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Abstract

By increasing the residual working horizon of employed individuals, pension reforms that raise minimum retirement age are likely to affect individual investment in health-promoting behaviors before retirement. Using the exogenous variation in minimum retirement age induced by the sequence of Italian pension reforms during the 1990s and 2000s, we show that middle-aged Italian makes who were close to regiment age reacted to the expected longer working horizon by increasing regular exercise and by reducing smoking, with positive consequences for obesity and self-reported satisfaction with health.

JEL Classifications: H55, I12, J26.

Keywords: Retirement, Working Horizon, Healthy Behaviors, Pension Reforms.

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Introduction

There is substantial research exploring the causal effects of retirement on individual health and health behaviors *after* retirement. This literature typically reports that the transition into retirement has positive effects both on self-reported health and on indices of physical health. Recent evidence includes Insler, 2014, for the U.S., Coe and Zamarro, 2011, and Eibich, 2015, for Europe and Zhao et al., 2013, for Japan.[1] Most studies argue that a mechanism behind these effects is the positive change in health-promoting behaviors – such as additional physical exercise and reductions in drinking and smoking – induced by retirement (see also Kaempfen and Maurer, 2016, and Celidoni and Rebba, 2016).[2]

The existing literature, however, has somewhat overlooked that exogenous changes in minimum retirement age can also affect behavior *before* retirement, by altering the residual

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[1] Godard, 2016, finds instead that the retirement transition has a positive effect on the incidence of obesity among European workers. This effect is particularly pronounced among those who were employed in blue-collar and physically demanding jobs. The studies on the effects of retirement on cognition generally find negative effects (see Rohwedder and Willis, 2010, Adam et al., 2012, Mazzonna and Peracchi, 2012, and Celidoni et al., 2013). The evidence is less clear-cut for mental health (see Charles, 2004, Börsch-Supan and Jürges, 2009, Johnston and Lee, 2009, Clark and Fawaz, 2009, Bonsang and Klein, 2012, Bertoni and Brunello, 2017, and Fonseca et al., 2015).

[2] These positive effects, however, may be short-lived and disappear with time (the so-called 'honeymoon effect' of retirement). For instance, Mazzonna and Peracchi, 2017, and Bertoni et al., 2017, estimate that - given age - a longer time spent in retirement has a negative effect on an index of overall physical health and on muscle strength, a robust predictor of disability, cardiovascular diseases and mortality.

working horizon of workers who – in the absence of constraints – would have chosen an optimal retirement age that falls below the minimum required by retirement rules. Changes in behavior due to a longer working horizon can occur, for instance, if earnings and employment in the additional period before retirement depend on health. In this case, affected individuals may have an incentive to keep fit so as to reap these benefits, and may therefore invest more in health-promoting behaviors.

To the best of our knowledge, only few contributions have examined the effects of a longer working horizon on individual behaviors before retirement, and none has considered the impact on health-promoting behaviors. Hairault et al., 2010, show that French workers exposed to an exogenous increase in their expected retirement age increase job search effort. They explain this finding by showing that the economic returns to jobs depend on their expected duration, which increases with retirement age. Similarly, Montizaan et al., 2010, and Brunello and Comi, 2015, use respectively Dutch and Italian data and show that policies that increase the residual working horizon have positive consequences on training participation by active older workers.[3] In a study close to ours, De Grip et al., 2012, find that

[3] Montizaan and Vendrik, 2014, find that the same policies negatively affected job satisfaction of treated Dutch workers. A longer working horizon may also have inter-generational consequences on the children of potential retirees. Manacorda and Moretti, 2006, find negative effects of a longer parental working horizon (and thus higher parental income) on the nest-leaving decisions of Italian youngsters. Battistin et al., 2014, find that policies raising the retirement age have negatively affected the supply of informal childcare provided by Italian grandparents, thereby reducing the number of grandchildren. Coda Moscarola et al., 2016, find that older Italian employed women reacted to the postponement of retirement induced by a recent reform by increasing their sick leave, and that this effect is

a Dutch reform reducing pension rights and postponing the minimum retirement age of public sector workers has reduced their mental health.

In this paper, we use data on several cohorts of Italian working men aged 42 to 51 during the period 1997 to 2011 to investigate whether changes in minimum retirement age – induced by reforms affecting eligibility conditions – have affected health-promoting behaviors before retirement. Within the selected time span, these reforms have increased minimum retirement age from 52 to 60.[4] By considering male workers aged 42 to 51, we focus on individuals who are generally too young to be retired but not too far from retirement.

Italy provides an interesting setup for the issue at hand, because of the sequence of pension reforms that occurred during the period under study (see e.g. Angelini et al., 2009). Before these reforms, eligibility for early retirement required in most cases that social security contributions be paid for at least 35 years. The sequence of reforms progressively tightened eligibility requirements in terms of both age and accrued years of contributions, thereby generating exogenous variation in the expected minimum retirement age for comparable workers belonging to different cohorts.

Our research design is based on instrumental variables. In Italy, the minimum time to retirement combines age requirements and years of paid social security contributions, that depend on individual careers and are likely to be endogenous, because negative health shocks affect both health behaviors and working careers. We instrument minimum *actual* years to retirement – computed using information on the actual years of paid contributions – with

stronger among low-income grandmothers living in areas with limited supply of childcare services.

[4] This increase refers to employees. For the self-employed, minimum retirement age increased from 56 to 61.

minimum *potential* years to retirement – obtained by replacing the number of years of paid social security contributions with potential experience, measured as age minus school leaving age, a pre-determined variable in our setup.

Conditional on survey year and age by school leaving age by sector dummies, the only remaining source of variation in the instrument is induced by changes in the retirement rules in place for individuals belonging to different cohorts – which are exogenous to individual choices. Conditional on age and time effects, however, younger cohorts may be more likely to adopt healthier lifestyles irrespective of their exposure to a longer working horizon, because of other factors that vary by cohort, including the changing general attitude toward smoking, drinking, dieting and exercising, and the improvement in general health conditions, as summarized by a longer life expectancy. Different cohorts may also be exposed to different education policies and initial labor market conditions.

Since the inclusion of survey year and age dummies forecloses the possibility of using cohort fixed effects, we control for cohort differences by using a "proxy variable" approach (see Heckman and Robb, 1985).[5] We show that our results hold irrespective of the inclusion of proxies for cohort effects, mitigating concerns about the presence of omitted cohort-level variables bias and supporting a causal interpretation of our estimates. As an alternative approach to identification, we replace survey year dummies with cohort dummies and use the proxy variable approach to capture time effects. Our qualitative results are unchanged.

We study the effects of changes in minimum retirement age on regular exercise, smoking and drinking alcohol. We also consider the impact on obesity, self-reported satisfaction with

[5] Since one cannot use simultaneously age, period and cohort effects, this approach suggests substituting for one of the three effects – in this case cohort dummies - with variables that pick up the underlying reasons for cohort-level changes in the outcome.

health, and a few indicators of nutritional habits.[6] Our results show that a one-year increase in the residual working horizon increases: the likelihood of exercising regularly by 2.33 percentage points (equivalent to about 12 percent of the mean value of the outcome for workers with median time to retirement – 10 years); the probability of refraining from smoking by 1.92 percentage points (equivalent to 2.9 percent of the mean value of the outcome for workers with median time to retirement); and the probability of having a body mass index below obesity by 1.81 percentage points (or 2.04 percent).[7] There is also evidence that a longer minimum time to retirement increases self-reported high satisfaction with own health, although this effect is imprecisely estimated.[8]

Potential mechanisms explaining why healthy behaviors change in the presence of pension reforms that increase minimum retirement age include the lifetime income effects associated to a longer working life and the need to keep fit longer. While the former mechanism applies to all workers, the latter is less likely to apply to public sector employees, who in Italy have stronger job guarantees than private sector workers, and therefore may be less concerned with preserving their health in order to work longer. Since we find that exercising regularly changes significantly for private sector workers – including the self-employed – but not for

[6] Our measures of investments in healthy behaviors are self – reported and drawn from a survey that does not include either time diaries or objective health measures. Additionally, the range of health behaviors that we observe in the available data does not exhaust all the important behaviors that may affect health. For instance, we do not have any information on the use of illicit drugs or unsafe sex.

[7] The instrumental variables effects on smoking and obesity are statistically significant at the 10 percent level of confidence.

[8] However, the reduced-form effect of potential years to retirement on satisfaction with health is statistically significant at conventional levels.

public sector employees, we conclude that the need to keep fit in the expectation of having to work longer is a plausible mechanism explaining our results.[9]

Assessing the effects of a longer working horizon on behaviors before retirement has relevant policy implications. Several OECD countries have recently introduced pension reforms that raise minimum retirement age in order to deal with the increased burden of population ageing on public finances. By delaying retirement and by increasing the residual working horizon of employed individuals, these reforms may generate unexpected costs and benefits. In this paper, we highlight that one benefit could be better health before retirement, as constrained individuals react to the longer horizon by investing in some healthy behaviors and reducing some unhealthy ones. Ceteris paribus, in countries with universalistic public health care, better health before retirement may generate important savings, and these savings should be accounted for when evaluating the impact of pension reforms.

The remainder of the paper is organized as follows. In Section 1 we introduce the institutional background and the sequence of pension reforms affecting minimum retirement age in Italy. Section 2 presents the data. We discuss the empirical setup in Section 3 and results in Section 4. Conclusions follow.

1. Institutional Background: Recent Italian Pension Reforms

[9] We have also explored whether effects are heterogeneous by education and type of job (physically demanding or not), but found that they are not.

In this section, we briefly describe the key features of the Italian social security reforms that were introduced between 1997 and 2011 – the period covered by our data – and repeatedly changed retirement eligibility rules.[10]

Before 1992, the minimum age for *old-age* pension for men was 60 for employees in the private sector and for self-employed workers, 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 instead possible at any age for workers who had paid social security contributions for at least 35 years.[11] The first social security reform (the so-called "Amato" reform – from the name of the Prime Minister at the time of its introduction) took place in 1992 and introduced a progressive increase in the requirements for eligibility to *old age* pensions, that were to reach at least 20 years of paid contributions and age 65 by 2001, as shown in Appendix Table A1.

In 1995, a second major reform (the "Dini" reform) tightened the eligibility requirements for *seniority* pensions, that were to raise gradually from 1997 to 2008 to reach either 40 years of paid contributions independently of age, or 57 years of age and 35 years of paid contributions.[12] The reform also prescribed a faster increase of eligibility requirements for the

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[10] We exclude the "Monti-Fornero" reform, that was effective from January 2012. See Angelini et al., 2009, and Bottazzi et al., 2011, among others for further details on the pension reforms occurring during our sample period.

[11] Since our empirical analysis is restricted to men, we do not discuss here the changes in pension eligibility rules for females.

[12] By introducing eligibility requirements for *seniority* pensions, this reform abolished the so-called "baby-pensions", that since 1973 allowed public employees with at least 20 years of paid contributions to retire independently of age. This requirement was set as low as 14 years,

self-employed, as documented in Appendix Table A2, where we summarize all changes in *seniority* pension eligibility rules introduced by the reforms of interest.

After only three years, in 1998, pension eligibility rules changed again with the "Prodi" reform, that accelerated the transition period and increased the minimum retirement age to 58 for the self-employed (from 2001 onwards).

The fourth reform took place in 2005, when Welfare Minister Roberto Maroni modified the eligibility requirements for *seniority* pensions, introducing a sharp 3-year increase in minimum eligibility age (the so-called "scalone" or "big jump"), 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 incoming left-wing government led by Romano Prodi (or "Prodi bis") postponed the proposed 3-year increase to 2011, introducing instead a gradual adjustment in the requirements, starting again from 2008, as documented in Appendix Table A2. For this reason, no worker has actually retired under the requirements prescribed by the "Maroni" reform. Yet, this reform is still relevant for our purposes, because it changed the expected minimum retirement age during the years from 2005 to 2008. 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 (see Appendix Table A2).

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

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6 months and 1 day for married women with children who were employed in the public sector.

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. Since this change could have altered health behaviors independently of the changes in minimum retirement age, we account for it in our empirical analysis.[13]

2. The Data

Our data consist of a main and an auxiliary sample. The main sample is from the survey "*Aspetti della Vita Quotidiana"* (*Aspects of Daily Life,* hereafter AVQ), carried out on a yearly basis by the Italian Bureau of Statistics (ISTAT), and the auxiliary sample is from the *Survey on Household Income and Wealth* (SHIW from now on), conducted on a bi-annual basis by the Bank of Italy.

AVQ is a cross-sectional annual survey of a representative sample of about 50,000 individuals. It covers several aspects of daily life, including behaviors such as exercising, smoking, drinking and several dietary habits. Since information on healthy behaviors is collected only since 1997, we use the 14 waves from 1997 to 2011. Our data do not include 2004, because the survey did not take place in that year, and the years from 2013 to 2015, because in these waves the information on individual age is only available in five-years brackets. We also exclude data for 2012. In that year, a new pension reform was introduced. With just one year of data, however, we cannot separately identify the effect of the new reform from a year fixed effect.

We focus on middle aged males between 42 and 51 years, who are generally too young to be retired but not too far from retirement. On the one hand, individuals aged 51 are the oldest

[13] In Italy, the average gross pension replacement rate for average earners was 69.8 percent during the period 2005-2011. Source: OECD, *Pensions at a Glance,* several years.

workers whose retirement eligibility status is not affected by the reforms that we consider. On the other hand, by choosing a 10-year age interval before age 51, we increase sample size and the precision of our estimates. We show, however, that our qualitative results are robust to limiting our sample to individuals aged 47 to 51, who are even closer to retirement age.

We exclude females because their labor market careers – a crucial aspect of our empirical exercise – are often more discontinuous than those of men, due to their childbearing responsibilities. After eliminating from the sample the very few who are retired, disabled or have never worked in their life, as well as those with missing values in the variables used in the analysis, we end up with a final sample of 38,439 individuals.[14]

We construct the following indicators of healthy lifestyles: a dummy equal to 1 if the individual exercises on a regular basis, and 0 otherwise; a dummy equal to 1 if he does not smoke, and 0 otherwise; a dummy equal to 1 if he does not drink alcohol regularly and 0 otherwise;[15] a dummy equal to 1 if his body mass index (BMI) is below 30 (not obese), and 0 otherwise. As indicator of health satisfaction, we define a dummy equal to 1 if the individual is very satisfied with his own health, and 0 otherwise.

We also consider as supplementary outcomes the following indicators of nutritional habits: a dummy equal to 1 if the individual refrains from eating red meat at least once a day and 0 otherwise; a dummy equal to 1 if he eats vegetables or fruit at least once a day, and 0

[14] This sample includes the currently employed and the unemployed with at least a previous job (1.5% of the sample). For each individual in the sample, we have information on the sector of current or last employment.

[15] We define regular drinking if the individual drinks at least 1 or 2 glasses of either wine or beer per day, or if he drinks alcohol outside meals on a daily basis. Our data do not have information on binge drinking.

otherwise; a dummy equal to 1 if he refrains from soft drinks at least once a day, and 0 otherwise.

The sequence of pension reforms illustrated in the previous section has generated exogenous variation in the minimum retirement age of workers with the same age, who have paid social security contributions for the same number of years and belong to the same sector, but were born in different years. To isolate this variation from endogenous changes in the length of working careers, we define potential years to retirement (*PYR)* at time *t* as the minimum number of years to retirement prescribed by the law in place at the time, under the assumption that the years of paid social security contributions are equal to the years of potential labor market experience, measured as the difference between age and school leaving age.[16] Potential years to retirement differ from actual years to retirement (*YR)* because the latter are based on actual rather than potential labor market experience.

We illustrate how *PYR* varies over time using the example shown in Table 1. In the table, we consider hypothetical private sector employees with high school education and aged 42, 47 and 51 in 1997, 1998, 2005 and 2008, the first year of application of each reform. We show that *PYR* increased from 15 years in 1997 to 17 in 2008 for those aged 42, from 10 to 12 for those aged 47 and from 3 to 8 for those aged 51.

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[16] We compute school leaving age as the canonical number of years required to complete the highest attained school degree plus six – the school starting age. This variable can take the following values: 15 for individuals with a lower secondary degree; 17 for those with a short-term high school degree; 19 for those with a regular high school degree; 22 for those with a bachelor degree and 24 for those with higher degrees. We use minimum working age (15) for those with less than or equal to lower secondary education.

We further document the variation of *PYR* over time – conditional on age, sector of activity and average educational attainment – by plotting in Figure 1 the residuals of the regression of *PYR* on school dummies separately by age (42, 47 and 51) and sector (private, public or self-employment). The observed jump between 2003 and 2005 corresponds to the steep rise in minimum retirement age introduced by the Maroni reform.

Using the survey "*Health Conditions and Use of Health Services*", carried out by the National Statistical Institute (ISTAT) in 2005 and 2013, and the information on the age when individuals started to smoke, we construct cohort trends in the share of individuals smoking at age 14. We also collect data on: life expectancy at birth by cohort from the *Human Mortality Database*; the share of individuals doing regular physical exercise at age 20[17] by cohort from the ISTAT survey "*The Spare Time of Italians*", carried out in 2000 and 2006; years of compulsory education by cohort - equal to five for the cohorts born from 1946 to 1948 and to eight for younger cohorts (see Brunello, Fort and Weber, 2009);[18] real GDP per capita at school leaving age. In Appendix Figure A1 we plot these indicators – with the exception of years of compulsory education – by birth cohort (from 1946 to 1969).

The AVQ survey includes variables that we use as covariates in our regressions: age, educational attainment, sector of employment in the current or previous job (private, public or self-employment), region of residence, marital status and presence of children.[19] Since the

[17] After finishing high school, as physical education is compulsory at school.

[18] We code this variable as a dummy equal to one for the cohorts with eight years of compulsory education and to zero for the cohorts with five years of compulsory education.

[19] We have also information on type of job (whether physically demanding or not) and type of accommodation (as a proxy of wealth). However, we only use these controls in a

survey does not include information on the total years of paid social security contributions, because of the lack of data on labor market histories, we can compute potential years to retirement *PYR* but not actual years *YR.*

To measure *YR* we turn to the auxiliary sample from the SHIW survey, that includes data on (self – reported) years of paid social security contributions at the time of the interview, but has no information on health-promoting behaviors. This sample consists of 8,670 males aged 42 to 51 from 1998 to 2010 (in 7 different waves), for whom we can compute both *PYR* and *YR*.[20]

Table 2 presents the summary statistics of the main variables introduced in this section. In our main sample, 19 percent of the individuals exercise regularly, 66 percent do not smoke, 45 percent do not drink alcohol regularly, 83 percent do not eat red meat at least once a day, 85 percent eat fruits or vegetables at least once a day, 87 percent do not drink soft drinks at least once a day and 89 percent is not obese. In addition, 19 percent are very satisfied with their health. While the actual minimum number of years to retirement *YR* is 13.21, the potential number *PYR* is about three years shorter at 10.06. Average school leaving age is just below 18; the share of self-employed and public sector employees is 29 and 21 percent respectively. Finally, the percent married and with no children is 85 and 18 percent respectively.

3. The Empirical Approach

Galama et al., 2013, have recently developed a structural model of consumption, leisure, health, health behaviors, wealth accumulation and retirement decisions using the human

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robustness test, because they are not available in the auxiliary sample that we use for our first-stage regressions.

[20] The main and the auxiliary sample are broadly comparable in terms of observable characteristics.

capital framework of health developed by Grossman, 2000. We present in the Appendix a simplified version of this model to illustrate how changes in minimum retirement age R_{min} affect investment in healthy behaviors.

In the model, the health stock is in the utility function both during working life and during retirement, and also affects earnings during working life. Individuals cannot modify their health stock directly, but can invest in costly healthy behaviors, which affect current and future health. The optimal health investment before retirement equalizes marginal benefits and costs. Optimal retirement age *R* is also subject to choice, and it is jointly determined with consumption and investment in healthy behaviors.

Exogenous changes in minimum retirement age R_{min} affect only the individuals with an optimal retirement age equal to or lower than R_{min}. As illustrated in the Appendix, a sufficient condition for an increase in R_{min} to promote investment in healthy behaviors before retirement is that the marginal benefits of better health during the additional period of active working life – in terms of higher utility and earnings – are higher than the marginal costs – in terms of the foregone benefits due to the shorter retirement period.

Let the *optimal* minimum time to retirement chosen by an individual aged *A* in the absence of any constraint be $YR^* = R - A$. Let instead *YR* be the *actual* minimum time to retirement, that depends on the exogenous retirement rules $\{R_{min}, SSC_{min}\}$ – where SSC_{min} is the minimum number of years of paid social security contributions required to access early retirement – on individual age and accumulated social security contributions at age *A*, defined as *SSC*. By modifying R_{min} and SSC_{min}, policy makers can alter the minimum time to retirement, which in turn affects the behavior of individuals for whom $YR^* \leq YR$ holds.

We model the empirical relationship between actual minimum years to retirement and healthy behaviors as follows

$$B_{it} = \alpha + \beta YR_{it} + \gamma X_{it} + \varepsilon_{it} \qquad (1)$$

where the indices *i* and *t* are for the individual and time, *B* is for healthy behaviors, *X* is a vector of controls, that includes age by school leaving age by sector of employment dummies, survey year dummies, regional dummies, and dummies for having kids and being married, and ε is the error term. Since our sample consists of male workers aged 42 to 51 and in Italy transitions from a sector to another are infrequent in this age range,[21] we treat both school leaving age and sector of employment as pre-determined variables.

The parameter β measures the marginal effect of a one-year increase in the actual minimum time to retirement on healthy behaviors. Denote with *s* the share of individuals with $YR^* \leq YR$ and assume that β_u and β_c are the marginal effects of *YR* on *B* for the sub-groups with $YR^* > YR$ and $YR^* \leq YR$, respectively. Then $\beta = \frac{\partial B}{\partial YR} = \beta_c s * \left[1 + \frac{\partial s}{\partial YR}\frac{YR}{s}\right]$, because $\beta_u = 0$. Therefore, the estimated marginal effect of *YR* in (1) compounds the effect on the sub-group with $YR^* \leq YR$ and the effect on the share of individuals who are constrained by the minimum retirement age.

As discussed in Section 1, the sequence of pension reforms introduced by Italian governments during the 1990s and 2000s repeatedly modified both the minimum retirement

[21] Using quarterly data from the Italian *Labor Force Survey*, we estimate the following year-to-year average transition rates across sectors for workers aged 42 to 51 during the years 2004 to 2011: 1.44 percent from self-employed to private sector employee; 0.23 percent from self-employed to public sector employee; 0.09 percent from private to public sector employee; 0.25 percent from private sector employee to self-employed; 0.08 percent from public to private sector employee and 0.03 percent from public sector employee to self-employed.

age R_{min} and the minimum number of years of paid social security contributions (SSC_{min}) required to access retirement with a *seniority* pension. These changes – that have been specific to the self-employed and to public and private sector employees – have generated variability among cohorts in the minimum number of years to retirement for workers of the same age, who have paid social security contributions for the same number of years and belong to the same sector (i.e. private, public, self-employed).

Since eligibility requires a minimum number of years of paid social security contributions, *YR* is shorter for workers with no employment gaps in their careers, even conditional on age, sector and school leaving age. One reason for observing discontinuous careers is the experience of negative health shocks – either currently or in the past – which in turn may depend on the adoption of unhealthy behaviors. These shocks generate reverse causality, as people who experience bad health – or adopt unhealthy behaviors – also end up having a longer working horizon. In this case, conditioning on vector *X* does not suffice in preventing OLS estimates of β in Eq. (1) from being inconsistent.

We address reverse causality and the possibility of recall bias in the years of contribution by instrumenting *YR* with *PYR*, the potential years to retirement, or the minimum residual working horizon under the assumption of continuous careers. Contrarily to *YR*, the selected instrument does not depend on individual careers and varies with age, retirement eligibility conditions and education, that are either exogenous or predetermined for the relevant age group (42 to 51). Conditional on the variables in vector *X*, the residual variation in *PYR* is due *exclusively* to changes in retirement rules, which we treat as exogenous to individual behavior. Although it is true that the instrument varies with age, education and sector, we never exploit variation *between* these groups for identification, that only hinges upon variation in *PYR* *within* these groups and between cohorts – see Table 1 and Figure 1.

In the reduced form equation,

$$B_{it} = \alpha_R + \beta_R PYR_{it} + \gamma_R X_{it} + \varepsilon_{Rit} \qquad (2)$$

the identification of parameter β_R as the intention to treat effect of *PYR* on *B* requires that, conditional on the vector *X*, *PYR* is as good as randomly assigned. A potential threat to our identification strategy is that individuals belonging to different cohorts differ because of omitted factors that correlate with both the length of their residual working horizon and their adoption of healthy behaviors.

One such factor is life expectancy. Younger cohorts share longer expectancy and longer working horizons, and the former is likely to correlate with healthy behaviors. Other factors are the changing propensities to smoke and exercise in modern societies, the former declining and the latter increasing, that clearly affect behaviors and correlate with longer working horizons, given age. To control for these factors, we include in vector *X* cohort trends in life expectancy at birth and in the shares of smokers at age 14 and of individuals engaged in physical activity at age 20. Since the cohorts in our data could have been exposed to different education policies and initial labor conditions, we also control for the cohort-specific years of compulsory education and for real GDP per capita at school leaving age.

The identification of parameter β in Eq. (1) as the Average Causal Response (*ACR*)[22] of behaviors *B* to minimum time to retirement *YR* requires two additional conditions: first, we need a significant first-stage relationship between *YR* and the selected instrument *PYR.* Visual evidence that such relationship exists is reported in Figure 2, where we plot the distribution of *YR* and *PYR* in our auxiliary sample, as well as the linear regression fit, showing a strongly positive association between the two. Formal evidence is discussed in the next section.

[22] *ACR* is a generalization of the Local Average Treatment Effect (*LATE*) when the treatment is continuous (Angrist and Imbens, 1995).

Second, we require that *PYR* influences *B* only via its effect on *YR*, a tenable exclusion restriction in this context.

As described in the previous section, our main data source – the AVQ survey – has detailed information on the adoption of healthy behaviors, but no information on the years of paid social security contributions. Since we cannot compute *YR* using these data, we only estimate the reduced form equation (2). To estimate parameter β in Eq. (1), we use our auxiliary SHIW sample and a two-sample instrumental variables estimator (TSIV – see Angrist and Krueger, 1992 and Inoue and Solon, 2010).[23] Letting π be the effect of *PYR* on *YR* in the first stage regression, the IV estimate of parameter *β* is obtained as the ratio $\beta = \frac{\beta_R}{\pi}$.[24] In all regressions, we cluster standard errors by cohort, sector and school leaving age – the level of variation of *PYR*.

In our baseline specification, we estimate equations (1) and (2) for each behavior separately. We also perform two robustness exercises: first, we estimate all equations jointly using seemingly unrelated estimation and test whether the coefficients associated to *PYR* are jointly equal to zero. We strongly reject this hypothesis (p-value of the test < 0.01). Second, using the stepdown methods for multiple testing based on re-sampling devised by Romano and

[23] We estimate the first stage using data for 1998, 2000, 2002, 2004, 2006, 2008 and 2010. Even if we observe in this dataset the accrued years of social security contributions, we still need to assume continuous careers from the time of observation until retirement (as done also by Battistin et al., 2009).

[24] Inference is carried out by bootstrapping. Notice that, since there is a single endogenous variable and the model is just-identified, our estimation procedure is equivalent to a two-sample two-stage least squares procedure, which involves computing the fitted values of *YR* in the AVQ data using the first-stage coefficients estimated in SHIW.

Wolf, 2005, we show that the statistical significance of our baseline estimated effects is confirmed even when we take into account the problem of multiple testing and the consequent over-rejection of the true null hypothesis.

4. Empirical Results

4.1 Intention-To-Treat (ITT) effects: baseline estimates

If pension reforms that raise minimum retirement age affect at least part of the relevant population, and workers understand the effects of these changes, average expected retirement age should raise with the increase of minimum retirement age. To document that this is the case, we use our auxiliary sample drawn from *SHIW* – where individuals are asked about their expected retirement age – and regress expected retirement age on minimum age and on the vector of controls *X*. We estimate that a one year increase in minimum retirement age raises expected age by about half a year (0.52, standard error 0.02).[25]

We estimate equation (2) using a linear probability model and report in Table 3 the estimated effects of potential minimum time to retirement *PYR* on the probability of exercising regularly, refraining from smoking and regular alcohol consumption, having a BMI lower than 30 and being satisfied with own health. The table reports both the estimated coefficients (multiplied by 100) and the percentage effects computed with respect to the mean of the outcome variable for workers with *PYR* = 10, the median value in the sample.

We find that a 1-year increase in *PYR* raises the probability of exercising regularly by 6.14 percent and reduces the probability of smoking and drinking by 1.07 and 0.86 percent. While

[25] See Bottazzi et al., 2006, and Baldini et al., 2015, for additional evidence for Italy.

the former two effects are statistically significant at the 1 and 10 percent level of confidence respectively, the last effect is imprecisely estimated.

Consistently with the increase in regular exercising, we find that a higher value of *PYR* increases the probability of not being obese by 0.92 percent per year – statistically significant at the 5 percent level. Also, we find that a longer time to retirement increases the probability of being very satisfied with own health by 3.66 percent per year.

Since changes in body weight are driven by the difference between calories intake and calories expenditure, we also look at the effects of changes in *PYR* on nutritional choices. Table 4 show that the estimated effects on the probability of refraining from eating red meat or drinking soft drinks and the probability of eating fruit and vegetables at least once a day are small and imprecisely estimated.

4.2 ITT effects: robustness tests

The estimated *ITT* effects presented in Table 3 are robust to several robustness checks.[26] First, Panel 1 of Appendix Table A3 shows that results are similar when we compute the marginal effects of *PYR* on behaviors using a Probit specification instead of a linear probability model.

Second, since we are simultaneously testing effects on multiple outcomes, there is a risk of over-rejecting some of the true null hypotheses because of pure chance. Reassuringly, as reported in Panel 2, adjusting the p-values of our estimates for multiple testing using the

[26] For brevity, the robustness tests on the outcomes presented in Table 4 are not presented here, but are available from the authors upon request.

stepdown method devised by Romano and Wolf, 2005, does not alter the statistical significance of our results.[27]

Next, Panels 3 to 7 in the table show that our baseline results hold also when: a) we exclude the proxies for cohort effects; b) we include additional covariates – a dummy for working in a physically demanding job and dummies for accommodation type (luxury apartment, standard apartment, social housing, country house, sheltered housing and villa or single house as the omitted category); c) we include age-specific time trends; d) we replace survey year dummies with a cubic time trend; e) we restrict our sample to individuals aged 47 to 51.

We also verify whether the linear specification of the relationship between *PYR* and the selected outcomes is overly restrictive by replacing *PYR* with dummies for each level of *PYR*. Appendix Figure A2 shows – for the outcomes in Table 3 – the estimated effects of the *PYR* dummies,[28] their 95 percent confidence intervals and the estimated linear trend. The figure suggests that the linear functional form is a good approximation of the data, as in nearly all cases the effects implied by the linear trend lie within the estimated confidence intervals.

[27] If a single test is performed at the 5% level of confidence and the null hypothesis being tested is true, we expect a 5% chance of incorrectly rejecting it. If N independent tests are simultaneously carried out and all corresponding null hypotheses are true, the probability of at least one incorrect rejection is $1\text{-}0.95^N$. In our case, N=5 and this probability is equal to 22.6%. Romano and Wolf, 2005, have devised a stepdown method for multiple testing based on resampling, which allows control over the Family Wise Error Rate – that is, the probability of incorrectly rejecting *one or more* true null hypothesis – and accounts for dependence across tests to improve power.

[28] In our data, *PYR* ranges from 1 to 20 years. Hence, we consider *PYR* = 1 as our baseline and include in the regressions 19 separate dummies $d_k = I(PYR = k)$, k = 2, …, 20.

Additionally, we show in Appendix Table A4 that our baseline results in Table 3 are qualitatively unchanged when – instead of controlling for year and age effects and using proxies for cohort effects – we adopt a different identification approach that controls for age and cohort effects and exploits the variation over time in *PYR*, using as proxies for year effects the real GDP per capita at the time of the survey, the relative price of each outcome of interest, a dummy for being surveyed after the introduction of the 2005 smoking ban and the variable $\overline{B_{65-75}}_t$, defined as the average value of *B* in year t for males aged 65 to 75, who are not affected by pension reforms.[29] Not only are our results confirmed but also we find that the estimated effect of a higher value of *PYR* on the probability of consuming alcohol regularly is statistically significant at the 5 percent level of confidence. Adding cohort-specific time trends to this specification does not alter our results.[30]

In a further robustness test – not reported but available upon request – we re-define our outcomes as ordinal variables. Again, our qualitative results are unchanged. For instance, in the case of exercising we distinguish between no exercising, light physical activity, irregular and regular exercising, and find that an additional year to retirement has a negative effect on the former two categories and a positive effect – of similar size – on the latter two categories.

[29] We cannot estimate a placebo regression for workers aged 65-75, because *PYR* – the potential years to retirement – cannot be computed for this group of individuals, who are already beyond minimum retirement age. We have estimated the effects of *PYR* on males aged 25 to 30, who are very far away from retirement and may therefore be less affected by changes in pension eligibility rules. We find that *PYR* has small and largely insignificant effects on the healthy behaviors of this group (results are available from the authors).

[30] Results are available from the authors upon request.

Last but not least, we consider the potential confounding effects on our estimates of changes in pension replacement rates, that could have modified healthy behaviors independently of changes in *PYR*. The relevant change during our sample period is the method of computation of pension benefits, that was altered starting in 2007 for those who could retire from 2010 onwards. To control for this effect, we add to our baseline specification a dummy equal to 1 for individuals observed in years 2007-2011 and eligible to retire since 2010, and to 0 otherwise, but find no change in our results (not shown but available upon request). To the best of our knowledge, during our observation period no other change in the pension system has affected our respondents with a variation over time and across cohorts that is consistent with the observed changes in *PYR*, confounding the identification of its effects.

4.3 Heterogeneous ITT effects

To investigate whether responses to changes in *PYR* are heterogeneous, we estimate separate regressions by sector of activity (public employees, private employees and self-employed workers). Results are presented in Table 5. We find that changes in minimum retirement age have increased regular exercise and reduced obesity only for private sector workers (employed and self-employed).

Mechanisms explaining why healthy behaviors change in the presence of pension reforms that increase minimum retirement age include income effects – when pension benefits are lower than earnings before retirement, additional years of working life raise expected lifetime earnings and the willingness to spend for the gym – and the need to keep fit longer. This need is less pressing for public sector workers, who in Italy have stronger job guarantees than private sector workers, and therefore may be less concerned with preserving their health in order to work longer. Therefore, while the former mechanism applies to all workers, the latter applies mainly to private sector employees and the self-employed. Since we find that regular exercise and the probability of being obese do not change significantly for public sector employees, as they should if changes in lifetime income is the key mechanism, we believe that the need to keep fit to be able to work longer is an important candidate mechanism that could explain our results.

An additional candidate is that workers anticipating a longer minimum working horizon compensate the future reduction of leisure by substituting current working time with leisure time and more exercising. Yet we find that weekly working hours are not negatively affected by changes in *PYR*.[31]

4.4 Average causal responses (TSIV effects)

We have presented so far the intention to treat effects of potential minimum time to retirement on healthy behaviors. We now turn to estimating the average causal responses of these behaviors to changes in the actual minimum time to retirement *YR*, using potential time *PYR* as the instrument for actual time and a two-sample instrumental variables estimator.[32] First, we regress actual time on potential time and the vector of control *X* in our auxiliary *SHIW* sample, and report the result at the bottom of Table 6. According to our estimates, a one-year increase in *PYR* raises *YR* conditional on *X* by 0.40 years. Since the value of the first-stage F statistic for instrument weakness is 48.5, well above the threshold of 10, our instrument is not weak. Second, we compute for each health behavior the two-sample IV

[31] We regress the log of the number of weekly working hours on *PYR* and the vector X and find that one additional potential year to retirement increases hours by 0.7 percent, a small, positive and significant effect.

[32] For the TSIV analysis we use a cubic trend in survey year instead of survey years dummies, because the AVQ and SHIW surveys were carried out in different years. As shown in Table A3, ITT effects are comparable when using cubic trends and dummies for survey year.

estimate of β as $\frac{\beta_R}{\pi}$, and show our results in the first row of the table (multiplied by 100).[33] It turns out that the IV effects of *YR* on *B* are sizable and about twice as large as the *ITT* estimates shown in Table 3. When evaluated at the mean value for those with median *PYR*, a one-year increase in the actual minimum time to retirement increases the likelihood of exercising regularly by 12.02 percent and reduces smoking by about 2.9 percent. We also confirm that a longer time to retirement induces a reduction in obesity (by about 2 percent). Finally, health satisfaction also increases, but the *TSIV* estimate is imprecise.

Conclusions

We have investigated the effects of postponing minimum retirement age on healthy behaviors before retirement using data for several cohorts of middle aged Italian working men observed during the period 1997 – 2011, when repeated pension reforms took place in an effort to contain public expenditure. Italy is an interesting laboratory because these reforms generated exogenous variation in minimum retirement age.

While much research has been devoted to establishing whether and how retirement affects the health of retired individuals, less has been done to understand whether policy measures that alter the length of the residual working horizon affect health and healthy behaviors before retirement.

We have estimated the causal effect of changes in the potential as well as actual minimum number of years to retirement on the health lifestyles of Italian workers aged 42 to 51 and found that – when evaluated at the mean value of each outcome for workers with median time to retirement – a one-year increase in minimum actual years to retirement has raised the

[33] Bootstrapped standard errors clustered by cohort, school leaving age and sector in parentheses (1,000 bootstrap replications).

likelihood of exercising regularly by 12 percent and reduced smoking by about 2.9 percent. Considering that minimum actual years to retirement have increased in our sample by 2.3 years between 2000 and 2010, these effects are not small. Probably due to the increase in calorie expenditures associated to the improvement in regular exercise rather than to a reduction in calorie intakes, the probability of being obese has also fallen. Consistently with these findings, we also estimate a positive effect on self-reported high satisfaction with health.

Our finding that regular exercise and smoking respond to economic and financial incentives is not new in the empirical literature. Mitchell et al, 2013, conduct a meta-analysis of empirical studies that have investigated the impact of financial incentives on exercise related behaviors in the US. They report that incentives have significant and positive effects on exercise in eight of the eleven studies they consider. Qualitatively similar conclusions are reached by Dishman et al, 2009, in their evaluation of the *Move to Improve* intervention in the US, and by Eibich, 2015, who uses German data to show that an important mechanism through which retirement affects health is an increase in physical activity.

On the other hand, Volpp et al, 2009, randomly assigned 878 employees of a multinational company based in the United States to receive information either about smoking-cessation programs or about programs plus financial incentives. They find that the incentive group had significantly higher rates of smoking cessation than did the information-only group both 9 or 12 months and 15 or 18 months after enrollment.

Pension reforms that raise minimum retirement age have been introduced in several OECD countries to deal with the increased burden of population ageing on public finances. By delaying retirement and by increasing the residual working horizon of employed individuals, these reforms may reap unexpected dividends. We have shown that one such dividend could be better health before retirement, as constrained individuals react to the longer expected horizon by investing in healthy behaviors (regular exercise) and reducing unhealthy ones (smoking and drinking). Better health before retirement may generate important savings to private and public expenditure, and these savings should be accounted for when evaluating the overall impact of pension reforms.[34]

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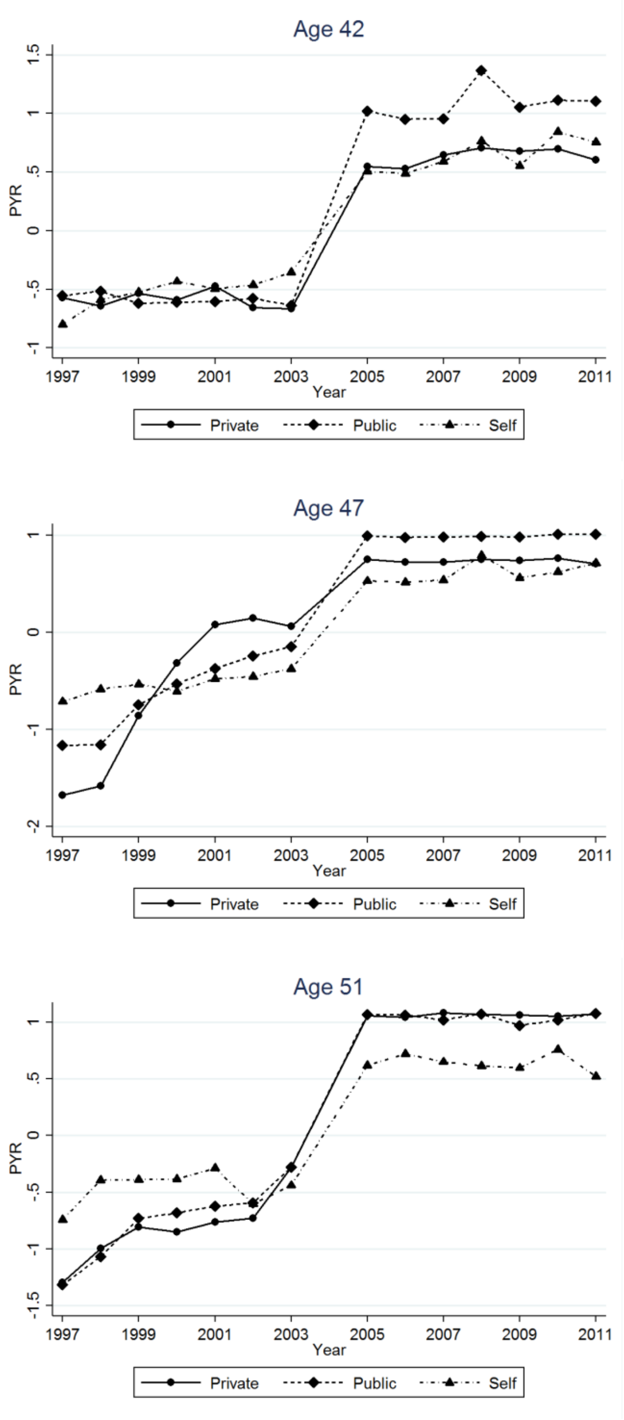
[34] One might argue that the adoption of healthy behaviors could increase longevity. However, given the abundant empirical evidence supporting the "compression on morbidity" hypothesis – see e.g. Felder et al., 2010 – we do not expect that a longer life will lead to increased health expenditure, as the relevant determinant of health expenditure is time-to-death, not age per se.

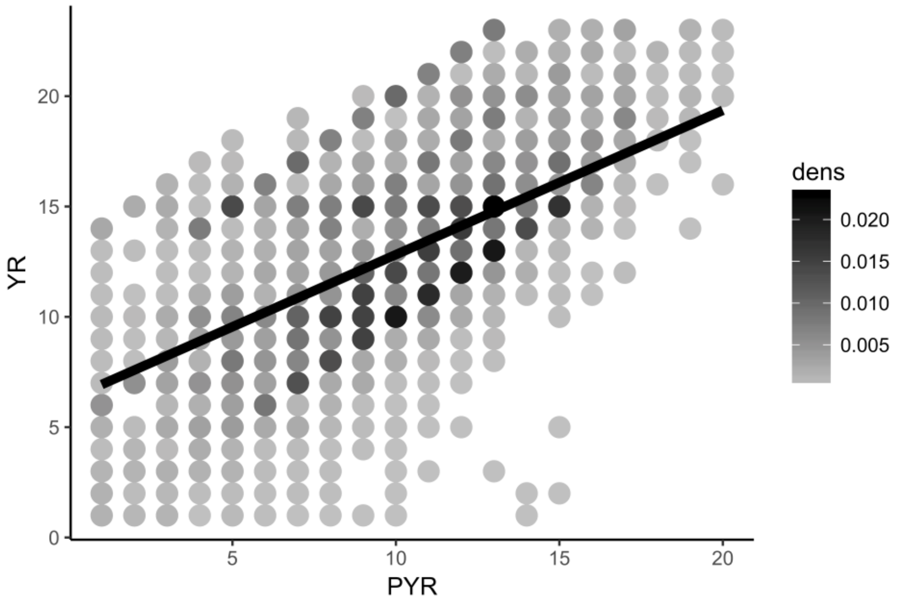
Figures and Tables

Figure 1. PYR residuals by age, sector of activity and survey year.



Notes: PYR residuals are obtained from the regression of PYR on school degree dummies.

Figure 2. Minimum years to retirement *YR* and *PYR* - Bank of Italy SHIW data 1998-2010.



Notes: the figure reports a scatterplot and a linear fit of *YR* on *PYR* using Bank of Italy SHIW data. Darker dots indicate cells with higher density.

Table 1. Potential years to retirement (PYR) for private sector employees with a high school degree and aged 42, 47 or 51, observed in the first year of application of each reform (1997, 1998, 2005, and 2008).

Age	Year	Reform	Contributions paid	PYR
42	1997	Dini	23	15
42	1998	Prodi	23	15
42	2005	Maroni	23	17
42	2008	Prodi Bis	23	17
47	1997	Dini	28	10
47	1998	Prodi	28	10
47	2005	Maroni	28	12
47	2008	Prodi Bis	28	12
51	1997	Dini	32	3
51	1998	Prodi	32	5
51	2005	Maroni	32	8
51	2008	Prodi Bis	32	8

Notes: see Tables A1 and A2 for pension eligibility requirements under the different reforms.

Table 2. Descriptive statistics

	Mean	Std. Dev.
Treatment Variable		
YR	13.21	4.51
Instrumental Variable:		
PYR	10.06	3.93
Outcomes:		
Exercises regularly	0.19	0.40
Does not smoke	0.66	0.47
Does not drink alcohol regularly	0.45	0.50
Not obese	0.89	0.32
Very satisfied with health	0.19	0.39
Does not eat red meat at least once a day	0.83	0.38
Eats fruit or vegetables at least once a day	0.85	0.35
Does not drink soft drinks at least once a day	0.87	0.34
Other covariates:		
Age	46.40	2.88
Survey year	2003.46	4.52
Birth year	1957.06	5.38
School leaving age	17.54	2.99
Public employee	0.21	0.41
Self-employed	0.29	0.45
No kids	0.18	0.39
Married	0.85	0.36

Notes: Data for YR are from the Bank of Italy SHIW survey, all other data are from the ISTAT AVQ survey. Both samples include male workers aged 42 to 51 who do not have missing values in the variables used in the analysis. Total number of observations in SHIW: 8,670. Total number of observations in AVQ: 38,439. “Does not drink soft drinks at least once a day” is only observed since 1998 (N = 35,339) and “Not obese” since 2001 (N = 25,392). YR is the actual work horizon, or the number of years before becoming eligible to retire according to the rules in place at the time of the interview and using the observed number of years of social security contributions. PYR is the potential work horizon, computed using the potential number of years of social security contributions. The excluded occupational sector is “private employee”.

Table 3. Intention-To-Treat effects of potential years to retirement (*PYR*) on healthy behaviors – OLS estimation – linear probability models.

	(1)	(2)	(3)	(4)	(5)
	Exercise regularly	No Smoking	No alcohol regularly	Not obese	Very satisfied with health
PYR/100	1.11***	0.71*	0.39	0.81**	0.68**
	(0.32)	(0.37)	(0.39)	(0.32)	(0.28)
% effect	6.14***	1.08*	.86	.92**	3.66**

Notes: the table reports the estimated effects of *PYR/100* on the outcome listed at the top of each column. Percentage effects are computed with respect to the mean value of the outcome for the group with median PYR (10 years). Total number of observations: 38,439; for "Not obese": 25,392. All models include age by school degree by sector dummies, survey year dummies, regional dummies, a dummy for having kids, a dummy for being married, life expectancy at birth by cohort, the percentage of people exercising regularly at age 20 by cohort, the percentage of people smoking at age 14 by cohort, compulsory years of education by cohort, GDP per capita at school leaving age. Standard errors clustered by cohort, school leaving age and sector in parentheses. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Table 4. Intention-To-Treat effects of potential years to retirement (*PYR*) on indicators of nutritional habits – OLS estimation – linear probability models.

	(1)	(2)	(3)
	No red meat at least once a day	Fruit or vegetables at least once a day	No soft drinks at least once a day
PYR/100	0.09	0.05	0.39
	(0.27)	(0.29)	(0.31)
% effect	.11	.05	.45

Notes: the table reports the estimated effects of *PYR/100* on the outcome listed at the top of each column. Percentage effects are computed with respect to the mean value of the outcome for the group with median PYR (10 years). Total number of observations: 38,439; for "No soft drinks at least once a day": 35,339. All models include age by school degree by sector dummies, survey year dummies, regional dummies, a dummy for having kids, a dummy for being married, life expectancy at birth by cohort, the percentage of people exercising regularly at age 20 by cohort, the percentage of people smoking at age 14 by cohort, compulsory education years by cohort, GDP per capita at school leaving age by cohort and educational level. Standard errors clustered by cohort, school leaving age and sector in parentheses. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Table 5. Effects of potential years to retirement (*PYR*) on healthy behaviors by sector of activity – OLS estimation – linear probability models.

	(1) Exercise regularly	(2) No Smoking	(3) No alcohol regularly	(4) Not obese	(5) Very satisfied with health
Public sector	-0.26	0.88	0.70	-0.41	0.43
	(0.75)	(0.88)	(0.97)	(0.89)	(0.81)
Private sector	1.27***	0.53	0.42	1.07***	0.74**
(including self-employed)	(0.37)	(0.40)	(0.44)	(0.35)	(0.30)

Notes: the table reports the estimated effects of *PYR/100* on the outcome listed at the top of each column. Split sample estimation for workers employed in the public sector and in the private sector (including the self-employed). Number of observations: 8,229 for the public sector and 30,210 for the private sector, except for "not obese" (5,128 and 20,264). All models include age by school degree by sector dummies, survey year dummies, regional dummies, a dummy for having kids, a dummy for being married, life expectancy at birth by cohort, the percentage of people exercising regularly at age 20 by cohort, the percentage of people smoking at age 14 by cohort, compulsory years of education by cohort, GDP per capita at school leaving age. Standard errors clustered by cohort, school leaving age and sector in parentheses. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Table 6. Average Causal Response of *YR* on healthy behaviors – Two-sample IV estimation

	(1)	(2)	(3)	(4)	(5)
	Exercise regularly	No Smoking	No alcohol regularly	Not obese	Very satisfied with health
YR/100	2.33**	1.92*	0.84	1.81*	1.18
	(1.04)	(1.10)	(1.16)	(1.05)	(0.86)
% effect	12.02**	2.90*	1.88	2.04*	6.28
First-stage *PYR/100*	0.40***				
	(0.05)				
First-stage F statistic	48.50				

Notes: the table reports the Two-sample IV estimates of the effects of years to retirement *YR* on the outcome listed at the top of each column. Percentage effects are computed with respect to the mean value of the outcome for the group with median PYR (10 years). The ITT is estimated in the AVQ sample. Total number of observations in AVQ: 38,439; for "Not Obese": 25,392. The first stage is estimated using the SHIW sample. Total number of observations in SHIW: 8,670. All models include age by school degree by sector dummies, a cubic trend in survey year, regional dummies, a dummy for having kids, a dummy for being married, life expectancy at birth by cohort, the percentage of people exercising regularly at age 20 by cohort, the percentage of people smoking at age 14 by cohort, compulsory years of education by cohort, GDP per capita at school leaving age. Bootstrapped standard errors clustered by cohort, school leaving age and sector in parentheses (1,000 bootstrap replications). ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Appendix

1. An illustrative model

Following Galama et al., 2013, we consider an individual in his forties who intends to spend his residual lifetime partly at work and partly in retirement. In each period before retirement, his utility is given by

$$U_{wt} = U_w(C_t, H_t) \tag{A.1}$$

where *C* is consumption and *H* is the health stock in period *t*.[35]

Let B_t be a strictly positive measure of health investment (or healthy behavior) and p_t its unit cost.[36] For instance, this investment can be an healthy diet or physical exercise. The relationship between health and health investment is given by the following law of motion

$$\frac{\partial H_t}{\partial t} = \mu B_t - \sigma H_t \tag{A.2}$$

By increasing B_t, the individual can compensate the natural decay of health. Using (A.2) we can write health at time *t* as a function of initial health and of the entire history $0 \le t' < t$ of

[35] Leisure is assumed equal to L_0 during work and to τL_0 during retirement, with $\tau > 1$. The utility function is separable so that $U(C_t, H_t, L_0) = g(L_0)U(C_t, H_t) = U_w(C_t, H_t)$ during work, and $U(C_t, H_t, \tau L_0) = g(\tau L_0)U(C_t, H_t) = U_r(C_t, H_t)$ during retirement. Letting $\gamma = \frac{g(\tau L_0)}{g(L_0)} > 1$, $U_r = \gamma U_w(C_t, H_t)$, where *γ>1* indicates that "…a dollar with leisure – while retired – is better than a dollar that is only had together with work…" (Stock and Wise, 1990, p.213).

[36] We broadly interpret the unit cost as including both monetary and non-monetary costs. We assume that there are no corner solutions in the optimal choice of health investments. See Galama et al., 2013, for a discussion.

health investment $B_{t'}$

$$H_t = H_0 e^{-\sigma t} + \int_0^t \mu B_x e^{-\sigma(t-x)} dx \qquad \text{(A.3)}$$

In the optimization problem, we consider the entire prior history of health investment $B_{t'}$ (Galama et al., 2008, p.5), that affect current health.

Denoting assets with A_t, the inter-temporal budget constraint is given by

$$\frac{\partial A_t}{\partial t} = \delta A_t + Y(H_t) - C_t - p_t B_t \qquad \text{(A.4)}$$

where income Y is equal to yearly earnings $W(H_t)$ before retirement and to Γ (pension benefits) after retirement. $W(H_t)$ is an increasing and concave function of *H*, the health stock. Better health affects earnings both by raising productivity and by increasing the probability of being gainfully employed. As in Galama et al., 2013, we do not distinguish further between these two channels.

Changes in minimum retirement age R_{min} affect individual choice only if R_{min} is binding, that is, if optimal retirement age is lower than or equal to R_{min}. We shall focus on this case. Moreover, we shall only consider the optimization problem faced by the individual before his retirement, as this is the situation studied in the current paper. The individual chooses consumption and healthy behaviors to maximize the following inter-temporal utility

$$\max_{C_t, B_t} \int_s^{R_{min}} [U_{wt}(C_t, H_t)] e^{-rt} dt + \int_{R_{min}}^T [\gamma U_{wt}(C_t, H_t)] e^{-rt} dt \qquad \text{(A.5)}$$

subject to (A.3) and (A.4), where T denotes total lifetime, that we assume to be independent of health, as in Galama et al., 2013, *s* is initial age and *r* is the interest rate. Following Galama et al., 2008, this is equivalent to maximizing

$$\max_{C_t, B_t} \int_s^{R_{min}} [U_{wt}(C_t, H_t)]e^{-rt}dt + \int_{R_{min}}^{T} [\gamma U_{wt}(C_t, H_t)]e^{-rt}dt$$

$$+\lambda_0 \int_s^T [\delta A_t + Y(H_t) - C_t - p_t B_t]e^{-\delta t}dt \tag{A.6}$$

where $\lambda_t = \lambda_0 e^{-\delta t}$ is the co-state variable associated to (A.4).

The first order necessary condition for optimal $B_{t'}$ when $t' \leq R_{min}$ is

$$\int_{t'}^{R_{min}} [\frac{\partial U_{wt}}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt + \int_{R_{min}}^{T} [\gamma \frac{\partial U_{wt}}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt$$

$$-\lambda_0 \int_{t'}^{T} [p_t - \frac{\partial Y(H_t)}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-\delta t}dt = \phi(B_{t'}, C_{t'}, \lambda_0, R_{min}) = 0 \tag{A.7}$$

For consumption, the first order condition is

$$\frac{\partial U_{wt'}}{\partial C_{t'}} - \lambda_0 e^{(r-\delta)t'} = 0 = \chi(B_{t'}, C_{t'}, \lambda_0) \tag{A.8}$$

Finally, by differentiating (A.7) with respect to λ_0 we obtain

$$\int_s^T [\delta A_t - C_t - p_t B_t]e^{-\delta t}dt + \int_s^{R_{min}} [Y(H_t)]e^{-\delta t}dt$$

$$+ \int_{R_{min}}^{T} [\Gamma_t]e^{-\delta t}dt = \theta(B_{t'}, C_{t'}, R_{min}) = 0 \tag{A.9}$$

At the optimum, health investment when $t' \leq R_{min}$ equalizes the marginal benefits during both active working life $\int_{t'}^{R_{min}} [\frac{\partial U_{wt}}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt$ and after retirement $\int_{R_{min}}^{T} [\gamma \frac{\partial U_{wt}}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt$ and the net marginal costs $\lambda_0 \int_{t'}^{T} [p_t - \frac{\partial Y(H_t)}{\partial H_t} \frac{\partial H_t}{\partial B_{t'}}]e^{-\delta t}dt$.

Since the contribution of health to wages ends with retirement, we can re-write (A.7) as follows

$$\int_{t'}^{R_{min}}[\frac{\partial U_{wt}}{\partial H_t}\frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt+\lambda_0\int_{t'}^{R_{min}}\frac{\partial Y(H_t)}{\partial H_t}\frac{\partial H_t}{\partial B_{t'}}]e^{-\delta t}dt$$

$$+\int_{R_{min}}^{T}[\gamma\frac{\partial U_{wt}}{\partial H_t}\frac{\partial H_t}{\partial B_{t'}}]e^{-rt}dt-\lambda_0\int_{t'}^{T}p_t e^{-\delta t}dt=\phi(B_{t'},C_{t'},R_{min},p_t)=0$$

Totally differentiating (A.7), (A.8) and (A.9) with respect to $R_{min}, B_{t'}$, $C_{t'}$ and λ_0 we obtain

$$\begin{aligned}
&\phi_1 dB_{t'}+\phi_2 dC_{t'}+\phi_3 dR_{min}+\phi_4 d\lambda_0=0\\
&\chi_1 dB_{t'}+\chi_2 dC_{t'}+\chi_3 d\lambda_0=0\\
&\theta_1 dB_{t'}+\theta_2 dC_{t'}+\theta_3 dR_{min}=0
\end{aligned}$$

where $\phi_i=\frac{\partial\phi}{\partial Z_i}$, $\chi_i=\frac{\partial\chi}{\partial Z_i}$ $\theta_i=\frac{\partial\theta}{\partial Z_i}$ and vector Z includes R_{min}, $B_{t'}$, λ_0 and $C_{t'}$.

By Cramer's rule, we get that

$$\frac{\partial B_{t'}}{\partial R_{min}}=\frac{\phi_3\theta_2\chi_3+\theta_3(\chi_2\phi_4-\chi_3\phi_2)}{\Delta} \qquad \text{(A.10)}$$

where the determinant of the bordered Hessian Δ is positive because of the second order conditions. We know that χ_3, θ_2, χ_2 and ϕ_4 are negative and that θ_3 is positive. Assuming that ϕ_2 is also positive, the second term on the right hand side of (A.10) is also positive. Since $\theta_3=(Y(H_t)-\Gamma_t)e^{-\delta t}$, this term is the effect of a higher minimum retirement age on health behaviors that operates via higher income. The sign of $\frac{\partial B_{t'}}{\partial R_{min}}$ depends on the sign of ϕ_3. A sufficient condition for this sign to be positive is

$$\phi_3=\lambda_0\frac{\partial Y(H_{R_{min}})}{\partial H_{R_{min}}}e^{(r-\delta)R}-\frac{\partial U_{wR_{min}}}{\partial H_{R_{min}}}(\gamma-1)>0$$

In words, postponing minimum retirement age increases optimal healthy behaviors before retirement if the benefits of a longer working life induced by better health are higher than the costs in terms of leisure due to a shorter retirement period. Since the second term on the right hand side of the above expression is positive, a sufficient condition for healthy behaviors to increase with minimum retirement age is that better health before retirement positively affect earnings, for instance because it increases the probability of being gainfully employed.

2. Tables and Figures

Table A1. Old-age pension eligibility during the sample period (1997-2011)

Sector: Retirement year:	Private Age & YContr	Public Age & YContr	Self-employed Age & YContr
1997	63+18	65+18	63+18
1998	64+18	65+18	64+18
1999	64+19	65+19	64+19
2000	65+19	65+19	65+19
2001 onwards	65+20	65+20	65+20

Note: Y Contr: years of paid contributions.

Table A2. Old-age pension eligibility according to the different reforms in place during the sample period (1997-2011)

a. "Dini" reform. Survey years of application: 1997

Sector:	Private		Public		Self-employed	
Retirement year:	Age & YContr	Only YContr	Age & YContr	Only YContr	Age & YContr	Only YContr
1997	52&35	36	52+35	36	56+35	40
1998	53&35	36	53&35	36	57&35	40
1999	53&35	37	53&35	37	57&35	40
2000	54&35	37	54&35	37	57&35	40
2001	54&35	37	54&35	37	57&35	40
2002	55&35	37	55&35	37	57&35	40
2003	55&35	37	55&35	37	57&35	40
2004	56&35	38	56&35	38	57&35	40
2005	56&35	38	56&35	38	57&35	40
2006	57&35	39	57&35	39	57&35	40
2007	57&35	39	57&35	39	57&35	40
2008 onwards	57&35	40	57&35	40	57&35	40

b. “Prodi” reform. Years of application: 1998-2004

Sector: Retirement year:	Private Age & YContr	Private Only YContr	Public Age & YContr	Public Only YContr	Self-employed Age & YContr	Self-employed Only YContr
1998	54&35	36	53&35	36	57&35	40
1999	55&35	37	53&35	37	57&35	40
2000	55&35	37	54&35	37	57&35	40
2001	56&35	37	55&35	37	58&35	40
2002	57&35	37	55&35	37	58&35	40
2003	57&35	37	56&35	37	58&35	40
2004	57&35	38	57&35	38	58&35	40
2005	57&35	38	57&35	38	58&35	40
2006	57&35	39	57&35	39	58&35	40
2007	57&35	39	57&35	39	58&35	40
2008 onwards	57&35	40	57&35	40	58&35	40

c. “Maroni” reform. Years of application: 2005-2007

Sector: Retirement year:	Private Age & YContr	Private Only YContr	Public Age & YContr	Public Only YContr	Self-employed Age & YContr	Self-employed Only YContr
2005	57&35	38	57&35	38	58&35	40
2006	57&35	39	57&35	39	58&35	40
2007	57&35	39	57&35	39	58&35	40
2008	60&35	40	60&35	40	61&35	40
2009	60&35	40	60&35	40	61&35	40
2010 onwards	61&35	40	61&35	40	62&35	40

d. “Prodi bis” reform. Years of application: 2008 onwards

Sector: Retirement year:	Private Age & YContr & (Age+ YContr)	Private Only YContr	Public Age & YContr & (Age+ YContr)	Public Only YContr	Self-employed Age & YContr & (Age+ YContr)	Self-employed Only YContr
2008	58&35	40	58&35	40	59+35	40
2009	59&35&95	40	59&35&95	40	60+35, 96	40
2010	59&35&95	40	59&35&95	40	60+35, 96	40
2011	60&35&96	40	60&35&96	40	61+35, 97	40
2012	60&35&96	40	60&35&96	40	61+35, 97	40
2013 onwards	61&35&97	40	61&35&97	40	62+35, 98	40

Notes: see Table A1. The requirement in terms of (age+ YContr) only applies since 2009

Table A3. The effects of potential years to retirement (PYR) on health behaviors. Alternative specifications.

	(1)	(2)	(3)	(4)	(5)
	Exercise regularly	No Smoking	No alcohol regularly	Not obese	Very satisfied with health
Panel 1: marginal effects from probit models					
PYR/100	0.99***	0.84**	0.41	0.76**	0.68**
	(0.33)	(0.37)	(0.41)	(0.33)	(0.29)
Panel 2: adjusting p-values for multiple hypothesis testing using Romano and Wolf (2005) stepdown method					
PYR/100	1.11***	0.71*	0.39	0.81*	0.68*
[p-values]	[0.01]	[0.09]	[0.38]	[0.06]	[0.09]
Panel 3: excluding proxies for cohort trends from baseline estimates in Table 3					
PYR/100	1.03***	0.55	0.17	0.69**	0.49*
	(0.32)	(0.36)	(0.37)	(0.33)	(0.28)
Panel 4: adding to the baseline in Table 3 a dummy for working in a physically demanding job and dummies for type of accommodation (luxury apartment, standard apartment, social housing, country house, sheltered housing – omitted category: villa or single house)					
PYR/100	1.15***	0.73**	0.39	0.83**	0.69**
	(0.32)	(0.36)	(0.39)	(0.32)	(0.28)
Panel 5 adding age-specific time trends to the baseline in Table 3					
PYR/100	1.05***	0.61*	0.67*	0.93***	0.68**
	(0.34)	(0.36)	(0.39)	(0.33)	(0.29)
Panel 6: using a cubic trend instead of dummies for survey year					
PYR/100	0.93***	0.77**	0.34	0.72**	0.47**
	(0.32)	(0.36)	(0.41)	(0.32)	(0.28)
Panel 7: keeping only individuals aged 47 to 51 years.					
PYR/100	1.08**	0.36	0.80*	1.06**	1.06**
	(0.42)	(0.48)	(0.48)	(0.48)	(0.41)

Notes: the table reports the estimated effects of *PYR/100* on the outcome listed at the top of each column. Total number of observations: 38,439; for "Not Obese": 25,392. For Panel 4 only, total number of observations: 18,533; for "Not Obese": 12,140. Unless stated differently, the specification adopted is the one of Table 3. Standard errors clustered by cohort, school leaving age and sector in parentheses. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

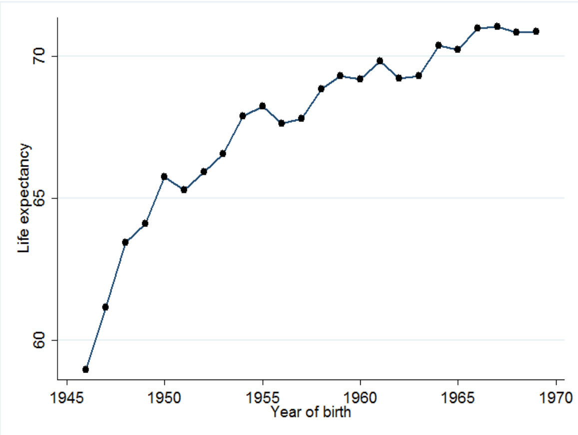
Table A4. The effects of potential years to retirement (PYR) on health behaviors. Alternative identification approach.

	(1)	(2)	(3)	(7)	(8)
	Exercise regularly	No Smoking	No alcohol regularly	Not obese	Very satisfied with health
PYR/100	1.11***	0.74**	0.67*	0.93***	0.75***
	(0.35)	(0.37)	(0.40)	(0.33)	(0.28)

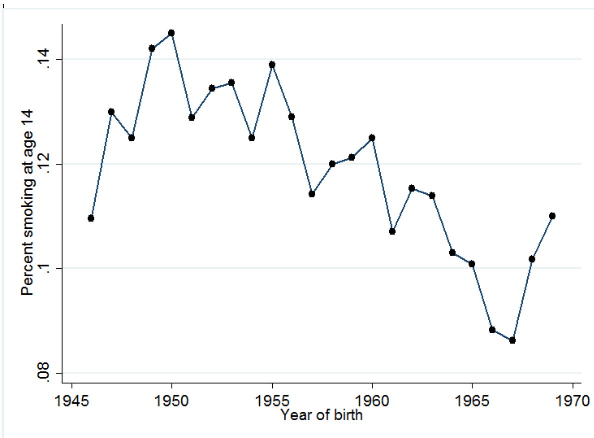
Notes: the table reports the estimated effects of *PYR/100* on the outcome listed at the top of each column. Total number of observations: 38,439; for “Not obese”: 25,392. All models include age by school leaving age by sector dummies, cohort dummies, regional dummies, a dummy for having kids, a dummy for being married, real GDP per capita by survey year, relative prices of the outcomes by survey year, national trends in the relevant outcome for males aged 65-75 by survey year, a dummy for being surveyed after the introduction of the 2005 smoking ban. Standard errors clustered by cohort, school leaving age and sector in parentheses. ***: significant at the 1% level; **: significant at the 5% level; *: significant at the 10% level.

Figure A1. Cohort trends in life expectancy at birth, the percentage of smokers at age 14 and of individuals engaged in physical activity at age 20.

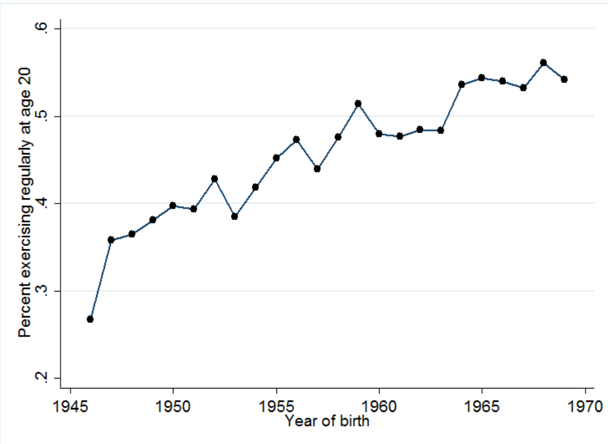
a. Life Expectancy at birth



b. Percent smoking at age 14



c. Percent exercising regularly at age 20



d. Average GDP per capita at school leaving age

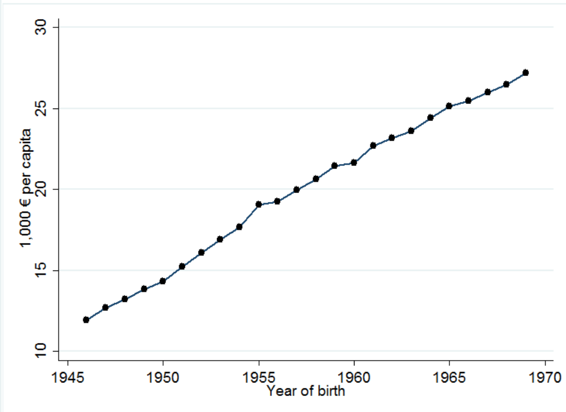
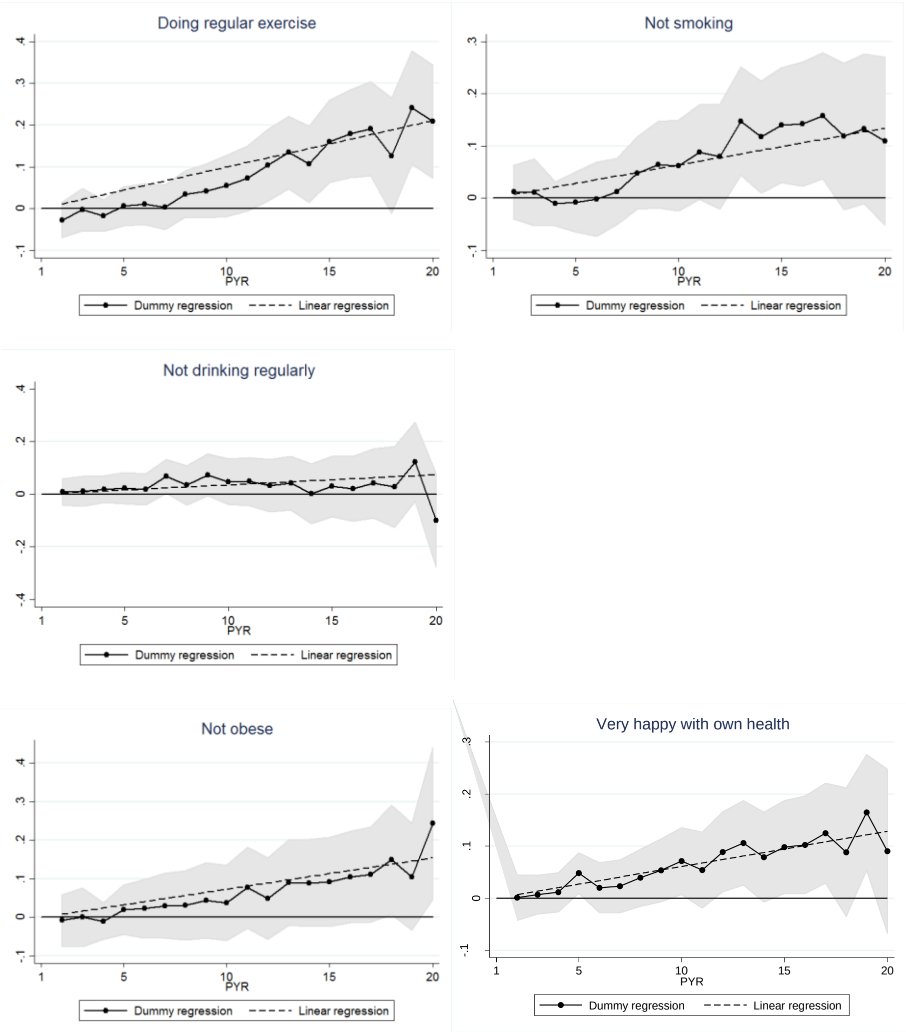


Figure A2. Intention-To-Treat non-linear effects of potential years to retirement (PYR) on health behaviors



Notes: the figures show the estimated effects of *PYR* dummies and their 95 percent confidence intervals, as well as the estimated linear effect of *PYR*. There are 19 *PYR* dummies, with *PYR* = *k*, k = 2, …, 20. The omitted category is *PYR=1*. The specification adopted is the one of Table 3.

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