Does A Higher Retirement Age Reduce Youth Employment?

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Abstract

Pension reforms that rise minimum retirement age increase the pool of senior individuals who are not eligible to retire from the labor market. We exploit the variation in this pool across local labor markets and an instrumental variable strategy to investigate whether these reforms could have negatively affected youth employment. Using Italian data, we find that delayed retirement eligibility has increased senior employment and reduced youth employment. The negative effects on youth employment are larger in the sample period dominated by the 2008 Recession.

Keywords: pension reforms; lump of labor; youth employment; local labor markets.
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Introduction

Over the last 20 years, the pension rules applying to workers in the private and public sector have changed in many OECD countries. Reforms have included measures such as increasing pension eligibility age, increasing contributions or duration of career needed for a full pension or reducing accrual rates. These measures were designed to improve the financial sustainability of the 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, who have criticized the so-called "lump of labor fallacy" (see Gruber and Wise, 2010 and the contributions therein). The fallacy is to assume that the total number of jobs is given. If it is, higher senior employment must imply lower youth employment.

However, output and the total number of jobs are not given in a modern economy, and youth employment can even raise following pension reforms that rise minimum retirement age if the total number of jobs increases sufficiently. Whether it does is an important empirical question. Yet, in spite of the policy relevance, empirical research on the causal effects of changes in retirement eligibility age on youth employment and unemployment is relatively scarce and with contrasting results.

[1] In Germany, the retirement age has been gradually increasing 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 the basic pension will reach 66 by 2018 and 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 increase 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, the early-retirement 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.

On the one hand, the country-specific papers in Gruber and Wise, 2010, show that greater labor market participation by older persons is associated with higher youth employment. On the other hand, Vestad, 2013, uses micro data from Norway and finds that a young worker replaces an additional early retiree, and Boeri, Garibaldi and Moen, 2016, use firm-level data for Italy and conclude that five additional older workers locked in employment by an increase in minimum retirement age generate the loss of employment by one young worker.[2]

Perhaps, one reason why empirical research in this area is relatively scarce is that the identification of causal effects - from pension reforms to youth employment - is complicated by the fact that policies changing minimum retirement age affect entire countries. Therefore, 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.

In this paper, we address this identification problem by observing that national labor markets typically consist of a collection of local labor markets characterized by less than perfect labor mobility. 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.

We measure the local intensity of the treatment induced by pension reforms with the change in the local pool of individuals older than 50 but younger than minimum retirement age. Define this pool at time *t* and in the local market *p* as PT_{pt}. By delaying retirement eligibility, pension reforms increase PT_{pt}. Depending on the local age structure, this increase varies across local labor markets. Adopting an approach similar to Card, 1992, and Acemoglu and

[2] Additional recent contributions include Munnell and Wu, 2012, for the United States; Bovini and Paradisi, 2017, for Italy and Kondo, 2016, for Japan.

Johnson, 2007, we use this within-country variation to study the causal effects of changes in minimum retirement age on youth employment.

Our empirical analysis focuses on Italy, an interesting laboratory to study the issue at hand because several national pension reforms from 1992 to 2011 have increased the minimum retirement age of employees in the private sector from 52 in 1996 to above 65 for females and above 66 for males in 2016. Following these reforms, the share of individuals aged 50 to 70 who report to be retired from work declined from 41 percent in 1996 to 28.9 percent in 2016 (Italian Labor Force Survey, several waves). During the same period, the employment of individuals aged 50 to 70 increased substantially, from 3.8 to 7.6 million, but youth employment steadily declined (see Figure 1).[3]

Since the current pool PT_{pt} is affected not only by pension reforms, but also by endogenous mobility and by demographic factors, we adopt an instrumental variable strategy and select as instrument the local pool PT_{pt}^{1991}, or the value of PT_{pt} computed by applying the changes in minimum retirement age not to the current but to the local population in the 1991 Population Census, well before the start of our sample period. By using the 1991 population, the instrument isolates the local changes induced by variations in minimum retirement age from demographic changes and mobility across local areas.

We show with a simple model that, when local firms can freely hire young labor but cannot dismiss senior labor because of protective employment protection legislation, a higher pool PT_{pt} increases senior employment and the average age of senior employees, but has ambiguous effects on youth employment, even though current output is not given. By focusing on local labor markets rather than on individual firms, our approach captures not only the employment changes at the firm level but also local changes associated to

[3] The number of employees aged 16 to 24 fell from 1.8 million in 1996 to 1 million in 2016, and employment in the age group 25 to 34 fell from 5.7 to 4.1 million.

between-firms mobility, firm turnover and local spill-overs.

We estimate both the intention to treat effect of PT_{pt}^{1991} and the IV effect of current PT_{pt} on employment and unemployment by age group. Using data from 102 provinces and the period 2005-2016, we find that – in our preferred IV specification – adding one thousand additional senior individuals to PT_{pt} reduces employment for the age groups 16 to 24, 25 to 34 and 35 to 49 by 46, 142 and 72 units respectively, and increases senior (age 50 to 70) employment by 214 units. Overall, aggregate employment falls by 46 units, a small and not statistically significant effect.

Since the selected period is dominated by the Great Recession of 2008, we ask whether our results are confirmed when we consider a longer time span, that includes the second part of the nineties, when the Italian economy was (slowly) growing. Our estimates based on data for 19 Italian regions and covering the longer period 1996 to 2016[4] confirm the negative effects of pension reforms raising minimum retirement age on the employment level of young and prime age Italians and the positive effect on senior individuals. The size of the negative effects, however, is generally smaller than for the period 2005-2016. These findings suggest that the employment costs of pension reforms rising retirement eligibility age may be lower when these reforms are introduced in a growing economy.

Our results apply to Italy. Yet, as more and more countries in the OECD have increased or are increasing minimum retirement age, we believe that the interest of this study goes beyond the specific Italian case. The paper is organized as follows. Section 1 briefly reviews pension reforms in Italy and their impact on minimum retirement age, and shows how these reforms have affected the supply of older workers in Italy. Section 2 introduces a theoretical framework that illustrates how increases in minimum retirement age can affect

[4] Province-level Labor Force Survey data are not publicly available before 2004.

youth employment. Section 3 presents the empirical specification and Section 4 illustrates the data. We show our baseline empirical findings in Section 5 and several extensions and sensitivities in Section 6. Conclusions follow.

1. Pension reforms in Italy and the local pool PT_{pt}

1.1. Pension Reforms in Italy – 1992/2016

In this Section, we briefly describe how the Italian social security reforms implemented between 1992 and 2011 changed retirement eligibility rules and minimum retirement age.[5] Before 1992, the minimum age for *old-age* pension for men was 60 for employees in the private sector and for the self-employed, and 65 for public sector employees – conditional on having paid social security contributions for at least 15 years. Early retirement with a *seniority* pension was however possible at any age for workers who had paid social security contributions for at least 35 years.

Motivated mainly by the need to contain pension expenditure in a rapidly ageing society, the first social security reform in 1992 introduced a progressive increase in the requirements for eligibility to *old age* pensions, that were to reach age 65 by 2001. In 1995, a second major reform tightened the eligibility requirements for *seniority* pensions, which were to rise gradually from 1996 to 2008 and reach either 40 years of paid contributions independently of age, or 57 years of age and 35 years of paid contributions. This reform also prescribed tighter eligibility requirements for the self-employed.

After only three years, pension eligibility rules changed again with a reform that accelerated the transition period and increased minimum retirement age to 58 years for the self-employed, starting in from 2001. An additional reform took place in 2005, when Welfare Minister Roberto Maroni modified again the

[5] See Brugiavini and Peracchi, 2010, Angelini et al, 2009, Bottazzi et al, 2011, Bertoni, Brunello and Mazzarella, 2016, among others for further details on the pension reforms occurring during our sample period.

eligibility requirements for *seniority* pensions, introducing a sharp 3-year increase in minimum eligibility age (the so-called "*scalone*"), from 57 to 60 years for public and private employees, and from 58 to 61 for the self-employed, starting from year 2008. However, in 2007, the new left-wing government led by Romano Prodi postponed the proposed three - year increase to 2011, introducing instead a gradual adjustment in the requirements, starting from 2008. For this reason, no worker has actually retired under the requirements prescribed by the "Maroni" reform. In addition, under the "Prodi bis" regime, eligibility to *seniority* pensions was made conditional to achieving a further threshold, defined as the sum of age and years of contributions – that also varied by year of retirement and sector.[6]

By abolishing seniority pensions starting in 2012, the Monti - Fornero reform further increased minimum retirement age, from 60 to 66 years for males and from 60 to 62 for females in the private sector.[7] For males, additional seven months were added to these new thresholds from 2013 to 2016. Since the gender retirement gap was expected to close in a few years, the minimum retirement age for females employed in the private sector increased by three years and seven month between 2013 and 2016 (see for instance Moscarola, Fornero and Strom, 2015).

Table 1 illustrates the changes in minimum retirement age for male and female employees in the private sector and for the self-employed during the years 1996 to 2016. Minimum age increased by more than ten years, from 52 to 66.58 years for male employees and from 52 to 65.58 for female employees.

[6] Pension reforms in Italy have also modified pension benefits. The major change occurred in 1995, before the start of our sample period, with the transition from a system based on defined benefits to a system relying on defined contributions. Another important change occurred within our sample period, when in 2007 the second Prodi government ("Prodi bis") reduced the coefficients used to transform accumulated contributions into pension benefits for workers retiring from 2010 onwards. Finally, the Monti - Fornero reform of 2011 applied a defined contributions formula to all workers and further increased the minimum number of year of social security contributions required to retire before minimum retirement age.

[7] Minimum age for females employed in the public sector also increased from 60 to 66 in 2012.

For the self-employed, it increased from 56 to 66.58 years for males and from 56 to 66.08 years for females.

1.2. The local pool PT_{pt}

To illustrate the effects of an increase in minimum retirement age on the composition of the senior population aged 50 to 70, define this population as $P_{pt}{}^{O} = \sum_{a=50}^{70} P^{O}_{a,pt}$, where the subscript *a* is for age, and let MRA_t be the minimum retirement age prescribed by law at time t. Then $P_{pt}{}^{O} = \sum_{a=50}^{MRA_t-1} P^{O}_{a,pt} + \sum_{a=MRA_t}^{70} P^{O}_{a,pt}$, where $PT_{pt} = \sum_{a=50}^{MRA_t-1} P^{O}_{a,pt}$ is the current pool of senior individuals aged 50+ who are not eligible to retire because they are younger than MRA_t and $NPT_{pt} = \sum_{a=MRA_t}^{70} P^{O}_{a,pt}$ is the group eligible to retire. Pension reforms that raise minimum retirement age increase PT_{pt} and reduce NPT_{pt}, thereby altering the composition of the population aged 50 to 70.

Since eligibility requirements vary by gender, sector and type of employment, in this paper we define gender-specific MRA_t as the gender – specific minimum among the available requirements in each year of the sample. For instance, the value of MRA_t in 2012 was 66 for males and 62 for females.

In local labor markets, the current pool PT_{pt} depends both on MRA_t, which is set at the national level, and on the age structure of the local population. Because of limited labor and population mobility and of differences in the composition of local population by age, the impact of national changes in MRA_t on local PT_{pt} varies across local areas. Therefore, while the treatment (higher MRA_t) is common across localities, the intensity of the treatment varies locally.

Changes in local PT_{pt} depend also on endogenous migration and on local demographic changes. To separate these changes from those induced by changes in MRA_t, we generate a shift-share variable à-la Bartik, 1991, and Card, 2007, by applying the MRA_t in place in any given year to the provincial population aged 50 to 70 as it is in the Population Census in 1991 – before our observation period begins. We call this new variable $PT_{pt}^{1991} = \sum_{a=50}^{MRA_t - 1} P^{O}_{a,p1991}$. Figure 2 shows the average values of PT_{pt} and PT_{pt}^{1991} from 2004 to 2016. Both variables exhibit a clear positive trend, which is driven for the latter solely by increases in MRA_t, and for the former by the combined effects of higher MRA_t and local demographic changes.

In this paper, we identify local labor market either with provinces (NUTS 3) or with regions (NUTS 2). In Italy, provinces are administrative areas that consist of several municipalities. Usually, several provinces together form a region. At present, there are 110 provinces in Italy, and 20 regions. Since a few provinces have been created during the sample period, we reclassify our data so as to have a common number of provinces (102) during the entire sample period.

Figure 3 shows the percentage change in PT_{pt}^{1991} between 2005 and 2016, by province. The dark blue areas are characterized by relatively high percent changes (between 103 and 117 percent) and the light blue areas by relatively low changes (between 84 and 97 percent). While dark blue areas are more frequent in Central Italy, light blue areas prevail in the North. However, no stark geographic pattern emerges.

2. 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 Cobb Douglas technology with two types of workers, the young N_y and the old N_o. The technology used by firm *i* is

$$Y_{it} = \left(e_{yi,t} N_{yi,t}\right)^{\alpha_i} \left(e_{oi,t} N_{oi,t}\right)^{1-\alpha_i} \tag{1}$$

where Y is output, $\alpha_i < 1$, and e_{zi}, $z = y, o$ are efficiency parameters. Product markets are imperfectly competitive and the demand for product variety *i* is given by

$$Y_{it} = Y_t \left(\frac{P_{it}}{P_t}\right)^{-\sigma_i} \tag{2}$$

where Y is aggregate demand, P the average price and $\sigma_i > 1$ the elasticity of demand (in absolute value) with respect to the relative price. In countries such as Italy, strict employment protection regulation makes terminating the employment relation of older workers very costly. Following Boeri et al, 2016, 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

$$N_{oi,t} = (1 - \delta_{it} + \omega_{it}) N_{oi,t-1} \tag{3}$$

where $\delta_{i,t}$ is the percentage of older workers who retire in each period and $\omega_{i,t}$ is the rate of change due to demographic factors, inclusive of mobility. In this setup, pension reforms raising minimum retirement age and the local pool PT_{pt} increases $N_{oi,t}$ by reducing δ_{it} ($\frac{\partial \delta_{it}}{\partial PT_{pt}} < 0$).

Local firms select youth employment by maximizing profits $\Pi_{it} = P_{it} Y_{it} - w_{yi,t} N_{yi,t} - w_{oi,t} N_{oi,t} - c_i$, where w_{ai}, $a = y, o$ are wages and *c* are idiosyncratic fixed costs. 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; Manacorda and

Petrongolo, 2006).[8] 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, $w_{ai} = w_a, a = y, o$.

Profit maximization with respect to youth employment yields

$$N_{yi,t} = \alpha_i (1 - \frac{1}{\sigma_i}) Y_{it}^{(1-1/\sigma_i)} \left(\frac{w_{yt}}{P_t} \right)^{-1} Y_t^{1/\sigma_i} \tag{4}$$

Total differentiation of Equation (4) with respect to local PT_t yields

$$\frac{\partial N_{yi,t}}{\partial PT_{pt}} = \left\{ \frac{(1-1/\sigma_i)}{1-\alpha_i(1-1/\sigma_i)} \right\} N_{yi,t} \left[(1-\alpha_i)\left(\frac{1}{N_{oi,t}} \frac{\partial N_{oi,t}}{\partial PT_{pt}} + \frac{1}{e_{oi,t}} \frac{\partial e_{oi,t}}{\partial PT_{pt}}\right) + \alpha_i \frac{1}{e_{yi,t}} \frac{\partial e_{yi,t}}{\partial PT_{pt}} \right] \tag{5}$$

The ratio within braces is positive. The sign of the expression within brackets on the right hand side of (5) depends instead on the sign of three components. Given our assumptions, the first component $\frac{\partial N_{oi,t}}{\partial PT_{pt}}$ has positive sign. The second component $\frac{\partial e_{oi,t}}{\partial PT_{pt}}$ has negative sign if the productivity of older workers decline with age.[9] The sign of the last component $\frac{\partial e_{yi,t}}{\partial PT_{pt}}$ is unclear, but could be negative if the interaction with a higher number of older workers has negative spill-over effects on the young. Overall, the effect of an increase

[8] 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 set both industry, that define both specific wage floors and employment rules. Local agreements that redistribute productivity gains occur mainly in large firms and can add to the national floors without undoing them. See Rosolia, 2015. Using regional data from the Bank of Italy's Survey on Household Income and Wealth (SHIW), we report in Table A1 in the Appendix the ITT effects of PT_{pt}^{1991} on local real wages, showing that the former are not sensitive to changes in the latter.

[9] There are relatively few studies examining the relationship between age and *individual* productivity. Skirbekk, 2004, reviews this literature and concludes that productivity follows an inverted U-shaped profile, with significant decreases taking place from around age 50. See also Bertoni et al, 2015, and Van Ours, 2009. Indirect measures of productivity such as numeracy skills also show and inverted U-shaped profile. See OECD, *The Survey of Adult Skills*, 2012.

in local PT_t on youth employment at the firm level cannot be signed.

Since youth employment in the local labor market *p* is obtained by aggregating firm-level employment over the number of existing firms - $N_{yp,t} = \int_0^{G_{pt}} N_{yip,t} di$, where G_{pt} is the number of local firms - changes in PT_{pt} could affect local employment also by changing the number of firms.[10] For instance, in the presence of very high separation costs for older workers and of centrally set wages, an increase in local PT_{pt} that keeps these workers longer in their jobs raises costs and may reduce profits below zero, forcing some firms to exit the market, with employment losses for both young and senior workers.

In addition, incumbent firms could pass through the higher costs due to the retention of older workers into higher prices, thereby encouraging new firms to enter the market. Changes in relative prices could also spill-over on other local firms because of the network of buyer-seller relationships. In summary, we expect the effects of changes in PT_{pt} on local youth employment to differ from those on youth employment at the level of a single firm. We also expect that a higher PT_{pt} influences local gross firm turnover, defined as the ratio of the sum of firm entries and exits to the initial stock of firms.

3. The Empirical Specification

We investigate the effects of local changes in PT_{pt} on employment and unemployment by age group by estimating on provincial data and the period 2005-2016 the following empirical specification

$$Y_{pt}^{Q} = \alpha + \theta X_{pt} + \gamma_T + \gamma_P + \delta Z^{Q}_{pt-\tau} + \varepsilon_{pt} \qquad (6)$$

10 $\frac{\partial N_{yp,t}}{\partial PT_{pt}} = \int_0^{G_{pt}} \frac{\partial N_{yip,t}}{\partial PT_{pt}} di + N_{yG_{pt}p,t} \frac{\partial G_{pt}}{\partial PT_{pt}}$

where Y is either employment or unemployment gross of inactivity (in thousand individuals); the superscript Q indicates the age group (very young: 16-24, young: 25-34, prime age: 35-49 and senior: 50-70); γ_T is for time dummies, that capture macroeconomic shocks, and γ_P is for province dummies, that capture time invariant province effects; X_{pt} is either the current pool PT_{pt} or PT_{pt}^{1991}, also measured in thousand individuals, and Z is a vector of lagged province by time controls, that we include to control for local shocks and trends. Finally, ε is an error term, that we allow to be clustered by province.

We use the first or the second lag rather than current values of Z to attenuate endogeneity concerns. 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 unconditional on the current level of output (see Banks et al, 2010). In baseline specification (6), vector Z includes the lags of: local *GDP* per head, the share of immigrants in the local population, the sectoral index S and, for each age group, average age, local population, the percent with at least high school and the percent of males.

The sectoral index S is constructed by selecting the production sectors that had in 1995 – well before the start of our sample – 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. Figure A1 in the Appendix illustrates how this index changes over time and across three selected provinces, Milan, Rome and Palermo.

We estimate (6) using either PT_{pt}^{1991} or PT_{pt} as our key explanatory variable. In the former case, we obtain "intention to treat" effects, estimated by OLS. In the latter case, we instrument PT_{pt} with $PT_{pt}{}^{1991}$. By so doing, we address the concern that, even after conditioning on time and province dummies, there may be unobservable concurring local employment shocks - the timing of which is correlated with changes in MRA_t - that affect both local employment and PT_{pt}. This is likely because the variation in PT_{pt} relies on the within-province variation in the age structure of the senior population over time, and there could be shocks affecting both this demographic component and local employment. For instance, employment shocks could trigger migration across localities. In this case, the OLS estimates of parameter θ in Equation (6) are inconsistent.

Since employment and unemployment are persistent over time, we also estimate a dynamic version of (6)

$$Y_{pt}^{Q} = \alpha \ + \beta Y_{pt-1}^{Q} + \theta X_{pt} + \gamma_T + \gamma_P + \delta Z^{Q}{}_{pt-\tau} + v_{pt} \qquad (7)$$

that includes the lagged dependent variable, which is endogenous because of the presence of the province fixed effects (Nickell, 1981). To address this, we implement the Arellano-Bond, 1991, GMM estimator, that uses employment lags of order two and higher to generate instruments for the first lag. If the random error *v* is not seriously correlated, these instruments are valid. We test for second order serial correlation and also report the Hansen over-identification test for instrument validity.[11]

As in Boeri et al, 2016, and Bovini and Paradisi, 2017, the dependent variable in (6) is either employment or unemployment rather than employment or unemployment population ratios. Our choice is motivated by the fact that changes in employment population ratios could be due to changes in the

[11] In general, we include only the first lag of the dependent variable on the right hand side of equation (7), and use all other available lags to generate internal instruments. However, in some specifications – in order to pass both tests – we include also additional lags of the dependent variable on the right-hand side or use only a limited set of lags to generate internal instruments. Details are available from the authors upon request.

numerator, denominator, or both. Moreover, the impact of *X* on age-specific employment population ratios N_{pt}^{Q}/P_{pt}^{Q}, where population P is the sum of employment and unemployment gross of inactivity, can be easily obtained from our estimates of (6) as follows

$$\frac{\partial \frac{N_{pt}^{Q}}{P_{pt}^{Q}}}{\partial \ln X_{pt}} = \frac{N_{pt}^{Q}}{P_{pt}^{Q}}\left(\frac{\partial \ln N_{pt}^{Q}}{\partial \ln X_{pt}} - \frac{\partial \ln P_{pt}^{Q}}{\partial \ln X_{pt}}\right) \tag{8}$$

4. The Data

Our data are drawn from the Italian Labor Force Survey (LFS). The LFS is a quarterly survey on labor market conditions covering a representative sample of almost 77,000 households and 175,000 individuals per quarter. We use the second quarter of waves 2004 to 2016, because information about the province of residence is publicly available in the LFS only from 2004.[12] Because some of the lagged controls in vector *Z*, including average age, the percentage of individuals with higher education and the index *S* are computed using these province-level data, our baseline regressions are based on the period 2005-2016.

We aggregate data by province and wave, using sampling weights to reproduce national aggregates. Due to changing boundaries over the sampling period, we use the definition of provinces in 2004 and reclassify our data accordingly. Since this re-classification is not feasible for the neighboring Provinces of Bari and Foggia, we treat them as a single province and end up with 102 provinces.

Data on local GDP per head are from Eurostat, and data on the provincial age structure in 1991 – used to construct PT_{pt}^{1991} – are drawn from the 1991

[12] Additionally, the survey was changed in 2004 with the transition to a quarterly continuous survey.

Population Census (source: National Statistical Institute). Finally, we use data on the existing stock of firms and the number of firms that entry and exit in each province and year to compute gross and net firm turnover. Data on firms are from the *MovImprese* database of the local Chambers of Commerce, which contains population-level data on the stock of active firms, as well as on firm entries and exits by province and year. Due to a change in the definition of firm exit, data can only be used from 2006 onwards. Descriptive statistics for the relevant variables are reported in Table 2.

5. Results

Table 3 reports the intention to treat (ITT) estimates of Equation (6). For each age group, we show the marginal effect of local PT_{pt}^{1991} on local employment. In the first row, we consider a parsimonious specification that excludes the vector of controls *Z*; in the second and third row, we add the variables in vector *Z* lagged either once or twice. In the fourth row, in an effort to further improve our ability to parsimoniously capture local trends, we add to the vector *Z* the interactions of a national quadratic time trend with the 2004 value of local *GDP* per head. Finally, in row five we consider a dynamic employment equation with lagged *Z*.[13]

Starting from the least rich specification in row (1), we find that adding one thousand individuals to the local pool PT_{pt}^{1991} of senior individuals who cannot retire because of lack of eligibility reduces local employment for the youngest group and for the group aged 25 to 34 by 59 and 278 units respectively. We also detect no statistically significant effect for prime age workers aged 35 to 49, and a positive effect equal to 335 units for senior workers aged 50 to 70.

[13] For this model, we always fail to reject the null hypothesis that the over-identifying restrictions are valid (Hansen J test) and that the error term in the dynamic equation is not serially correlated (Arellano-Bond AR(2) test). Results are available from the authors upon request.

Adding to the regressions the variables in vector *Z* lagged either once or twice has almost no effect on the very young (16-24) but somewhat reduces the absolute value of the marginal effect for young workers aged 25 to 34 (from 278 units to the range 201 to 262 units). Conversely, the marginal effect for prime age workers becomes negative and statistically significant (in the range between 98 to 100 units) and the effect on senior workers declines in absolute value, from 335 to the range 264 to 279 units. On the one hand, the addition of a national quadratic time trend interacted with local 2004 GDP does not affect qualitatively our results. On the other hand, the estimates based on the dynamic specification in row (5) are broadly similar to those in row (2), with the exception of prime age and older workers. For the former group, the marginal impact of PT_{pt}^{1991} is not significantly different from zero (in a statistical sense). For the latter group, the estimated effect of a 1,000 units increase in PT_{pt}^{1991} declines in absolute value from 264 to 149 units.

Independently of the selected specification, these results suggest that the increase in local senior employment induced by higher local PT_{pt}^{1991} is not sufficient to compensate for the reduction of local youth and prime-age employment.[14] Using the specification in row (2), we estimate that a 1,000 units increase in PT_{pt}^{1991} reduces overall employment by 102 units.[15]

Table 4 reports both OLS and IV estimates of (6) when we use the current pool of senior workers PT_{pt} rather than the pool based on the population in 1991, PT_{pt}^{1991}, as the key explanatory variable, and instrument the former with the latter. For the static specifications in rows (1) to (4), we find that the first

[14] Following Oster, 2016, we have tested whether the sign of our estimates is affected by the assumption that un-observables are at least as important as all observables included in the specification in row (2), under the additional assumption that maximum achievable R Squared is equal to 1. We have found no evidence that this is the case. Results available from the authors upon request.

[15] This effect is statistically significant at the 5 percent level of confidence (standard error: 0.051).

stage F - test for instrument relevance ranges between 598 to 1095, well above the critical value of 10, suggesting that the instrument PT_{pt}^{1991} is not weak.[16]

Since the conditional first stage effect of PT_{pt}^{1991} on current PT_{pt} is positive and larger than 1 –most likely because the size of the population aged 50 to 70 has increased in the last two decades for demographic reasons – we find that the IV estimates have the same sign as the ITT estimates but are smaller in absolute value. Focusing on the specification in row (2), we find that adding one thousand additional senior individuals to PT_{pt} reduces employment for the age groups 16 to 24, 25 to 34 and 35 to 49 by 46, 142 and 72 units respectively, and increases senior (age 50 to 70) employment by 214 units. The overall employment effect is negative (-46 units) but not statistically significant.[17]

Finally, tables A2 and A3 report the ITT, OLS and IV estimates of the effects of PT_{pt}^{1991} and current PT_{pt} on unemployment gross of inactivity. We find that higher minimum retirement age increases the unemployment of young and prime age individuals and reduces the unemployment of senior individuals.

6. Extensions and sensitivities

In this section we consider several sensitivities. Unless otherwise stated, we focus on ITT effects for the simplicity in carrying out inference with respect to IV estimates. Since the qualitative results in rows (2) to (5) of Table 3 are similar, we shall consider hereafter only the baseline specification in row (2).

6.1 Effects on the employment to population ratio

By using (8) and combining the estimates in the second row of Tables 3 and A2, we find that raising the pool PT_{pt}^{1991} by ten percent reduces the

[16] For the dynamic model in row (5), we use the weak identification diagnostics proposed by Bazzi and Clemens, 2013, and fail to detect weak instrument problems.

[17] The estimated standard error is 51 units (0.051).

employment / population ratio for individuals aged 16-24, 25-34 and 35-49 by 1.06, 1.35 and 0.49 percent respectively, and increase the employment / population ratio for senior individuals aged 50 to 70 by 1.19 percent (see Table 5). The change in the employment / population ratio is smaller in absolute value than the change in employment because the effects of PT_{pt}^{1991} on employment and unemployment (gross of inactivity) have opposite signs.

6.2 Changes in the definitions of the dependent variable

We have considered so far the effects of PT_{pt}^{1991} and PT_{pt} on total employment and unemployment. Rows (1) to (3) in Table 6 consider as alternative dependent variables: employment in the private sector; employment net of the self-employed; full time employment. Results are broadly in line with those in the second row of Table 3.

6.3 Two Alternative Definitions of PT_{pt}

We propose here two alternative definitions of the current pool PT_{pt}. The first alternative takes into account the fact that pension eligibility requirements in Italy combine minimum age and a minimum number of years of paid social security contributions, and that minimum age requirements can be by-passed if the number years of paid contributions attains a time varying threshold.

Table A4 in the Appendix shows how the number of years of social security contributions required to retire before minimum retirement age has changed over the sample period, by gender and employment type. To assess whether retirement age is determined by minimum retirement age or by the requirement on contributions, we need information on the number of years of contributions actually paid.[18] This information is available in the Survey on the Income and Wealth of Italian Households. We select the 2000 wave of this

[18] To illustrate, if minimum retirement age is 55 and the number of years of paid contributions is 35, somebody with a discontinuous career may reach age 55 without having completed 35 years of contributions, and therefore must retire at a later age. This case is particularly frequent among females.

survey, before our sample period, compute the median years of paid contributions by age, gender, and education, and use these years to estimate whether retirement can be earlier or later than minimum retirement age. Since many Italian women have discontinuous careers, we expect the pool PT_{pt} that considers also paid social security contributions to be larger than baseline PT_{pt}, which is constructed using only minimum retirement age requirements. Figure A2 in the Appendix confirms that this is indeed the case.

The second alternative excludes from population PT_{pt} the inactive, many of whom could be permanently disengaged from the labor market. For this group, changes in eligibility conditions for retirement are unlikely to affect labor supply decisions. This second alternative is also shown in Figure A2.

We replicate our estimates using these alternative measures of PT_{pt} and instrumenting them with PT_{pt}^{1991}. The results in Table A5 confirm our qualitative results – positive effects on senior employment and negative effects on young and prime age employment.

6.4 Heterogeneous Effects by Gender

In Table A6 in the Appendix we ask whether the effects reported in Table 3 vary by gender. We find that senior female employment benefits slightly more than male employment from the increase in local PT_{pt}^{1991}, and that youth and prime age female employment decline less than male employment in the same age groups. In particular, we estimate that a one thousand increase in local PT_{pt}^{1991} increases senior employment by 103 units for males and by 168 units for females, and reduces youth employment in the age group 25 - 34 by 136 units for males and by 76 units for females.

6.5 Estimates based on the years 1996 to 2016

We have found that local increases in PT_{pt} and PT_{pt}^{1991} raise senior employment but reduce youth and prime-age employment. These results are based on a sample period, 2005-2016, dominated by the 2008 recession and its consequences. As shown in Figure 4, real Italian GDP grew more or less continuously from 1996 to 2007, declined sharply in 2009 and remained below its 2007 level until the end of the sample period in 2016.

We ask whether the direction and size of the negative impact on youth employment of pension reforms rising minimum retirement age are affected by the selection of a sample period characterized by declining GDP and total employment. To address this question, we consider the longer time period 1996 – 2016, that includes the interval 1996 to 2007, when both GDP and total employment show positive growth. Since the disaggregation of Labor Force data at the province level is not publicly available before 2004, we use regions rather than provinces as local areas.

As in the case of provinces, we collapse individual Labor Force data at the regional level using sampling weights, and estimate our baseline ITT specification.[19] Since we have data on 19 regions[20] and cluster standard errors by region, we account for the small number of clusters by using wild bootstrap techniques for inference (see Cameron, Gelbach and Miller, 2008).[21]

The ITT estimates for both employment and unemployment are reported in Table 7.[22] Comparing the results in this table with those in Tables 3 and A2 (second row), we confirm the positive (negative) effect of PT_{pt}^{1991} on employment (unemployment) for the senior age group and the negative (positive) effect on employment (unemployment) for the youngest age

[19] Results using the other specifications are qualitatively similar and available from the authors upon request.
[20] The tiny Val d'Aosta is pooled with Piedmont.
[21] The vector *Z* includes the same controls used for provincial data, with the exception of the share of immigrants, that is not available for the late 1990s.
[22] Not reported here, the first stage F statistics for instrument relevance - obtained by wild bootstrap - are always way above the critical value of 10.

group.[23] For the remaining two age groups, however, there are differences worth noting. For the young (25-34), we estimate that a one thousand increase in local PT_{pt}^{1991} reduces employment by 27 units in the longer sample, much less than with provincial data covering the shorter period 2005 to 2016 (-201 units). For the prime-aged (35-49), the estimated decline in the longer sample is 45 individuals, smaller than the effect reported in Table 3 (-100). For the unemployed or inactive, we find that an increase in local PT_{pt}^{1991} has negative rather than positive effects on youth (25-34) unemployment and much larger negative effects on senior unemployment than those found using provincial data and a shorter time period.

A source of concern here is that these differences may not be driven by the sample period but by the definition of local labor markets (regions rather than provinces). To dispel this concern, we replicate our estimates using regions on the shorter period (2005-2016) and show that for this period estimated effects are much more similar to those presented in Tables 3 and A2, especially for young and prime-aged individuals.[24]

We conclude that, while the direction of the effect of a higher local PT_{pt}^{1991} on local youth employment does not vary across different sample periods, the size of the effect does and is larger in absolute value for the period 2005 to 2016 than for the longer period including the years 1996 to 2004. On the one hand, employment responses to changes in minimum retirement age could vary with business cycle conditions. While the period from 2005 to 2016 is dominated by the 2008 recession and its consequences, the period 1996 to 2016 also

[23] The estimated increase for senior workers is 310 units, larger than the one estimated in Table 3 (264). The quantitative effects for the very young (16-24) are similar (a decline of 78 rather than 65 units).

[24] We test whether the estimated differences of coefficients across periods are statistically significant and find that they are for the age groups 25-34 and 35-49 (p-values of the test equal to 0.000 and 0.002 in the case of employment, and to 0.000 and 0.062 in the case of unemployment).

contains a period of moderate economic growth. On the other hand, the individuals affected by pension reforms in the late nineties were younger than those affected in the second part of the 2000s and early 2010s. For these individuals, retention into employment could have had smaller negative effects on labor productivity than for those forced to stay at work at higher ages, and therefore smaller negative consequences on youth employment.

6.6 Spill-over effects

Conditional on local changes, aggregate changes in the pool of senior individuals who are too young to retire could also affect local employment. For instance, nationally set wages may change. The wages of senior workers w_o could fall as the national supply of these workers increases, and the wages of young workers w_y could increase if the aggregate demand for youth labor rises because of complementarities with senior labor.

To clarify the implications of this for our estimates, it is useful to write the following stripped down version of Eq. (6)

$$N_{pt}^{Y} = \alpha + \theta X_{pt} + \lambda X_{-pt} + \omega_{pt} \tag{9}$$

where *p* is for a given province and *"–p"* is for all the other provinces in the country. In this model, θ is the effect of a change in the local pool of senior workers, either PT_{pt}^{1991} or PT_{pt}, on local youth employment, and λ is the spill-over effect from all other provinces. Since the national value of X is given by $X_t = X_{pt} + X_{-pt}$, Equation (9) can be re-written as:

$$N_{pt}^{Y} = \alpha + (\theta - \lambda) X_{pt} + \lambda X_t + \omega_{pt} \tag{10}$$

By aggregating (10) over the *K* provinces, we obtain $N_t^Y = K\alpha + [(\theta - \lambda) + K\lambda] X_t + \omega_t$. The overall effect of *X* on N^Y is $[(\theta - \lambda) + K\lambda]$.

When Equation (10) is estimated using unrestricted time dummies, as we do, the parameter (θ-λ) is identified by the province-by-time variation of local *X,* but parameter λ cannot be separately estimated. If λ is negative, our estimates are a lower bound of the total effect of *X* on youth employment. If λ is positive, however, the direction of the effect estimated using the local variation of *X* may be different from the direction of the overall effect.

Although λ cannot be estimated, we can evaluate the importance of *local* spill-overs by asking whether local employment and unemployment are affected by changes of PT_{pt}^{1991} in neighboring areas. For this purpose, we augment our baseline ITT specification with the pool of senior workers too young to retire in the provinces that share their border with the selected province. The results in Table A7 indicate that, conditional on the local PT_{pt}^{1991}, the value of PT_{pt}^{1991} in neighboring areas has very small and imprecisely estimated effects on local employment.

6.7 Firm turnover

Changes in local PT_{pt} could affect firm turnover by facilitating firm entry and / or exit. We regress gross and net turnover on PT_{pt}^{1991} (measured in million units) and the controls used in Table 3 (row 2). Results shown in Table 8 indicate that an increase in PT_{pt}^{1991} raises both gross and net turnover. The estimated effect, however, is very small.

Conclusions

Estimating the causal effects of pension reforms rising minimum retirement age on youth employment is important but difficult because these reforms are typically national policies, the effects of which cannot be easily separated from concurring macroeconomic shocks. Following Card, 1992, and Acemoglu and Johnson, 2007, we have addressed this identification problem by using the fact that, although every local labor market in a country 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 2005 to 2016, we have estimated the causal effect of a local increase in the pool of senior workers who are too young to retire on the local employment and unemployment of four age groups, the very young, the young, the prime aged and the seniors. In all these estimates, we do not condition on current output. We have found that – in our preferred specification – raising the local pool PT_{pt} by one thousand additional senior individuals reduces youth employment by 188 individuals and increases senior employment by 214 individuals. These estimates indicate that to each additional 10 senior workers locked into employment correspond 9 young individuals out of work, a large negative effect.

We have shown that the size of this effect varies significantly with the sample period. In particular, estimates based on an extended sample that includes also a period when the economy is growing indicate that to each additional 10 senior workers locked into employment correspond slightly more than 3 young individuals out of work. These results suggest that the employment costs of pension reforms that delay the retirement of older workers may be lower if these reforms are implemented – at least in part – when the economy is in relative good shape.

We are not the first to find that higher retirement eligibility age negatively affects youth employment. Using administrative data on Italian firms, Boeri, Garibaldi and Moen, 2016, find a smaller ratio between employment losses for the young and employment gains for the old because of the 2011 pension reform. On the one hand, by looking at local labor markets rather than at firm level data, our approach captures the additional effects on firm turnover as well as spill-over effects from one local firm to another. On the other hand, while their approach focuses on the short-term impact of a single pension reform, our method considers several changes in minimum retirement age occurring over a longer span of time.[25]

Our empirical approach works well under the assumption that the effects on local labor markets of national changes in the pool of senior workers who are too young to retire are negligible once we condition for local changes in the pool. It works less well when these effects are not negligible. Unfortunately, verifying whether this assumption holds would require perhaps even stronger assumptions to distinguish the aggregate effects of pension reforms from concurring aggregate economic shocks.

Even without imposing that the number of jobs in the economy is given, we have found that pension reforms increasing minimum retirement age contribute to reducing youth employment. Does this suggest that these reforms introduced in Italy or in other OECD countries should be reversed and minimum retirement age be reduced so as to favor the young? Although many politicians seem to consider this option seriously, we believe that a cost-benefit analysis should also carefully consider the implications of reducing minimum retirement age on the public budget and social security expenditures, especially in ageing societies that are already burdened by a high public debt.

[25] Boeri et al also consider slightly different age groups.

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Tables and Figures

Table 1. Minimum retirement age in the private sector, by year and gender.

Year	Male employees	Male self-employed	Female employees	Female self-employed
1996	52	56	52	56
1997	52	56	52	56
1998	53	57	53	57
1999	53	57	53	57
2000	54	57	54	57
2001	55	57	55	57
2002	55	58	55	58
2003	56	58	56	58
2004	57	58	57	58
2005	57	58	57	58
2006	57	58	57	58
2007	57	58	57	58
2008	58	59	58	59
2009	59	60	59	60
2010	59	60	59	60
2011	60	61	60	61
2012	66	66	62	63.5
2013	66.25	66.25	62.25	63.75
2014	66.25	66.25	63.75	64.75
2015	66.25	66.25	63.75	64.75
2016	66.58	66.58	65.58	66.08

Source: national legislation.

Table 2**.** Descriptive statistics. Province-by-year variables. Observations: 1,224

	Mean	Std. Dev.
# Employees/1,000 – age 16-24	11.13	12.10
# Employees/1,000 – age 25-34	47.83	57.20
# Employees/1,000 – age 35-49	103.11	124.42
# Employees/1,000 – age 50-70	61.36	72.37
# Unemployed and Inactive/1,000 – age 16-24	42.09	51.37
# Unemployed and Inactive/1,000 – age 25-34	25.69	35.42
# Unemployed and Inactive/1,000 – age 35-49	35.14	47.51
# Unemployed and Inactive/1,000 – age 50-70	90.22	99.58
Total population/1,000 – age 16-24	53.23	61.87
Total population/1,000 – age 25-34	73.53	85.13
Total population/1,000 – age 35-49	138.26	161.11
Total population/1,000 – age 50-70	151.57	169.71
PT/1,000	90.60	108.96
PT/1,000 – excluding the inactive	55.51	67.60
PT/1,000 – using info on contributions paid and required	97.77	114.13
PT 1991/1,000	80.30	96.43
Gross firm turnover	0.12	0.14
Net firm turnover	0.01	0.01

Notes: 102 provinces observed for 12 years (2005-2016). *Source*: ISTAT Labor Force Survey. Data on firm turnover refer to the period 2006-2016. *Source*: MovImprese archive.

Table 3. Estimated ITT effects of the pool PT_{pt}^{1991} on employment by age group. Provincial data 2005-2016

	(1)	(2)	(3)	(4)
Age group	16-24	25-34	35-49	50-70
Estimation method	OLS	OLS	OLS	OLS
Without Z	-0.059***	-0.278***	0.025	0.335***
	(0.008)	(0.032)	(0.053)	(0.028)
With lagged Z	-0.065***	-0.201***	-0.100***	0.264***
	(0.006)	(0.017)	(0.026)	(0.035)
With Z lagged twice	-0.053***	-0.262***	-0.098***	0.279***
	(0.006)	(0.021)	(0.026)	(0.031)
With lagged Z and a national quadratic trend interacted with provincial GDP per head in 2004	-0.063***	-0.197***	-0.095***	0.237***
	(0.005)	(0.018)	(0.027)	(0.034)
With lagged N and lagged Z (Arellano Bond GMM estimates)	-0.058***	-0.220***	0.006	0.149***
	(0.010)	(0.021)	(0.029)	(0.028)

Notes: number of observations: 1,224 in rows (1), (2) and (4), 1,122 in rows (3) and (5). All specifications include province and wave dummies. The vector Z includes the first lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 4. Estimated effects of the current pool PT_{pt} on employment by age group. OLS, IV and Arellano Bond estimates. Static and dynamic specifications. Provincial data 2005-2016

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age group	16-24	16-24	25-34	25-34	35-49	35-49	50-70	50-70
Estimation method	OLS	IV	OLS	IV	OLS	IV	OLS	IV
Without Z	-0.043***	-0.042***	-0.192***	-0.201***	0.014	0.018	0.246***	0.243***
	(0.006)	(0.006)	(0.028)	(0.028)	(0.040)	(0.038)	(0.026)	(0.023)
With lagged Z	-0.046***	-0.046***	-0.133***	-0.142***	-0.072***	-0.072***	0.219***	0.214***
	(0.004)	(0.004)	(0.014)	(0.016)	(0.018)	(0.019)	(0.024)	(0.026)
With Z lagged twice	-0.038***	-0.038***	-0.176***	-0.186***	-0.067***	-0.071***	0.227***	0.217***
	(0.004)	(0.004)	(0.018)	(0.020)	(0.017)	(0.019)	(0.021)	(0.024)
With lagged Z and quadratic	-0.045***	-0.045***	-0.131***	-0.140***	-0.068***	-0.069***	0.199***	0.193***
trend * provincial 2004 GDP	(0.004)	(0.004)	(0.014)	(0.017)	(0.019)	(0.020)	(0.025)	(0.027)
With lagged N and lagged Z	-	-0.040***	-	-0.151***	-	0.001	-	0.131***
(Arellano Bond GMM estimates)		(0.005)		(0.018)		(0.021)		(0.019)

Notes: number of observations: 1,224 in rows (1), (2) and (4), 1,122 in rows (3) and (5). All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table 5. Estimated ITT effects of a ten percent increase in the pool PT_{pt}^{1991} on employment shares by age group. Percentage points.

Age group	(1) 16-24	(2) 25-34	(3) 35-49	(4) 50-70
Without Z	-1.118*** (0.081)	-1.460*** (0.078)	-0.398** (0.134)	1.315*** (0.219)
With lagged Z	-1.061*** (0.066)	-1.353*** (0.088)	-0.488*** (0.117)	1.194*** (0.211)
With Z lagged twice	-0.930*** (0.066)	-1.511*** (0.074)	-0.533*** (0.087)	1.142*** (0.205)
With lagged Z and quadratic trend * provincial 2004 GDP	-1.011*** (0.064)	-1.331*** (0.087)	-0.491*** (0.112)	1.044*** (0.210)

Notes: ***: p<.01; **: p<.05; *: p<.10.

Table 6. Estimated ITT effects of the pool PT_{pt}^{1991} on employment by age group. Sensitivities. Provincial data 2005-2016

	(1)	(2)	(3)	(4)
Age group	16-24	25-34	35-49	50-70
Estimation method	OLS	OLS	OLS	OLS
Private sector employment	-0.061***	-0.171***	-0.067***	0.177***
	(0.006)	(0.022)	(0.022)	(0.030)
Net of self-employment	-0.053***	-0.156***	-0.016	0.231***
	(0.004)	(0.019)	(0.031)	(0.032)
Full time employment	-0.044***	-0.167***	-0.020	0.221***
	(0.006)	(0.021)	(0.030)	(0.029)

Notes: number of observations: 1,224. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10

Table 7. Estimated ITT effects of the pool of potential senior workers PT_{pt}^{1991} on employment and unemployment by age group. Baseline specification with lagged X. Regional data. Full period 1996-2016 and sub-period 2005-2016

Age group	(1) 16-24	(2) 25-34	(3) 35-49	(4) 50-70
Employment 1996-2016	-0.078***	-0.027	-0.045	0.310***
	[0.000]	[0.393]	[0.141]	[0.000]
Employment 2005-2016	-0.075***	-0.120***	-0.122**	0.293***
	[0.000]	[0.010]	[0.015]	[0.000]
Unemployment 1996-2016	0.090***	-0.028	0.028	-0.267**
	[0.000]	[0.410]	[0.114]	[0.031]
Unemployment 2005-2016	0.085**	0.078***	0.067***	-0.234***
	[0.011]	[0.025]	[0.004]	[0.004]

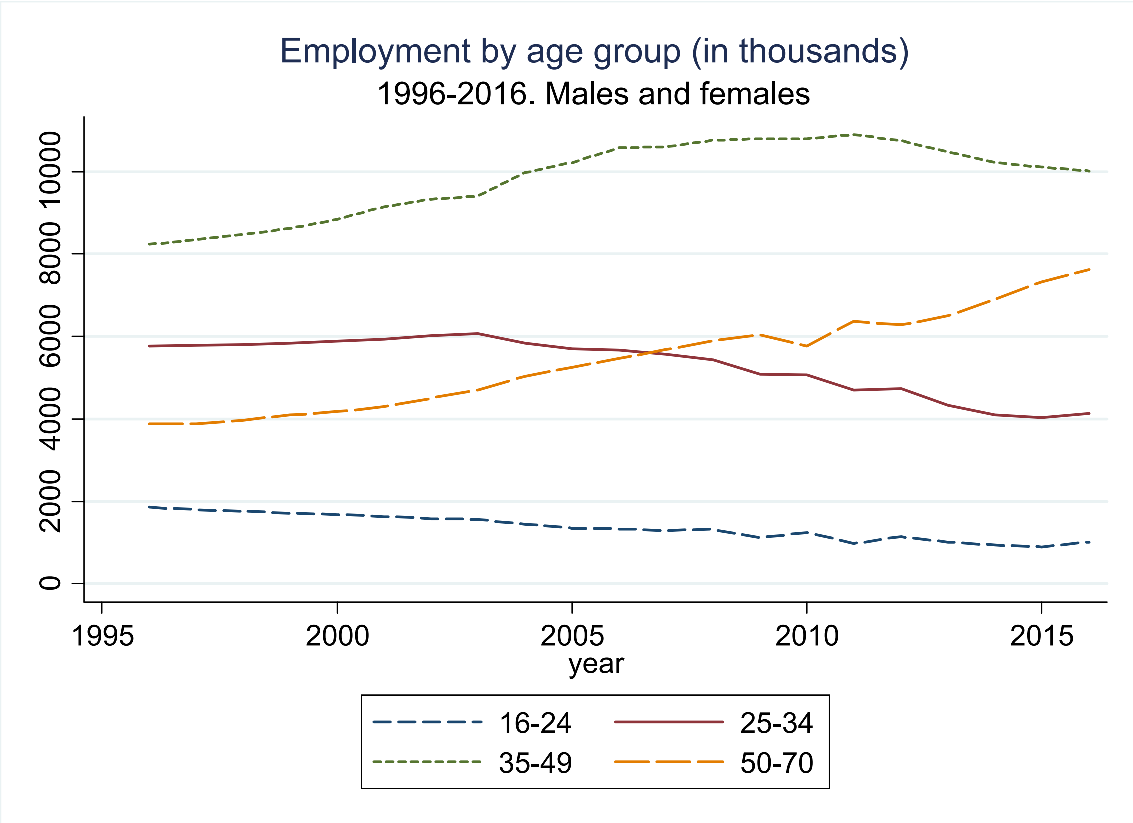
Notes: number of observations for the full period: 399. All models include region and wave dummies. The vector Z includes the firs lag of regional GDP per capita; the first lags of the index of sectoral composition S by region; total population by region and age group; the percentage of workers with high school or higher degree by region and age group; the percentage of males by region and age group; average age by region and age group; the share of immigrants by region and age group. Standard errors are clustered by region and adjusted using wild bootstrap. P-value of test that the estimated coefficient is statistically different from zero within brackets. ***: p<.01; **: p<.05; *: p<.10

Table 8. Estimated ITT effects of a higher pool of potential senior workers (PT_{pt}^{1991}) on gross and net firm turnover. Estimates of the static model with lagged Z. Provincial data 2006-2016.

Estimation method	OLS
Gross turnover – baseline specification	0.028**
	(0.012)
Net turnover – baseline specification	0.027*
	(0.014)

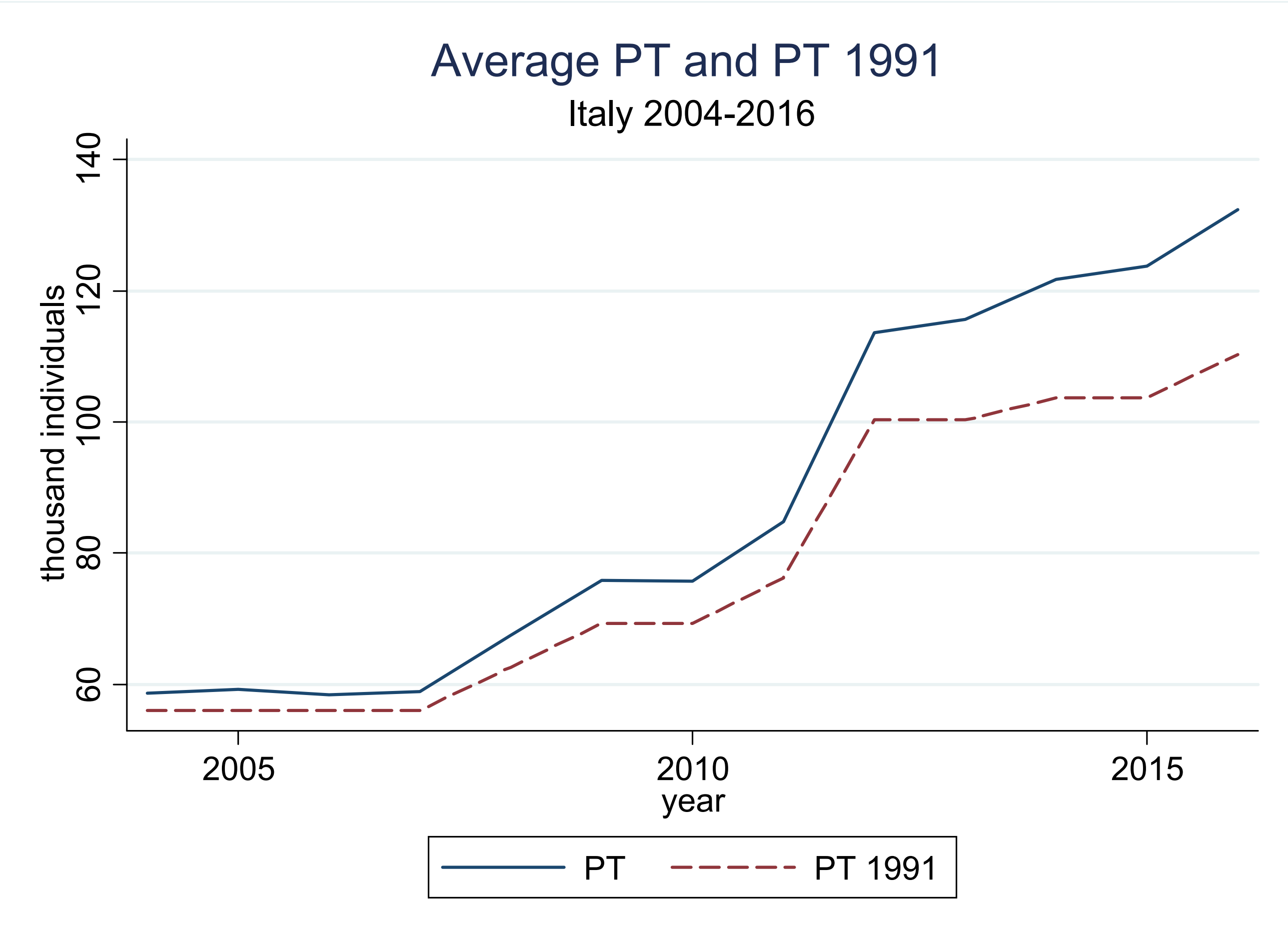
Notes: number of observations: 1,122. *PT 1991* is measured in million individuals. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; stock of firms by province; the percentage of workers with high school or higher degree by province; the percentage of males by province; average age by province; the share of immigrants by province. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10

Figure 1. Employment by age group. Italy 1996 to 2015



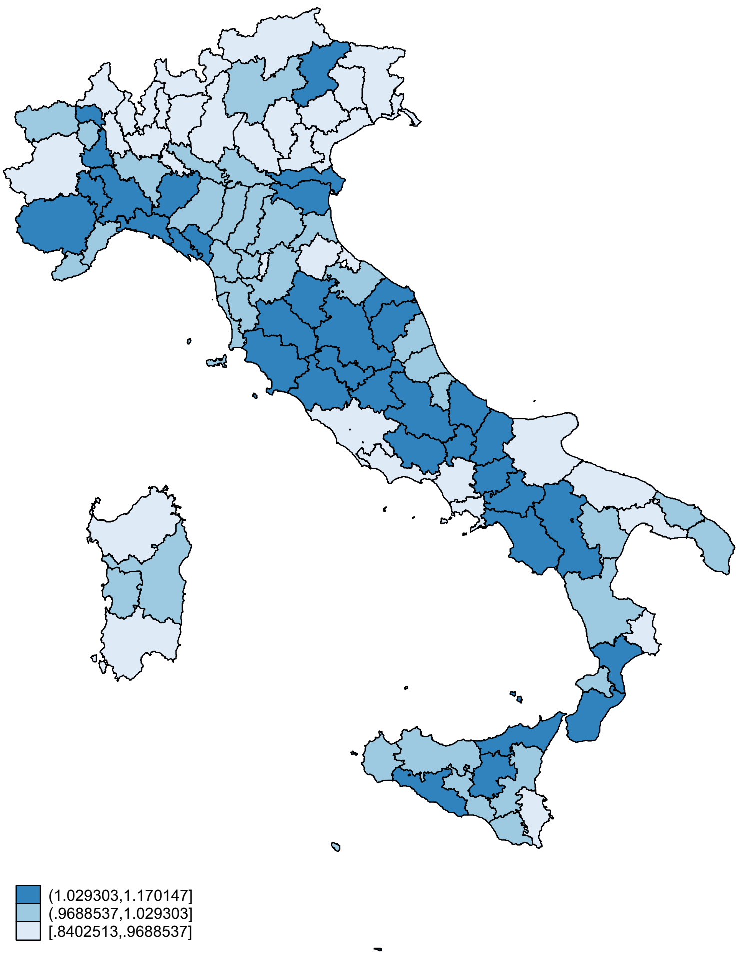
Source: Italian Labor Force Survey

Figure 2. The ratio of national PT to the population aged 50 to 70



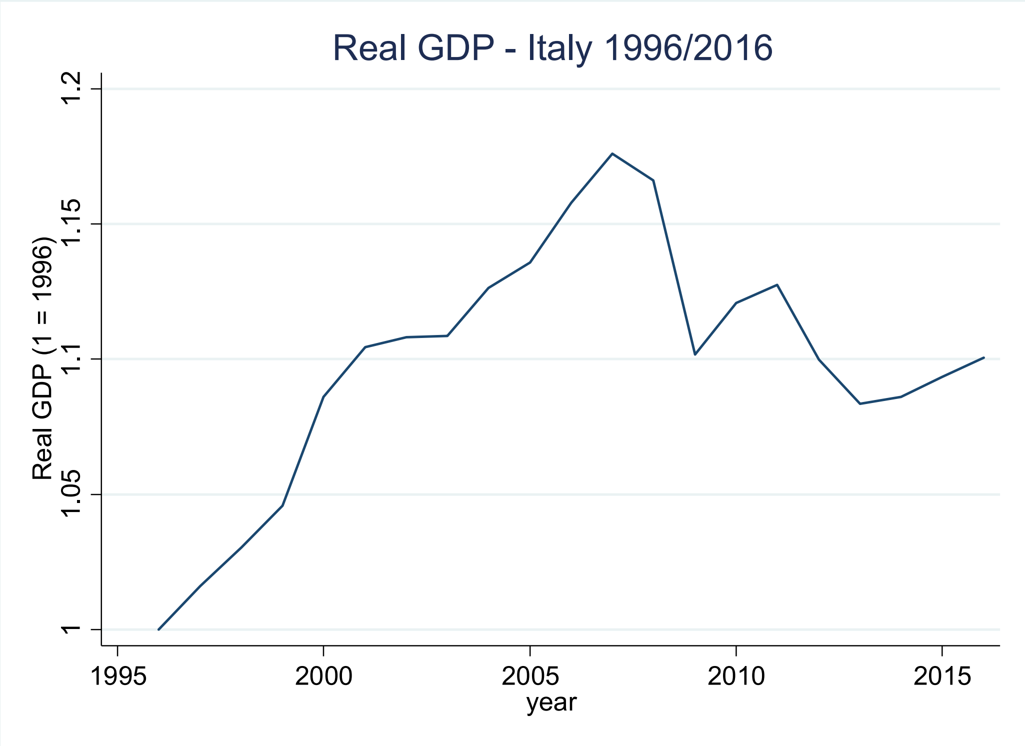
Source: Italian Labor Force survey.

Figure 3. Changes of PT_{pt}^{1991} between 2005 and 2016 by province.



Source: Italian 1991 Census

Figure 4. Italian Real GDP 1996-2016



Source: Eurostat data

Appendix

Table A1. Estimated ITT effects of the pool PT_{pt}^{1991} on regional real wages by age group. 1998-2014

	(1)	(2)	(3)	(4)
Age group	16-24	25-34	35-49	50-70
Real log wages	0.010	0.022	0.006	0.001
	(0.686)	(0.421)	(0.564)	(0.957)

Notes: wage data from the Bank of Italy SHIW survey. Number of observations: 171. All models include region and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group. Inference robust to clustering by region is carried out using wild bootstrap. P-values of the test that the estimated coefficient is statistically different from zero within parentheses. ***: p<.01; **: p<.05; *: p<.10

Table A2. Estimated ITT effects of the pool PT_{pt}^{1991} on unemployment (gross of inactivity) by age group. Provincial data 2005-2016

	(1)	(2)	(3)	(4)
Age group	16-24	25-34	35-49	50-70
Estimation method	OLS	OLS	OLS	OLS
(1) Without Z	0.116***	0.065***	0.100***	-0.058
	(0.024)	(0.012)	(0.015)	(0.088)
(2) With lagged Z	0.091***	0.083***	0.078***	-0.169***
	(0.011)	(0.012)	(0.020)	(0.049)
(3) With Z lagged twice	0.093***	0.072***	0.090***	-0.122**
	(0.018)	(0.014)	(0.016)	(0.052)
(4) With lagged Z and a national quadratic trend interacted with provincial GDP per head in 2004	0.082***	0.082***	0.081***	-0.138***
	(0.011)	(0.012)	(0.018)	(0.049)
(5) With lagged N and lagged Z (Arellano Bond GMM estimates)	0.092**	0.001	0.031**	-0.034
	(0.037)	(0.020)	(0.012)	(0.074)

Notes: number of observations: 1,224 in rows (1), (2) and (4), 1,122 in rows (3) and (5). All specifications include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table A3. Estimated effects of the current pool PT_{pt} on unemployment by age group. OLS, IV and Arellano Bond estimates. Static and dynamic specifications. Provincial data 2005-2016.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age group	16-24	16-24	25-34	25-34	35-49	35-49	50-70	50-70
Estimation method	OLS	IV	OLS	IV	OLS	IV	OLS	IV
(1) Without Z	0.079***	0.084***	0.046***	0.047***	0.070***	0.073***	-0.019	-0.042
	(0.020)	(0.018)	(0.010)	(0.008)	(0.008)	(0.009)	(0.067)	(0.065)
(2) With lagged Z	0.064***	0.065***	0.056***	0.059***	0.054***	0.057***	-0.126***	-0.137***
	(0.009)	(0.009)	(0.009)	(0.007)	(0.013)	(0.014)	(0.042)	(0.039)
(3) With Z lagged twice	0.064***	0.066***	0.050***	0.051***	0.057***	0.065***	-0.080*	-0.094**
	(0.015)	(0.014)	(0.013)	(0.009)	(0.010)	(0.010)	(0.046)	(0.040)
(4) With lagged Z and quadratic trend * provincial 2004 GDP	0.057***	0.059***	0.056***	0.058***	0.055**	0.059***	-0.100**	-0.113***
	(0.009)	(0.008)	(0.009)	(0.007)	(0.012)	(0.012)	(0.043)	(0.039)
(5) With lagged N and lagged Z (Arellano Bond GMM estimates)	-	0.061**	-	0.002	-	0.018*	-	-0.013
		(0.026)		(0.016)		(0.010)		(0.056)

Notes: number of observations: 1,224 in rows (1), (2) and (4), 1,122 in rows (3) and (5). Row 5: Arellano Bond estimates. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10.

Table A4. Years of social security contributions required to retire before minimum retirement age. By sector and gender.

Year	Male employees	Male self-employed	Female employees	Female self-employed
1996	36	40	36	40
1997	36	40	36	40
1998	36	40	36	40
1999	37	40	37	40
2000	37	40	37	40
2001	37	40	37	40
2002	37	40	37	40
2003	37	40	37	40
2004	38	40	38	40
2005	38	40	38	40
2006	39	40	39	40
2007	39	40	39	40
2008	40	40	40	40
2009	40	40	40	40
2010	40	40	40	40
2011	40	40	40	40
2012	42.08	42.08	41.08	41.08
2013	42.41	42.41	41.41	41.41
2014	42.5	42.5	41.5	41.5
2015	42.5	42.5	41.5	41.5
2016	42.83	42.83	41.83	41.83

Source: national legislation.

Table A5. Estimated effects of the current pool of potential senior workers PT_{pt} on alternative measures of employment by age group. OLS and IV. Provincial data 2005-2016.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age group	16-24	16-24	25-34	25-34	35-49	35-49	50-70	50-70
Estimation method	OLS	IV	OLS	IV	OLS	IV	OLS	IV
Baseline PT	-0.046***	-0.046***	-0.133***	-0.142***	-0.072***	-0.072***	0.219***	0.214***
	(0.006)	(0.004)	(0.014)	(0.016)	(0.019)	(0.019)	(0.024)	(0.026)
PT defined using rules	-0.051***	-0.055***	-0.139***	-0.165***	-0.078***	-0.083***	0.253***	0.260***
on social security contributions	(0.004)	(0.005)	(0.012)	(0.019)	(0.021)	(0.023)	(0.030)	(0.033)
PT net of the inactive	-0.087***	-0.092***	-0.237***	-0.277***	-0.109***	-0.149***	0.512***	0.418***
	(0.009)	(0.011)	(0.040)	(0.046)	(0.041)	(0.042)	(0.022)	(0.038)

Notes: number of observations: 1,224. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10

Table A6. Estimated ITT effects of a higher pool of potential senior workers PT_{pt}^{1991} on employment by age group. Static model with lagged X. Heterogeneities by gender. Provincial data 2005-2016.

	(2)	(4)	(6)	(8)
Age group	16-24	25-34	35-49	50-70
Males	-0.045***	-0.136***	-0.111***	0.103***
	(0.005)	(0.019)	(0.026)	(0.014)
Females	-0.018***	-0.076***	0.035**	0.168***
	(0.003)	(0.008)	(0.017)	(0.022)

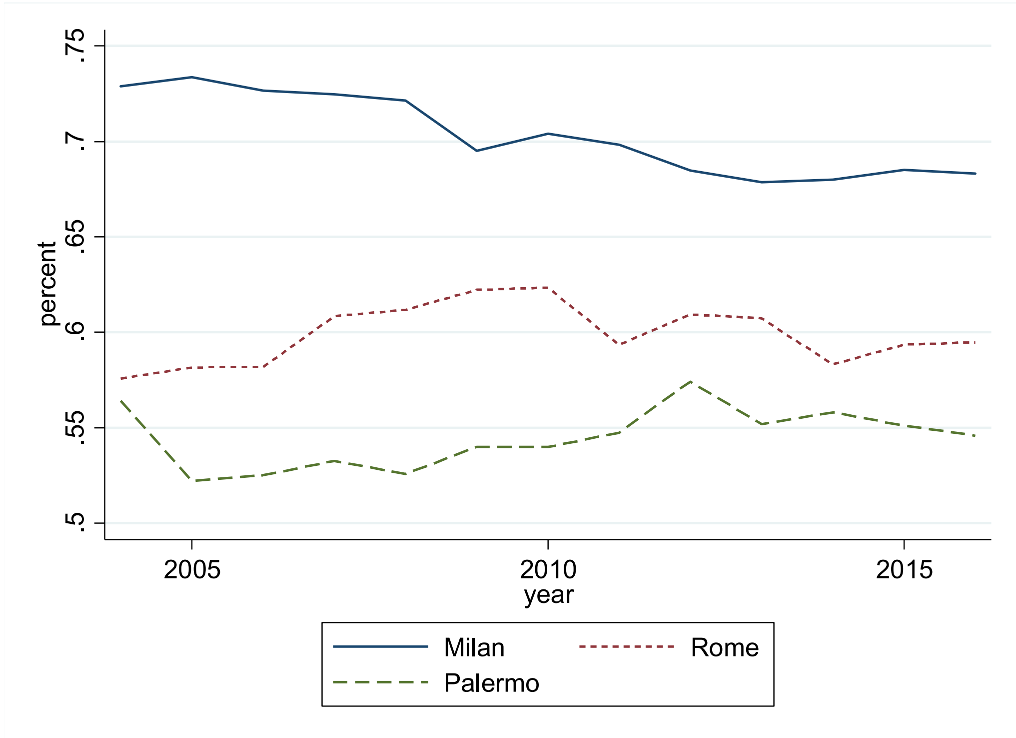
Notes: number of observations: 1,224. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10

Table A7. Estimated ITT effects of the pool of potential senior workers PT_{pt}^{1991} on employment by age group. Spill-over effects

Age group	(1) 16-24	(2) 25-34	(3) 35-49	(4) 50-70
Provincial PT	-0.063***	-0.195***	-0.103***	0.253***
	(0.005)	(0.017)	(0.027)	(0.032)
PT of neighboring provinces *1000	-0.002	-0.005*	0.003	0.009**
	(0.002)	(0.003)	(0.003)	(0.003)

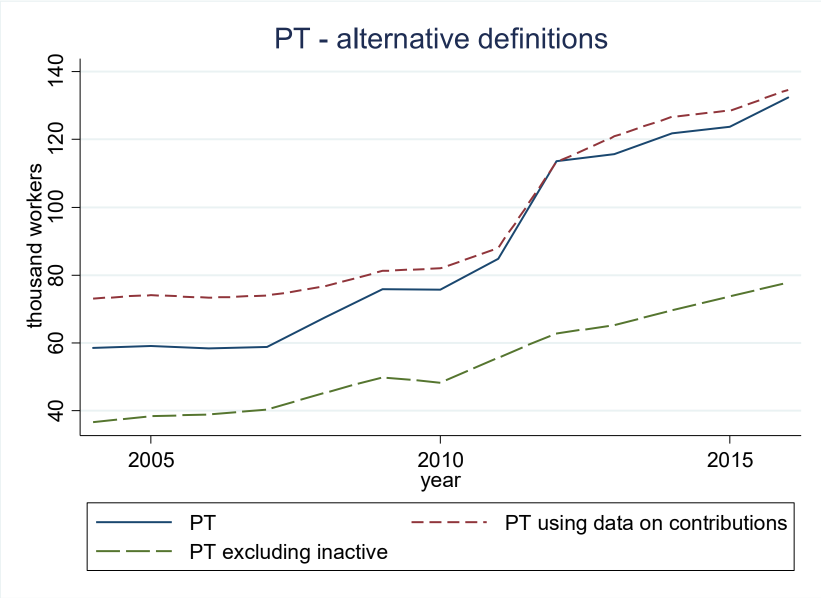
Notes: number of observations: 1,224. All models include province and wave dummies. The vector Z includes the firs lag of provincial GDP per capita; the first lags of the index of sectoral composition S by province; total population by province and age group; the percentage of workers with high school or higher degree by province and age group; the percentage of males by province and age group; average age by province and age group; the share of immigrants by province and age group. Standard errors are clustered by province. ***: p<.01; **: p<.05; *: p<.10

Figure A1. The sectoral shocks index S in three Italian provinces.



Source: Italian Labor Force Survey

Figure A2. Alternative definitions of the pool PT_{pt}



Source: Italian Labor Force Survey