# Does a Higher Retirement Age Reduce Youth Employment?

Marco Bertoni^
University of Padova and IZA

Giorgio Brunello*
University of Padova and IZA

February 2020

## Abstract

Pension reforms rising minimum retirement age force some senior workers to retire later than originally expected. We evaluate the impact of a 2011 Italian reform, implemented during a recession, on youth and prime age employment. Our research design is based on difference-in-differences, and exploits the variations in the intensity of the treatment across local labor markets due to differences in the age structure of the population. We estimate that, for any 1,000 local senior workers locked into employment by the reform, local youth and prime age employment declined by 273 (-0.86%) and 199 (-0.12%) workers, and senior employment increased by 833 (+2.70%) individuals. The estimated reduction in youth employment is broadly similar to the one induced by earlier reforms, implemented when the economy was growing. We estimate that an important part of the total employment change induced by the 2011 reform is due to higher firm turnover.

**Keywords:** pension reforms; lump of labor; youth employment; local labor markets.
**JEL codes**: J26; H55; J21; J14; J11.

---

^ Corresponding author. Department of Economics and Management "M. Fanno", University of Padova, via del Santo 33, 35123 Padova - Italy. Email: marco.bertoni@unipd.it. * Department of Economics and Management "M. Fanno", University of Padova, via del Santo 33, 35123 Padova - Italy. Email: giorgio.brunello@unipd.it. The authors are grateful to the Editor, Tommaso Monacelli, our discussants Roland Rathelot and Alexandra Roulet, three anonymous reviewers and Oriana Bandiera, Axel Börsch-Supan, Giulia Bovini, Francesca Carta, Niccolò Cattadori, Emanuele Ciani, Francesco D'Amuri, Stephen Gibbons, Eric Hanushek, Marcello Messori, Raymond Montizaan, David Neumark, Lorenzo Rocco, Jan Van Ours, Enrico Zaninotto, Guglielmo Weber and the audiences at the Bank of Italy, Cosenza (AIEL), European Commission JRC – Ispra, LUISS – Rome, Maastricht, Padova, Tbilisi State University (ISET), Trento and Venice (CINTIA workshop) for helpful discussions, comments and suggestions. Financial support from the CARIPARO Foundation Starting Grant on "Education, Retirement and Household Behavior" is gratefully acknowledged. The usual disclaimer applies.

## Introduction

Over the last 20 years, pension rules have changed in many OECD countries. Reforms have increased pension eligibility age and the requirements to qualify for a full pension, in terms of years of social security contributions, and have reduced accrual rates. These measures were designed mainly to improve the financial sustainability of pension systems (OECD, 2016).[1]

By forcing senior workers to retire later, pension reforms that rise retirement eligibility age increase senior labor supply and employment. A concern often voiced in policy circles is that higher senior employment automatically reduces the number of jobs available to the young. This view has been forcefully opposed by professional economists (see Gruber and Wise, 2010 and the contributions therein, and Kalwij et al, 2010), who have criticized the so-called "lump of labor fallacy".

The fallacy is to assume that output and the total number of jobs in an economy are given. If they are, higher senior employment must imply lower youth employment. "Taken literally, …the theory says that if an additional older worker is employed, one younger worker must be displaced. The implication is that economies are boxed and that the box cannot be enlarged." (Gruber and Wise, 2010, p. 4). However, output and the total number of jobs are typically not given, and youth employment can even go up following a pension reform raising minimum retirement age if the total number of jobs increases sufficiently.

Whether it does is an important empirical question. In this paper, we investigate this question by looking at the Italian experience. We focus mainly on the employment effects of the pension reform implemented in 2011 (or the Monti-Fornero reform, after the names of the prime and labor ministers at the time), which significantly increased minimum retirement age in an economic environment

[1] In Germany, retirement age has been gradually increased by one month a year from 65 years and four months for the 1950 cohort to reach 67 in the future as a general rule for individuals with less than 45 years of contributions. In the Netherlands, the retirement age for basic pensions will reach 67 by 2021. In the United Kingdom, the pension age will increase to 66 in 2026 and 67 by 2028. In Belgium, the age for early retirement benefits has increased from 60.5 years in 2013 to 62 years in 2016, and the necessary contribution period has increased as well from 38 years to 40 years. In Denmark, the early retirement age is currently being increased from 60 years to 64 years in 2023. In Spain, this age is increasing in line with the change in legal retirement age from 61 to 63 by 2027 in cases of registered unemployment. See OECD, 2015.

characterized by the Great Recession of 2008 and by virtually absent economic growth, but consider also the effects of earlier reforms implemented in the 1990s, which also increased minimum retirement age when the economy was moderately growing.

In spite of the policy relevance, empirical research on the causal effects of changes in retirement eligibility age on youth employment is relatively scarce and with contrasting results, partly due to differences in the empirical approach. Studies using firms as the unit of analysis (see for instance Martins et al, 2009; Vestad, 2013; Boeri et al, 2017; Bovini and Paradisi, 2019) tend to find evidence of crowding out: firms that are exposed to increases in the retirement age because of the composition of their existing workforce reduce the employment of younger workers with respect to unexposed firms.[2] In contrast, studies using states or countries as the unit of analysis find little evidence of crowding out (e.g. Gruber and Wise, 2010; Kalwij et al., 2010; Munnell and Wu, 2012).[3]

On the one hand, firm-level studies do not capture both the employment effects associated with reform-induced firm entries and exits – because they essentially consider only surviving firms – and the indirect effects on unexposed firms which interact with exposed firms in local labor markets. On the other hand, studies using aggregate data face the challenge of finding sources of exogenous variation at that level. Since policies changing minimum retirement age affect entire countries, it is difficult to disentangle the consequences of these policies from those of concurring macroeconomic shocks, including technical progress affecting the level and composition of employment.

This paper bridges the gap between these two approaches by using local labor markets as the unit of analysis. A key feature of our strategy is that we are able to estimate local employment effects that include direct effects on treated firms,

[2]A recent exception is Carta, D'Amuri and von Watcher, 2017, who use data on a subsample of large Italian firms and find no evidence of crowding out effects.
[3] Earlier studies considering whether workers of different age groups are substitutes or complements are reviewed by Hamermesh, 1993, who concludes that the empirical evidence is rather inconclusive on the degree of substitution.

indirect effects on untreated firms and effects due to firm turnover.[4] We show that the latter contributes in a non-negligible way to local employment changes. By considering indirect effects as well as firm turnover, our design uncovers negative effects on youth employment that are larger than the ones detected by firm-level studies.

Our identification strategy relies on the fact that, while each local market is affected by national changes in minimum retirement age, the intensity of the treatment varies across markets because of differences in the local age structure. Our empirical analysis focuses on Italy, an interesting laboratory to study the issue at hand because of the several national pension reforms taking place in this country since 1992, which have increased minimum retirement age from 52 in 1996 to 66 in 2016. Following these reforms, the share of individuals aged 55 to 70 reporting to be retired from work declined from 49.4 percent in 1996 to 36.2 percent in 2016 (Italian Labor Force Survey, several waves). During the same period, the employment of individuals in the same age group increased substantially, from 2.01 to 4.27 million, but youth employment (aged 16 to 29) steadily declined, from 4.62 to 2.79 million.

We evaluate the employment effects of the Monti-Fornero reform, introduced in 2011 and effective since 2012, which abruptly increased minimum retirement ages. We measure the local intensity of the reform with *LS*, the local stock of senior workers who were eligible to retire before it but lost eligibility after its introduction. By construction, *LS* is equal to zero until 2011 and exhibits a positive discontinuity in 2012. Depending on the local age structure, the size of the discontinuity varies across local labor markets.

Adopting an approach similar to Acemoglu and Johnson, 2007, we leverage this within-country variation in a difference-in-differences setup and study the causal effects of a higher local supply of senior workers on local employment by age group. To avoid reverse causality from local employment to local *LS*, due for

---

[4] By focusing on local labor markets, we implicitly assume that the bulk of spillover effects are contained within these markets. Yet we cannot exclude that these effects are more relevant at the sectoral level, independently of location.

instance to endogenous migration, we compute the latter using local employment by age in 2011, the year before the reform. At that time the Monti-Fornero reform, which was voted by a new government as part of a package of emergency measures in December, was clearly unexpected.

We evaluate the effects of the reform on local employment, unemployment and firm turnover using data for 102 Italian provinces. We dispel concerns about the presence of non-parallel trends using pre-reform data, and condition on an extensive set of controls, which includes age group by time effects, age group by province effects and province-specific time trends.

We estimate that, for any 1,000 senior workers locked into local employment by the reform, local youth (16-29) and prime age (30-54) employment declined by 273 (-0.86%) and 199 workers (-0.12%), and senior employment (55-70) increased by 833 individuals (+2.70%). Since the average annual local increase in the stock of workers losing retirement eligibility between 2012 and 2015 because of the reform was equal to 6,782 individuals, we conclude that – for the average province – the reform caused local youth and prime age employment to decline with respect to the pre-reform period (2008-2011) by 5.83% and 0.81%, and senior employment to increase by 18.31%. Overall local employment (16-70) increased (+1.09%) but so did overall local unemployment (+0.13%). We evaluate that the employment changes induced by firm entries and exits contributed 6.6% of the estimated overall local employment change. Employment adjustments took place also along the intensive margin. In particular, we show that, for prime aged workers, full time employment decreased but part time employment increased.

We document that the increase of local *LS* because of the reform triggered a decline in the profits of local (manufacturing) firms, consistent with the higher labor costs associated with the retention of older workers and with the lower productivity of an older workforce (see e.g. Bertoni et al., 2015). Local (real) consumption also declined, mainly among households with young and prime aged heads.

Because the post-reform period in our data lasts four years, we interpret our estimates as short to medium – term effects. In countries such as Italy, terminating the employment relation is very costly, at least in the short run and especially for senior workers (see Boeri et al, 2017). This institutional feature may have hampered the ability of firms to adjust, and could be partly responsible for the positive effect of a higher local *LS* both on firm exits and on total employment.

We investigate whether the pattern of employment adjustment varied when we take a longer term view and extend the post-treatment period to seven years, until 2018. We find that firms retained fewer older workers and reduced both youth and prime aged employment to a larger extent than in the short-run.

A candidate reason why the Monti-Fornero reform has had a negative effect on youth employment is that it was implemented during a period characterized by negative GDP and employment growth, when the Italian economy was effectively a “boxed economy”. We inquire whether our findings are confirmed when we consider the effects of earlier reforms, also increasing minimum retirement age, which were implemented in the second part of the 1990s, when macroeconomic conditions were different and the Italian economy was growing, albeit at a slow pace. Compared to the 2011 reform, the changes in minimum retirement age occurring between 1996 and 1999 affected mainly prime aged individuals in their forties or early fifties and employed in the public sector. In spite of these differences, however, we find that the negative effect on youth employment was similar and not statistically different from the one detected after the 2011 reform.

The paper is organized as follows. Section 1 briefly describes the effects of the Monti-Fornero reform on minimum retirement age. Section 2 presents the data, Section 3 introduces the empirical approach and Section 4 describes the results. Conclusions and an appendix with an illustrative economic model and additional material follow.

## 1. The Monti-Fornero Reform

Before 1992, eligibility for *old-age* pension for Italian males and females was reached at 60 for employees in the private sector and the self-employed, and at 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.[5] We discuss earlier reforms in sub-section 4.3 of the paper.

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, a 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. After the reform, a male worker could retire either with 66 years of age and 20 years of contributions or with 42.08 years of contributions at any age. A female could instead retire in 2012 four years earlier, at 62, or with 41.08 years of contributions at any age.

Tables 1a and 1b illustrate for males and females the eligibility criteria in place before and after the reform, by sector and for the period 2012 to 2015. For instance, the first rows in the table report for private employees the alternative retirement requirements in 2012 before and after the reform. Our identification strategy combines these nation-wide changes in retirement eligibility with the variation in the age structure across local labor markets in a difference-in-differences design.

## 2. The Data

### *2.1 Data Sources*

We draw data from three main sources.[6] Labor market data are taken from the Italian Labor Force Survey (LFS), a quarterly survey of labor market conditions covering a representative sample of almost 77,000 households and 175,000

---

[5] See Angelini et al, 2009; Battistin et al, 2009; Brugiavini and Peracchi, 2010; Bottazzi et al, 2011; Bertoni et al, 2018, among others, for further details on these reforms.

[6] We also use data on local GDP per head from Eurostat (see https://ec.europa.eu/eurostat/data / database), individual firm data from AIDA Bureau Van Dijk, data on regional consumption from National Accounts and from the Bank of Italy’s Survey of Household Income and Wealth (SHIW). With a slight abuse of language, in this paper we use GDP and value added as synonymous.

individuals per quarter. For each year, we use the second quarter and aggregate data by local labor market and age group, using sampling weights to reproduce national aggregates.

Our sample consists of a pre-treatment period, running from the second quarter of 2008 to the second quarter of 2011, and a treatment period, from the second quarter of 2012 to the second quarter of 2015. We start from 2008 because a different pension eligibility regime was in place until 2007. By ending our sample in early 2015, we avoid the bulk of confounding effects induced by a major labor market reform, the Jobs Act, which was progressively introduced during 2015 (see Pinelli et al, 2015).

We consider three age groups: youth (16-29), prime age (30-54) and senior (55-70) individuals.[7] We identify local labor markets with provinces (NUTS 3). For each individual, we have information on the province of residence and on whether he/she works within or outside the province.[8] In Italy, provinces are administrative areas that consist of several municipalities. Typically, several provinces together form a region. At present, there are 107 provinces in Italy, and 20 regions. Since a few provinces have been either created or eliminated during the sample period, we reclassify our data to have a common number of provinces (102) during the entire sample period. In total, we have 2,448 year-by-province-by-age group observations (8 years, 102 provinces, 3 age groups).

Individual data on years of paid social security contributions, which we require to compute pension eligibility and therefore the number of locked-in workers by province, are not available in the LFS. Therefore, we construct the number of workers locked into employment by the Monti-Fornero reform in each province using the 2011 wave of ISFOL PLUS (Participation Labour Unemployment Survey), a national representative survey of the labour supply of individuals aged 18 to 74, that contains this information.

[7] Boeri et al, 2017, use the same definition of youth. Bovini and Paradisi, 2019, use instead a slightly different grouping: 16-34 for youth, 35-54 for prime age and 55 or older for the senior group.
[8] Travel to work areas are not available in the public-use Italian Labor Force Survey data used in this paper. On average, only 10 percent of employed individuals work outside of their province of residence.

Finally, the data on firm turnover are drawn from *MovImprese*, a database of the local Chambers of Commerce, which provides population-level information on the number of firms that enter and exit the register in each province and year.

### 2.2 *The local pool of locked - in individuals LS*

We define local *LS* as the stock of employed individuals in a local labor market and in a given year who were eligible to retire before the reform but lost eligibility after the reform. The computation of *LS* uses information on individual age and actual years of paid social security contributions. Since retirement rules vary by sector of employment, we assign employed individuals to their sector and the unemployed with a previous job to the sector of that job. We exclude the inactive as we cannot assign them to a sector.

To avoid the problems associated with reverse causality running from employment in the period 2012-2015 to the stock of locked - in workers during the same years, induced for instance by migration between local labour markets, we compute *LS* by using the individuals employed in 2011 - the year before the reform - and by assigning to them the age and years of contributions completed between 2012 and 2015.[9]

Descriptive statistics for *LS,* as well as for the other variables used in the analysis, are reported in Table 2. The average number of locked in workers in a province is 6.78 thousand. With more than 100 provinces, this corresponds to a total of 698 thousand workers per year between 2012 and 2015, or about 20.3% of employment in the age group 55 to 70. Average local employment during the sample period is equal to 29.39 thousand for the young age group, 160.29 thousand for the prime aged and 33.34 thousand for those aged 55 to 70. Average local unemployment (population) in the three age groups is equal to 8.92 (86.63), 13.55 (220.51) and 1.48 (111.28) thousand respectively. Finally, average local firm entries and exits are equal to 3.81 and 3.46 thousand respectively.

Further details on the composition of local *LS* are provided in Table 3. Considering

[9] We treat the sector of employment as fixed within the selected sample period and – as in Carta and De Philippis, 2019 - assume continuous employment after 2011. These choices may produce some measurement error, which is unlikely to vary across provinces.

the population of individuals aged 55-64, we find that locked – in workers are aged on average 60.3 years and are mainly males (88%) with less than high school (44.1%), employed in the private sector (44.0%) and in white collar jobs (51.8%). Compared to individuals in the same age group who are not locked in, locked-in individuals are more likely to be males, have a college degree and work as private employees.

Figure 1 shows the time series variation of aggregate $LS$ during the period 2008-2015. By definition, $LS$ is equal to zero until 2011, and exhibits a stark positive discontinuity in 2012, when the reform was implemented. The longitudinal variation after 2012 is due to the changes in eligibility requirements introduced by the Monti–Fornero reform (for instance, the minimum retirement age for females employed in the private sector or self-employed increased by one year in 2014 and 2015 with respect to 2012 and 2013 - see Tables 1a and 1b) and – to a marginal extent – to changes over time in the size of the cohorts belonging to each age group.

Figure 2 displays instead the distribution of average local $LS$ for the years 2012-2015. Different colors are for different tertiles. The dark areas are for relatively high values (between 6.2 and 51.2 thousands), and the light ones are for relatively low values (between 0 and 2.6 thousands). While dark areas are somewhat more frequent in the North, no stark geographic pattern emerges.

## 3. The Empirical Specification

We study the short to medium term effects of the local changes in $LS$ induced by the Monti-Fornero reform on local employment and unemployment by age group using the following difference-in-differences model:

$$Y_{gpt} = \sum_{g=1}^{3} \beta_g D_g LS_{pt} + \alpha_{gt} + \gamma_{pg} + \sum_{p=1}^{102} \delta_p \times t + \theta X_{gpt} + \varepsilon_{gpt} \qquad (1)$$

where $Y$ is the selected outcome, D is a binary variable identifying the age group, the subscripts $g$, $p$ and $t$ are for the group (young, prime aged and senior), the province and time; $\beta_g$, g = 1, 2, 3 are the parameters of interest; $\alpha_{gt}$ are time-by-age group dummies, which capture the impact of aggregate effects, which might differ by age group; $\gamma_{pg}$ are province-by-age group dummies; and $\delta_p \times t$ are province-specific linear time trends. By including these trends, we avoid that, for instance, when a large number of seniors in a province is symptomatic of economic decline, the observed variation of local *LS* captures the fact that the province is on a downward trend.

Since our treatment period lasts from 2012 to 2015, our estimates can detect the short and medium term effects of the 2011 reform but are silent on the long term effects. As discussed in the previous section, our sample choice is motivated by the need to avoid important concurrent factors, such as the implementation of the Jobs Act during 2015. In an extension, and with the drawback of concurrent factors in mind, we will nonetheless include data up and until 2018 – the most recent available LFS wave – and estimate longer run effects.

Following Peri and Sparber, 2011, and Moretti and Thulin, 2013, we use a specification with constant marginal effects rather than a log-log specification, which assumes constant elasticities.[10] We estimate the model on stacked data – where each observation refers to a year-by-province-by-age group combination – using Ordinary Least Squares, and allowing errors to be clustered by province. We control for time-varying province-specific confounders with the vector *X*, which includes the first lag of: local *GDP* per capita, the share of immigrants in the local population, an index of sectoral composition, *S*[11], and, for each age group, average age, local population aged 16 to 70, the percent with at least high school and the percent of males. We use lagged rather than current values of *X* to attenuate endogeneity concerns.[12]

As in Boeri et al, 2017, and Bovini and Paradisi, 2019, we use employment levels

---

[10] We compare the linear with the log-log specification using the Box-Cox transformation $Y^* = \frac{Y^\delta - 1}{\delta}$ for both the outcome and *LS*. The linear and the log-log model correspond to δ=1 and δ=0 respectively. We estimate that the linear specification is closer than the log-log one to the one which provides the best fit, obtained by setting δ=1.22 for the outcome and δ=0.43 for *LS*. In addition, the estimated semi-elasticities obtained using the values of δ which best fit the data are quite close to those reported in Table 4 (-0.74 for the young, -0.12 for the prime aged and 1.66 for senior workers).

[11] The sectoral index *S* is constructed by selecting the production sectors that in 2011 had a higher than average share of youth employment (manufacturing, construction, commerce, hotels and restaurants, business services and other services), and by computing - for each year and province in our data - the employment share of these sectors. An increase in the local index indicates a local change in the production structure that favors youth employment.

[12] Another reason for using lagged rather than current *GDP* per head is that the question we are addressing – whether pension reforms affect youth employment – is not conditional on the current level of output (see Banks et al, 2010).

rather than shares. A potential concern with using levels is that both local employment and local LS are affected by province size. Following Wozniak and Murray, 2012, we control for scale effects by using both province-by-age group dummies and time-varying (lagged) local working-age population. We prefer this specification to the one using shares (local employment over local population) because in the latter case the estimated marginal effect of a change in *LS* combines the effects on local employment with those on local population, which may also vary with *LS*. In addition, the effect of changes in local *LS* on the local employment share N (employment) /P (population) can be recovered from (1), because $\frac{\partial \ln(\frac{N}{P})}{\partial LS} = \frac{\partial \ln N}{\partial LS} - \frac{\partial \ln P}{\partial LS}$.

## 4. Results

### *4.1. The effects of the 2011 Monti-Fornero reform*

#### *Main results: employment effects by age group*

Table 4 reports the estimates of (1) and shows, for each age group, the marginal effect of local *LS* on local employment, as well as the semi-elasticity with respect to the 2008-11 mean. Our results indicate that adding one thousand individuals to the local pool of senior individuals – who could not retire because of the changes in eligibility requirements introduced by the 2011 reform – reduced local youth and prime age employment in the post-reform period by 273 and 199 units respectively. In percentage terms, that we compute with reference to the pre-treatment mean of the dependent variable, the estimated reduction was much larger for the young age group (-0.86%) than for the prime aged (-0.12%).[13] Senior local employment increased instead by 833 units (+2.70%), more than enough to compensate for the reduction in local youth and prime-age employment. Therefore, overall local

[13] If we use the classification proposed by Bovini and Paradisi, 2019, we find that employment in the age group 16-34 decreased by 1.1% and remained virtually unchanged in the age group 35-54. The smaller (larger) percent negative change experienced by the intermediate (younger) group when we reallocate the 30-34 to the youngest group may be driven by differences in labor adjustment cost. As their younger peers, these workers are more likely to be on temporary contracts and have lower employment protection than adults, who are often employed in open-ended contracts with higher guarantees. These differences may hamper the substitutability between senior and adult workers, and facilitate the substitution of the former with the young, at least in the short-run.

employment increased by 360 units (+0.16%).[14]

To assess the contribution of the reform to local age group-specific employment changes, we multiply average local *LS* between 2012 and 2015, equal to 6.78 thousand workers, by the semi-elasticities reported in Table 4. We estimate that – for the average province – local youth and prime age employment declined by 5.83% (6.78 x 0.86) and 0.81% (6.78 x 0.12) respectively, senior employment increased by 18.31% (6.78 x 2.70) and total employment increased by 1.09% (6.78 x 0.16) with respect to the pre-treatment period.

Between 2011 and 2015, Italian local employment declined on average by 13.7% for the young, by 7.8% for the prime aged and increased by 18.1% for the senior, with an overall 4.6% decline. The comparison of actual changes with the changes induced by the reform suggests that the latter accounted for very little of the former when considering the prime age group, for less than 50 percent of the former with regard to the young and for about 100 percent of the former for seniors.

*Robustness tests*

In Equation (1) we compare provinces which experienced different shocks to their labour supply due the pension reform. A source of concern is that provinces with higher *LS* could have experienced stronger economic downturns around the time of the reform and therefore be on a different economic track compared with other provinces, independently of the change in *LS*. To address this concern, we plot in Figure 3 the log of real GDP per capita between 2002 and 2011 in the provinces with *LS* above and below its median value, after partialling out province fixed effects. Reassuringly, we do not detect relevant differences across groups in the years before the treatment.

Although we will never be able to observe the counterfactual evolution of local GDP after 2011 had the reform not been implemented, this evidence supports the assumption that provinces would have continued to experience the same macroeconomic trends in the absence of the Monti-Fornero reform. In addition, we

---

[14] Interacting the sectoral index *S* with year dummies to capture differential trends across sectors does not alter these results.

also run separate regressions for provinces with either positive or negative per capita GDP growth during 2008-2011. As reported in Table A1 in the Appendix, the estimated effects are statistically similar across groups, suggesting that differences in pre-treatment GDP growth are not driving our results.

A key assumption for our empirical approach is that employment trends by age group do not vary before the treatment between provinces which are more or less exposed to changes in *LS* (the parallel trends hypothesis). To support our identification strategy, we present in Figure 4 employment trends by age group for provinces with *LS* above and below median, after partialling out province specific fixed effects. We show that, with the exception of the year 2008, when we detect a statistically significant difference between areas for youth employment, these trends are broadly parallel before the reform, but show clear differences after it. In particular, senior employment grew faster in provinces with higher than median *LS*, and youth and prime age employment declined faster than in other provinces (although the gap is not statistically significant for the last group).

We formally test for parallel pre-reform trends by using an "event study" specification and a regression format that exploits the continuous variation in *LS* while controlling for time effects, observable covariates, as well as province-specific time trends. We construct this study by fixing the level of local *LS* to its 2012 value and by interacting it with dummies for the periods 2008-09 and 2012-15, leaving 2010-11 as the baseline period (Draca et al., 2011, use a similar placebo experiment). We find (see Table 5) that the estimated parameter associated with the interaction with the 2008-09 dummy is always not statistically significant at the conventional levels of confidence. Therefore, we do not reject the null for all age groups.

Table 6 shows that our findings are qualitatively similar to the baseline results in Table 4 when we drop province-specific time-varying controls (Panel A) or include quadratic trends (Panel B).[15] Excluding province-specific linear trends (Panel C)

[15] Excluding the self-employed from both the outcome and *LS* increases the estimated impact on young workers (-1.26 rather than -0.86%) and reduces the one on prime aged workers (from -0.12 to 0%), rising marginally the overall employment effect (from 0.16 to 0.20%).

affects the impact on total employment, which becomes negative but is imprecisely estimated and not statistically different from the total effect shown in Table 4. Since we compare employment levels by age group across provinces with relatively large and small numbers of workers locked in employment by the 2011 reform, our identification strategy would be invalid if the employment trends across small and large provinces were not parallel. In support to the validity of our strategy, we show in Panel D that our results are robust to substituting the province-specific time trends with time trends that vary both with the age group and with the deciles of population in 2008, which capture how employment by age group evolves over time in provinces of different size.

In Panel E of the table we exclude the year 2011 to purge our estimates from potential anticipation effects and show that results are unaffected. In Panel F we compute both provincial employment by age and provincial *LS* by considering only the individuals who live and work in the same province. Since this information is not available in the PLUS survey, we construct *LS* using LFS data, and approximate actual years of social security contributions with potential years (age minus the estimated age of completion of highest education).[16] Although total employment effects turn out to be negative, they are not statistically different from the ones presented in Table 4.

Finally, following Chetty et al., 2009, we use random permutations of *LS* across provinces as an alternative to clustered standard errors to carry out inference in difference-in-differences. We find that the effects of *LS* on youth, prime age and senior employment are significant at better than the 1 percent level of confidence and conclude that the cluster-robust variance estimator is conservative.

*Heterogeneous effects*

Specification (1) assumes that the effects of *LS* on local outcomes vary by age group. These effects, however, could also vary along other dimensions, including gender, type of job (public versus private), sector of activity and occupation. We classify sectors into two groups, tradables and non-tradables, with the former

[16] We further discuss this approximation in sub-section 4.2.

including manufacturing and the latter including all other sectors with the exception of agriculture, mining and government (see Moretti, 2010, for a similar classification). We classify occupations into white collar (ISCO groups 1-4: managers and entrepreneurs, professionals, technicians and clerks) and blue collar and service jobs (ISCO groups 5-9).

Table 7 shows the effects of changes in local *LS* on employment by gender (Panel A), public or private employment (Panel B), sector (Panel C), occupation (Panel D) and part or full time jobs (Panel E). We find that the negative impact of higher *LS* on youth employment is larger among females than males, while the opposite is true for the age group 30 to 54. It also turns out that the overall employment effect detected in Table 4 (+0.360) is entirely driven by the private sector, where the bulk of the increase of employment in the senior age group takes place (654 versus 178 individuals per thousand *LS*). Employment losses for the young age group are similar in levels (118 and 155) in the public and private sectors but much larger in percentage term in the former, due to its smaller size.

Changes in the stock of locked-in public (private) sector employees are likely to affect directly employment by age in the public (private) sector, but can also indirectly affect private (public) sector employment because of spill-over effects. For instance, a decline in public sector vacancies can have important consequences on labour supply to the private sector (see Gomes, 2015, and Caponi, 2017). We try to disentangle direct and indirect effects by estimating a version of Equation (1) where we replace aggregate *LS* with *LS* for private and public sector employees. However, due to the strong positive correlation between these two variables (0.67), we obtain imprecise results.[17]

Although we may expect employment in the tradeables sector to be less exposed to local economic conditions than in the non-tradeables sector, our estimates show little evidence that this is the case: the overall employment effect – measured as semi-elasticity – is identical for tradeables and non-tradeables, and the lower response of youth employment in the tradeables sector (0.53 percent versus -0.96

[17] These estimates are available from the authors upon request.

percent for non-tradeables) is accompanied by the higher response of senior employment (2.54% versus 0.09%).

There is also evidence that white collar occupations are more sensitive to changes in *LS* than blue collar jobs, especially among young and senior workers. Finally, we find that the effect of local *LS* on employment by age varies according to the employment contract (part versus full time). This effect is particularly evident in the case of prime aged workers, for whom we find that part time employment increased with higher *LS*, while full time employment declined, suggesting that employment adjustments to local labor supply shocks involved also the intensive margin.

*Additional outcomes*

Table 8 present the estimated effects of the 2011 reform on the number of retired individuals (Panel A), local unemployment (Panel B), labor force (Panel C), population (Panel D) and net internal migration flows (Panel E). We estimate that a one thousand increase in local *LS* reduced both the number of retired seniors in the age range 55 to 70 – by 190 individuals – and the unemployment of senior individuals – by 169 individuals, but raised the number of unemployed young and prime aged workers (by 120 and 338 individuals respectively).

The observed decline of senior unemployment is consistent with the work by Hairault et al, 2010, who documents for France that an increase in the distance to retirement stimulates job search and the acceptance of job offers. On the one hand, senior workers who face delayed retirement may find it convenient to take up a bridge job rather than facing a longer unemployment period. On the other hand, firms may experience higher resistance to dismiss workers who, in the case of dismissals, face a longer period before retirement.

An increase in *LS* reduces labour force participation by the young and raises it for the prime aged and senior. Population is broadly unaffected, as is mobility between provinces, measured by net migration flows.[18] The effects on population - small and imprecisely estimated - can be used in combination with the semi-elasticities reported in Table 4 to compute the effects of the reform on employment shares (see Section 3). We estimate that adding one thousand workers to local *LS* because of the reform reduced local youth and prime age employment shares in the post-reform period by 0.78% (-0.86+0.08) and 0.11 (-0.12+0.01) respectively, and increased the share of both senior employment – by 2.64% (2.70-0.06) – and total employment – by 0.15% (0.16-0.01).

In a similar fashion, we compute the effects of changes in local *LS* on local unemployment rates by age group and find that adding one thousand senior workers to the local pool increased youth and prime age unemployment rates in the post-reform period by 1.77 (1.69+0.08) and 3.25% (3.24+0.01), and reduced the senior unemployment rate by 16.41% (-16.40-0.01).

*4.2. Mechanisms*

We have shown that the higher local supply of senior workers caused by the 2011 reform triggered an increase in local senior employment and a decline in local youth and prime age employment during the years 2012-15. Total local employment increased, youth and prime age unemployment went up, and senior unemployment declined.

In this sub-section, we investigate the mechanisms which could help explaining these effects. The simple model presented in the Appendix illustrates the effects of changes in local *LS* on youth employment and highlight that potentially relevant mechanisms include: a) substitution effects among age groups; b) productivity changes by age group; c) local demand effects; d) firm turnover.

We discuss some of these mechanisms by presenting four pieces of evidence. First, we use the data on firm entries and exits from the *Movimprese* dataset to evaluate the contribution to total employment of the changes induced by local firm turnover.

[18] We define internal mobility as change of province of residence. This measure is likely to under-estimate true mobility if individuals who move to another province do not immediately change their official residence.

Second, we employ individual firm data from the *AIDA* databank (Bureau Van Dijk) to investigate the effects of changes in local *LS* on (manufacturing) firm profits. Third, we use National Accounts data to look at the effects of *LS* on local (regional) consumption, both total and divided into durables and non-durables. We complement these data with Bank of Italy data on consumption per capita by age group. Finally, we ask whether the uncovered short to medium term effects would change were we to consider a longer post-treatment period.

*Firm turnover*

When separation costs for older workers are very high and wages are centrally set, as in the case of Italy (see e.g. Rosolia, 2015), an increase in local *LS* that keeps these workers longer in their jobs could raise labor costs and reduce profits below zero, forcing some treated firms to exit the market, with employment losses for both young and senior workers.

Untreated firms, that have no locked-in worker, could also be affected by the reform, for instance because treated firms transfer part of the higher costs due to the retention of older workers into higher prices, thereby encouraging new firms (with younger workers) to enter the market, or because of networks of buyer-seller relationships that transmit changes of relative prices from treated to untreated firms.

Table 9 reports our estimates of the effects of local *LS* on firm turnover. We find that a thousand workers increase in *LS* raised firm exits by 32 units (0.91 percent of the 2008-11 average) and firm entries by 12 units (0.29 percent), although only the former effect is precisely estimated.[19] We evaluate that the observed average increase in local *LS* during the post-treatment period – equal to 6.78 thousand workers – raised local firm entries and exits by 81.36 (6.78 x 0.012) and 216.96 units (6.78 x 0.032) respectively. Consequently, net firm turnover – defined as the difference between entries and exits (see Davis and Haltiwanger, 1999) – declined.

Using data from Eurostat business demography (https://ec.europa.eu/Eurostat/data/database), we compute the average number of persons employed in

[19] Table A2 in the Appendix shows the results of a placebo test similar to the one discussed in the previous sub-section, which broadly support the assumption of pre-treatment parallel trends.

newborn and dead firms during the years 2012-2015 – equal to 1.25 and 1.21 respectively – and multiply this number by the estimated number of local firm entries and exits induced by the change in local *LS*. We find that the employment gains and losses associated with firm entries and exits were equal to +101.6 (81.38 x 1.25) and to -262.5 (216.96 x 1.21) workers, respectively, which corresponds to a net effect of -160.9 units. This is equivalent to 6.6 percent of the total employment gain estimated in Table 4 (6.78 x 0.360=2.441 thousand workers). These numbers suggest that the contribution of firm turnover to the employment changes associated with changes in local *LS* during 2012-15 was not negligible.

We document in Table 10 the heterogeneous effects of firm turnover by sector. First, in Panel A we find that net firm turnover responded to local *LS* shocks much less in the tradeables than in the non-tradeables sector (0.22 percent, not statistically significant versus -2.78 percent, statistically significant). Second, we investigate whether the firm turnover triggered by changes in local *LS* affected average local productivity, perhaps because firms that exited due to the higher costs were the least productive in the local economy. We address this question by using 2007 national account data on value added per head by industry to classify industries into two groups, a “high productivity” and a “low productivity” group. We treat manufacturing, energy and water, information and communication, finance, professional and technical services as high productivity sectors and the remaining sectors – net of agriculture and government – as low productivity.

Panel B of Table 10 shows that the impact of changes in local *LS* on firm exits was negative for the high productivity sector and positive for the low productivity sector (-1.14% versus 1.52%), which should have increased average productivity. We also find, however, that a higher local *LS* increased firm entry in the low productivity sector and decreased for highly productive industries, with opposite effects on average productivity.[20] We conclude that the overall effects on the latter are unclear.

*The profits of firms*

We investigate the effects of local *LS* on the profits of local firms by using data on

[20] The entry of low productive firms hiring young workers could have been facilitated by the fact that they did not have to face the higher costs of retaining older workers.

incorporated manufacturing enterprises for the period 2008 to 2015 (data from the AIDA databank). We consider both firms always present in the sample and firms present only for part of the sample. Overall, we have data on the operating profits of 26,950 manufacturing firms, which we aggregate at the province level. We estimate a version of the difference-in-differences model in Eq. (1) that includes as explanatory variables local *LS*, province and time dummies, provincial trends and the controls in vector *X*, using operating profits as the dependent variable. Results (not tabulated for brevity) show that a one thousand units increase in local *LS* triggered a 2.84 percent (standard error: 1.47 and p-value: 0.056) decline in local real profits.

*Local consumption*

Changes in local *LS* could affect local employment because of changes in local consumption. On the one hand, since fewer individuals were retiring because of a higher *LS*, the retirement consumption puzzle (see Banks et al, 1998) suggests that local consumption (by seniors) increased because of the treatment. On the other hand, employment increased among the old but declined among the young and the prime aged, with potentially negative effects on local consumption (by the young and prime aged).

We use the same empirical specification adopted for profits to estimate the effect of local changes in *LS* on local consumption, using regions rather than provinces as our unit of observation because the territorial accounts data that we use are only available at the regional level. As shown in Table 11, our estimates indicate that an increase of local *LS* by one thousand individuals triggered a decline in total local real consumption (-0.12%), the consumption of durables (-0.26%), non-durables (-0.14%) and services (-0.09%). During the pre-treatment period – between 2008 and 2011 – total real regional consumption was equal on average to 51.67 billion (real) euro. Our estimates indicate that the observed increase in *LS* during the period 2012 to 2015 – equal to 6.78 thousand workers – reduced average regional total consumption by 0.42 billion (real) euro, less than 1 percent of the pre-treatment average.

As expected, consumption per capita did not decline equally across individuals belonging to different age groups. We document this by using data on the consumption of non-durables drawn from the Bank of Italy survey SHIW (Survey on the Income and Wealth of Italian Households). Table A3 compares the change in real consumption per capita by age group between 2010 and 2014 for the regions with a share of *LS* on the population above and below the median. We find that real consumption for the young and prime aged declined more in the areas with a more intense treatment, and that real consumption for the senior group increased more in the areas less exposed to the treatment.

*Institutional constraints and longer term effects*

The observed increase in total local employment – documented in Table 4 – in the presence of increased firm exits, lower (manufacturing) profits and lower consumption, could reflect institutional constraints limiting the extent of employment adjustment, especially in the short run. These constraints include the high costs of terminating employees on open ended contracts (see Boeri et al, 2017).

If these constraints were operating mainly in the short term, we should observe larger employment changes in the longer run, as firms adjust to the negative local demand effects documented above (lower local profits and lower local consumption), which should affect all age groups. To investigate this possibility, we add to our sample period three additional years, from 2016 to 2018, and report our results in Table 12. Compared to our baseline results in Table 4, we find that the effect of a higher local LS is positive but smaller for senior employment (2.16 vs. 2.70%), negative and larger in absolute value for youth (-1.58 vs -0.86%) and prime age (-0.26 vs. -0.12%) employment. In addition, the effect on total employment turns over time from positive to negative (0.16 vs -0.11%). Although we are cautious when interpreting these estimates because of the potential confounding effects of other labor market policies implemented in 2015 and after, they confirm that in the longer run the impact of short term institutional constraints is less binding, which allows firms to retain fewer older workers and further reduce youth and prime aged employment.

*4.3 The effects of previous reforms*

It is natural to ask whether the findings presented in Table 4 are specific to the selected period and reform or could be extended to other periods and reforms. After all, the period under study – 2008 to 2015 – includes the Great Depression of 2008 and has been characterized by poor or absent economic growth. While real Italian GDP grew more or less continuously from 1996 to 2007, it declined sharply in 2009 and remained below its 2005 level until the end of the sample period in 2015.[21] Perhaps our findings can be explained with the fact that Italy during the period 2008 to 2015 was indeed rather similar to a "boxed" economy – or an economy with constant or even declining total employment.[22]

If our results were specific to a rather special period when GDP growth was absent or even negative, they should change when we consider a different period, characterized by positive, albeit moderate, income growth, as the one spanning the years 1992 to 1999. During this period, two reforms tightened the eligibility requirements for seniority pensions, trapping workers too young to retire with the new criteria.

The combined Dini reform of 1995 and Prodi reform of 1997 introduced minimum retirement age for *seniority* pensions, which was set in 1996 at 52 for private and public employees and at 56 for the self-employed (all with 35 years of contributions) and was due to rise to 55 (and 35 years of contributions) in 1999 for private sector employees, to 53 for public sector employees and to 57 for the self-employed. Retirement at younger ages was guaranteed to employees in the private and public sector with 36 to 37 years of paid contributions and to the self-employed with 40 years of paid contributions. Table 13 presents the eligibility rules for seniority pensions between 1996 and 1999 as they apply to males and females before and after the reforms.[23]

[21] Real GDP growth between 2011, the year before the implementation of the Monti-Fornero reform, and 2015, the end of our sample period, was negative (-3.5%). In contrast, real GDP grew by 6.3% between 1995 and 1999.

[22] We thank, without implicating, Axel Börsch-Supan for suggesting this explanation.

[23] The Dini reform also changed the rules for pension benefit calculations applied to newly hired workers and to workers with less than 18 years of contributions at the beginning of 1996. Since these changes were applied on a national basis, their potential effects on employment are captured by the time-by-age group effects included in specification (1).

In order to compare the employment effects of the 1995/97 and 2011 reforms, we need to make a few adjustments to our data. First, the disaggregation of the Italian Labor Force Survey data at the province level that we have used for the 2011 reform is not available in the data before 2004, which only contain information on regions. There are 20 regions in Italy, each consisting of a number of provinces. We aggregate tiny Val d'Aosta to Piemonte and end up with 19 regions.

Second, since the questionnaire used to collect the survey information was changed in October 1992 (see ISTAT, 2006), comparability across years suggests that we consider the period from the second quarter of 1993 to the second quarter of 1999. In a similar fashion, we restrict the period considered for the 2011 reform between the second quarter of 2009 and the second quarter of 2015.

Third, the ISFOL PLUS survey was not available in the 1990s. Therefore, we approximate the actual years of paid security contributions, which are not available in the Labour Force Survey, with the potential years, computed as the difference between age and age at completion of the highest education degree.[24] Figure A1 in the Appendix compares average *LS* computed using PLUS and LFS data for the period when both measures are available. Clearly, the LFS measure based on potential experience over-estimates *LS*.

As for the 2011 reform, we compute for the 1995/97 reform the stock of locked-in individuals by using the number of workers in 1995, the year before these reforms actually started to operate. Figure 5 illustrates the aggregate value of *LS* during the sample period and Table 14 compares the key characteristics of locked – in workers in the two reforms (1995/97 and 2011).[25] The table shows that locked-in workers in the earlier reforms were younger than those affected by the 2011 reform (average age: 48.3 versus 57.9), less likely to be males (59.1 versus 72.7 percent), less educated (61.5 versus 52.3 percent with less than high school education) and more

[24] This is known as potential experience in the labor literature. Angelini et al, 2009; Battistin et al, 2009 and Bertoni et al, 2018, also use this approximation. In practice, we compute potential experience as age minus 15 for those with less than high school; age minus 17 for those with a short term high school diploma; age minus 19 for those with a high school diploma and age minus 23 for those with a college degree or higher education.
[25] The age range considered for the two reforms is the one where locked in workers are present. This corresponds to age 40 to 57 for the earlier reform and to age 55 to 64 for the later reform.

likely to be employed in the public sector (60.9 versus 28.6 percent).

We estimate Eq. (1) for two periods, 1993 to 1999 for the earlier reforms and 2009 to 2015 for the latter reform, and cluster standard errors by region, accounting for the small number of clusters with wild bootstrap inference techniques (see Cameron, Gelbach and Miller, 2008).[26] When interpreting the estimates, one should keep in mind that locked-in workers belong entirely to the senior age group after the 2011 reform and almost entirely to the prime-age group after the 1995/97 reforms.[27]

The results reported in Table 15 show that the effect on youth employment of raising local *LS* by one thousand individuals is very similar across reforms (-121 workers in the earlier reform and -168 workers in the latter reform).[28] Since the bulk of locked-in individuals in the earlier reforms belongs to the prime age group, we find as expected that prime age employment rised (by 229 individuals) and that senior employment declined (by 114 workers). The overall employment effect was virtually zero (-6 workers) in the earlier reforms and positive (+317 workers) in the latter reform.[29]

Although the 2011 and 1995/97 reforms are not directly comparable because they affected different groups of workers, this evidence suggests that pension reforms forcing treated prime aged or senior individuals to remain employed longer reduced youth employment, independently of whether they were implemented when the economy was slightly growing or not growing at all.

## Conclusions

Previous research on the effects of higher retirement age on youth employment has considered either the firm or the country as the unit of analysis. In this paper, we

[26] The vector $X$ includes the same controls used for provincial data, with the exception of the share of immigrants, that is not available for the late 1990s.
[27] Redefining the cutoff age separating prime-aged and senior workers to 50 does not help, because close to 60 percent of locked-in workers would end up in the former group and close to 40 percent would belong to the latter group.
[28] A test of the hypothesis that these effects are statistically different rejects the null (p-value=0.464). The test is computes using seemingly unrelated estimation. Table A4 shows the results of a placebo test for the period 1993-99, which does not reject the null hypothesis of pre-treatment parallel trends.
[29] The difference between these effects is statistically different from zero (p-value of the test: 0.049).

have chosen to focus instead on local labor markets. Our strategy implies that the employment effects that we estimate include both those occurring in local surviving treated firms and the ones induced by firm exits and entries or by firm spillovers affecting also local untreated firms.

We have exploited for identification the fact that, while each local market is affected by national changes in minimum retirement age, the intensity of the treatment varies across markets because of differences in the local age structure. Using Italian data for the period 2008 to 2015 and a difference-in-differences approach, we have estimated the causal effect of a local increase in the pool of senior workers who are too young to retire on local youth, prime age and senior employment and unemployment. Our evidence suggests that senior employment increased and youth (and prime-age) employment decreased in the post-reform period. We have estimated that adding 1,000 individuals to the local pool of senior workers too young to retire reduced youth and prime age employment in the post – treatment period by 472 individuals and increased senior employment by 833 individuals.

Therefore, overall employment increased, possibly because of institutional constraints that make redundancy in Italy costly, especially in the short run and for senior workers. To confirm this, we have used a longer post-treatment period and found evidence of stronger substitution between senior and younger workers. In the longer run, total employment declined.

Considering only young and senior employees, we have found that the estimated ratio between employment losses (in absolute value) and gains is 0.33 (273/833), larger than the one estimated by Boeri et al, 2017, (about 0.2). Although both studies apply to Italy and consider the same reform, they are not directly comparable, mainly because of differences in the unit of analysis – surviving firms in Boeri et al's study vs. local labor markets in our research. We believe that the larger effects that we have uncovered are accounted by the fact that our approach considers also the effects of firm turnover and the employment changes occurring in local non-treated firms because of lower local consumption and the consequent decline in product demand, especially for local firms operating in the non-tradeables sector.

We have compared the effects on youth employment of the 2011 reform, which occurred during an economic recession, with those of the combined 1995 and 1997 reforms, which also increased minimum retirement age but did so during a period of moderate economic growth. We have found that the effects are similar, suggesting that the negative consequences of pension reforms on youth employment are invariant with respect to the economic context when the reform is implemented. This evidence does not confirm the OECD policy recommendation that – in order to avoid negative side effects – structural reforms such as those discussed in this paper should better be implemented during periods of output and employment growth (Caldera et al, 2016).

One of the main justifications for implementing pension reforms is to reduce pension-related public deficits. We conclude by considering the 2011 reform and evaluate the gains for the public budget associated with higher senior employment as well as the costs due to lower youth and prime-age employment, using the following back of the envelope calculations.

On the one hand, individuals aged 55 to 70 who are induced to work by the pension reform contribute to the budget by paying taxes (including social security contributions). In addition, we assume that those induced to work who are aged 60 plus do not draw pensions. On the other hand, individuals aged 16 to 54 who lose employment opportunities because of the reform do not pay taxes but draw unemployment benefits. To consider the effect on the current budget, we use the simplifying assumption that changes in the level and composition of employment by age do not affect average productivity and wages.

We focus on the effects on the 2015 public budget and compute taxes and social security benefits per-capita using the EUSILC (European Union Survey on Income and Living Conditions) survey, which provides information on both gross and net earnings (including those from self-employment). Furthermore, we estimate average gross pensions as 79.7 percent of net earnings before retirement (source: OECD, 2015)[30] and average unemployment benefits as 76 percent of net earnings

[30] We compute net average earnings before retirement as average net earnings in the age bracket 58 to 64.

(source: OECD, 2018). We set the percent of unemployed covered by unemployment benefits at roughly 30 percent (source: OECD, 2018). Finally, we compute the share of seniors (age 55-70) aged 60 to 70, who were eligible to retire before the reform (49.5 percent).

We estimate that each senior individual in the age bracket 55 to 70 who works because of the reform adds to the annual public budget 16,478 euro (8,660.1 euro in taxes and social security contributions plus the saved net pension, computed as 15,797 euro times the share of individuals aged 60 to 70, 49.5 percent). On the other hand, each individual aged 30 to 54 who does not work because of the reform contributes negatively to the budget 11,650.8 euro (7,326.4 euro in foregone taxes and social security contributions plus 4,324.4 euro in unemployment benefits). Finally, each individual aged 16 to 29 who is displaced from work by the reform contributes negatively 5,830.8 euro (3,515.9 euro in foregone taxes plus 2,314.9 euro in unemployment benefits).

Using the estimated annual number of locked in workers (698 thousand) and the results in Table 4, we evaluate the positive contribution to the 2015 public budget by the oldest age group at 9.581 billion euro, and the negative contribution by the young and prime-age groups at 3.838 billion euro. Therefore, the estimated overall annual gain for the 2015 public budget associated with the employment effects of the reform is 5.742 billion euro.[31]

Our estimates in Table 12 need to be interpreted with caution because of confounding labor market shocks. Yet they suggest that in the longer run the overall gain is likely to be much smaller, as the net benefits for the public budget associated with higher senior employment (7.683 billion euro) are foreseen to be slightly above the costs of higher youth and prime-age employment (7.465 billion euro), mainly because of the higher expected employment losses experienced by the young and the prime-age labor force.

[31] Other effects of the reform are not included. For instance, we exclude in these calculations the effects of changes in the pension benefit formula for those who were still covered by the defined benefit method of calculation (individuals with at least 18 years of accrued contributions by January 1996). For these individuals, the 2011 reform introduced the notional defined contribution method for all working years after 2011.

Although the Monti-Fornero reform was introduced during an economic crisis, the underlying problem it addressed is long-term. Italy is rapidly ageing, with a fertility rate well below the OECD average and life expectancy among the highest in the OECD. In 2011, public pension spending was 9% of GDP, almost the highest in the OECD and well above the OECD average of 5.2% (see Barr and Diamond, 2015). In part, this high cost reflects the fact that Italy's average retirement age is one of the lowest in the OECD. In their evaluation of this reform, Barr and Diamond, 2015, concluded that it "…put Italy's long-term interests ahead of short-term political advantage. Thus they were not only right; they were brave. At a strategic level, these reforms were both right and necessary. They were right in that they brought pensions in Italy closer to best practice internationally – desirable reform directions even in the absence of a fiscal crisis…".

## References


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

Angelini, V, Brugiavini, A, and Weber, G, 2009. Ageing and unused capacity in Europe: is there an early retirement trap? *Economic Policy*, 24(59), 463-508.

Banks, J, Blundell R. and Tanner, S, 1998, Is there a retirement-savings puzzle?, *American Economic Review* 88(4): 769-788.

Banks, J, Blundell, R, Bozio, A, and Emmerson, C, 2010, Releasing jobs for the Young? Early Retirement and Youth Unemployment in the UK, in Gruber, J, and Wise, D, *Social Security Programs and Retirement around the World. The Relationship to Youth Employment*, University of Chicago Press.

Barr, N, and Diamond, P, 2015, Italy's pension reforms: facing the facts, *Il Sole 24 Ore*, 15 May

Battistin, E, Brugiavini, A, Rettore, E and Weber, G, 2009, The Retirement Consumption Puzzle: Evidence from a Regression Discontinuity Approach, *The American Economic Review*, 2209-26.

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

Bertoni, M, Brunello, G and Rocco, L, 2015. Selection and the age-productivity profile. Evidence from chess players, *Journal of Economic Behavior & Organization*, 110(C), 45-58.

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

Boeri, T, Ichino, A, Moretti, E, and Posch, J. 2019, Wage Equalization and Regional Misallocation: Evidence from Italian and German Provinces, NBER WP 25612

Bottazzi, R, Jappelli, T, and Padula, M, 2011. The portfolio effect of pension reforms: evidence from Italy. *Journal of Pensions Economics and Finance*. 10 (1), 75-97.

Bovini, G, and Paradisi, M, 2019. *Labor Substitutability and the Impact of Raising the Retirement Age*, mimeo.

Brugiavini, A, and Peracchi, F, 2010, Youth Unemployment and Retirement of the Elderly: The Case of Italy, in Gruber, J and Wise, D, *Social Security Programs and Retirement around the World. The Relationship to Youth Employment*, University of Chicago Press.

Brunello, G, Lupi, C, and Ordine ,P, 2000. Regional Disparities and the Italian NAIRU, *Oxford Economic Papers*, 52

Caldera Sánchez, A, de Serres, A and Yashiro, N, 2016. *Reforming in a difficult macroeconomic context: A review of the issues and recent literature*, OECD Economics Department Working Papers, No. 1297, OECD Publishing, Paris.

Caponi, V, 2017, The effect of public sector employment on the economy, *IZA World of Labor*, 1-10.

Cameron, A C, Gelbach, J B, and Miller, D L, 2008. Bootstrap-based improvements for inference with clustered errors. *Review of Economics and Statistics*, *90*(3), 414-427.

Carta, F, D'Amuri, F, and Von Wachter, T, 2017. *Workforce Aging, Pension Reforms, and Firm Dynamics*. Presentation held at the *NBER Summer Institute 2017,* "Aging and Social Security" Session.

Carta, F, and De Philippis, M, 2019. Working horizon and labour supply: the effect of raising minimum retirement age on middle-aged individuals. *Banca d'Italia, Temi di discussione*, *26*.

Chetty, R, Looney, A, and Kroft, K, 2009. Salience and taxation: Theory and evidence. *American Economic Review*, *99*(4), 1145-77.

Davis, S, and Haltiwanger, J, 1999. Gross Job Flows, in Ashenfelter, J. and Card, D, eds., *Handbook of Labor Economics*, North Holland, 2712-57.

Draca, M., Machin, S., and Van Reenen, J., 2011. Minimum wages and firm profitability. *American Economic Journal: Applied Economics*, *3*(1), 129-51.

Gomes, P, 2015. Optimal public sector wages. *The Economic Journal* 125, 587, 1425–1451

Gruber, J, and Wise, D, 2010. *Social Security Programs and Retirement around the World. The Relationship to Youth Employment*, University of Chicago Press.

Hairault, J, Langot, F, and Sophaseuth, T, 2010. Distance to Retirement and Older Workers' Employment: the Case for Delaying the Retirement Age, *Journal of the European Economic Association*, 1034-76.

Hamermesh, D, 1993. *Labor Demand*, Princeton University Press.

Kalwij, A, Kapteyn, A, and De Vos, K, 2010. Retirement of older workers and employment of the young, *De Economist*, 158, 341-359.

ISTAT, 2006. *La rilevazione sulle forze di lavoro: contenuti, metodologie, organizzazione,* Rome

Martins, P S, Novo, A A, and Portugal, P, 2009. *Increasing the Legal Retirement Age: The Impact on Wages, Worker Flows and Firm Performance.* IZA Discussion Paper 4187.

Moretti, E, 2010. Local multipliers. *American Economic Review, 100*(2), 373-77.

Moretti, E, and Thulin, P, 2013. Local multipliers and human capital in the United States and Sweden. *Industrial and Corporate Change*, 22(1), 339-362.

Munnell, A, and Wu, Y, 2012. *Will Delayed Retirement by the Baby Boomers lead to Higher Unemployment among Younger Workers?* Boston College, mimeo.

OECD, 2015. *Pensions at a Glance*, OECD, Paris.

OECD, 2016. *OECD Pensions Outlook*, OECD Paris.

OECD, 2018, Employment Outlook, OECD Paris.

Peri, G, and Sparber, C, 2011. Assessing inherent model bias: An application to native displacement in response to immigration. *Journal of Urban Economics*, 69 (1), 82-91.

Pinelli, D, Torre, R, Pace, L, Cassio, L, and Arpaia, A, 2017. *The Recent Reform of the Labour Market in Italy: A Review*, European Commission Discussion Paper n.072.

Rosolia, A, 2015. *On the Response of Italian Wages to the Unemployment Rate*, Bank of Italy Occasional Papers

Skirbekk, V, 2004. Age and Individual Productivity: a Literature Survey, in Feichtinger, G, ed., *Yearbook of Population Research*, Vienna.

Van Ours, J, 2009. Will You Still Need Me – When I'm 64? *De Economist*, 157 (4), 441-460

Vestad, O, 2013. *Early Retirement and Youth Unemployment in Norway*, mimeo.

Wozniak, A, and Murray, T J, 2012. Timing is everything: Short-run population impacts on immigration in US cities. *Journal of Urban Economics*, 77, 60-78.

## Tables and Figures

Table 1a. Retirement rules before and after the 2011 reform. Males, by sector of employment.

| Year | Sector | Before the 2011 reform | After the 2011 reform |
|---|---|---|---|
| 2012 | Private employees | $a \geq 65$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 60$ & $yc \geq 36$<br>$a \geq 61$ & $yc \geq 35$ | $a \geq 66$ & $yc \geq 20$<br>$yc \geq 42.08$ |
| 2012 | Public employees | $a \geq 65$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 60$ & $yc \geq 36$<br>$a \geq 61$ & $yc \geq 35$ | $a \geq 66$ & $yc \geq 20$<br>$yc \geq 42.08$ |
| 2012 | Self employed | $a \geq 65$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61$ & $yc \geq 36$<br>$a \geq 62$ & $yc \geq 35$ | $a \geq 66$ & $yc \geq 20$<br>$yc \geq 42.08$ |
| 2013-15 | Private employees | $a \geq 65.25$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61.25$ & $yc \geq 36$<br>$a \geq 62.25$ & $yc \geq 35$ | $a \geq 66.25$ & $yc \geq 20$<br>$yc \geq 42.42^{a}$ |
| 2013-15 | Public employees | $a \geq 65.25$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61.25$ & $yc \geq 36$<br>$a \geq 62.25$ & $yc \geq 35$ | $a \geq 66.25$ & $yc \geq 20$<br>$yc \geq 42.42^{a}$ |
| 2013-15 | Self employed | $a \geq 65.25$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 62.25$ & $yc \geq 36$<br>$a \geq 63.25$ & $yc \geq 35$ | $a \geq 66.25$ & $yc \geq 20$<br>$yc \geq 42.42^{a}$ |

Source: National legislation. Notes: $a$ and $yc$ are for age and minimum number of years of paid social security contributions. Each line reports different combinations of requirements that allow to be eligible for retirement. Since we do not observe in our data age and contributions in months, in the empirical analysis all requirements have been rounded to the lower integer.

[a] 42.5 in $2014-15$

Table 1b. Retirement rules before and after the 2011 reform. Females, by sector of employment.

| Year | Sector | Before the 2011 reform | After the 2011 reform |
|---|---|---|---|
| 2012 | Private employees | $a \geq 60$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 60$ & $yc \geq 36$<br>$a \geq 61$ & $yc \geq 35$ | $a \geq 62$ & $yc \geq 20$<br>$yc \geq 41.08$ |
| 2012 | Public employees | $a \geq 65$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 60$ & $yc \geq 36$<br>$a \geq 61$ & $yc \geq 35$ | $a \geq 66$ & $yc \geq 20$<br>$yc \geq 41.08$ |
| 2012 | Self employed | $a \geq 60$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61$ & $yc \geq 36$<br>$a \geq 62$ & $yc \geq 35$ | $a \geq 63.5$ & $yc \geq 20$<br>$yc \geq 41.08$ |
| 2013-15 | Private employees | $a \geq 60.25^a$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61.25$ & $yc \geq 36$<br>$a \geq 62.25$ & $yc \geq 35$ | $a \geq 62.25^b$ & $yc \geq 20$<br>$yc \geq 41.42^c$ |
| 2013-15 | Public employees | $a \geq 65.25$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 61.25$ & $yc \geq 36$<br>$a \geq 62.25$ & $yc \geq 35$ | $a \geq 66.25$ & $yc \geq 20$<br>$yc \geq 41.42^c$ |
| 2013-15 | Self employed | $a \geq 60.25$ & $yc \geq 20$<br>$yc \geq 40$<br>$a \geq 62.25$ & $yc \geq 36$<br>$a \geq 63.25$ & $yc \geq 35$ | $a \geq 63.75^d$ & $yc \geq 20$<br>$yc \geq 41.42^c$ |

Source: National legislation. Notes: $a$ and $yc$ are for age and minimum number of years of paid social security contributions. Each line reports different combinations of requirements that allow to be eligible for retirement. Since we do not observe age and contributions in months, in the empirical analysis all requirements have been rounded to the lower integer.

[a] 60.33 in 2014 and 60.5 in 2015

[b] 63.75 in 2014 and 2015

[c] 41.5 in $2014-15$

[d] 64.75 in 2014 and 2015

Table 2. Descriptive statistics. Province-by-wave data, 2008-2015. Thousand individuals or firms.

| | Mean | Std. Dev. |
|---|---|---|
| <u>*Treatment intensity*</u> – waves 2012-15 | | |
| # locked in workers - *LS* (PLUS data) | 6.78 | 9.42 |
| <u>*Main outcomes*</u> – waves 2008-15 | | |
| # Employed – age 16-29 | 29.39 | 33.12 |
| # Employed – age 30-54 | 160.29 | 191.84 |
| # Employed – age 55-70 | 33.34 | 39.75 |
| <u>*Additional outcomes*</u> – waves 2008-15 | | |
| # Unemployed – age 16-29 | 8.92 | 12.49 |
| # Unemployed – age 30-54 | 13.55 | 18.47 |
| # Unemployed – age 55-70 | 1.48 | 2.21 |
| # In the labor force – age 16-29 | 38.32 | 42.95 |
| # In the labor force – age 30-54 | 173.84 | 206.49 |
| # In the labor force – age 55-70 | 34.86 | 41.55 |
| Population – age 16-29 | 86.63 | 100.30 |
| Population – age 30-54 | 220.51 | 254.58 |
| Population – age 55-70 | 111.28 | 124.20 |
| <u>*Firm level outcomes*</u> – waves 2008-15 | | |
| # firm entries | 3.81 | 4.68 |
| # firm exits | 3.46 | 3.61 |
| Net turnover level (entries minus exits) | 0.35 | 1.30 |

Notes: 102 provinces observed for 8 years (2008-2015). *Source*: ISFOL PLUS for treatment intensity, ISTAT *Italian Labor Force Survey* for main and additional outcomes, and Italian Chambers of Commerce *MovImprese* for firm-level outcomes. Treatment intensity: equal to zero before the treatment and to *LS* during the treatment period

Table 3. Average characteristics of locked-in and not locked-in workers due to the 2011 reform. Year: 2012.

| | Locked-in | Not locked-in |
|---|---|---|
| Age | 60.3 | 58.0 |
| Female (%) | 12.0 | 42.7 |
| Below high school degree (%) | 44.1 | 43.2 |
| High school degree (%) | 32.0 | 39.6 |
| College degree (%) | 23.9 | 17.2 |
| Lives in Northern Italy (%) | 51.0 | 45.2 |
| Unemployed (%) | 5.3 | 9.2 |
| Public employee (%) | 37.0 | 38.3 |
| Private employee (%) | 44.0 | 40.1 |
| Self-employed (%) | 19.0 | 21.5 |
| Works in a tradeable industry (%) | 19.5 | 13.9 |
| Works in a non-tradeable industry (%) | 44.1 | 44.9 |
| Works in a white collar job (%) | 51.8 | 55.2 |
| Works in a blue collar or service job (%) | 42.9 | 35.6 |

Notes: estimates from PLUS micro-data for 2011 using 2012 retirement rules. The age range considered is 55-65, where locked in workers are present. Tradeable industries: manufacturing. Non-tradeable industries: everything else except agriculture and government (see Moretti, 2010). White collar professions: ISCO-08 codes 1-4. Service and blue collar jobs: ISCO-08 codes 5-9.

Table 4. The effect of the local number of locked-in workers *LS* on employment by age group (thousand individuals). Provincial data, 2008-15. Baseline.

| Outcome: # employed (thousand individuals) | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Age group | 16-29 | 30-54 | 55-70 | 16-70 |
| Estimated effect | -0.273** | -0.199* | 0.833*** | 0.360** |
| | (0.118) | (0.120) | (0.108) | (0.162) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.86 | -0.12 | 2.70 | 0.16 |
| Time-varying province-specific lagged controls | Yes | Yes | Yes | Yes |
| Wave-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-specific time trends | Yes | Yes | Yes | Yes |

Notes: Each observation is a province-by-age group-by-wave combination. Number of observations: 2,448 (102 provinces, 3 age groups, 8 waves). The employment effect for the age group 16-70 is obtained as linear combination of the effects for the other three age groups. Time-varying province-specific lagged controls are: GDP per capita; the index of sectoral composition *S*; total population by age group; the percentage of workers with high school or higher degree by age group; the percentage of males by age group; average age by age group; the share of immigrants by age group. *LS* is computed in the PLUS data using self-reported information on paid years of contributions. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 5. Test for parallel pre-reform trends. Provincial data, 2008-15.

| Outcome: # employed (thousand individuals) | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Age group | 16-29 | 30-54 | 55-70 | 16-70 |
| *LS2012*(Year=2008-09)* – Placebo test | 0.170 | -0.223 | 0.224 | 0.171 |
| | (0.132) | (0.369) | (0.261) | (0.179) |
| Semi–elasticity with respect to mean outcome in 2010-11 (%) | 0.56 | -0.14 | 0.72 | 0.08 |
| *LS2012*(Year=2012-15*) – Treatment effect | -0.338 | -0.410* | 0.971*** | 0.222 |
| | (0.210) | (0.218) | (0.231) | (0.516) |
| Semi–elasticity with respect to mean outcome in 2010-11 (%) | -1.11 | -0.25 | 3.13 | 0.10 |
| Time-varying province-specific lagged controls | Yes | Yes | Yes | Yes |
| Wave-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-specific time trends | Yes | Yes | Yes | Yes |

Notes: Each observation is a province-by-age group-by-wave combination. The baseline specification is as in Table 4. Total number of observations: 2,448 (102 provinces, 3 age groups, 8 waves). *LS2012* is the province-specific level of the shock in 2012, that we interact with dummies for the periods 2008-09 and 2012-15, leaving 2010-11 as the baseline period. The employment effect for the age group 16-70 is obtained as linear combination of the effects for the other three age groups. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 6. The effect of the local number of locked-in workers *LS* on employment by age group (thousand individuals). Provincial data, 2008-15. Sensitivities.

| Outcome: # employed (thousand individuals) | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Age group | 16-29 | 30-54 | 55-70 | 16-70 |
| *Panel A. Without time-varying controls* | | | | |
| Estimated effect | -0.301** | -0.154 | 0.833*** | 0.379 |
| | (0.146) | (0.162) | (0.111) | (0.194) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.95 | -0.09 | 2.70 | 0.17 |
| *Panel B. Including province-specific quadratic time trends* | | | | |
| Estimated effect | -0.318** | -0.242** | 0.785*** | 0.225 |
| | (0.130) | (0.106) | (0.133) | (0.199) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -1.01 | -0.15 | 2.55 | 0.01 |
| *Panel C. Excluding province-specific linear time trends* | | | | |
| Estimated effect | -0.460*** | -0.366** | 0.671*** | -0.155 |
| | (0.094) | (0.144) | (0.093) | (0.132) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -1.46 | -0.22 | 2.18 | -0.07 |
| *Panel D. Including age group by decile of 2008 population-specific linear time trends* | | | | |
| Estimated effect | -0.194* | -0.090 | 0.634*** | 0.349** |
| | (0.115) | (0.146) | (0.133 | (0.148) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.61 | -0.05 | 2.05 | 0.15 |
| *Panel E. Excluding 2011 to avoid anticipation effects* | | | | |
| Estimated effect | -0.372*** | -0.187 | 0.854*** | 0.295 |
| | (0.131) | (0.189) | (0.116) | (0.180) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -1.14 | -0.11 | 2.83 | 0.13 |
| *Panel F. Employment and LS computed using only individuals who live and work in the same province* | | | | |
| Estimated effect | -0.250* | -0.555 | 0.460*** | -0.344 |
| | (0.129) | (0.374) | (0.173) | (0.402) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.89 | -0.37 | 1.60 | -0.17 |

Notes: Each observation is a province-by-age group-by-wave combination. The baseline specification is as in Table 4. Total number of observations: 2,448 (102 provinces, 3 age groups, 8 waves). In Panel F, *LS* is computed using LFS data by approximating actual years of contributions with potential years, measured by potential experience. The employment effect for the age group 16-70 is obtained as linear combination of the effects for the other three age groups. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 7. The heterogeneous effects of the number of locked-in workers *LS* on employment by gender, public and private employment, tradeable and non-tradeable industries, white and blue collar jobs, part and full time jobs (thousand individuals). Provincial data, 2008-15.

| Outcome: # employed in thousand individuals | (1) | | (2) | | (3) | | (4) | |
|---|---|---|---|---|---|---|---|---|
| Age group | 16-29 | | 30-54 | | 55-70 | | 16-70 | |
| *Panel A. Gender* | Males | Females | Males | Females | Males | Females | Males | Females |
| Estimated effect | -0.066 | -0.222*** | -0.284** | 0.097 | 0.527*** | 0.288*** | 0.178** | 0.163 |
| | (0.062) | (0.067) | (0.113) | (0.106) | (0.057) | (0.054) | (0.085) | (0.160) |
| Semi–elasticity (%) | -0.35 | -1.71 | -0.29 | 0.14 | 2.69 | 2.57 | 0.13 | 0.18 |
| *Panel B. Public and private employment* | Public | Private | Public | Private | Public | Private | Public | Private |
| Estimated effect | -0.118*** | -0.155 | -0.067 | -0.132 | 0.178*** | 0.654*** | -0.007 | 0.367 |
| | (0.034) | (0.130) | (0.054) | (0.107) | (0.066) | (0.069) | (0.079) | (0.158) |
| Semi–elasticity (%) | -4.42 | -0.53 | -0.21 | -0.10 | 2.13 | 2.91 | -0.00 | 0.16 |
| *Panel C. Tradeable and non-tradeable industries* | Trade | Non-trade | Trade | Non-trade | Trade | Non-trade | Trade | Non-trade |
| Estimated effect | 0.035 | -0.201** | -0.037 | -0.158 | 0.205** | 0.384*** | 0.203 | 0.026 |
| | (0.049) | (0.101) | (0.031) | (0.107) | (0.084) | (0.077) | (0.142) | (0.120) |
| Semi–elasticity (%) | 0.53 | -0.96 | -0.11 | -0.18 | 4.38 | 2.54 | 0.09 | 0.01 |
| *Panel D. White and blue collar jobs* | White | Blue | White | Blue | White | Blue | White | Blue |
| Estimated effect | -0.150** | -0.123 | -0.129 | -0.070 | 0.539*** | 0.294*** | 0.260 | 0.100 |
| | (0.065) | (0.077) | (0.090) | (0.124) | (0.110) | (0.046) | (0.232) | (0.129) |
| Semi–elasticity (%) | -1.29 | -0.62 | -0.17 | -0.08 | 3.33 | 2.00 | 0.12 | 0.04 |
| *Panel E. Part and full time jobs* | Part | Full | Part | Full | Part | Full | Part | Full |
| Estimated effect | -0.051 | -0.222** | 0.265** | -0.464*** | 0.031 | 0.801*** | 0.116 | 0.245 |
| | (0.041) | (0.102) | (0.113) | (0.049) | (0.039) | (0.097) | (0.181) | (0.129) |
| Semi–elasticity (%) | -0.88 | -0.86 | 1.12 | -0.33 | 0.79 | 2.98 | 0.05 | 0.11 |

Notes: Each observation is a province-by-age group-by-wave combination. The specification adopted is the same as in Table 4. Total number of observations: 2,448 (102 provinces, 3 age groups, 8 waves). Effects on employment for ages 16-70 are obtained as linear combinations of the effects on the other three age groups. Panel C: Tradeable industries: manufacturing. Non-tradeable industries: everything else except agriculture and government (see Moretti, 2010). Panel D: White collar jobs: ISCO-08 codes 1-4. Blue collar and service jobs: ISCO-08 codes 5-9. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 8. The effect of the number of locked-in workers *LS* on retirement, unemployment, population, labor force and net migration flow by age group (thousand individuals). Provincial data, 2008-15.

| Age group | (1)<br>16-29 | (2)<br>30-54 | (3)<br>55-70 | (4)<br>16-70 |
|---|---|---|---|---|
| *Panel A. Outcome: # retirees* | | | | |
| Estimated effect | - | - | - | -0.190** |
| | | | | (0.092) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | | | | -0.39 |
| *Panel B. Outcome: # unemployed* | | | | |
| Estimated effect | 0.120 | 0.338*** | -0.169*** | 0.289 |
| | (0.084) | (0.097) | (0.041) | (0.198) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | 1.69 | 3.24 | -16.4 | 0.13 |
| *Panel C. Outcome: # in the labor force* | | | | |
| Estimated effect | -0.153* | 0.139 | 0.663*** | 0.649** |
| | (0.082) | (0.208) | (0.099) | (0.280) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | 0.21 | 0.12 | 0.31 | 0.29 |
| *Panel D. Outcome: Population* | | | | |
| Estimated effect | -0.066 | -0.028 | 0.070 | -0.024 |
| | (0.072) | (0.147) | (0.079) | (0.094) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.08 | -0.01 | 0.06 | -0.01 |
| *Panel E. Outcome: Net internal migration flow* | | | | |
| Estimated effect | 0.005 | -0.001 | -0.002 | 0.003 |
| | (0.012) | (0.009) | (0.008) | (0.019) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | 0.00 | 0.00 | 0.00 | 0.00 |

Notes: Each observation is a province-by-age group-by-wave combination. The specification adopted is the same as in Table 4. Total number of observations: 2,448 (102 provinces, 3 age groups, 8 waves), except for Panel A, where N=816 because we pool retirees across all age group and only report the total effect, and Panel D, where N=2,142 because data on internal migrations are not available for 2008. The employment effect for the age group 16-70 is obtained as linear combination of the effects for the other three age groups. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 9. The effect of the local number of locked-in workers *LS* on firm turnover (thousand workers and firms). Provincial data, 2008-15.

| | (1) | (2) | (3) |
|---|---|---|---|
| Outcome: | Net turnover | Firm entries | Firm exits |
| Estimated effect | -0.020*** | 0.012 | 0.032** |
| | (0.007) | (0.012) | (0.016) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -4.65 | 0.29 | 0.91 |
| Time-varying province-specific lagged controls | Yes | Yes | Yes |
| Wave dummies | Yes | Yes | Yes |
| Province dummies | Yes | Yes | Yes |
| Province-specific time trends | Yes | Yes | Yes |

Notes: Each observation is a province-by-wave combination. Number of observations: 816 (102 provinces, 8 waves). Time-varying province-specific lagged controls include GDP per capita; the index of sectoral composition *S*; total population; the percentage of workers with high school or higher degree; the percentage of males; average age; the share of immigrants. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 10. The heterogeneous effects of the number of locked-in workers *LS* on firm turnover (thousand workers and firms). Provincial data, 2008-15.

| | (1) | | (2) | | (3) | |
|---|---|---|---|---|---|---|
| Outcome: | Net turnover | | Firm entries | | Firm exits | |
| *Panel A. Tradeable and non-tradeable industries* | Trade | Non-trade | Trade | Non-trade | Trade | Non-trade |
| Estimated effect | -0.000 | -0.021*** | 0.004*** | 0.005 | 0.005*** | 0.026* |
| | (0.001) | (0.007) | (0.001) | (0.010) | (0.001) | (0.014) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | 0.22 | -2.78 | 1.98 | 0.15 | 1.38 | 0.98 |
| | | | | | | |
| *Panel B. High and low productivity industries* | High-prod | Low-prod | High-prod | Low-prod | High-prod | Low-prod |
| Estimated effect | -0.001 | -0.021*** | -0.007** | 0.016 | -0.006* | 0.037** |
| | (0.002) | (0.007) | (0.003) | (0.011) | (0.003) | (0.015) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | 0.51 | -2.68 | -1.64 | 0.51 | -1.14 | 1.52 |

Notes: Each observation is a province-by-wave combination. The specification adopted is the same as in Table 4. Number of observations: 816 (102 provinces, 8 waves). Panel A: Tradeable industries: manufacturing. Non-tradeable industries: everything else except agriculture and government (see Moretti, 2010). Panel B: Productivity by industry is measured in 2007 from national accounts. High (above median) productivity industries: manufacturing, energy and water, information and communication, finance, professional and technical services. Low (below median) productivity: everything else except agriculture and government. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 11. The effects of the number of locked-in workers *LS* (thousand workers) on local consumption (billion euro). Regional data, 2008-15.

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Outcome: | Total consumption | Durable goods | Non-durable goods | Services |
| Estimated effect | -0.064*** | -0.011*** | -0.030*** | -0.023*** |
| Wild cluster bootstrap p-values | [<0.01] | [0.012] | [<0.01] | [<0.01] |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -0.12 | -0.26 | -0.14 | -0.09 |
| Time-varying region-specific lagged controls | Yes | Yes | Yes | Yes |
| Wave dummies | Yes | Yes | Yes | Yes |
| Region dummies | Yes | Yes | Yes | Yes |
| Region-specific time trends | Yes | Yes | Yes | Yes |

Notes: Each observation is a region-by-wave combination. Number of observations: 152 (19 regions, 8 waves). Total consumption is the sum of consumption in durables, non-durables and services. Time-varying region-specific lagged controls include GDP per capita; the index of sectoral composition *S*; total population; the percentage of workers with high school or higher degree; the percentage of males; average age; the share of immigrants. OLS estimates. P-values reported within square brackets are obtained using wild cluster bootstrap by region (19 clusters). ***: p<.01; **: p<.05; *: p<.10.

Table 12. The long-run effect of the local number of locked-in workers *LS* on employment by age group (thousand individuals). Provincial data, 2008-18.

| Outcome: # employed (thousand individuals) | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| Age group | 16-29 | 30-54 | 55-70 | 16-70 |
| Estimated effect | -0.500*** | -0.418*** | 0.668*** | -0.251* |
| | (0.138) | (0.116) | (0.053) | (0.148) |
| Semi–elasticity with respect to mean outcome in 2008-11 (%) | -1.58 | -0.26 | 2.16 | -0.11 |
| Time-varying province-specific lagged controls | Yes | Yes | Yes | Yes |
| Wave-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-by-age group dummies | Yes | Yes | Yes | Yes |
| Province-specific time trends | Yes | Yes | Yes | Yes |

Notes: Each observation is a province-by-age group-by-wave combination. Number of observations: 3,366 (102 provinces, 3 age groups, 11 waves). The employment effect for the age group 16-70 is obtained as linear combination of the effects for the other three age groups. Time-varying province-specific lagged controls are: GDP per capita; the index of sectoral composition *S*; total population by age group; the percentage of workers with high school or higher degree by age group; the percentage of males by age group; average age by age group; the share of immigrants by age group. *LS* is computed in the PLUS data using self-reported information on paid years of contributions. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 13. Retirement with a seniority pension rules before and after the reforms of 1995 and 1997. Males and females, by sector of employment.

| Year | Sector | Before the 1995 reform | After the 1995/1997 reforms |
|---|---|---|---|
| 1996-97 | Private employees | $yc \geq 35$ | $a \geq 52$ & $yc \geq 35$<br>$yc \geq 36$ |
| 1996-97 | Public employees | $yc \geq 25$ | $a \geq 52$ & $yc \geq 35$<br>$yc \geq 36$ |
| 1996-97 | Self employed | $yc \geq 35$ | $a \geq 56$ & $yc \geq 35$<br>$yc \geq 40$ |
| 1998 | Private employees | $yc \geq 35$ | $a \geq 54$ & $yc \geq 35$<br>$yc \geq 36$ |
| 1998 | Public employees | $yc \geq 25$ | $a \geq 53$ & $yc \geq 35$<br>$yc \geq 36$ |
| 1998 | Self employed | $yc \geq 35$ | $a \geq 57$ & $yc \geq 35$<br>$yc \geq 40$ |
| 1999 | Private employees | $yc \geq 35$ | $a \geq 55$ & $yc \geq 35$<br>$yc \geq 37$ |
| 1999 | Public employees | $yc \geq 25$ | $a \geq 53$ & $yc \geq 35$<br>$yc \geq 37$ |
| 1999 | Self employed | $yc \geq 35$ | $a \geq 57$ & $yc \geq 35$<br>$yc \geq 40$ |

Source: National legislation. Notes: $a$ and $yc$ are for age and minimum number of years of paid social security contributions. Each line reports different combinations of requirements that allow to be eligible for retirement. Post-reform requirements for the years 1996 and 1997 are drawn from the 1995 Dini reform and requirements for the years 1998 and 1999 are from the 1997 Prodi reform. Requirements for old-age pensions are not reported. Under the assumption of uninterrupted working lives, the requirements for old-age pensions are never binding.

Table 14. Average characteristics of locked-in and not locked-in workers across the 2011 and 1995/97 Reforms. Years: 1996 and 2012.

| | 1995/97 Reform | | 2011 Reform | |
|---|---|---|---|---|
| | Locked-in | Not locked-in | Locked-in | Not locked-in |
| Age | 49.1 | 47.2 | 57.9 | 58.4 |
| Female (%) | 40.9 | 31.8 | 27.3 | 42.9 |
| Below high school degree (%) | 61.5 | 61.5 | 52.3 | 48.7 |
| High school degree (%) | 26.2 | 27.9 | 24.5 | 34.6 |
| College degree (%) | 12.3 | 10.6 | 23.2 | 16.7 |
| Lives in Northern Italy (%) | 40.2 | 49.0 | 47.0 | 47.4 |
| Unemployed (%) | 2.8 | 7.2 | 3.8 | 3.8 |
| Public employee (%) | 60.9 | 13.9 | 28.6 | 28.7 |
| Private employee (%) | 8.7 | 57.1 | 44.6 | 42.0 |
| Self-employed (%) | 30.4 | 29.0 | 26.7 | 29.2 |
| Works in a tradeable industry (%) | 8.8 | 26.9 | 17.2 | 14.6 |
| Works in a non-tradeable industry (%) | 22.8 | 44.7 | 42.9 | 45.8 |
| Works in a white collar job (%) | 48.6 | 37.7 | 51.9 | 50.4 |
| Works in a blue collar or service job (%) | 48.6 | 55.5 | 44.1 | 45.8 |

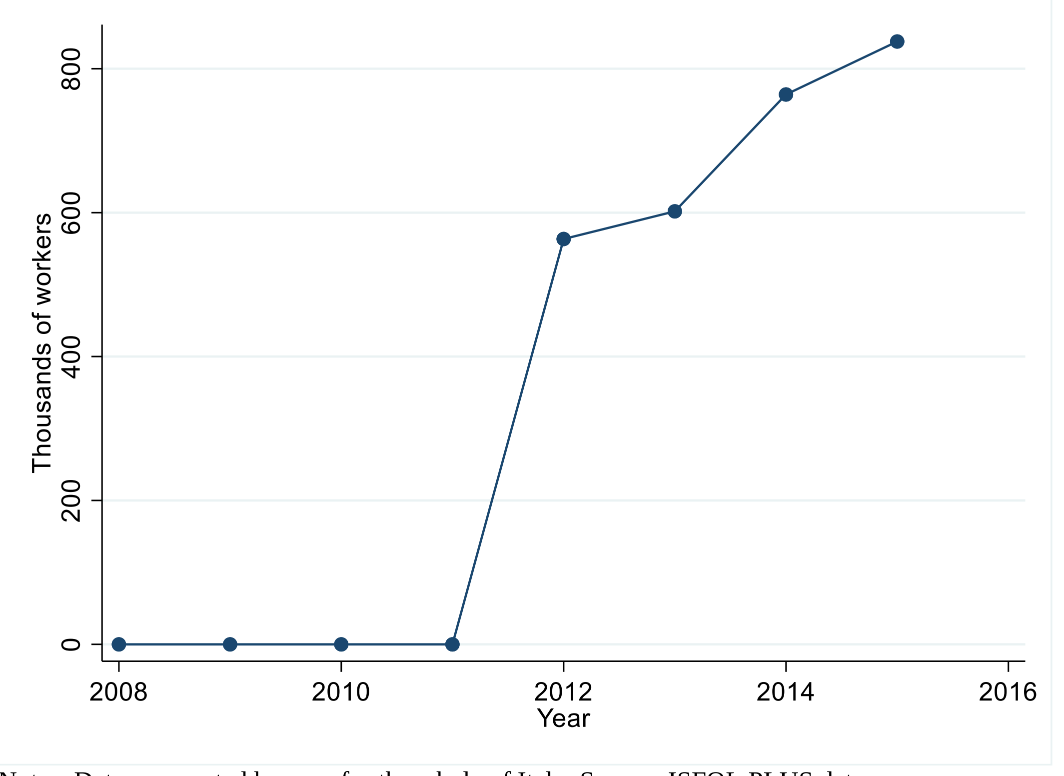
Notes: *LS* is computed using LFS microdata for 1995 using 1996 retirement rules for the 1995/97 reforms and for 2011 using 2012 retirement rules for the 2011 reform. The age range considered is the one where locked in workers are present in each reform. This corresponds to the age range 40-57 for the 1995/97 reforms and to the age range 55-64 for the 2011 reform.

Table 15. The estimated effect of the number of locked-in workers *LS* on employment by age group – comparing the 1995/97 and the 2011 Reforms (thousand individuals). Regional data, 1993-99 and 2009-15.

| Outcome: #employed (thousand individuals)<br>Age group | (1)<br>16-29 | (2)<br>30-54 | (3)<br>55-70 | (4)<br>16-70 |
|---|---|---|---|---|
| *Panel A. 1995/97 Reform* | | | | |
| Estimated effect | -0.121*** | 0.229* | -0.114** | -0.006 |
| Wild cluster bootstrap p-value | [0.001] | [0.052] | [0.045] | [0.909] |
| Semi–elasticity with respect to mean outcome in 1993-95 (%) | -0.05 | 0.03 | -0.10 | -0.00 |
| *Panel B. 2011 Reform* | | | | |
| Estimated effect | -0.168*** | -0.275*** | 0.761*** | 0.317 |
| Wild cluster bootstrap p-value | [0.019] | [<0.01] | [<0.01] | [0.368] |
| Semi–elasticity with respect to mean outcome in 2009-11 (%) | -0.10 | -0.03 | 0.45 | 0.03 |
| Time-varying region specific lagged controls | Yes | Yes | Yes | Yes |
| Wave-by-age group dummies | Yes | Yes | Yes | Yes |
| Region-by-age group dummies | Yes | Yes | Yes | Yes |
| Region-specific time trends | Yes | Yes | Yes | Yes |

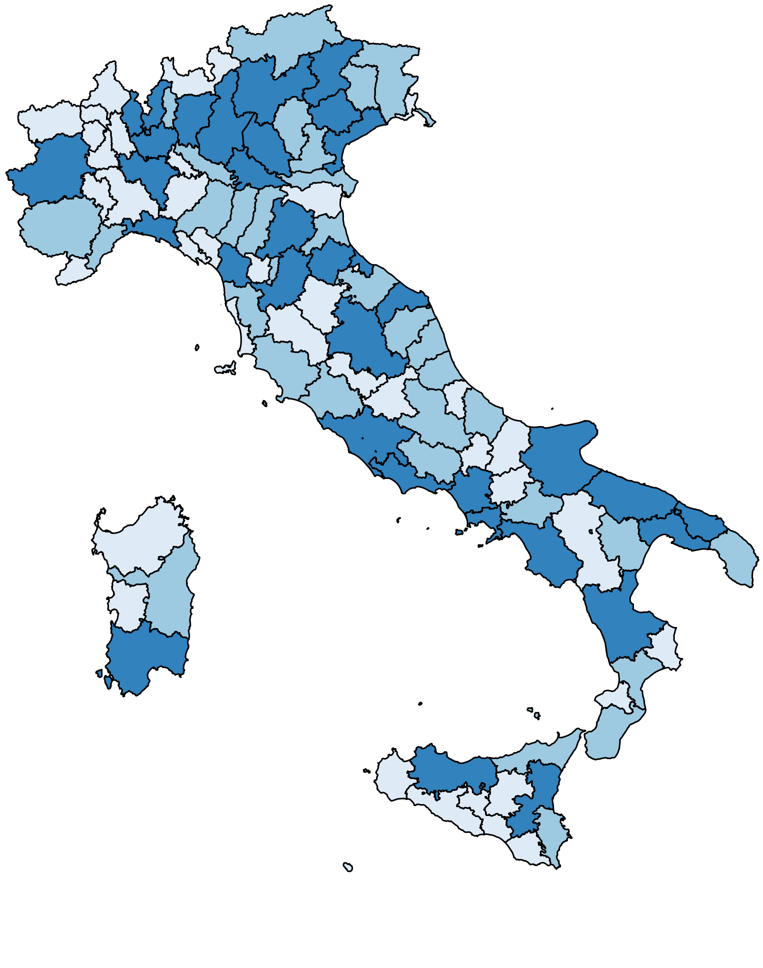
Notes: for both reforms, *LS* is computed in the LFS data using potential experience (age – school leaving age) as proxy for years of social security contributions. Each observation is a region-by-age group-by-wave combination. Number of observations: 399 (19 regions, 3 age groups, 7 waves). Effects on employment for ages 16-70 are obtained as linear combinations of the effects for the other three age groups. Time-varying region-specific lagged controls are: GDP per capita; the index of sectoral composition *S*; total population by age group; the percentage of workers with high school or higher degree by age group; the percentage of males by age group; average age by age group. OLS estimates. P-values reported within square brackets are obtained by wild cluster bootstrap by region (19 clusters). ***: p<.01; **: p<.05; *: p<.10.

Figure 1. Change of aggregate *LS*. Italy 2008-15



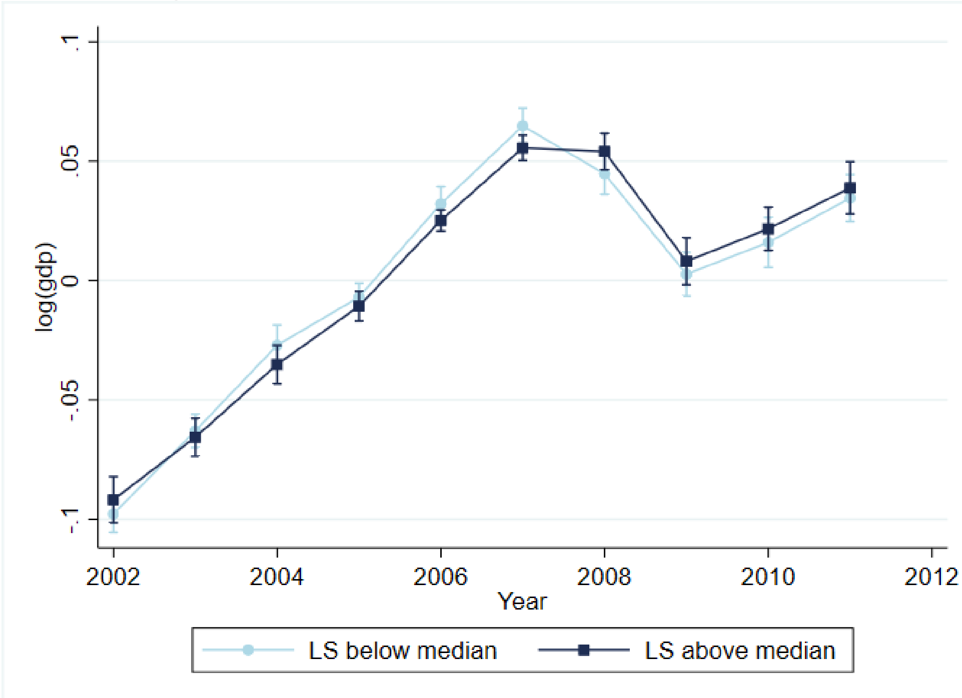

Notes: Data aggregated by year for the whole of Italy. Source: ISFOL PLUS data.

Figure 2. Distribution of average *LS* between 2012 and 2015 across provinces



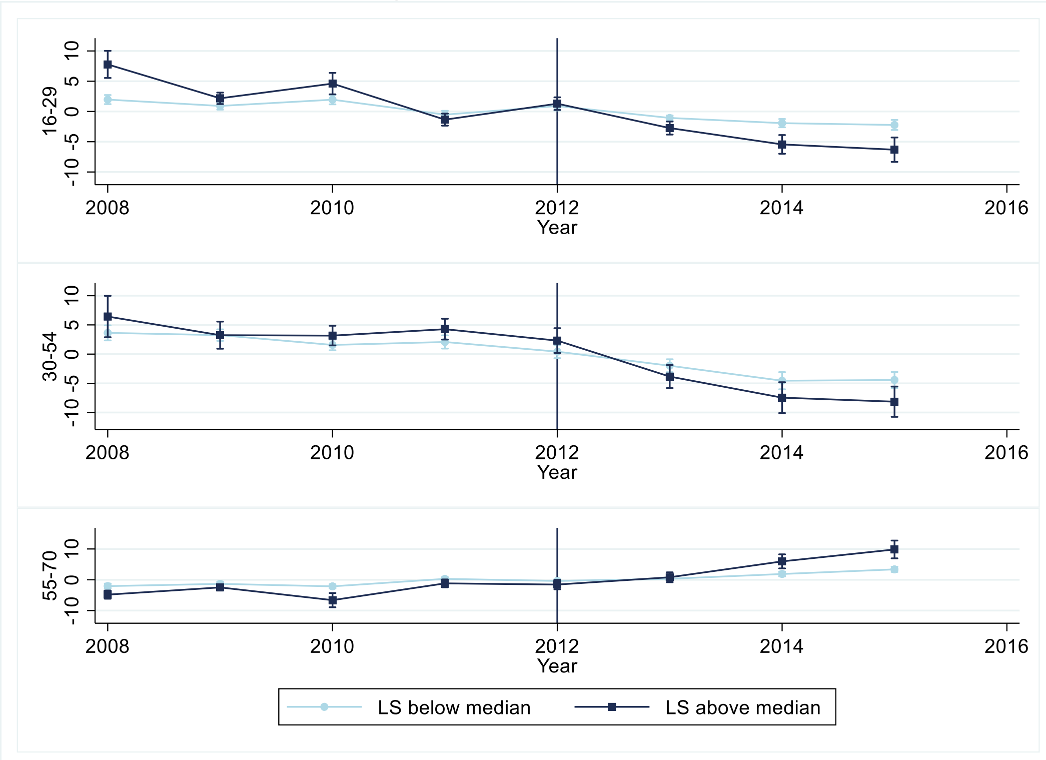
Source: ISFOL PLUS. Dark blue: top third, blue: middle third, light blue: bottom third.

Figure 3. Trends of log(GDP) for provinces with *LS* above and below the median. Residuals after partialling out province fixed effects. Period: 2002-2011



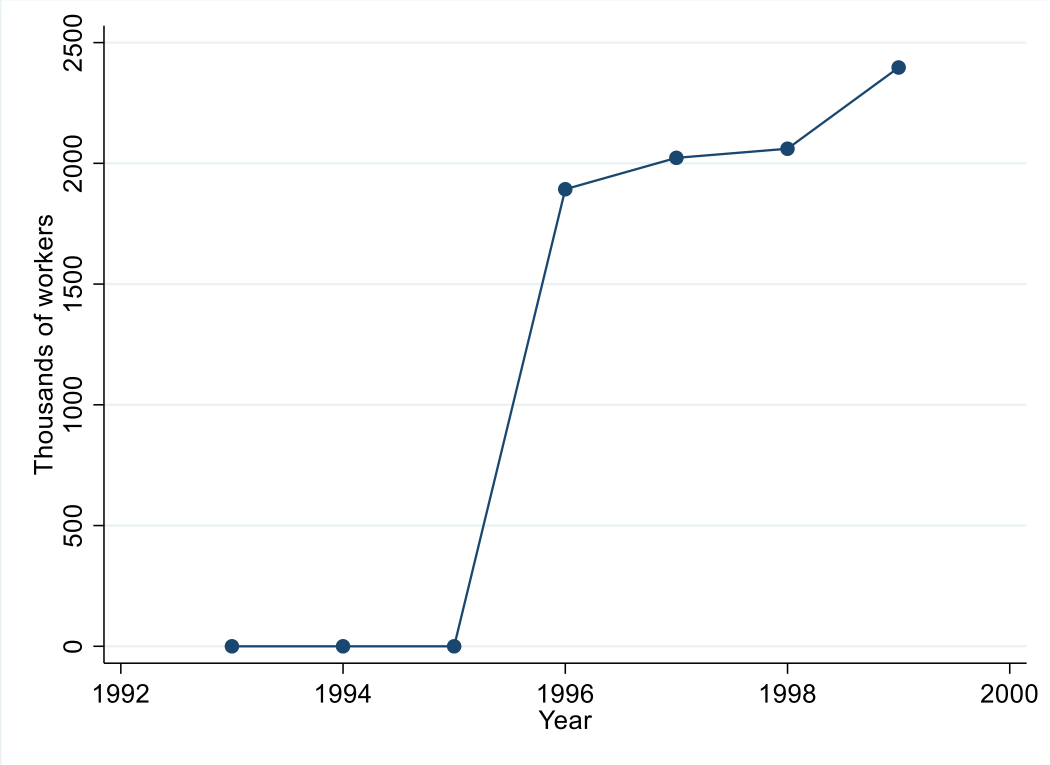

Notes: Dependent variable: residuals of log(GDP) (and 95 percent confidence intervals robust to clustering by province) after partialling out province fixed effects. Wave specific averages for provinces with *LS* above and below median are computed using the regression of the residuals on wave by *LS* group dummies.

Figure 4. Trends of employment by age group for province with *LS* above and below the median. Residuals after partialling out province fixed effects. Period: 2008-2015




Notes: Each panel is for a different age group. Dependent variable: residuals of employment after partialling out province fixed effects (and 95 percent confidence intervals robust to clustering by province). Wave specific averages for provinces with *LS* above and below median are computed using the regression of the residuals on wave by *LS* group dummies.

Figure 5. Change of aggregate *LS*. Italy 1993-1999




Source: Italian Labor Force Survey data. *LS* is computed in the LFS data using potential experience (age – school leaving age) to impute years of contributions.

**Appendix**

## 1. An illustrative theoretical framework

Consider a local economy – indexed with $j$ - where a continuum of $G_j$ firms produce goods and services by operating a concave technology with two types of workers, the young $N_y$ and the old $N_o$. The technology used by firm *i* at time *t* is

$$Y_{it} = F(E_{yi,t}, E_{oi,t}) \tag{A.1}$$

where Y is output, $e_{ai}$, with $a$=$y$,$o$, are efficiency parameters and $E_{yi,t} = e_{yi,t}N_{yi,t}$; $E_{oi,t} = e_{oi,t}N_{oi,t}$ are junior and senior employment in efficiency units.

In countries such as Italy, strict employment protection regulation makes terminating the employment relation of older workers very costly. Following Boeri et al, 2017, we capture this institutional feature in a rather extreme way by assuming that local firms cannot dismiss older employees, at least in the short run. Thus, senior employment evolves according to the following simple law of motion

$$N_{oi,t} = (1 - \delta_{it} + \omega_{it})N_{oi,t-1} \tag{A.2}$$

where $\delta_{it}$ is the percentage of old workers who retire in each period and $\omega_{it}$ is the rate of change due to demographic factors. In this setup, pension reforms rising minimum retirement age and the local pool $LS$ increases $N_{oi,t}$ by reducing $\delta_{it}$ $(\partial\delta_{it}/\partial LS_{pt})$.

Local firms select youth employment by maximizing profits $\pi_{i,t} = P_{i,t}Y_{i,t} - w_{yi,t}N_{yi,t} - w_{oi,t}N_{oi,t}$, where $w_{ai,t}$, $a = y, o$ are wages and P is the product price. We assume that

$$P_{it} = \bar{Y}_t^{\sigma} \tag{A.3}$$

implying that the inverse demand function of the firm is horizontal but shifts upwards or downwards as aggregate demand $\overline{Y_t}$ rises or falls.

Wage determination in Italy is characterized by a centralized structure, with wages responding mainly to the economic conditions prevailing in the industrialized North of the country rather than to local conditions (see for instance Brunello et al, 2000; Boeri et al, 2019).[32] We capture this institutional feature by assuming that wages are set at the national level, and that local firms take wages as given when setting employment. Therefore, $w_{ai,t} = w_{a,t},\ a = y, o$.

Profit maximization with respect to youth employment yields

$$\bar{Y}_t^{\,\sigma} \frac{\partial F(E_{yi,t}, E_{oi,t})}{\partial E_{yi,t}} e_{yi,t} - w_{y,t} = 0 \tag{A.4}$$

Total differentiation of Equation (A.4) with respect to local $LS_{pt}$, where p is for the local labour market, gives

$$\frac{\partial N_{yi,t}}{\partial LS_{pt}} = \frac{\frac{\partial^2 F}{\partial E_{yi,t}\partial E_{oi,t}}}{\left|\frac{\partial^2 F}{\partial E_{yi,t}^2}\right| e_{yi,t}} \left[ e_{oi,t} \frac{\partial N_{oi,t}}{\partial LS_{pt}} + N_{oi,t} \frac{\partial e_{oi,t}}{\partial LS_{pt}} \right] +$$
$$+ \frac{\left\{ \frac{\partial F}{\partial E_{yi,t}} \frac{1}{e_{yi,t}} - \left| \frac{\partial^2 F}{\partial E_{yi,t}^2} \right| N_{yi,t} \right\}}{\left|\frac{\partial^2 F}{\partial E_{yi,t}^2}\right| e_{yi,t}} \frac{\partial e_{yi,t}}{\partial LS_{pt}} + \frac{\frac{\partial F}{\partial E_{yi,t}} \frac{\sigma}{\bar{Y}_t}}{\left|\frac{\partial^2 F}{\partial E_{yi,t}^2}\right| e_{yi,t}} \frac{\partial \bar{Y}_t}{\partial LS_{pt}} \tag{A.5}$$

Given our assumptions, $\frac{\partial N_{oi,t}}{\partial LS_{pt}}$ has a positive sign. An increase in LS due to higher minimum retirement age rises the average age of senior workers locked in employment. If productivity declines with age, $\frac{\partial e_{oi,t}}{\partial LS_{pt}}$ is negative.[33] Thus, the sign of the expression within brackets on the right hand side of (A.5) is ambiguous. This expression is pre-multiplied by the ratio of the cross partial derivative to the absolute value of the second derivative of the function *F* with respect to youth employment (in efficiency units). The sign of this ratio is positive or negative

[32] Since the early 90s, wage determination in Italy has taken place at two complementary levels. The backbone consists of multi-year contracts negotiated at the central level by sectorial employer associations and trade unions that define both specific wage floors and employment rules at the industry level. Local agreements that redistribute productivity gains occur mainly in large firms and can add to the national floors without undoing them. See Rosolia, 2015.

[33] Skirbekk, 2004, reviews the literature on the relationship between age and productivity and concludes that productivity follows an inverted U-shaped profile, with significant decreases occurring from around age 50. See also Bertoni et al, 2015, and Van Ours, 2009. Indirect measures of productivity such as numeracy skills also show an inverted U-shaped age profile. See OECD, *The Survey of Adult Skills*, 2012.

depending on whether youth and senior employment are complements or substitutes in production.

The sign of the second component on the right hand side of (A.5) is also ambiguous, and depends both on the effect of a higher local supply of senior labor on the efficiency of youth labor and on the sign of the expression within curly brackets. Turning to the third and final component, $\frac{\partial \bar{Y}_t}{\partial LS_{pt}}$, the sign of this final component is positive (negative) if an increase in the local stock of workers who cannot retire due to the reform raises (reduces) local demand, for instance by raising (reducing) local consumption or by affecting local profits.

Expression (A.5) refers to individual firms. Youth employment in the local labor market $p$ is obtained by aggregating firm-level employment over the number of existing firms and is equal to $N_{yp,t} = \int_0^{G_{pt}} N_{yip,t} di$, where $G_{pt}$ is the number of local firms. Changes in $LS_{pt}$ could affect local employment not only by modifying employment in existing firms but also by changing the number of firms.[34]

All these mechanisms suggest that the effects of changes in local $LS$ on local youth employment can differ from the simple addition of the effects on local surviving treated firms.

[34] The differentiation of $N_{yp,t}$ with respect to $LS_{p,t}$ yields

$$\frac{\partial N_{yp,t}}{\partial LS_{pt}} = \int_0^{G_{pt}} \frac{\partial N_{yip,t}}{\partial LS_{pt}} + N_y G_{pt} \frac{\partial G_{pt}}{\partial LS_{pt}}$$

## 2. Additional tables and figures

Table A1. The heterogeneous effects of the number of locked-in workers *LS* on employment for provinces with positive and negative GDP growth in 2008-11 (thousand individuals). Provincial data, 2008-15.

| Outcome: #employed (thousand individuals) | (1) | | (2) | | (3) | | (4) | |
|---|---|---|---|---|---|---|---|---|
| Age group | 16-29 | | 30-54 | | 55-70 | | 16-70 | |
| *GDP growth in 2008-11* | Positive | Negative | Positive | Negative | Positive | Negative | Positive | Negative |
| Estimated effect | -0.310*** | -0.248 | -0.349*** | -0.131 | 0.837*** | 0.840*** | 0.177 | 0.461** |
| | (0.103) | (0.158) | (0.069) | (0.148) | (0.255) | (0.074) | (0.321) | (0.201) |
| Semi–elasticity (%) | -0.98 | -0.78 | -0.21 | -0.08 | 2.71 | 2.72 | 0.08 | 0.20 |
| P-value for equal effect | 0.72 | | 0.11 | | 0.98 | | 0.43 | |

Notes: Each observation is a province-by-age group-by-wave combination. The specification adopted is the same as in Table 4. Total number of observations: 2,448 (102 provinces, 3 age groups, 8 waves). Effects on employment for ages 16-70 are obtained as linear combinations of the effects on the other three age groups. OLS estimates. Standard errors reported within parentheses are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table A2. The effect of the local number of locked-in workers *LS* on firm turnover (thousand workers and firms). Test for parallel trends. Provincial data, 2008-15.

| | (1) | (2) |
|---|---|---|
| Outcome: | Firm entries /1,000 | Firm exits /1,000 |
| *LS2012*(Year=2008-09)* – Placebo test | -0.004 | 0.029* |
| | (0.007) | (0.016) |
| Semi–elasticity with respect to mean outcome in 2010-11 (%) | -0.09 | 0.88 |
| *LS2012*(Year=2012-15*) – Treatment effect | 0.001 | 0.049*** |
| | (0.009) | (0.014) |
| Semi–elasticity with respect to mean outcome in 2010-11 (%) | 0.04 | 1.46 |

Notes: Each observation is a province-by-wave combination. The baseline specification is as in Table 5. Number of observations: 816 (102 provinces, 8 waves). *LS2012* is the province-specific level of the shock for 2012. It is interacted with dummies for the periods 2008-09 and 2012-15, leaving 2010-11 as the baseline period. OLS estimates with standard errors clustered by province (within parentheses). ***: p<.01; **: p<.05; *: p<.10.

Table A3. Ratio of consumption per capita by age group in 2014 with respect to 2010, in regions with higher than median or lower than median *LS* per capita. Regional data.

| Outcome: Consumption per capita | (1) | | (2) | | (3) | |
|---|---|---|---|---|---|---|
| Age group | 16-29 | | 30-54 | | 55-70 | |
| *Regional LS per capita:* | Above median | Below median | Above median | Below median | Above median | Below median |
| 2014/2010 ratio | 0.926 | 0.999 | 0.961 | 0.996 | 1.030 | 1.104 |

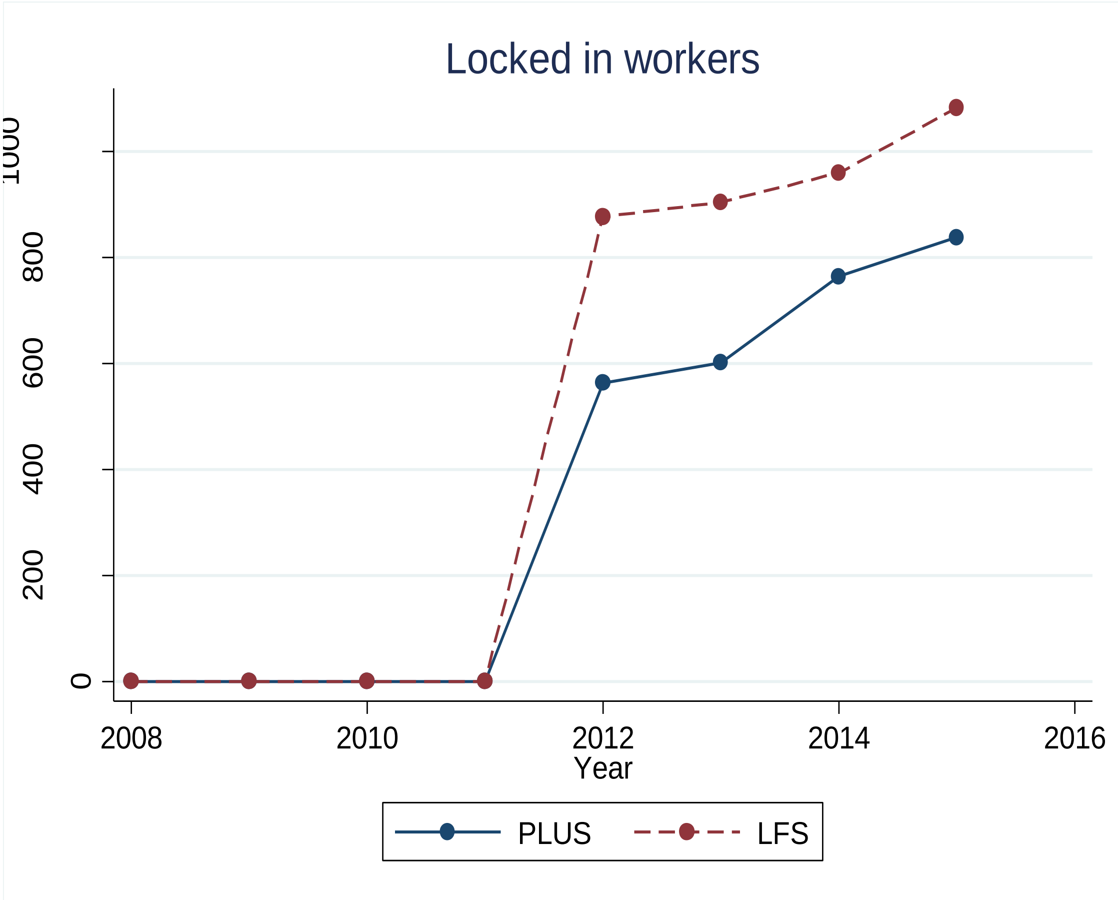
Notes: Bank of Italy regional data.

Table A4. The estimated effect of the number of locked-in workers *LS* on employment by age group – the 1995/97 reform (thousand individuals). Test for parallel trends. Regional data, 1993-99.

| Age group<br>Outcome: employment (thousand individuals) | (1)<br>16-29 | (2)<br>30-54 | (3)<br>55-70 | (4)<br>16-70 |
|---|---|---|---|---|
| *LS1996*(Year=1993)* – Placebo test | 0.071 | 0.050 | 0.094* | 0.117 |
| Wild cluster bootstrap p-value | [0.117] | [0.297] | [0.080] | [0.117] |
| Semi–elasticity with respect to mean outcome in 1994-95(%) | -0.03 | 0.00 | 0.09 | 0.02 |
| *LS1996*(Year=1996-99*) – Treatment effect | -0.131*** | 0.159* | -0.143** | -0.115*** |
| Wild cluster bootstrap p-value | [<0.01] | [0.099] | [0.014] | [<0.01] |
| Semi–elasticity with respect to mean outcome in 1994-95 (%) | -0.06 | 0.02 | -0.13 | -0.01 |

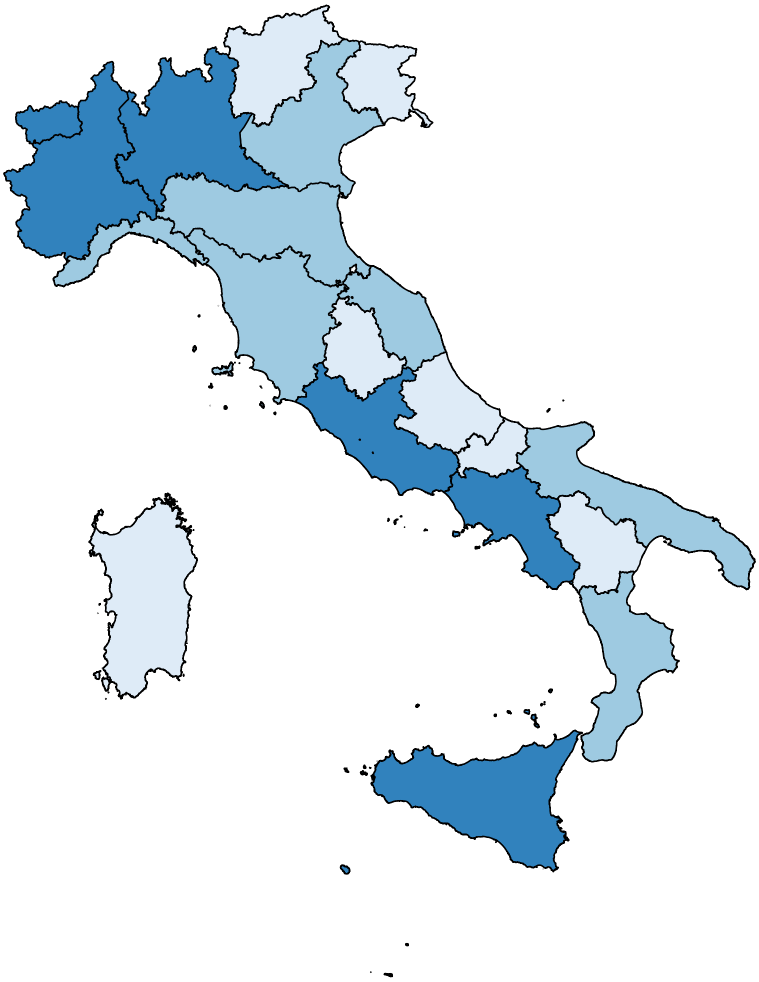
Notes: Each observation is a region-by-age group-by-wave combination. *LS* is computed in the LFS data using potential experience (age – school leaving age) as proxy for years of social security contributions. Number of observations: 399 (19 regions, 3 age groups, 7 waves). *LS1996* is the region-specific level of the shock for 1996. It is interacted with dummies for the periods 1993 and 1996-99, leaving 1993-94 as the baseline period. The employment effect for ages 16-70 is obtained as linear combination of the effects for the other three age groups. Time-varying region-specific lagged controls are: GDP per capita; the index of sectoral composition *S*; total population by age group; the percentage of workers with high school or higher degree by age group; the percentage of males by age group; average age by age group. OLS estimates. P-values reported within square brackets are obtained by wild cluster bootstrap by region (19 clusters). ***: p<.01; **: p<.05; *: p<.10.

Figure A1. The stock of locked-in workers *LS* computed using ISFOL PLUS and LFS data.




Notes: Data aggregated by year for the whole of Italy. *LS* is computed using reported years of contributions in PLUS, and using potential experience (age – school leaving age) to impute years of contributions in the LFS.

Figure A2. Distribution of average *LS* between 1996 and 1999 across regions



Source: Italian Labor Force Survey data. *LS* is computed in the LFS data using potential experience (age – school leaving age) to impute years of contributions. Dark blue: top third, blue: middle third, light blue: bottom third.