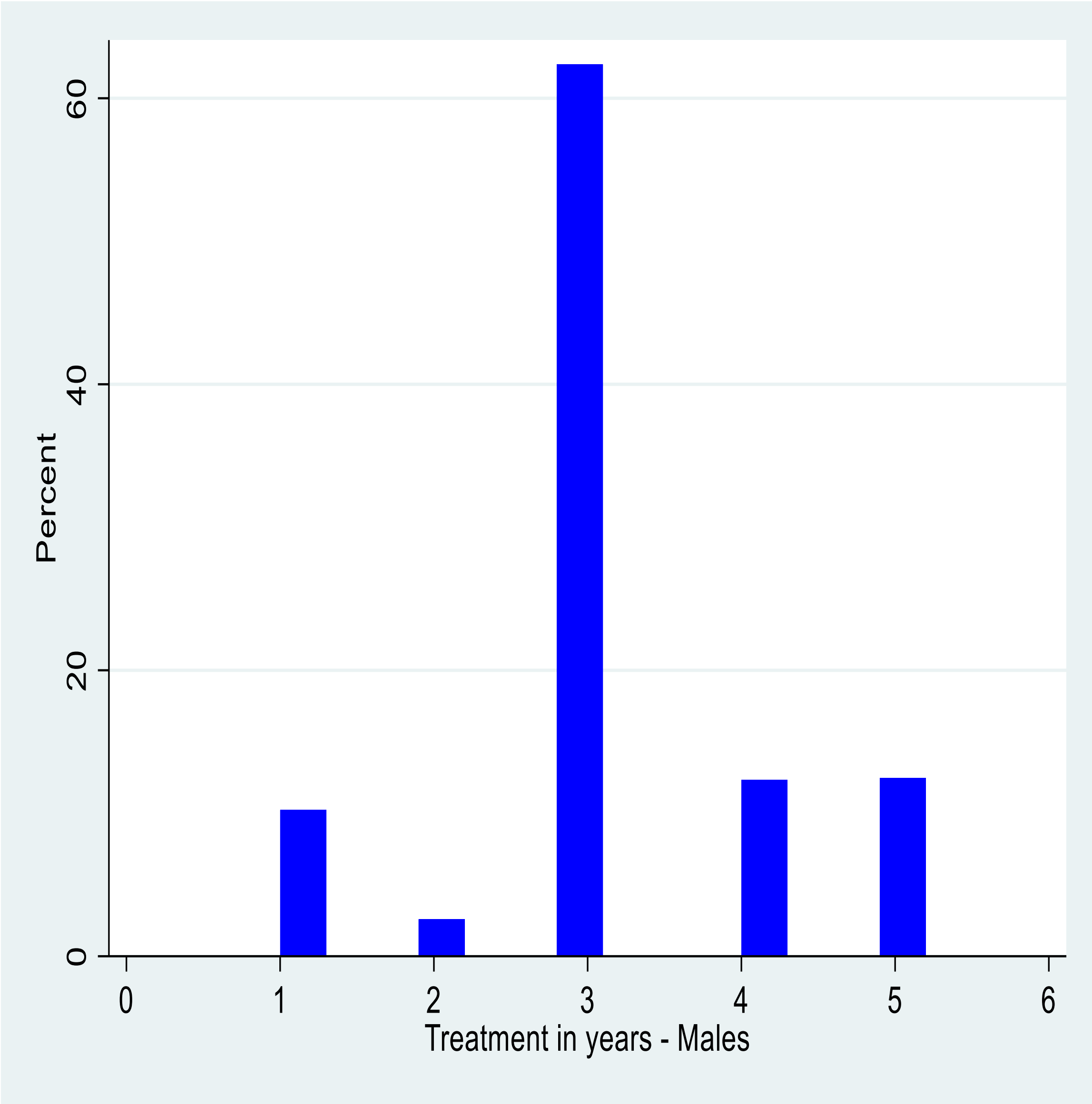
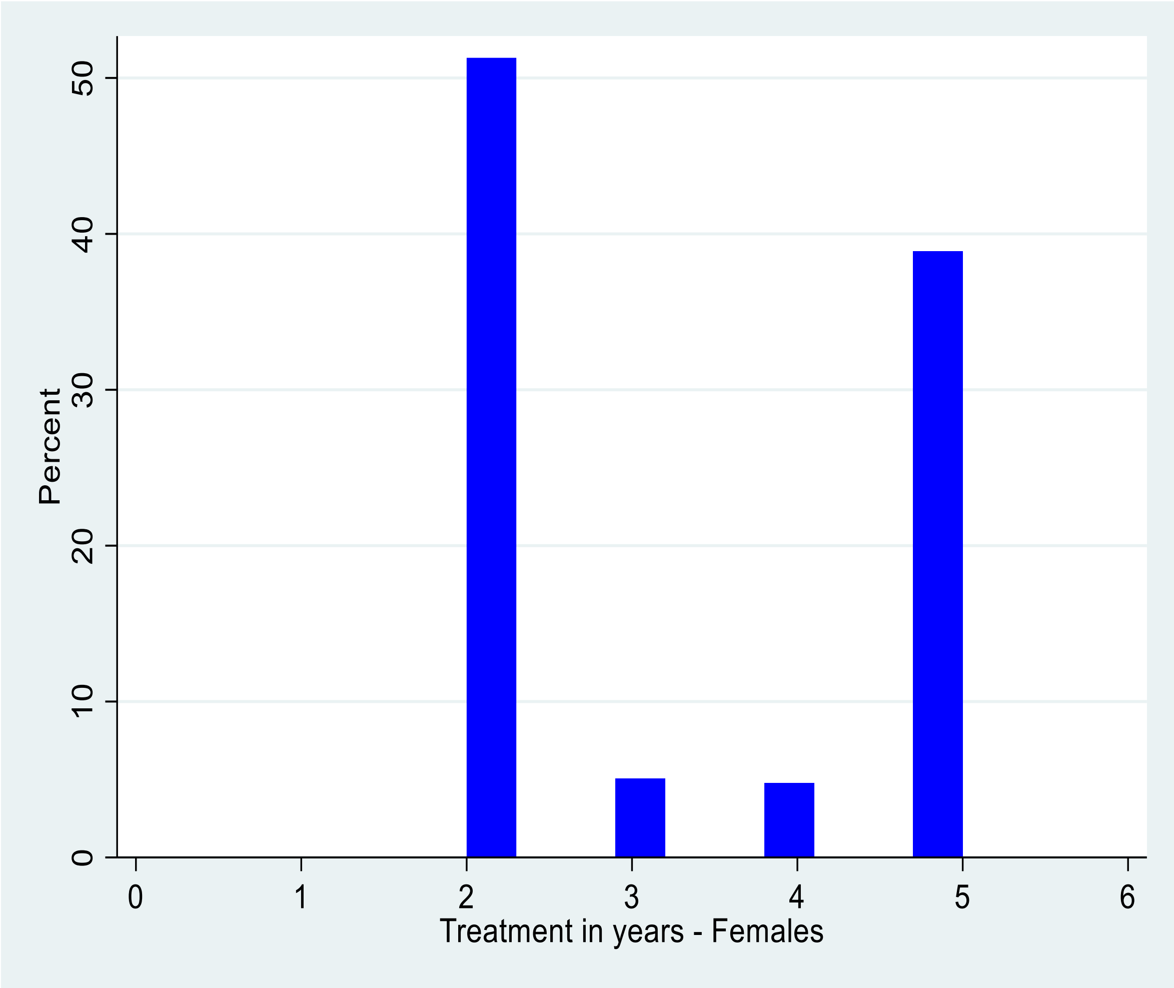


Figure 5. Distribution of the treatment T in years - Females



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Appendix

Table A1 Appendix. Eligibility rules for seniority and old age pensions.
Sacconi Reform, 2011

Year	Seniority pension: minimum retirement age	Minimum years of paid social security contributions	Minimum years of paid social security contributions with no age requirements	Old age pension: minimum age
Males				
2011	60 or 61	35 or 36	40	65
2012	60 or 61	35 or 36	40	65
2013	61 +3m or 62+3m	35 or 36	40	65 + 3m
2014	61 +3m or 62+3m	35 or 36	40	65 + 3m
2015	61 +3m or 62+3m	35 or 36	40	65 + 3m
2016	61 + 7m or 62 + 7m	35 or 36	40	65 + 7m
2017	61 + 7m or 62 + 7m	35 or 36	40	65 + 7m
2018	61 + 7m or 62 + 7m	35 or 36	40	65 + 7m
Females				
2011	60 or 61	35 or 36	40	60
2012	60 or 61	35 or 36	40	60
2013	61 +3m or 62+3m	35 or 36	40	60 + 3m
2014	61 +3m or 62+3m	35 or 36	40	60 + 4m
2015	61 +3m or 62+3m	35 or 36	40	60 + 6m
2016	61 + 7m or 62 + 7m	35 or 36	40	61 + 1m
2017	61 + 7m or 62 + 7m	35 or 36	40	61 + 5m
2018	61 + 7m or 62 + 7m	35 or 36	40	61 + 10m

Note: m: month. Source: national legislation.

Table A2 Appendix. Eligibility rules for seniority and old age pensions. Monti-Fornero Reform, 2012

Year	Old age pension: minimum age	Minimum years of paid social security contributions	Minimum years of paid social security contributions with no age requirements
Males			
2012	66	20	42 + 1m
2013	66 + 3m	20	42 + 5m
2014	66 + 3m	20	42 + 6m
2015	66 + 3m	20	42 + 6m
2016	66 + 7m	20	42 + 10m
2017	66 + 7m	20	42 + 10m
2018	66 + 7m	20	42 + 10m
Females			
2012	62	20	41 + 1m
2013	62 + 3m	20	41 + 5m
2014	63 + 9m	20	41 + 6m
2015	63 + 9m	20	41 + 6m
2016	65 + 7m	20	41 +10m
2017	65 + 7m	20	41 +10m
2018	66 + 7m	20	41 + 10m

Note: m: month. Source: national legislation.

Table A3. Effect of treatment on sick leaves and percent of days of sick leave per year- Males aged 45 to 52 with at least 3 years to retirement before the reform

	Days sick leave – binary males	Days sick leave – binary females	Percent days sick leave – males	Percent days sick leave – females
Pre - treatment: 2010-2011				
2010 x T	0.006	-0.013	0.001	-0.002
	(0.009)	(0.008)	(0.001)	(0.001)
Post - treatment: 2012-2016				
2012 x T	-0.013	0.008	-0.001	0.001
	(0.008)	(0.008)	(0.001)	(0.001)
2013 x T	-0.019*	0.020**	-0.003**	0.002*
	(0.010)	(0.009)	(0.001)	(0.001)
2014 x T	-0.006	0.029***	-0.000	0.0030**
	(0.010)	(0.009)	(0.001)	(0.001)
Mean dependent variable	24.15	26.65	1.391	1.541
Number of observations	790,840	490,715	790,840	490,715
R Squared	0.12	0.117	0.029	0.042

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, the interaction of YRS with a linear trend, months of treatment, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.

Table A.4. Effect of the treatment T on the number of days spent in sick leave. Males and females aged 45 to 50 with at least five years to retirement before the reform.

	Males	Females
2010 x T	0.006	-0.001
	(0.004)	(0.004)
Post - treatment: 2012-2016	-0.001	0.014***
	(0.003)	(0.004)
Mean dependent variable	4.705	5.362
Number of observations	827,104	525,285
R Squared	0.040	0.051

Note: T: treatment in months. Each regression includes a constant, year, region and sector dummies, months of treatment, the interaction of time to retirement before the reform with a linear trend, labor market experience in 2011, citizenship in 2011, qualification in 2011, fulltime status in 2011, number of days worked in 2011 and dummies gender x age x years of contribution in 2011). We exclude from each regression the interaction of the year before the treatment (2011) with T. Standard errors are clustered by gender x age x years of contributions in 2011. One, two and three stars for statistical significance at the 10, 5 and 1 percent level of confidence.