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Giorgio Brunello
University of Padua and IZA

Maria De Paola
INPS, University of Calabria and IZA

Lorenzo Rocco
University of Padua and IZA

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IZA – Institute of Labor Economics

Schaumburg-Lippe-Straße 5–9
53113 Bonn, Germany

Phone: +49-228-3894-0
Email: publications@iza.org

www.iza.org

ABSTRACT

Pension Reforms, Longer Working Horizons and Absence from Work*

Using matched employer-employee data for Italy and newly available information on sick leaves certificates, we study the effect of an exogenous increase in the length of the residual work horizon – triggered by a pension reform that increased minimum retirement age - on middle-aged employees' absence from work due to 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 to males. Sick leaves increased only for females working in firms paying smaller wage premia to female than to male workers, suggesting that, in these firms, females exchange lower pay with higher flexibility in their work schedule.

JEL Classification: J22, J26
Keywords: absences from work, retirement, Italy

Corresponding author:
Giorgio Brunello
Department of Economics and Management
University of Padua
via del Santo 33
35123 Padova
Italy
E-mail: giorgio.brunello@unipd.it

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Introduction

Labor supply decisions in middle age are affected by the residual working horizon, or the time remaining until retirement from the labor market. Pension reforms that increase minimum retirement age provide an ideal natural experiment to study the effects of changes in the horizon on individual labor supply. By delaying retirement, these reforms affect not only individuals who would have retired in the absence of changes, but also forward looking middle-aged individuals who are relatively far from retirement. They also affect employers by altering the expected composition of employment by age.

In a recent contribution, Carta and De Philippis, 2022, study the effect of a longer working horizon on the labor supply of workers aged 45 to 64 by focusing on the extensive margin. In this paper, we consider instead the intensive margin and ask whether a longer working horizon increase sick leaves by forward looking middle aged workers aged 45 to 50. These leaves are the largest cause of absence from work, which negatively influence both individual careers (Hansen, 2000; Markussen, 2011; Henningsen and Hægeland, 2008) and firm productivity (Grinza et al, 2020).

A priori, the effect of an extension of the residual working horizon on sick leaves is uncertain and depends on the interplay between workers' and firms' responses. On the one hand, middle aged workers might compensate the expected reduction of future leisure[1] with additional current leisure, by

[1] Following standard practice in economics, leisure is the time not devoted to paid work, and comprises both unpaid work, including the time spent in caring activities and other domestic tasks, and free time.

taking opportunistic sick leaves, or leaves that are not motivated by poor health conditions[2] (Ferman et al, 2021; Moscarola et al, 2016).

On the other hand, firms where the pay of middle aged workers increases with age faster than productivity – to compensate for the faster increase of productivity at earlier ages (see Lazear, 1979) – are likely to be negatively affected by the postponement of the retirement age, which forces them to keep "expensive" workers for a longer period of time. They may react by increasing dismissals of these workers or by pressuring them – directly or indirectly via the fear of job loss - to work harder and increase productivity, for instance by reducing sick leaves. Sick leaves could also fall if individual health improves, for instance because middle-aged workers adopt healthier lifestyles to stay healthy longer and manage to work longer (see Bertoni et al., 2018).

Our empirical study focuses on the "Monti-Fornero" (MF hereafter) pension reform, that increased significantly minimum retirement age and was implemented in Italy at the end of 2011. We combine matched administrative employer-employee data with individual data on sick leave certificates, that are available from 2010. We examine the effects of the reform on private sector employees aged 45 to 50, who at the time of its implementation had at least five years to go before reaching the minimum retirement age, by following them for five years, from 2012 to 2016. Adopting a difference-in-

2 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. 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. 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 requires.

differences strategy, we look at the changes in the probability of taking sick leave and in the percent of total working days spent in sick leave by workers exposed to different treatment doses, where a single dose is a one-month increase in the residual working horizon, measured at the end of 2011.

We find that, following an exogenous increase in the horizon, both the probability of taking sick leaves and the percent of total working days spent in sick leave increased for females but declined for males. Since these effects are driven by relatively short episodes of sick leave, it is unlikely that changes in health have played a major role.

We explain the increase of females' sick leaves after the pension reform with the higher demand for current leisure, driven by gender norms and lower costs. These norms place the bulk of informal care within the household in the hands of females, who need to combine work and care and may have reacted to the longer expected working horizon by demanding more current leisure. Since opportunistic sick leaves are paid, their cost is the expected sanction if caught cheating on health conditions, which depends on the wage and is lower for females than for males because the former have typically lower wages.

Consistent with these explanations, we find that: a) females living in geographical districts characterized by more traditional gender norms reacted to the reform by increasing sick leave absences more than females leaving in places with more egalitarian attitudes; b) the probability of taking sick leaves after the reform increased only for females with lower than gender-specific median wages.

We also find evidence that the increase of sick leaves is observed only among females working in firms that pay smaller wage premia[3] to female than to male workers with similar work experience. We conjecture that, in these firms, female workers exchange lower pay with more flexibility in their work schedule (see for instance Goldin and Katz, 2016; Bolotnyy and Emanuel, 2022), and therefore have ampler opportunities to adjust their labor supply at the intensive margin in response to the pension reform.

For males, the reduction of future leisure due to the reform did not trigger the same increase in the demand of current leisure observed for females. On the one hand, gender roles imply that males are less concerned with informal care within the household. They also have on average higher wages (see Casarico and Lattanzio, 2021), and higher expected sanctions associated with opportunistic behavior. Higher wages also mean that males who are expected to stay longer in firms are likely to be under heavier pressure than females to raise productivity and more at risk of losing their job.

This stronger pressure, and the associated higher fear of job loss, may explain why male workers have reduced their sick leaves after the reform to increase their productivity and avoid dismissal. In support of this interpretation, we find that the longer residual working horizon induced by the pension reform raised the hazard of dismissals for males but not for females, and that this increase concentrated mainly in firms paying higher wage premia to male than to female workers.

Our paper contributes to the literature that studies the effects of a longer residual working horizon on several choices made by middle aged individuals, including job search behavior and employment (Hairault et al,

[3] By wage premium, we mean the firm fixed effect in a gender-specific regression of wages on individual and firm fixed effects. See Abowd, Kramartz and Margolis, 1999.

2010), labor supply (Carta and De Philippis, 2022; Lalive et al., 2017, Cribb et al. 2016), unemployment benefit claims and the demand of disability pensions (Duggan et al 2007; Staubli and Zweimüller, 2013), human capital investment (Montizaan et al, 2009, Brunello and Comi, 2015 and Chinetti, 2019), mortality (Bloemen et al, 2017), informal care (Fischer and Muller, 2020) and, most importantly, sick leaves (Ben Halima et al, 2021 and Moscarola et al, 2016).[4]

In the study closest to ours, Moscarola et al, 2016, also examine the impact of the MF reform on sick leaves in Italy. This earlier research differs from ours because it considers only sick leaves lasting 7 days or longer, that in 2013 contributed to less than 35 percent of all sickness episodes in the private sector (see Figure 1), and looks only at females in 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 – year horizon.

By investigating how the relationship between a longer residual working horizon and sick leaves varies with firm wage premia, our paper also contributes to the literature on time flexibility and gender wage gaps (see Bertrand, Goldin, and Katz, 2010; Goldin, 2014; Goldin and Katz, 2016). Mas and Pallais, 2017, for instance, show that females value schedule regularity and flexibility more than males. Bolotnyy and Emanuel, 2022, using administrative data on bus and train operators, argue that the gender gap in earnings can be explained by the fact that females are both less likely than males to accept working longer hours and more likely to take unpaid time

[4] 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).

off. We show that in firms paying a lower pay premium to female workers, these workers react to an increase in the residual working horizon by taking more sick leaves.

Our paper is also related to the literature that examines the impact of a longer working horizon on individual health-promoting behaviors and health. 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, reducing obesity and raising self-reported satisfaction with health. On the other hand, a longer horizon can increase stress and strain, with negative effects on mental health (see De Grip et al., 2012; Carrino et al., 2020; Bertoni et al, 2022).

Finally, this paper speaks to the literature that investigates the determinants of absenteeism and emphasizes the role of working conditions (Osborg Ose, 2005, De Paola, 2010), 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 emphasize instead pension reforms that increase residual working horizons.

The rest of the paper is organized as follows: in Section 1 we present the key features of the MF 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 reported in Section 4. The discussion of candidate mechanisms that could explain our results is in Section 5. Conclusions follow.

1. Institutional Background

1.1 The Monti Fornero (MF) 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 MF 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 at age 60 and with 36 years of paid contributions, or at age 61 with 35 years of contributions, or finally with 40 years of contributions at any age. In the year immediately after the MF 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 The 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 three 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 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.[5] 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 residual working horizon, our treatment variable, which is 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.

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

The time invariant treatment variable T – the change in the residual working horizon - is the difference between the expected number of months to retirement eligibility (MR) after the MF law was introduced and the MR before the reform, when the Sacconi law was in place. The treatment is time invariant because the number of months is always computed at the end of 2011, under the assumption that individuals remain employed until

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

eligibility.[6] To compute the treatment T for everyone, the rules about minimum retirement age set 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 MF laws did not set a time invariant minimum retirement age but allowed this minimum to increase over the period 2012 to 2016 (and after), in line with the increase of life expectancy.

To take this into account, we compute the individual treatment T as time invariant but contingent 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 MF 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 MF and the Sacconi rules.

We exclude from our working sample individuals younger than 45 in 2011, who are too far away from retirement, and individuals older than 50 in the same year, because the share of those with less than five years to 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.[7] Since days of sick leave can be zero either because the employee is never absent or because the

[6] This assumption is regularly made in this literature. See for instance Carta and De Philippis, 2022, who compute their treatment variable by considering 2010 as the pre-treatment year. By so doing, they evaluate the joint impact of the Sacconi and the MF reforms.

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

individual is not employed, we further restrict our working sample to those who were employed in 2011 and spent at least one month per year in employment between 2011 and 2016.[8]

We also define a secondary sample by selecting individuals who were aged 45 to 52 in 2011, with a residual working horizon of at least 3 years, whom 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 males and 24.53 for females; mean age was 47.24 for males and 47.27 for females; mean years to retirement before and after the reform were 11.79 and 14.59 for males and 11.20 and 12.82 for females; and mean years of treatment were equal to 3.05 for males and 3.15 for females, or 25.8 and 28.1 percent of pre-treatment mean years.

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

3. The empirical setup

We investigate empirically the effect of exogenous changes in the length of the residual 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

$$\mathrm{Y_{iqt}} = \beta_0 + \beta_1 \mathrm{X_{iqt}} + \beta_2 \mathrm{T_{iq}} * \mathrm{D2010_t} + \sum_{y=2}^{6} \beta_{3y}\, T_{iq} D201y + \gamma_\mathrm{q} + \gamma_\mathrm{t} + \varepsilon_\mathrm{iqt} \qquad (1)$$

[8] Although we cannot identify in our data employees with a high degree of disability, who are exempt from the norms introduced by the MF reform, our estimates based on ISTAT, *Indagine multiscopo sulle famiglie* suggest that only 2 percent of employees aged 45 to 64 are in this condition.

where Y is the outcome, the indices i, t and q are for the individual, the time and 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, 2022, 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 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, computed as 20 times the number of months with at least one day of work.[9] Estimates of pre - and post-treatment effects are relative to the baseline year 2011, and standard errors are clustered by cell q.

In our setup, the treatment T is not binary but multi-valued. Hence, β_{3y} can be interpreted as the unit causal response of the treatment $\mathrm{Y_{iqt}}(\mathrm{T_{iq}} = x) - \mathrm{Y_{iqt}}(\mathrm{T_{iq}} = x - 1)$ if there is no selection-on-gains, implying that the effect of treatment $\mathrm{T_{iq}} = x$ on the outcome is the same for the group receiving dose x and for the other groups in the sample (see Callaway et al, 2021).

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* had at least one day of sick leave,

[9] We consider the potential rather than the actual number of days worked as the latter is affected by sick leaves.

and to zero otherwise; b) the percent PY, that is the ratio between the number of sick leave days in year t and the potential number of days worked in the same year, defined above. The second measure controls for the fact that individuals with a higher number of potential days of work are more likely to be absent because of sick leave simply because they have a longer exposure to the risk. Table 2 shows the summary statistics of the variables used in the estimates, by gender. On average, PY is equal to 2.13 percent for males and 2.37 percent for females, and DY is equal to 24.71 percent for males and 27.44 percent 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 different values of T were not on different trends before the MF reform was introduced, we always find that the interaction of T with the pre-treatment year 2010 is not statistically significant.

Our evidence indicates that females (males) with a higher reform-induced increase in the residual working horizon took more (fewer) sick leaves than females (males) with smaller increases. For females, the positive effect of a higher T on sick leaves was persistent and statistically significant from 2013 onwards. When we consider the last year in the sample, 2016, we estimate that adding one year to the expected residual working horizon raised both the probability of taking at least a day of sick leave and the percent of days of sick leave by 1.5 percent.[10] Since the average increase of the treatment for females is just above three years, this corresponds to a 4.5 percent increase of either outcome.

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

For males, we find that one additional year of treatment reduced 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, however, faded away in later years. One more year of T also reduced 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 was present also between 2014 and 2016 but was smaller and not statistically significant at the conventional levels of confidence.[11]

To further assess the size of the estimated effects, Table A4 in the Appendix shows the impact of the treatment T on the number of days of sick leave, assuming that this impact is constant over the post-treatment period. We estimate that one year of treatment increased the sick leaves of females by 0.168 days (or 3.1 percent with respect to the sample average) and decreased the sick leaves of males by 0.012 days (or -0.25 percent with respect to the sample average). Considering that the MF reform increased the treatment T by close to 3 years on average, the estimated effect of this increase on sick leaves is +0.529 days for females – or 9.8 percent with respect to the mean - and -0.036 days for males, or -0.76 percent with respect to the mean. We conclude from this that the estimated effect of the pension reform on sick leaves is negligible for males but not for females.

A potential concern with the estimates in Table 3 is that, because of lack of data, the pre-treatment period is too short for a credible test of the parallel trend hypothesis. We therefore consider an alternative dataset, the Italian

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

Labor Force Survey, which asks individuals to self-report absences and to the absent from work whether this absence was due to either sickness or injury.[12] Using annual data for the period 2006 to 2016, we test whether the interactions of the treatment T with the year dummies covering the period 2006 to 2010 are jointly different from zero, which would support the assumption of parallel trends.[13] Encouragingly, 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 this result.

5.1. Health

One reason why sick leave vary with the treatment is that a longer residual 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

[12] A drawback of these data is that there is no 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 when the highest education degree is attained.

[13] Since the sample size of the Labor Force Survey is much smaller than 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 because the variables T_{iq}*D20yy for y=06…10 are correlated.

short-term, leaves due to poor health have longer durations (see for instance Campolieti, 2004 and Ziebarth, 2013).[14]

If poorer health is driving our results, we should find that a longer residual working horizon increases long sick leave spells more than short spells. To verify this, we estimate the effect of the treatment T on the probabilities of having at least one short spell – lasting 1 to 9 days – and at least one long spell (lasting 10 days or more) of sick leave. Table 4 reports our estimates by gender. For females, a higher T has no effect on long spells but a positive and statistically significant effect on short spells. For males, both types of sick leaves decline with T. While the effect on short spells persists over time, the one on long spells fades away quickly. We conclude from this that health conditions are unlikely to be the story driving our results.[15]

5.2 Demand for leisure

A longer working horizon may increase the demand for current leisure as forward looking workers try to smooth their consumption over the life cycle. Additional leisure can be consumed by increasing the use of opportunistic sick leaves.[16] Since the time devoted to paid work and the demand for leisure are strongly related to gender norms, we investigate whether the individual response to the reform varies across areas with more or less traditional gender attitudes. In traditional societies, where females bear the main burden of housework and informal care, we expect that the perspective of

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

[15] Moscarola et al, 2016, consider the effect of the same reform on the weeks of sick leave 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.

[16] An alternative option that does not incur in the risk of sanctions is a leave motivated by blood donation. In Italy, workers donating blood are entitled to a day of leave with full pay. Table A5 in the Appendix shows that the probability of donating blood increases with the treatment for females but not for males.

having to stay in the labor market longer – and add work to care and housework for a longer spell - may have generated higher dissatisfaction with the reform among females, inducing them to reduce their current working days by taking opportunistic sick leaves.[17]

Using data from the fourth wave of the European Values Study, we allocate the regions of Italy in two groups, depending on whether the percentage of individuals who agree with the statement: "A job is right but what most women want is a home and children" is above/at or below the median. The former group is considered to have more traditional gender roles than the latter. We run separate regressions by group and find that the probability of taking sick leaves increases with the treatment to a higher extent among females living in areas with traditional gender roles (see Table 5) than among other females. There is also evidence (not reported here but available upon request) that the effect of the treatment disappears among females living in the regions belonging to the bottom quartile of the distribution. Our results are also robust to changes in the definition of traditional gender roles.[18]

We have mentioned above that the cost of taking opportunistic leaves 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

[17] An increase in the residual working horizon could have induced some middle-aged working females involved in informal care to re-arrange their organization of tasks within the household. For instance, those who before the reform were intensively relying on external services (or on the care provided by other members of the family) for support of their elderly, planning to take over after retiring, may respond to delayed retirement and to the expected higher costs by partially replacing these services with their own time.

[18] Table A6 in the Appendix of the paper shows our results when we consider the answers to the statement: "A man's job is to earn money; a woman's job is to look after the home and family". The table confirms that the impact of the reform on females' sick leaves is larger in magnitude in areas with more traditional gender roles.

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 turn to opportunistic sick leaves when the residual working horizon increases.

We investigate this hypothesis by running separate regressions for males and females, distinguishing between workers who earn median or below median wages and those who earn above median wages. For females, we find supportive evidence because the effect of a higher T 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 (See Table 6, column (3) and (4)). For males, we find that sick leaves decline when T increases, 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 6).

5.3 The role of gender-specific firm wage premia

In imperfect labor markets, firms pay differently workers with similar observable skills and education. Females, for instance, earn on average lower firm wage premia than males (see Casarico and Lattanzio, 2021, for the Italian case). As discussed above, a source of these differences might be the higher demand for flexibility in the time schedule that females have, mainly due to the different burden of caring. By accepting lower wage premia, female workers compensate the employer for the higher flexibility in the work schedule (see Goldin and Katz 2016). When the residual work horizon increases, females in firms that pay them lower wage premia (relative to males) can exploit this higher flexibility and consume additional leisure by using opportunistic sick leaves. If this hypothesis is correct, we should find that the sensitivity of sick leaves to changes in the working horizon is higher in firms with lower female wage premia (relative to male premia).

To verify this hypothesis, we estimate firm wage premia by gender as firm fixed effects in gender-specific wage regressions that include worker and firm fixed effects (see Abowd et al., 1999), using administrative employee-employer data for the pre-reform period 2005 to 2011 that include information on gross annual earnings and several firm and employee features.[19] Following Casarico and Lattanzio, 2021, we compute wage gaps as differences between gender-specific wage premia.[20]

We merge these gaps with our working sample and run separate regressions for the workers in firms characterized by a gender gap in pay premia in favor of males and in firms where such gender gap is either not observed or is in favor of females. The results reported in Table 7 show that a higher residual working horizon caused an increase in the probability that females take sick leaves, but only in firms paying lower pay premia to female than to male workers. The effect on males is to reduce sick leaves, and is often imprecisely estimated, independently of whether the firm pays a high or a low wage premium. These results confirm that female workers can exploit the higher flexibility of their work arrangements associated with the lower pay premia to increase their consumption of leisure when the expected time to retirement increases.

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[19] We restrict the sample to employees in the age range 19-65 with at least two years of work experience. We exclude firms that employ only individuals of the same sex for the entire period considered. For employees holding more than one job within the same year, we consider the job with the highest number of weeks worked.

[20] To compare male and female fixed effects we consider a double connected set of workers and firms and since firm fixed effects should be zero when firms do not share rents with their workers, we normalize firm fixed effects with respect to the average firm effect in the accommodation and food sector.

5.4 Fear of job loss

Because of the MF reform, senior employees aged 45 to 50 need to stay longer in the labor market to attain retirement eligibility. Earnings profiles typically increase with labor market experience and tend to overtake productivity profiles late in the career to compensate for the productivity premium that characterizes the earlier part. According to Lazear, 1979, the optimal retirement age of each individual worker from the employer's perspective 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.

We expect the pressure on employers' profits to be higher in the case of male workers, who are typically better paid and more likely to be on open ended contracts than females. This higher pressure, and the associated 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, and avoid dismissal.[21]

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

[21] Bovini and Paradisi, 2019, show that delayed retirement due to pension reforms has increased dismissals in Italy. An alternative view is that firms could have encouraged middle aged workers who have to stay longer because of the reform to take sick leaves so as to save on labor costs. This could happen because INPS partially pays workers on sick leave starting from the fourth day of leave. This story, however, does not explain why sick leaves have increased for females but not for males.

dismissals and for separations (dismissals plus quits).[22] Conditional on age and contributions paid until 2011, we find that a longer residual working horizon increases both dismissals and separations 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.[23]

Furthermore, we classify firms according to the wage premia paid to males and females and, similarly to what done in the previous sub-section, we run our survival model separately for workers employed in firms characterized by a gender gap in pay premia in favor of males and workers employed in firms where such gender gap is either not observed, or is in favor of females. As shown in Table 9, we find that a longer residual working horizon increases the hazard of being dismissed only for males working in firms favoring males in their wage policy, consistent with the idea that in these firms the threat of job loss is more salient.

Conclusions

In this paper, we have estimated for middle aged workers aged 45 to 50 the causal effect of a longer residual working horizon on their absence from work due to sick leaves. For identification, we have relied on a pension reform that occurred in Italy in 2011, which substantially increased minimum retirement age and the residual working horizon of employed males and females. Using matched employer-employee data for the period 2010 to 2016, we have found that, because of the expected longer horizon, sick leaves increased among females but decreased among males. The estimated effects are small for males and not negligible for females: by

[22] We notice that voluntary and involuntary separations are often not distinguishable empirically.

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

raising the residual working horizon by close to 3 years on average, the reform increased the average number of days of sick leave taken by females by about 9.8 percent with respect to the mean, and reduced the days taken by males by less than 1 percent.

We have argued that these results are not due to changes in individual health, because the longer working horizon affected relatively short sick leave spells (1 to 9 days) but did not impact on longer spells, which are more likely to signal poor health conditions. Our preferred explanation combines two mechanisms, the substitution of future with current leisure, which increases sick leaves, and the fear of job loss, which reduces them. While the first mechanism prevails for females, the latter is more salient for males.

By inducing forward looking economic agents to modify their choices, pension reforms designed to enhance financial sustainability can generate unintended side effects. In this paper, we have shown that reforms that lengthen the residual working horizon by delaying eligibility can induce even middle-aged workers, who are relatively far from retirement, to reconsider the intensive margin of their labor supply decisions. This is especially true for female workers earning below-median wages, who face relatively low costs of taking opportunistic sick leaves.

In their analysis of the same pension reform, Carta and De Philippis, 2022, find that the increase in the residual working horizon encouraged middle aged Italian females to participate more to the labor market, by raising both their employment and their unemployment. Our paper complements their evidence by showing that this increase also affected in a non-negligible way the willingness of middle aged females, who were employed before the reform, to take additional sick leaves.

Tables and Figures

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

	$YR_s > 5,\ 45 \le age \le 50$, Post-treatment: 2012-16 (1)	$YR_s > 3,\ 45 \le age \le 52$, Post-treatment: 2012-14 (2)
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 (1)	Females (2)
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 taking 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.

	Probability of sick leave - Males (1)	Probability of sick leave – Females (2)	Percent days of sick leave Males (3)	Percent days of sick leave Females (4)
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 taking 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.

	Probability of sick leaves lasting 1 to 9 days – Males (1)	Probability of sick leaves lasting 10+ days – Males (2)	Probability of sick leaves lasting 1 to 9 days – Females (3)	Probability of sick leaves lasting 10+ days – Females (4)
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

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 5. Heterogeneous effects of the treatment T on the probability of taking sick leaves. Females aged 45 to 50 with at least five years to retirement before the reform: Geographical area with more and less traditional gender norms. European Value Survey (A job is right but what most women want is a home and children)

	More traditional gender norms		Less traditional gender norms	
	Males (1)	Females (2)	Males (3)	Females (4)
Pre - treatment: 2010-2011				
2010 x T	-0.021	-0.015	0.021	0.021*
	(0.013)	(0.015)	(0.019)	(0.012)
Post - treatment: 2012-2016				
2012 x T	-0.029***	0.011	-0.003	0.016
	(0.011)	(0.014)	(0.015)	(0.013)
2013 x T	-0.026*	0.044***	-0.026*	0.016
	(0.015)	(0.014)	(0.015)	(0.013)
2014 x T	-0.005	0.053***	0.006	0.033**
	(0.014)	(0.015)	(0.018)	(0.014)
2015 x T	-0.023	0.043**	-0.008	0.023*
	(0.015)	(0.017)	(0.017)	(0.014)
2016 x T	-0.027*	0.044***	0.016	0.036**
	(0.015)	(0.017)	(0.017)	(0.015)
Mean dependent variable	24.03	27.24	25.27	27.41
Number of observations	440,715	236,289	413,161	305,726
R Squared	0.1080	0.1177	0.1233	0.1083

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 taking 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 median wage

	Males		Females	
	Below median (1)	Above/at median (2)	Below Median (3)	Above/at Median (4)
Pre - treatment: 2010-2011				
2010 x T	-0.003	0.011	0.007	-0.021
	(0.014)	(0.015)	(0.015)	(0.015)
Post - treatment: 2012-2016				
2012 x T	-0.030**	-0.009	0.032**	-0.007
	(0.013)	(0.013)	(0.015)	(0.014)
2013 x T	-0.042***	-0.022	0.049***	0.010
	(0.016)	(0.014)	(0.015)	(0.014)
2014 x T	-0.012	-0.001	0.059***	0.012
	(0.016)	(0.015)	(0.016)	(0.014)
2015 x T	-0.024	-0.007	0.041**	0.005
	(0.017)	(0.016)	(0.017)	(0.014)
2016 x T	-0.016	-0.002	0.047**	-0.001
	(0.015)	(0.017)	(0.018)	(0.015)
Mean dependent variable	29.09	20.33	30.40	24.49
Number of observations	413,548	413,556	262,630	262,655
R Squared	0.0577	0.2043	0.0766	0.1803

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 7. Effects of the treatment T on the probability of taking sick leaves. Males and females aged 45 to 50 with at least five years to retirement before the reform. By employer wage policy.

	Wage premium males>Wage premium females		Wage premium males≤Wage premium females	
	Males (1)	Females (2)	Males (3)	Females (4)
Pre - treatment: 2010-2011				
2010 x T	0.0053	-0.004	0.003	-0.019
	(0.0125)	(0.011)	(0.019)	(0.018)
Post - treatment: 2012-2016				
2012 x T	-0.018	0.013	-0.033*	0.020
	(0.011)	(0.011)	(0.018)	(0.017)
2013 x T	-0.032**	0.030**	-0.028	0.028
	(0.013)	(0.012)	(0.022)	(0.018)
2014 x T	0.001	0.044***	-0.008	0.022
	(0.012)	(0.013)	(0.023)	(0.019)
2015 x T	-0.021	0.033***	-0.005	0.014
	(0.014)	(0.013)	(0.024)	(0.020)
2016 x T	-0.014	0.047***	0.007	0.010
	(0.013)	(0.014)	(0.025)	(0.021)
Mean dependent variable	25.40	27.66	25.24	26.98
Number of observations	597,088	274,255	230,116	151.030
R Squared	0.116	0.107	0.135	0.156

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

	Wage premium males>Wage premium females		Wage premium males≤Wage premium females	
	Males	Females	Males	Females
	(1)	(2)	(3)	(4)
Treatment in months	0.002**	0.001	0.001	-0.003
	(0.001)	(0.002)	(0.001)	(0.002)
Contributions paid in 2011	-0.004***	-0.002***	-0.003***	-0.002***
	(0.000)	(0.000)	(0.000)	(0.000)
Age in 2011	0.067***	0.029***	0.047***	0.027**
	(0.007)	(0.010)	(0.008)	(0.012)
Experience in 2011	-0.028***	-0.017***	-0.028***	-0.012**
	(0.004)	(0.005)	(0.004)	(0.006)
Foreigner in 2011	0.3669***	0.2476***	0.3819***	0.3074***
	(0.048)	(0.077)	(0.067)	(0.098)
Days worked in 2011	0.0020**	-0.0002	0.0008	0.0007*
	(0.001)	(0.000)	(0.001)	(0.000)
Number of observations	63,598	36,973	35,286	22,901

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.

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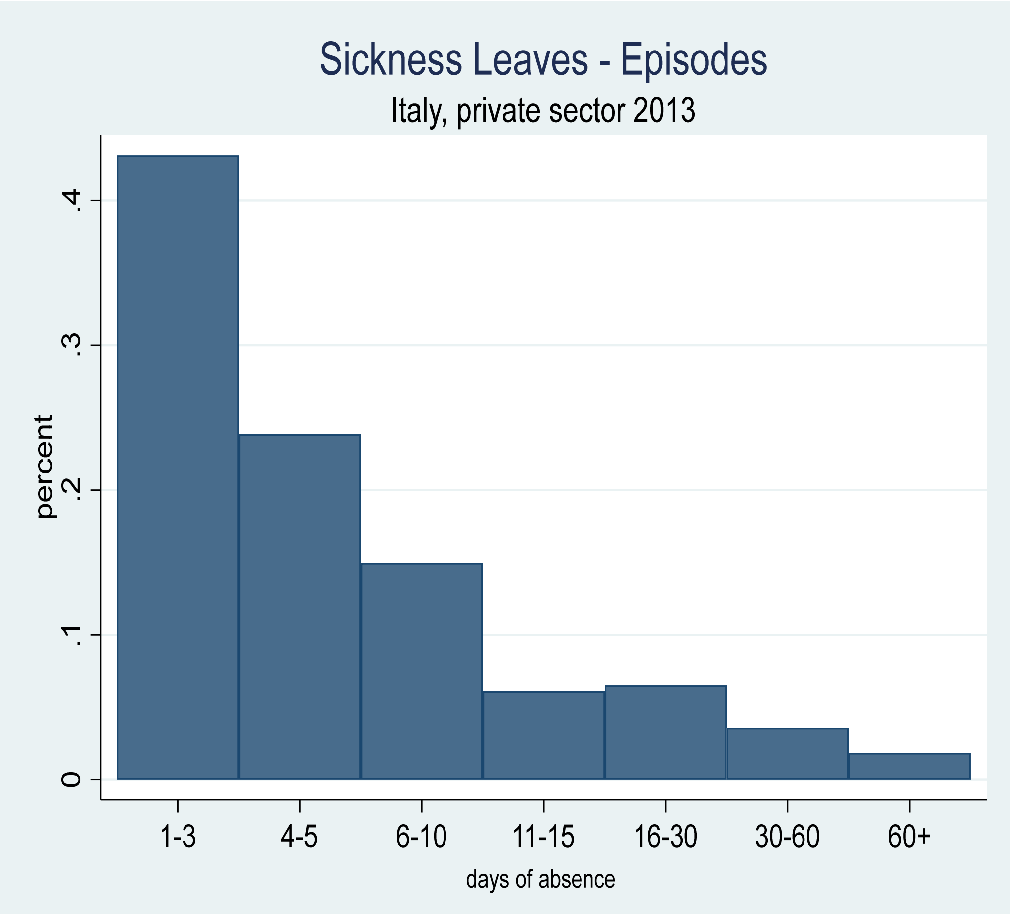
Figure 1. The distribution of sickness episodes, by duration of episodes



Source: INPS

Figure 2. Years to retirement before the reform. Males

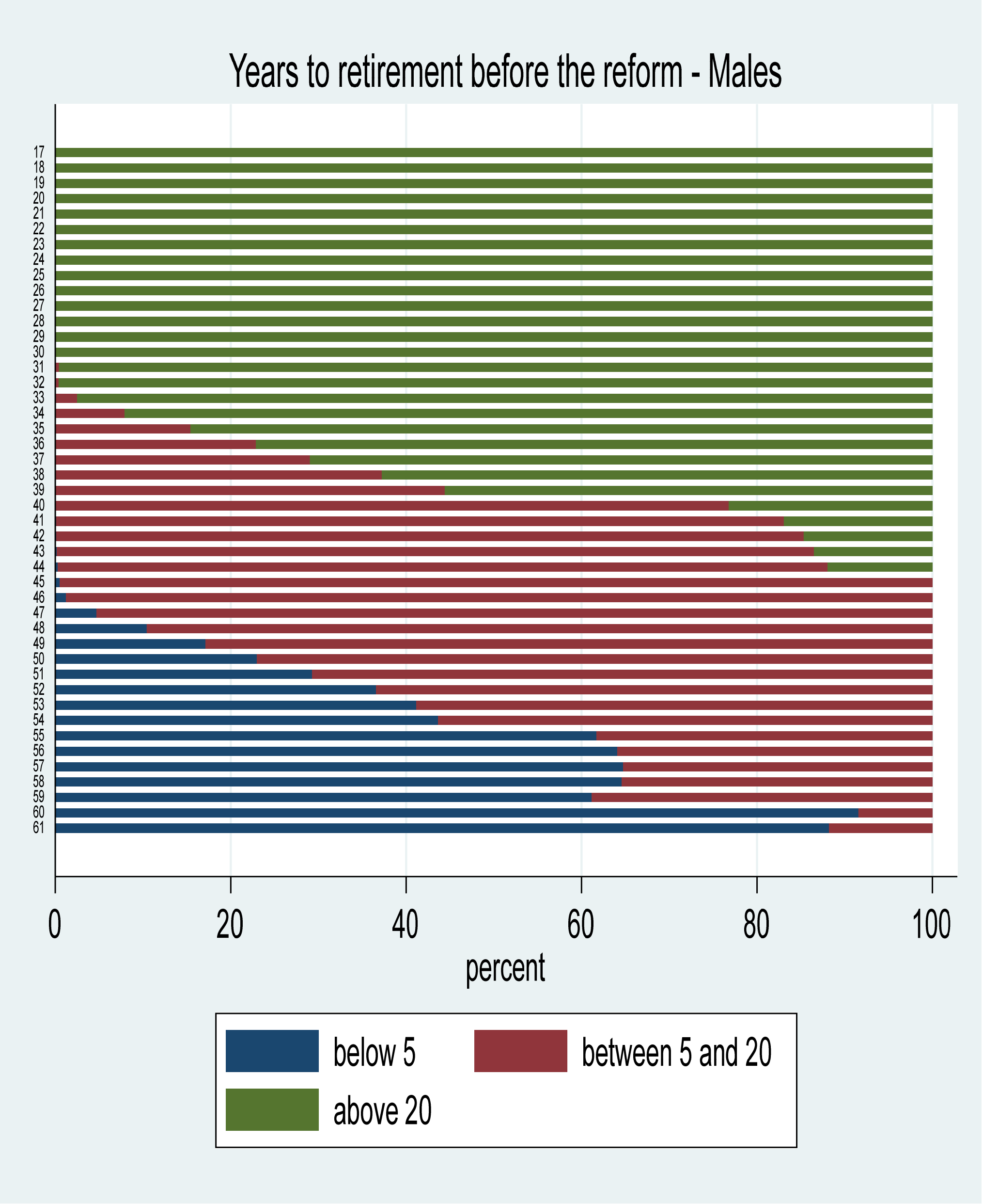


Figure 3. Years to retirement before the reform. Females

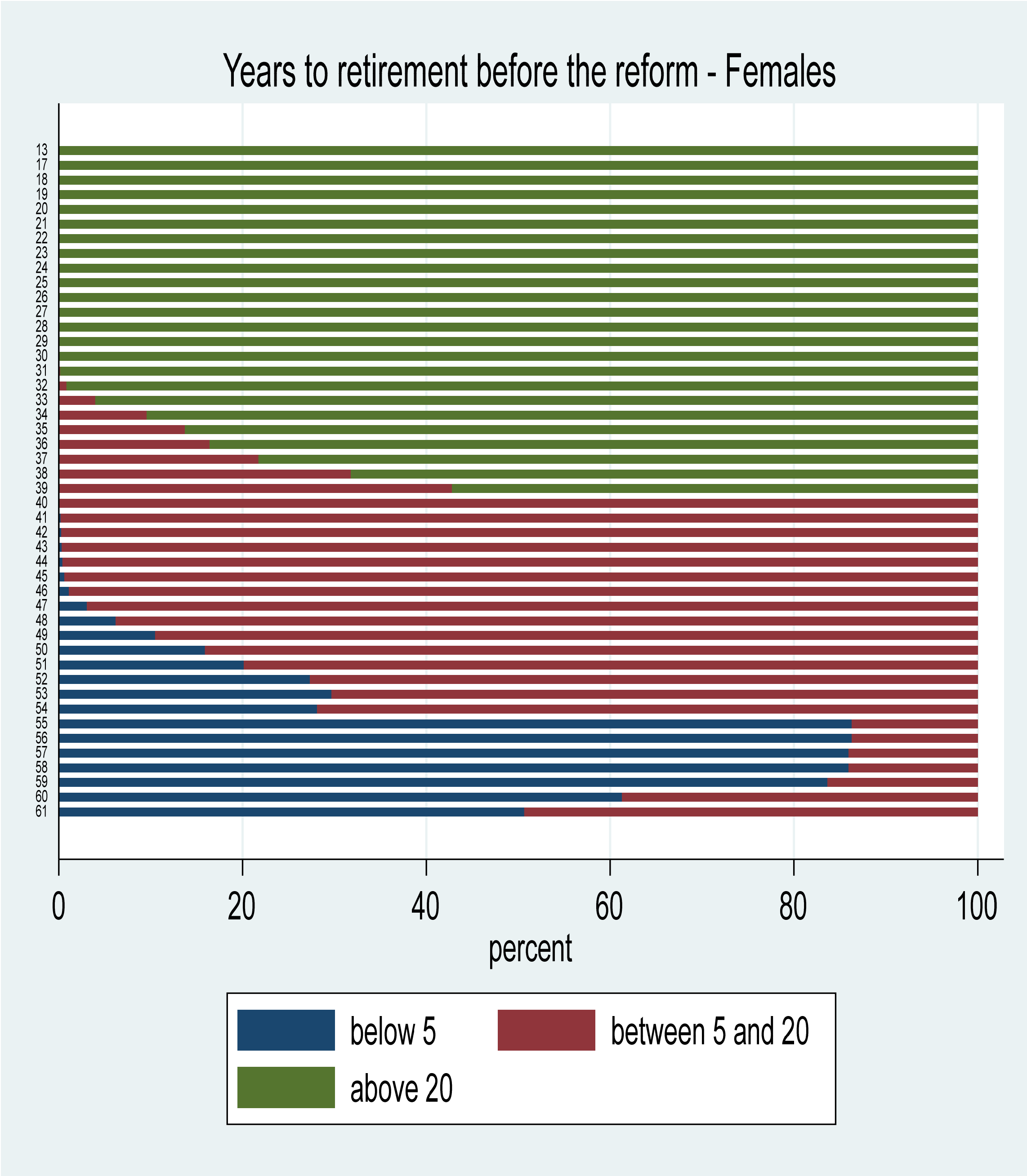


Figure 4. Distribution of the treatment T in years - Males

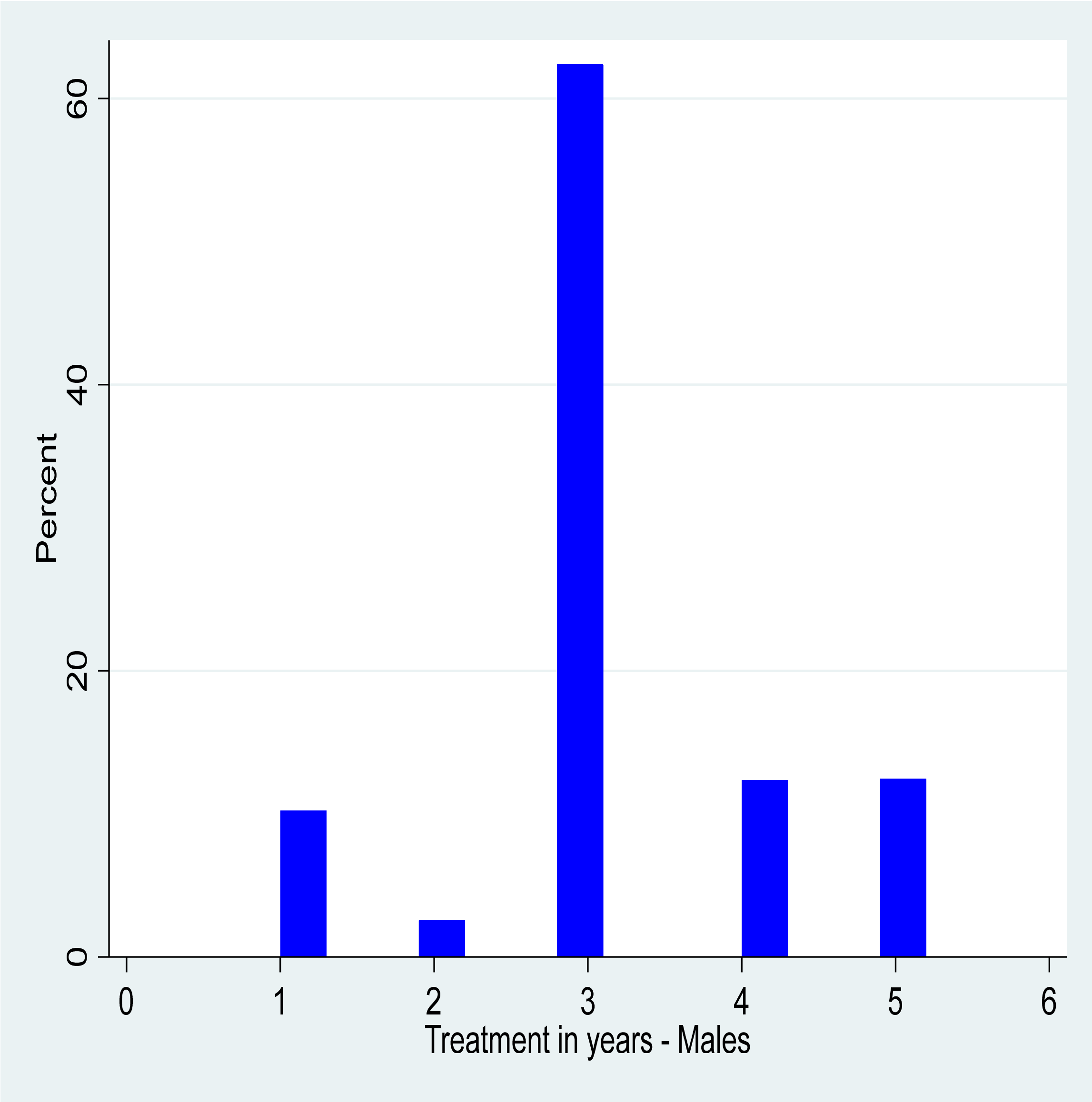
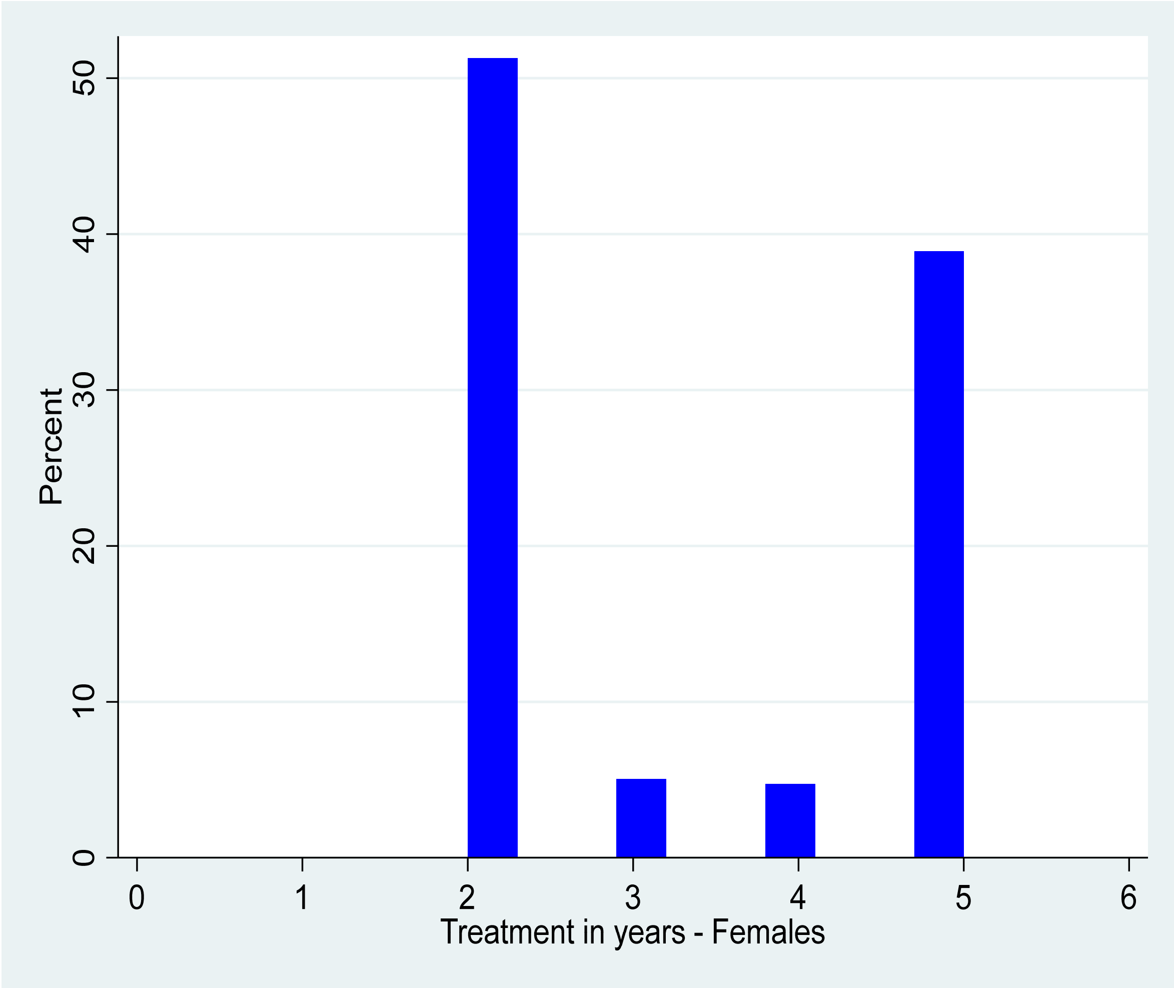


Figure 5. Distribution of the treatment T in years - Females



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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

	Probability of sick leave - males (1)	Probability of sick leave - females (2)	Percent days of sick leave – males (3)	Percent days of sick leave – females (4)
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 A4. 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 (1)	Females (2)
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.

Table A5. Effect of the treatment T on the probability of absence because of blood donations. Males and females aged 45 to 50 with at least five years to retirement before the reform.

	Probability of absence due to blood - Males (1)	Probability of absence due to blood - Females (2)
Pre - treatment: 2010-2011		
2009 x T	-0.003	-0.003
	(0.003)	(0.002)
2010 x T	0.001	-0.003
	(0.003)	(0.002)
Post - treatment: 2012-2016	0.001	0.004**
	(0.003)	(0.002)
Mean dependent variable	4.853	1.698
Number of observations	949,773	601,644
R Squared	0.023	0.007

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 A6. Heterogeneous effects of the treatment T on the probability of positive sick leaves. Females aged 45 to 50 with at least five years to retirement before the reform: Geographical area with more and less traditional gender norms. European Value Survey (A man job is to earn money, a woman job is to look after home and family).

	More traditional gender norms		More traditional gender norms	
	Males	Female	Males	Females
	(1)	(2)	(3)	(4)
Pre - treatment: 2010-2011				
	-0.0105	0.0101	0.0142	-0.0128
2010 x T	(0.0163)	(0.0177)	(0.0122)	(0.0106)
Post - treatment: 2012-2016				
2012 x T	-0.0255	-0.0028	-0.0171	0.0212*
	(0.0157)	(0.0169)	(0.0116)	(0.0116)
2013 x T	-0.0317*	0.0390**	-0.0255**	0.0247**
	(0.0186)	(0.0191)	(0.0125)	(0.0112)
2014 x T	-0.0099	0.0545***	0.0023	0.0345***
	(0.0179)	(0.0198)	(0.0128)	(0.0127)
2015 x T	-0.0153	0.0377*	-0.0193	0.0255**
	(0.0213)	(0.0208)	(0.0131)	(0.0126)
2016 x T	-0.0246	0.0355*	-0.0019	0.0364***
	(0.0207)	(0.0205)	(0.0130)	(0.0139)
Mean dependent variable	23.62	25.67	25.12	27.94
Number of observations	281,146	144,725	397,290	397,290
R Squared	0.1030	0.1087	0.1143	0.1143

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.