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Is there a Retirement-Health Care utilization puzzle? Evidence from SHARE data in Europe

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Abstract[1]

We investigate the causal impact of retirement on health care utilization. Using SHARE data (from 2004 to 2013) for 10 European countries, we show that health care utilization increases when individuals retire. This is true both for the number of doctor's visits and for the intensity of medical care use (defined as the probability of going more than 4 times a year to the doctor's). This increase turns out to be driven by visits to general practitioners', while specialists' visits are not affected. We also find that the impact of retirement on health care utilization is significantly stronger for workers retiring from jobs characterized by long hours worked - more than 48 hours a week and/or being in the 5th quintile of the distribution of hours worked. This suggests that at least part of the increase in medical care use following retirement is due to the decrease in the opportunity cost of time faced by individuals when they retire.

JEL classification: J26, I10, C26

Keywords: Retirement, Health, Health Care Utilization

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1 Introduction

Retirement is a milestone event in the life-cycle of individuals, often associated with significant changes in lifestyles and behaviors of those involved. The effects of retirement on a number of health and economic outcomes have been extensively studied in the literature and, despite the fact that the age of retirement is a fairly predictable event, many studies have reported unexpected jumps around the age of retirement in the health status of individuals, as well as in consumption (and savings) patterns. The fact that such jumps are hard to reconcile with the standard life-cycle theory - under the assumption that agents are forward looking -, combined with the mixed evidence available from empirical studies, have fostered a debate as to whether health investments and consumption (savings) should vary smoothly over the life cycle or, in contrast, experience discontinuities at the time of retirement.

The standard conceptual framework for analyzing health care utilization and health investment is the Grossman's human-capital model (Grossman 1972; 2000). In this model, individuals invest in health for both 'consumption' (health provides utility) and 'production' motives (healthy individuals get higher earnings). Within this framework, the stock of health is assumed to depreciate with age (increasingly in old age) but individuals can increase their health stock by investing in health inputs (e.g. medical care use, healthy lifestyles). Health care utilization is expected to increase smoothly with the aging process, to preserve the health stock, until it becomes too costly doing so: death occurs when the health stock falls below a given threshold. In this context, both the health stock and investments in health care are optimized over time and, consistent with the life-cycle theory, no jump should be observed.

In a recent paper, however, Galama et al. (2013) challenge the assumption that health is always at the 'optimal' level and investigate the implications of corner solutions. They extend the standard Grossman model introducing retirement as a permanent transition from employment to non-employment. They show that the 'optimal' level of consumption, health investment and health stock can be discontinuous at the age of retirement, as retired individuals reallocate away from health investment toward more consumption goods. The jump in health investment arise from the fact that health does not influence income after retirement and from the difference in the marginal utility of leisure after retirement. The existence of a discontinuous change in health investment at the time of retirement is what we investigate in this paper.

Early studies on consumption and saving patterns have shown that consumption (savings) significantly decreases (increase) at the time of retirement. This is known as the "retirement-consumption puzzle" since such changes are hard to reconcile with the consumption-smoothing pattern predicted by the standard life-cycle theory when agents are forward looking (Banks et al., 1998; Battistin et al., 2007).

More recent studies have put forward an explanation for the observed drop in consumption expenditures after retirement, based on substitution across categories of goods. In particular, consumption would be shifted from work-related goods (e.g. clothes, transportation) that are typically bought on the market, toward time-intensive consumption goods (e.g. recreation, sports, shopping aiming at finding good deals, cooking) that are home-produced (Aguila et al., 2011; Hurst, 2008; Miniaci et al., 2010). This change in the composition of consumption after retirement is likely to be optimal, since retirees have more leisure time as compared to employed individuals.

This paper investigates whether health care utilization and, in particular, the number of doctor's visits change in a discontinuous way at the time of retirement and whether this "puzzling" jump in health care utilization is associated with the drop in the opportunity cost of time induced by retirement, as suggested by the theory. To do so, we use all four waves of the Survey of Health, Ageing and Retirement in Europe (SHARE) which provides information on labor market status, health care utilization (in particular GP's visits and visits to specialists) and a wide range of socio-economic and demographic characteristics (Bolin et al., 2009; Dunlop et al., 2000).

Providing causal evidence of the impact of retirement on health care utilization is not straightforward. First, retirement decisions are likely to depend on individual (observable and unobservable) characteristics and second, health shocks may also affect the decision of individuals to retire early, further confounding the identification of the effects of retirement on medical care use. We tackle these issues by estimating a fixed-effect IV model. Our identification strategy exploits legal retirement ages (i.e. early and ordinary retirement ages) across European countries to instrument the probability of retirement (Bonsang et al., 2012; Coe and Zamarro, 2011, 2015).

Our results suggest that health care utilization increases when individuals retire. This is true both for the number of doctor's visits and for the intensity of medical care use (defined as the probability of going more than 4 times a year to the doctor's). This increase turns out to be driven by visits to general practitioners', while specialists' visits are not affected. We also find that the impact of retirement on health care utilization is significantly stronger for workers retiring from jobs characterized by long hours worked - more than 48 hours a week and/or being in the 5th quintile of the distribution of hours worked. This suggests that at least part of the increase in medical care use following retirement is due to the decrease in the opportunity cost of time.

Despite the growing relevance of medical care use for both the understanding of individuals' behavior, and for the design of suitable fiscal policies aimed at the 50+ population, only a limited number of studies have addressed the issue of health care utilization after retirement (Boaz and Muller, 1989; Jiménez-Martín et al., 2004). Their results have been quite mixed so far. Using respectively US and German data, Gorry et al. (2015) and Eibich (2014) hardly find any significant impact of retirement on health care utilization. Regarding Europe, Celidoni and Rebba (2015) do not find any effect of retirement on doctor's visits. In contrast, Coe and Zamarro (2015) show that the number of doctor's visits decreases when individuals transit from employment to retirement, unemployment or inactivity. In this paper, we focus on a sample of

EU countries and on retirement rather than other forms of exits from employment. We improve on the existing literature by providing causal evidence that medical care use increases when individuals retire, and that at least part of this increase is due the sharp reduction in the opportunity cost of time taking place at retirement.

The paper is structured as follows. Section 2 presents our empirical strategy. The data are described in Section 3. Section 4 presents our results and conclusive remarks are provided in section 5.

2 Empirical strategy

In order to assess the impact of retirement on health care utilization, we specify the following model:

$$V_{it} = \gamma R_{it} + X'_{it}\beta + \alpha_i + T_t + u_{it} \qquad (1)$$

where V_{it} is the number of doctor's visits of individual i in the 12 months preceding time t, R_{it} is a dummy variable equal to 1 if individual i is retired at time t, X'_{it} is a vector of demographics, household and job characteristics, α_i is an individual fixed-effect (including country fixed-effects) and T_t are wave dummies.

A key problem in estimating the impact of retirement on the pattern of health-care utilization comes from the fact that retirement is also correlated with individuals' health stock. The potential changes in health status upon retirement have been much studied in the literature but the evidence remains somewhat mixed. As regards physical health, a few studies find a negative effect of retirement (Behncke, 2012; Dave et al., 2008; Godard, 2016), while others find evidence of a substantial improvement following retirement, as leisure allows more time for health-preserving activities - i.e. sport, healthier lifestyles - (Celidoni and Rebba, 2015; Coe and Zamarro, 2011; Insler, 2014). As regards mental health and cognitive functions, they are generally found to be negatively affected by retirement (Bonsang et al., 2012; Mazzonna and Peracchi, 2012), although a few studies do not find any significant impact (Eibich, 2014; Lindeboom and Kerkhofs, 2009; Thompson and Streib, 1958). Overall, it is not clear whether one should expect an increase or a decrease in health status at the time of retirement (Bassanini and Caroli, 2015). In order to take into account the potential impact of retirement on health, we introduce a vector of health controls, H'_{it}, in our regression:

$$V_{it} = \gamma R_{it} + X'_{it}\beta_1 + H'_{it}\beta_2 + \alpha_i + T_t + u_{it} \qquad (2)$$

The number of visits an individual pays to the doctor is also likely to depend on several unobservable characteristics (e.g. individual time preference, latent health status etc.), that may also impact her retirement decision R_{it}. As long as individual unobservable heterogeneity is time-invariant, the inclusion of individual fixed effects in the regression, α_i, delivers consistent estimates of γ. This condition, however, is unlikely to hold in general. Time-varying omitted characteristics, such as own or partner's negative health shocks, are indeed likely to affect both the number of doctor's visits and retirement decisions. They may even generate some reverse causality if the doctor recommends that, given her health conditions, the individual should stop working.

To tackle these problems, we also estimate a FE-IV (*Fixed-Effects Instrumental Variable*) model. Our identification strategy exploits the fact that, as individuals reach legal retirement age, the financial incentive they have to retire strongly increases. This allows us to estimate the causal impact of retirement on health care utilization.

In most European countries, retirement occurs either at the Ordinary Retirement Age (ORA) - that is the age at which workers are eligible for full old-age pensions - or at an Early Retirement Age (ERA) - which represents the earliest age at which retirement benefits can be claimed conditional on a given number of years of social security contributions. This is evidenced in Appendix Figures A1 and A2 that display the shares of retirees and newly retired workers as well as the 'Early' and/or 'Ordinary' retirement-age thresholds (computed as averages over the period), by age, for each gender and country. The age-retirement profiles appear quite different across gender and countries. In many countries the early-retirement-age threshold is lower for females, as compared to males, which means that on average - conditional on social security contributions - females are eligible for retirement at an earlier age. In general, the trend in the share of newly retired workers reveals an increase in the probability of retirement as individuals approach the Early-Retirement-age threshold (particularly in Austria, Belgium, France and Italy), while trends tend to diverge from that point until the Ordinary Retirement Age: the share of newly retired workers further increases in some countries (i.e. Germany, The Netherlands, Spain, Switzerland, Denmark) while it declines in others (i.e. Belgium, France and Austria). In other words, depending on the interaction between the retirement rules in each country and the structure of financial incentives, the 'Early' and 'Ordinary' retirement-age thresholds account for most of the changes in individuals' probability of retirement.

In practice, we take advantage of the variability of both ERA and ORA across countries and use 2 instrumental variables, Y_{ict} and Z_{ict}, respectively defined as dummy variables equal to 1 if the individual is above the gender-specific Early - resp. Ordinary - retirement-age threshold in country c in year t.[2]

$$Y_{ict} = 1 \qquad if \qquad age_{ict} > ERA_{c,t}$$

$$Z_{ict} = 1 \qquad if \qquad age_{ict} > ORA_{c,t}$$

Note that identification here does not rely on the change in early or ordinary retirement ages across cohorts, but on the increase in the individual probability of retiring as individuals become eligible for pension benefits in their country of residence. In other words, the discontinuity in eligibility rules generates an

[2]See table A1 in the Appendix.

exogenous shock to retirement decisions which is what we use to instrument individual retirement status.

One problem with our data is the lack of information on the actual number of years of social security contributions. Hence, in countries where the ERA threshold is particularly low (Italy for example), and for those groups of workers with incomplete career spells (e.g. due to non-participation or long unemployment spells), the conditionality of a minimum number of years of contributions is likely to be binding and the ERA cutoff is unlikely to be a good predictor of the probability of being retired. In this case, the ORA cutoff is likely to do a better job in instrumenting retirement decision.

Note that, given the panel structure of our data, our identification relies on individuals who retire across waves. This approach identifies the local average treatment effect (LATE) - i.e. the effect of retirement on the subpopulation whom the instrument has induced to retire. So, the point estimate on the retirement variable, γ, in the FE-IV model, has to be interpreted as the local average short-term effect of retirement on health care utilization.

3 Data and descriptive statistics

3.1 Data and sample selection

We use data from the Survey of Health, Ageing and Retirement in Europe (SHARE). This is a multidisciplinary and cross-national survey collecting information on individuals' health, socio-economic status, household's income and wealth, as well as family networks. More than 85,000 individuals, aged 50 or older - and their partners - from 20 European countries have been interviewed. Five waves of data, from 2004 to 2013, are currently available (one of which is a retrospective survey). We use the 2004, 2006, 2011 and 2013 waves keeping only countries that were present in all waves.[3]

Our sample includes individuals that we observe for at least two consecutive

[3]Austria, Belgium, Germany, Sweden, Spain, Italy, France, Denmark, Switzerland and the Netherlands.

waves, aged 50 to 69 and who are either employed or retired in each wave.[4] We further restrict our sample excluding individuals permanently living in nursing homes and those who retired because of own ill health. After dropping all individuals with missing values on the variables of interest, our final sample consists of 2,883 individuals (9,266 observations).[5]

3.2 Variables

Health care utilization

Our measure of health care utilization is drawn from a question asking the respondent how many times she has seen a doctor over the past 12 months.[6] A breakdown of the total number of doctor's visits between general practitioner's and specialist's visits is also available up to 2011.

We also construct a measure of "high intensity" in the use of physician services, as a binary indicator which takes value one when the number of doctor's visits is larger than the median value for retirees (i.e. more than 4 visits in the past 12 months).

Retirement

The definition of retirement we adopt in our analysis follows that commonly used in the literature (Mazzonna and Peracchi, 2012; Rohwedder and Willis, 2010), where an individual is considered to be retired when leaving the labor force permanently (i.e. retirement is an absorbing state). In practice, retirement is coded using respondents' self-reported employment status in each wave. Note that SHARE also provides information about the exact year and month the individual left the labor market, which we use in some robustness

[4]So, our sample contains individuals who were employed all years, retired all years or who retired between waves. Retirement is therefore an absorbing state.

[5]Among our 2,883 individuals, about 51% are observed in all 4 waves, 19% in 3 consecutive waves and 30% are present in only two consecutive waves.

[6]The exact wording is: "About how many times in total have you seen or talked to a medical doctor about your health (last 12 months)?". Dentist's visits and hospital stays are excluded, but emergency room or outpatient clinic visits are included. We consider as outliers, and therefore exclude, individuals reporting a number of visits higher than three times the standard deviation.

checks.[7]

Our definition of retirement departs from that of Coe and Zamarro (2015) who consider as retired all individuals who report to be either retired or homemakers or sick and disabled or separated from the labor force or unemployed. The argument they give in favor of such a comprehensive definition of retirement is that individuals often report to have retired when they have left their "career" job even if they are still working part-time or full-time. However, this does not seem to be a major problem in the SHARE data since self-assessed retirement appears to be strongly correlated with the retirement status retrieved from the year/month-of-retirement variable - with a coefficient of correlation as high as 0.96 - as well as consistent with ordinary and early retirement eligibility ages in the country of residence. Given that our identification strategy relies on the exogenous shock to retirement decisions generated by the discontinuity in pension eligibility rules at ERA and ORA, we adopt a restricted definition of retirement that only includes individuals who report to have retired from work. We exclude individuals who have left employment for unemployment or inactivity since, in our identification strategy, they are likely to be non-compliers.[8] This should make our instrument stronger if anything. However, to account for possible biases in self-reported retirement, we run a robustness test where we exclude from the sample individuals who report to be retired but did nevertheless some paid work in the previous two weeks.

Control variables

In our empirical analysis we include a large set of controls, capturing individual, household and job characteristics. Demographic controls include age (continuous), education (primary/lower-secondary, upper/post-secondary and

[7]Since the retirement date (the exact year and month) as reported by the respondent is likely to be affected by measurement error (due to 'recall bias'), in the main analysis we choose to rely on the self-reported employment status. We use the information on the date of retirement only after performing some cleaning and consistency checks, and we then retrieve a binary retirement variable that takes value one if the distance between the year/month of interview and the year/month of retirement is non-negative.

[8]As a consequence, as mentioned in Section 3.1 above, our sample only contains individuals who are either employed or retired since, for most of those who are either unemployed or inactive there is probably little change in the opportunity cost of time upon retirement.

tertiary), a binary indicator for living with a spouse or partner, household size and a dummy variable for having children. Job characteristics are captured by industry (1-digit NACE classification) and occupational dummies (1-digit ISCO-88 classification).

Among the other personal characteristics of the sampled individuals that we control for, we include household income and ability to make ends meet. Household income is recoded into deciles and includes annual income from employment, self-employment or pension, as well as regular payments received by any member of the household. We also use a subjective measure of household financial distress drawn from a question that asks individuals: "Thinking of your household's total monthly income, would you say that your household is able to make ends meet [with great difficulty ... easily]", and recode it into a dummy variable taking value 1 if they make ends meet with difficulty or great difficulty and 0 otherwise.

Finally, since health status has been found to be a relevant predictor of health care utilization in the literature, we condition on individuals' health in our analysis (Bolin et al., 2009; Redondo-Sendino et al., 2006; Solé-Auró et al., 2012).[9] We include in our regressions a set of controls for self-assessed health, diagnosed conditions and an indicator of mental health. Self-assessed health is captured with a binary indicator taking value 1 if the individual reports to be in poor or fair health, based on a 5-point scale variable (ranging from "poor" to "excellent").[10] In order to get a more objective measure of individuals' health status, we build a summary indicator defined as the sum of all medical conditions that have ever been diagnosed by a doctor,[11] and an index of mental

[9]While we account for time-invariant unobserved heterogeneity in health conditions across individuals, we cannot rule out that health controls might be endogenous in our model. Hence, to check the robustness of our results, we also estimate the FE-IV specification without health controls.

[10]In the first wave of SHARE (2004), half of the sample was randomly assigned to a 5-point scale for self-reported health status defined as: "excellent", "very good", "good", "fair" and "poor". The other half was assigned to a 5-point scale defined as: "very good", "good", "fair", "bad" and "very bad". Looking at the distributions across categories for the two subsamples, we include in the binary indicator of poor health status those individuals that have answered "fair" and "poor" on the one hand, and "very bad", "bad" and "fair" on the other hand.

[11]These conditions include: heart attack, high blood pressure or hypertension, high cholesterol, stroke, diabetes, chronic lung disease, arthritis, cancer, stomach or duodenal ulcer, Parkinson disease, cataracts and hip or femoral fractures.

health (EURO-D depression scale ranging from 0 to 12, where higher values mean more depressed).

3.3 Descriptive statistics

Table 1 reports summary statistics for respondents' demographic characteristics, health status and doctor's visits, for the whole sample and separately for employed and retired individuals (pooling the four waves over 2004-2013).

The upper panel of Table 1 shows that there are no relevant differences between employees and retirees as regards demographic characteristics: the size of the household, the probability of having at least one child and the share of individuals living with a spouse or partner appear to be homogeneous across both groups, while retirees are on average older (64 versus 58 years-old) and have a lower level of education. Conversely, the lower panel of the table presents significant differences in terms of health status and health care utilization. The share of individuals reporting to be in poor health is 20% among retirees, as compared to only 12% among employed individuals. Using a more objective measure of health status, we see that, on average, retirees present more than one diagnosed chronic condition (1.22), as compared to 0.80 only for the sample of employed individuals.

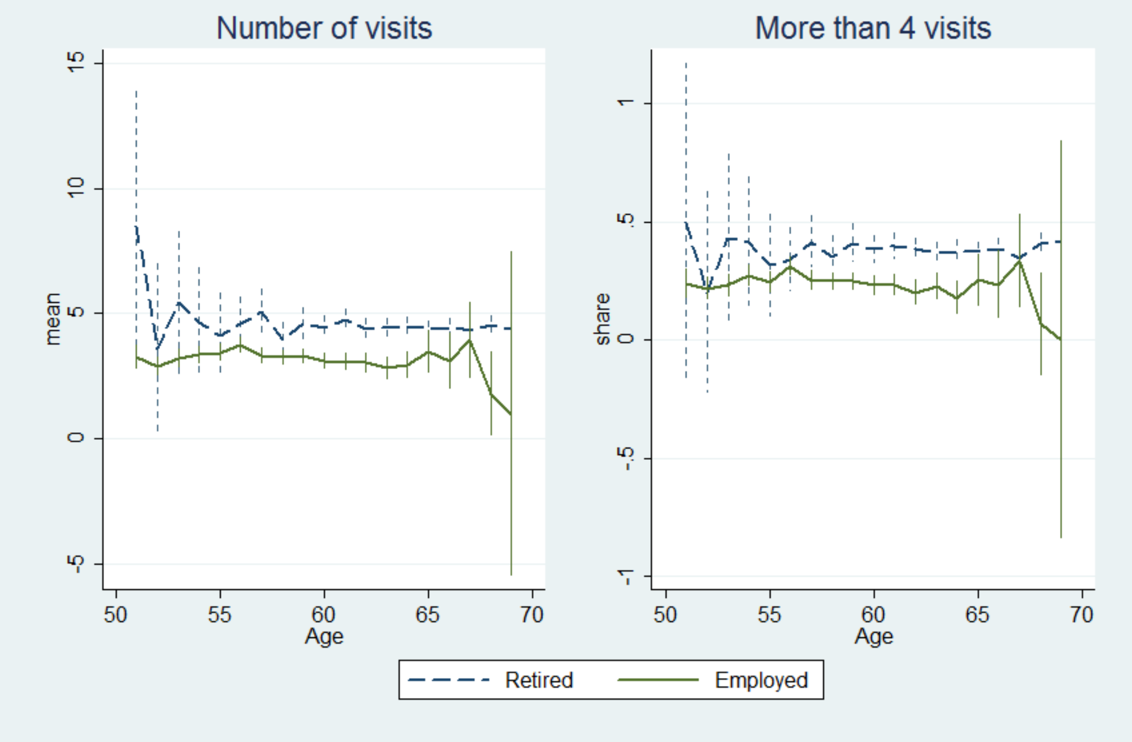
Retirees show a greater use of medical care, with an average of 4.5 visits over the past 12 months (as compared to 3.2 for employed individuals) and a median of 4 (as compared to 2 for employed individuals). 38% of them went to see a doctor at least 5 times in the past year (as compared to only 24% of employed individuals).[12]

The fact that retirees have a more intensive use of physician services than employed individuals is observed at all ages. As shown on the left-hand-side panel of Figure 1, the number of visits does not vary much with age, but it is significantly lower for employees as compared to retirees (particularly from age 59 onward). The sharp drop for employees aged 67 and over is likely to reflect

[12]Description and means of all other control variables can be found in Table A2 in the Appendix.

a selection effect: only individuals with better-than-average health are still in employment at older ages. A similar pattern is observed for the intensity of medical care use (see right hand side panel of Figure 1).

Figure 1: Age profile of doctor's visits



4 Results

4.1 Retirement and health care utilization

We first estimate the impact of retirement on health care utilization by pooled OLS and then with individual FE. The main results are reported in Tables 2 and 3 respectively. We consider different specifications of our baseline model. First we only include demographic controls along with wave and country fixed effects (columns 1 and 4). Second, we add to the previous specification a set of industry and occupational dummies as well as income decile dummies (columns 2 and 5). Finally, we also control for the health status of individuals (columns 3 and 6). As dependent variables we consider both the number of doctor's visits (columns 1 to 3) as well as our 'high intensity' health care utilization indicator (columns 4 to 6).

The number of doctor's visits and the probability of intensive use of medical care both increase with age and are found to be higher for females as compared to males (columns 1-2 and 4-5 in Table 2). In line with previous findings in the literature, health care utilization increases as the individual's health status deteriorates, being higher for individuals who report poor self-assessed health, a larger number of diagnosed conditions and a higher depression index (columns 3 and 6). Interestingly, the effect of age on the number of visits and on the probability of paying more than 4 visits a year becomes insignificant once controlling for health indicators, thus suggesting that the reason why older people visit the doctor more often is because their overall health is poorer. This is not the case for gender: its effect remains significant even after conditioning on health controls, suggesting that women are more concerned about their health than men are.

Coming to the the relationship between retirement and the number of doctor's visits, we find a positive and statistically significant coefficient suggesting that retirees go more often to the doctor's as compared to employed individuals (columns 1 and 4 of Table 2). The retirement gap in the number of visits is around 0.6 per year, which represents a 20% difference computed at the median of the sample. Retired individuals also show a higher intensity of medical care utilization, as their probability to see the doctor more than 4 times a year is about 6% higher than for individuals who are in employment. These differences are, if anything, larger when income, occupation and industry dummies are also taken into account (columns 2 and 5). When individual health is controlled for, the gap in the number of visits between retirees and employed individuals goes down to 14%, suggesting that part of the effect found in columns (1)-(2) and (4)-(5) was due to the fact that unhealthy individuals both tend to retire earlier and to go more often to the doctor's.

The estimated retirement gap is somewhat smaller when time-invariant unobserved characteristics are controlled for (Table 3): retirement induces a 9% increase in the number of doctor's visits and a 4% higher probability of 'high intensity' in health care utilization. As before, this may be due to the fact that poor health - along unobserved dimensions - is correlated both with the probability of retiring and with health care utilization. However, it should be stressed that if the retirement status is measured with error, the estimated effect is likely to suffer from attenuation bias, and particularly so when a FE estimator is used. Despite such downward bias, our results do suggest that doctor's visits tend to increase when individuals retire.

In the upper panel of Table 4, we report the estimates of the two-stage least square within estimator (FE-IV), while first-stage estimates are presented in the lower panel. In the first stage we estimate the probability of retiring between waves by means of a linear fixed-effect probability model.

The results reported in columns (1) and (4) correspond to a specification that uses minimum statutory ages for early retirement (i.e. ERA) as the only instrument. Alternatively, columns (2) and (5) use the age threshold at which individuals become eligible for ordinary pension benefits (i.e. ORA) as the sole instrument. Eventually, in columns (3) and (6) both ORA and ERA are used to instrument individuals' retirement decisions. In each specification, the instruments show a sizable and statistically significant effect on the probability of retiring. When both instruments are included, the effect of ERA on retirement appears to be larger than that of ORA. The F-stat for the joint significance of the instruments shows that both variables are significant predictors of the probability of retiring and the overidentification test supports the validity of the exclusion restrictions.

Estimates of the effect of retirement on the number of doctor's visits in the upper panel of Table 4 display statistically significant coefficients when ORA or both ERA and ORA are used as instruments. In contrast, results are only marginally significant when using ERA alone. Whatever the instrument we use, the point estimates in the second-stage are much larger than when estimated by OLS or FE. They turn out to be slightly larger when instrumenting retirement decisions with ORA than with both ORA and ERA and they are smallest when using ERA alone. In all cases, however, they suggest that moving from being employed to retirement causes an increase in the number of doctor's visits and in the probability of intensive use of physician services (i.e. more than 4 visits a year).

The reason why we find different results when using either ERA alone, ORA alone or both instruments together is likely to be due to differences in the groups of compliers. As previously discussed, our FE-IV estimates retrieve a Local Average Treatment Effect, that is the effect of retirement on those who effectively retire at the corresponding age threshold (Angrist and Pischke, 2009). When using only ORA to instrument retirement decisions, we capture the effect on those individuals who retire at a later age and would not have done so earlier. By contrast, when we use both ERA and ORA, we are more likely to get an average effect on all those individuals who have retired between the earliest and the latest age thresholds. In other words, we may expect that using both instruments allows to account for the effect of retirement on a larger and more heterogeneous pool of individuals. The latter includes those who have been reducing their working time before retirement, those who have been laid-off and offered an early exit and, last but not least, all the selected groups who, for various reasons, have been granted generous early retirement schemes. For most of those individuals, the effect of retiring on the number of doctor's visits is likely to be negligible as compared to full-time workers and to employees who used to work long hours, and experience a sharp change in the work-leisure trade-off after retirement.

To explore whether the reduction in the opportunity cost of time generated by retirement is actually a driving factor of the increase in health care utilization, we investigate whether the effect of retirement varies for those individuals retiring from a job characterized by long hours worked. In Table 5, we interact our retirement variable with a working-time dummy taking value 1 when the number of hours worked by the individual in his last job before retirement was above 48 hours a week[13] (columns 1 and 3) or alternatively, with a binary indi-

[13] 48 hours is the maximum average weekly working time set by the 2003 EU Working Time Directive.

cator for being in the upper quintile of the distribution of weekly hours worked (columns 2 and 4). By doing so, we again retrieve a LATE estimate of the heterogeneous impact of retirement on health care utilization for individuals who are likely to experience a sharp decrease in the opportunity cost of time upon retirement.

Whatever the indicator we use, our results suggest that the increase in the number of doctor's visits at the time of retirement is largest for individuals who used to work long hours before retiring: the point estimates on the interaction between retirement and each of the two dummy variables for long hours are indeed positive and significant at the 5% level. Similar results are found when considering the probability of going to the doctor's more than 4 times a year: the impact of retirement turns out to be stronger for individuals who used to work more than 48 hours a week and for those in the upper quintile of the distribution of hours worked. This suggests that the increase in health care utilization observed at the time of retirement is at least partly explained by the reduction in the opportunity cost of time.

One could be concerned however that the differential impact of retirement on health care utilization estimated in Table 5 could capture gender-specific behaviors to the extent that men tend to work longer hours than women. To control for this potentially confounding factor, we re-estimate the specification used in Table 5 also interacting retirement with gender. As shown in Table 6, the increase in the number of doctor's visits at the time of retirement - as well as, in the probability of high intensity' health care utilization - is mostly driven by male behavior while the estimated effect for females is essentially zero. Notice, however, that despite this differential effect by gender, the retirement gap in doctor's visits remains stronger for those individuals who used to work long hours before retiring: the point estimates on the interactions of retirement with the dummy variables for working more than 48 weekly hours and being in the upper quintile of the distribution of hours worked are still positive and statistically significant at least at the 10% level.

In Table 7, we estimate our model separately for the number of visits paid to a general practitioner and to a specialist. Our results suggest that the strong increase in the number of doctor's visits reported in Tables 5 and 6 for those individuals retiring from long-hour jobs - either more than 48 hours or in the upper quintile of the distribution of hours worked - is essentially driven by visits to the GP's (columns 2 and 3), whereas the impact is not statistically significant for specialists (columns 5 and 6). As can be seen when considering columns (1)-(3) and (4)-(6), poor self-rated health and a higher number of diagnosed conditions raise the number of visits to both types of doctors. In contrast, depression has a positive effect, but only on visits to the general practitioner's (GP) while it does not seem to affect specialists' visits.

4.2 Robustness checks

In order to test the robustness of our main findings, in this section we perform a number of sensitivity checks, experimenting with different specifications and samples, altering the definition of the retirement variables as well as estimating a dynamic specification.

One could be concerned that health controls could be endogenous if e.g. individuals who invest in preventive care both tend to go more to the doctor's and end up being in better health than others. This would generate a bias when estimating our FE-IV model to the extent that we cannot instrument the health status because of the lack of proper instrumental variables. To tackle this issue, we estimate a FE-IV specification without health controls - i.e. with a second stage corresponding to equation (1). As shown in Table A3 in the Appendix, the results are virtually unchanged. The same holds for all our key results: when re-estimating the models corresponding to Tables 5, 6 and 7 without health controls, the point estimates and standard errors are almost identical.[14] This suggests that endogeneity of health variables is not a major

[14]Results available upon request

issue in our main results.

In the analyses conducted so far, retirement has been defined on the basis of the employment status reported by individuals. One problem with this definition is that individuals may report to be retired even though they are still working. To account for this, we replicate our analysis using two alternative definitions of retirement. First, exploiting the available information on the declared year and month of retirement, we construct an indicator of retirement status where individuals are considered to be retired if the difference between the year/month of interview and the year/month of retirement is non-negative. Second, relying on self-reported retirement status, we exclude from the group of "retired" all individuals who performed any paid work during the two weeks preceding the interview.

Table A4 in the Appendix shows the results of these exercises. When re-running our FE-IV estimates using an indicator of retirement computed on the basis of the year/month in which individuals retired (Panel A), we find very similar results to those obtained in Tables 4 and 5. When instrumenting retirement status with both ORA and ERA, the number of doctor's visits and the intensity of medical care use turn out to increase significantly at the time of retirement (columns 1 and 4). This increase also appears to be significantly stronger for individuals retiring from long-hour jobs - whatever the measure we use for long hours (columns 2-3 and 5-6).
When we restrict the definition of retirement to individuals who report to be retired and have not performed any work in the two preceding weeks (Panel B), the effect of retirement on medical care use turns out to be even stronger. However, for retirees who used to work long hours before retiring, the increase in leisure time associated to retirement only affects the probability of high-intensity of health care utilization (i.e. more than 4 doctor's visits in the preceding 12 months). As regards the number of doctor's visits, it is positively impacted by retirement but not more so for retirees who used to work long hours.

Given the cross-country dimension of our data, one additional concern could be that our results are driven by a particularly large effect of retirement on health-care utilization in a specific country. However, when re-estimating our model excluding one country at a time, results are virtually unchanged - see Appendix Table A5 - Line 1.

Since a number of unobservable country-specific characteristics may change over time, we also augment our FE-IV specification including country-specific time trends (i.e. wave*country interaction terms) and find very similar results (see Appendix Table A5 - Line 2). Finally, results are also virtually unchanged if we introduce country-specific age instead of time trends (see Appendix Table A5 - Line 3).

We also investigate the dynamic properties of health care utilization. Following Coe and Lindeboom (2008) and Godard (2016), we estimate a simple dynamic IV version of our model by 2SLS, on the restricted sample of individuals who were employed at baseline interview.[15] We find that the lagged number of visits is a significant predictor of doctor's visits in the next period (i.e. in wave $t+1$) and that the effect of retirement is still positive and statistically significant (at the 10% level) on both the number of visits and the probability of intensive health care utilization (see Appendix Table A6).[16]

5 Conclusions

In this paper, we have established that health care utilization increases at the time of retirement: both the number of doctor's visits and the probability of intensive health care utilization increase when individuals move from employment to retirement. We argue that this increase can be explained by

[15]We specify the dynamic model as follows:

$$V_{i,t+1} = \delta V_{it} + \gamma R_{i,[t,t+1]} + X'_{it}\beta_1 + H'_{it}\beta_2 + T_t + \varepsilon_{it}$$

where$V_{i,t+1}$ is the number of doctor's visits in wave $t+1$, V_{it} is the number of doctor's visits in wave t and $R_{i,[t,t+1]}$ is a dummy variable taking value 1 if the individual retired between waves t and $t+1$.

[16]We instrument retirement with a binary indicator taking value 1 when individuals reach the ERA/ORA thresholds between waves t and $t+1$. All the other controls are the same as reported in Table 4.

the sharp drop in the opportunity cost of time that retirees experience upon retirement. The latter is likely to change their consumption pattern away from work-related goods and in favor of more time-intensive activities, such as medical care use. Consistent with these hypotheses, we find that the effect of retirement on health care utilization is significantly larger for individuals retiring from jobs characterized by long hours worked. We also find that the increase in health care utilization at the time of retirement is essentially due to men rather than women's behavior and that it is driven by visits to the GP's while specialists' visits are not significantly affected.

One interpretation of the above findings is that the optimal amount of health care utilization increases, at the time of retirement, due to the implicit decrease in the cost of time inputs. However we cannot exclude that alternative mechanisms be at work. For example, it may well be the case that, when still employed, individuals are rationed in terms of leisure time allocation and thus unable to reach their optimal choice of health care utilization. Moreover, consistent with our results, this effect is expected to be stronger for individuals working long hours before retiring.

While our identification strategy does not allow us to distinguish between these alternative hypotheses, it is clear that their implications differ substantially. In the first case, conditional on health status, the number of visits to the GP is likely to be optimal both before and after retirement. Conversely, in the second case, the number of times individuals visit their GP is suboptimal before retirement and jumps towards the optimal level only after they retire. The implications for health policy also differ substantially under the two alternative hypotheses. If the increase in the number of visits after retirement is the result of individuals' short-run response to a sudden change in the opportunity cost of leisure time there seems to be little scope for policy intervention. Conversely, if the increase in health care utilization upon retirement reflects a jump to the optimal level of health care for previously time-rationed individuals, concerns for public-health may arise. For example, it may imply that senior workers do not get enough health care towards the end of their career, which may further generate health problems particularly in a context in which the official retirement age has been raised. Of course, at this stage, whether the level of health care utilization of older workers is optimal or not remains an open issue.

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6 Tables

Table 1 Descriptive statistics

Variables	***Whole sample***	***Employed***	***Retired***
Demographics			
Age	60.8	57.8	64.3
Females	0.48	0.50	0.46
Males	0.52	0.50	0.54
Primary/Lower-secondary education	0.30	0.25	0.36
Secondary and upper-secondary education	0.36	0.36	0.36
Tertiary education	0.34	0.39	0.28
Living with a spouse or partner	0.77	0.77	0.78
Household size	2.2	2.3	2.0
Having at least 1 child	0.91	0.90	0.91
Health status			
Poor self-rated health	0.16	0.12	0.20
Diagnosed conditions (total)	0.99	0.80	1.22
Depression index (1-12)	1.68	1.67	1.69
Mean doctor's visits (median)	3.81 (3)	3.23(2)	4.47(4)
More than 4 visits	0.31	0.24	0.38
N.	9,266	4,953	4,313

Note: Figures reported are averages.

Table 2 Doctor's visits and retirement status - Pooled OLS

	Number of visits			***More than 4 visits***		
	(1)	(2)	(3)	(4)	(5)	(6)
retired	0.569***	0.601***	0.424***	0.0649***	0.0675***	0.0490***
	(0.124)	(0.124)	(0.108)	(0.0154)	(0.0154)	(0.0141)
age	0.0377***	0.0417***	0.00303	0.00398**	0.00459***	0.000531
	(0.0140)	(0.0140)	(0.0124)	(0.00169)	(0.00170)	(0.00156)
female	0.648***	0.556***	0.332***	0.0694***	0.0575***	0.0327***
	(0.102)	(0.114)	(0.102)	(0.0122)	(0.0134)	(0.0124)
poor health			1.642***			0.180***
			(0.130)			(0.0157)
diagnosed conditions (sum)			0.927***			0.0965***
			(0.0430)			(0.00518)
depression index			0.154***			0.0176***
			(0.0258)			(0.00309)
constant	0.550	0.438	1.436*	-0.0650	-0.0658	0.0374
	(0.909)	(0.932)	(0.812)	(0.110)	(0.114)	(0.105)
Demographics	✓	✓	✓	✓	✓	✓
Industry and Occupation		✓	✓		✓	✓
Income		✓	✓		✓	✓
Wave and Country dummies	✓	✓	✓	✓	✓	✓
R^2	0.0962	0.103	0.238	0.0721	0.0772	0.172
N	9,266	9,266	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01.

Table 3 Doctor's visits around retirement - Fixed-Effects

	Number of visits			***More than 4 visits***		
	(1)	(2)	(3)	(4)	(5)	(6)
retired	0.272**	0.280**	0.278**	0.0362**	0.0372**	0.0370**
	(0.124)	(0.125)	(0.121)	(0.0172)	(0.0172)	(0.0169)
age	-0.00330	0.00675	0.0319	0.00467	0.00671	0.00946
	(0.142)	(0.142)	(0.141)	(0.0207)	(0.0207)	(0.0205)
poor health			1.150***			0.124***
			(0.144)			(0.0186)
diagnosed conditions (sum)			0.544***			0.0627***
			(0.0574)			(0.00750)
depression index			0.127***			0.0144***
			(0.0296)			(0.00383)
Demographics	✓	✓	✓	✓	✓	✓
Industry and Occupation		✓	✓		✓	✓
Income		✓	✓		✓	✓
Wave and Country dummies	✓	✓	✓	✓	✓	✓
Individual fixed-effects	✓	✓	✓	✓	✓	✓
N	9,266	9,266	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.

Table 4 Doctor's visits around retirement - FE-IV

	Number of visits			***More than 4 visits***		
	(1)	(2)	(3)	(4)	(5)	(6)
retired	0.579*	1.101**	0.723***	0.0564	0.136**	0.0784**
	(0.303)	(0.428)	(0.274)	(0.0426)	(0.0600)	(0.0394)
age	0.0230	0.00760	0.0188	0.00888	0.00652	0.00823
	(0.141)	(0.142)	(0.141)	(0.0205)	(0.0206)	(0.0205)
poor health	1.151***	1.153***	1.152***	0.125***	0.125***	0.125***
	(0.143)	(0.144)	(0.143)	(0.0186)	(0.0186)	(0.0186)
diagnosed conditions (sum)	0.542***	0.537***	0.540***	0.0625***	0.0618***	0.0623***
	(0.0571)	(0.0573)	(0.0571)	(0.00748)	(0.00748)	(0.00747)
depression index	0.130***	0.133***	0.131***	0.0145***	0.0151***	0.0147***
	(0.0297)	(0.0299)	(0.0297)	(0.00384)	(0.00387)	(0.00384)
Demographics	✓	✓	✓	✓	✓	✓
Industry and Occupation	✓	✓	✓	✓	✓	✓
Income	✓	✓	✓	✓	✓	✓
Wave and Country dummies	✓	✓	✓	✓	✓	✓
Individual fixed-effects	✓	✓	✓	✓	✓	✓
First stage results						
Above ERA	0.368***		0.313***	0.368***		0.313***
	(0.0146)		(0.0153)	(0.0146)		(0.0153)
Above ORA		0.264***	0.161***		0.264***	0.161***
		(0.0148)	(0.0146)		(0.0148)	(0.0146)
R^2	0.4281	0.3812	0.4458	0.4281	0.3812	0.4458
F-stat of excluded instruments	634.33	317.18	411.05	634.33	317.18	411.05
N	9,266	9,266	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.

Hansen J statistic for overidentification is 1.369 ($p = 0.2419$) for column 3 and 1.625 ($p = 0.2024$) for column 6.

Table 5 Doctor's visits around retirement and pre-retirement hours worked - FE-IV[a]

	Number of visits		***More than 4 visits***	
	(1)	*(2)*	*(3)*	*(4)*
retired	0.607**	0.605**	0.0605	0.0641
	(0.285)	(0.289)	(0.0412)	(0.0417)
$\geq$ 48hours	-0.140		-0.0331	
	(0.171)		(0.0240)	
retired $\times \geq$ 48hours	1.009**		0.175***	
	(0.407)		(0.0619)	
5^{th}quintile[b]		-0.150		-0.0323
		(0.169)		(0.0239)
retired$\times 5^{th}$quintile		0.882**		0.128**
		(0.377)		(0.0588)
N	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.

[a] Retirement is instrumented with both ERA and ORA.

[b] Quintiles of weekly hours worked.

Table 6 Doctor's visits around retirement and pre-retirement hours worked - Gender differences - FE-IV[a]

	Number of visits			***More than 4 visits***		
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
retired	1.312***	1.182***	1.186***	0.134***	0.111**	0.116**
	(0.318)	(0.333)	(0.337)	(0.0446)	(0.0470)	(0.0477)
retired×female	-1.263***	-1.189***	-1.196***	-0.121***	-0.106**	-0.110**
	(0.297)	(0.299)	(0.300)	(0.0434)	(0.0436)	(0.0438)
$\geq$ 48hours		-0.109			-0.0303	
		(0.172)			(0.0240)	
retired× $\geq$ 48hours		0.748*			0.152**	
		(0.413)			(0.0625)	
5^{th}quintile			-0.110			-0.0287
			(0.170)			(0.0239)
retired×5^{th}quintile			0.638*			0.106*
			(0.382)			(0.0593)
N	9,266	9,266	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.

[a] Retirement is instrumented with both ERA and ORA.

Table 7 Number of General practitioner's and Specialist's visits - FE-IV[a]

	General Practitioner			*Specialist*		
	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
retired	0.333	0.218	0.235	0.0711	0.0761	0.0702
	(0.254)	(0.268)	(0.272)	(0.0557)	(0.0591)	(0.0597)
age	0.00665	0.00856	0.00550	-0.0331	-0.0332	-0.0330
	(0.118)	(0.118)	(0.117)	(0.0237)	(0.0238)	(0.0237)
$\geq$ 48hours		-0.135			0.0272	
		(0.151)			(0.0334)	
retired$\times\geq$ 48hours		1.069**			0.00669	
		(0.416)			(0.0830)	
5^{th}quintile			-0.212			0.0246
			(0.142)			(0.0324)
retired$\times 5^{th}$quintile			0.839**			0.0141
			(0.391)			(0.0811)
poor health	0.546***	0.541***	0.540***	0.104***	0.104***	0.104***
	(0.125)	(0.125)	(0.125)	(0.0224)	(0.0224)	(0.0224)
diagnosed conditions (sum)	0.278***	0.277***	0.278***	0.0528***	0.0528***	0.0528***
	(0.0492)	(0.0492)	(0.0492)	(0.00927)	(0.00928)	(0.00928)
depression index	0.0912***	0.0917***	0.0914***	0.00594	0.00598	0.00596
	(0.0264)	(0.0264)	(0.0264)	(0.00468)	(0.00469)	(0.00469)
N	7,486	7,486	7,486	7,486	7,486	7,486

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2 and the restricted sample from waves 1, 2 and 4.

[a] Retirement is instrumented with both ERA and ORA.

7 Appendix

Figure A1 Share of retirees and newly retired across countries - Females

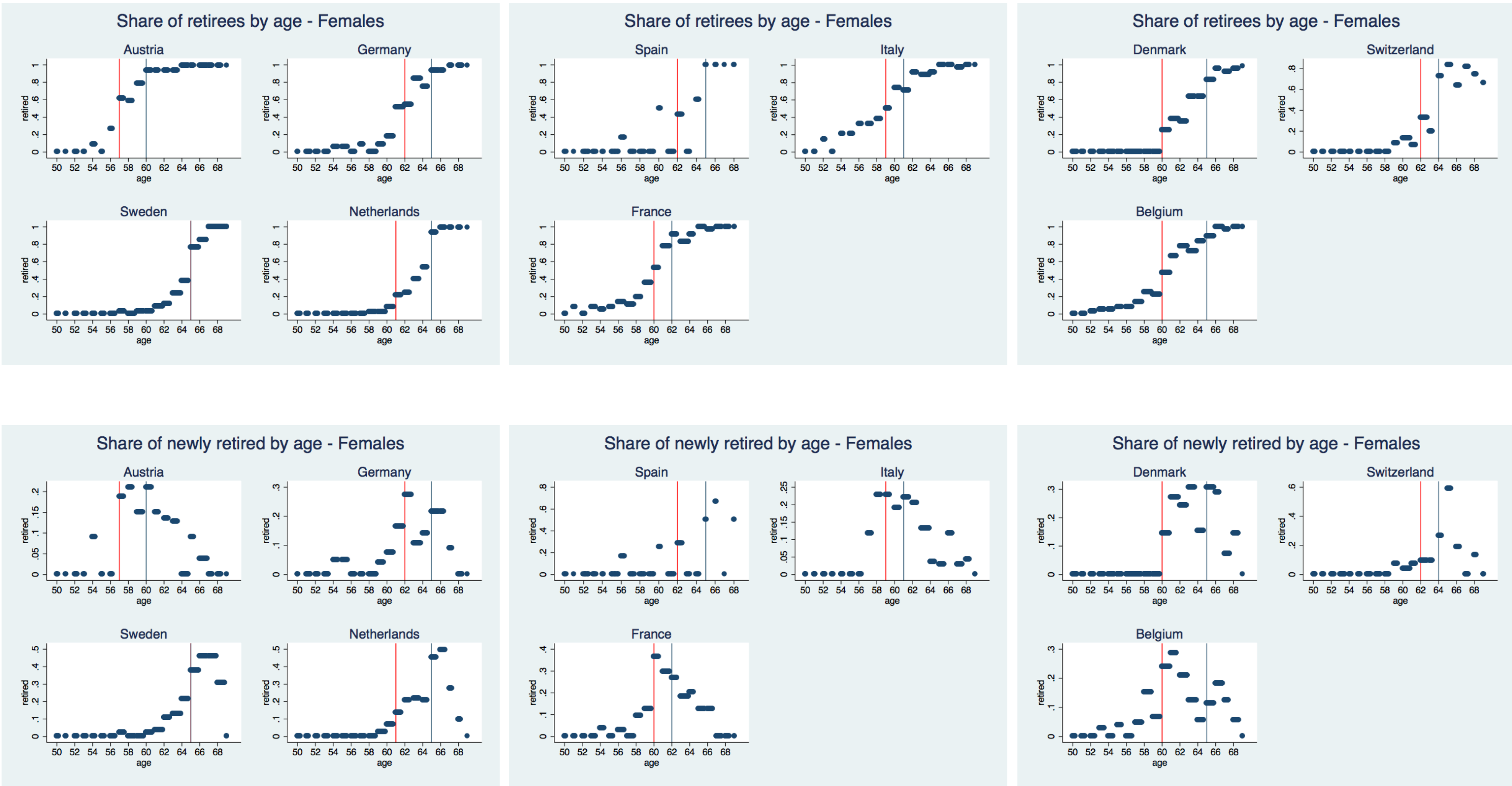

Figure A2 Share of retirees and newly retired across countries - Males

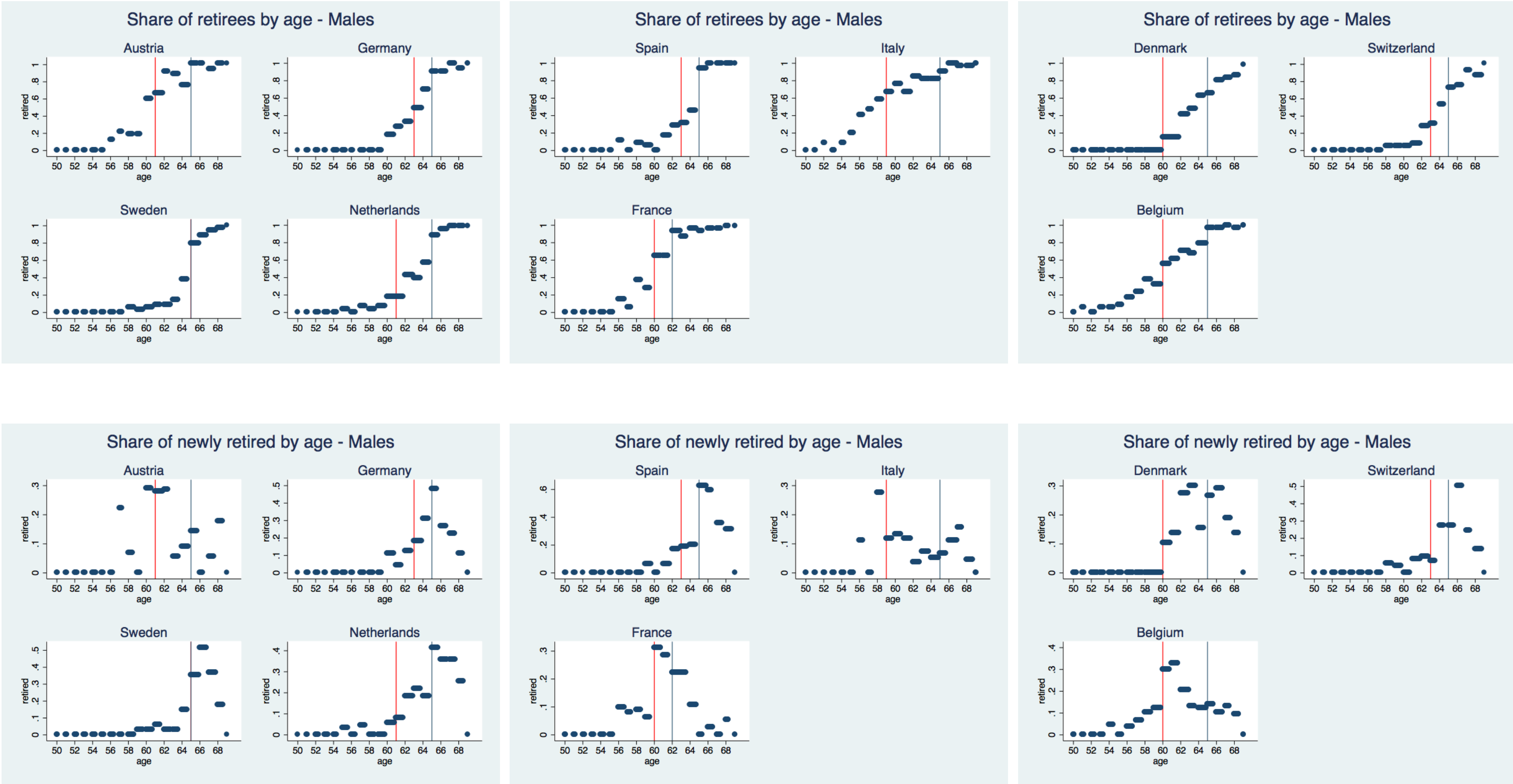

Table A1 Early and Ordinary Retirement ages across countries

Country	***ERA***[a,b]		***ORA***[a]		***Effective Retirement Age***	
	Men	Women	Men	Women	Men	Women
Austria	62	57	65	60	62.1	59.6
Germany	63	63	65	65	64.8	62.7
Sweden	65	65	65	65	66.0	65.9
Netherlands	65	65	65	65	63.8	64.1
Spain	63	63	65	65	65.1	62.9
Italy	61	61	65	62	62.4	61.6
France	60	60	65	65	61.1	61.9
Denmark	60	60	65	65	63.9	63.7
Switzerland	63	62	65	64	65.3	64.5
Belgium	60	60	65	65	61.8	61.0

[a] Early and Ordinary Retirement ages refer to individuals retiring in 2012 and are provided by OECD (2013), Gruber and Wise (2010); Hamblin (2013). ERA and ORA are defined for each country according to the year of retirement of individuals and the minimum and full statutory ages in force at that time; early retirement ages are obtained from minimum statutory ages and, for some countries (see country profiles below), they are adjusted according to contribution requirements and actuarial reductions.
[b] Country profiles for ERA:
Austria: early retirement in Austria was possible from the age of 55 for women and 60 for men until 2004. The 2004 reform introduced a gradual increase in ERA for both men and women born before 1943 and 1948, respectively. For individuals belonging to those cohorts and affected by the reform, we set ERA accordingly.
Germany: early retirement in Germany was possible from the age of 60 until 2003; after the 2004 reform, early retirement is possible from age 63 for men and 62 women (63 from 2006).
Sweden: the 1994 reform of the pension system in Sweden introduced a notional defined contribution scheme with no fixed-retirement age. Retirement is possible from age 61 in the public pension scheme, but with automatic actuarial reductions depending on the age of retirement, on individual contributions and on the demographic and macroeconomic environment. So, the minimum pension age by itself is not binding enough to generate a discontinuity in the probability of retirement. Moreover, the income-tested guaranteed pension cannot be claimed before age 65. Finally, in our sample less than 1% of retirees actually retire at 61 or earlier (with a median above 65). For these reasons we consider ERA as not binding in Sweden.
The Netherlands: early retirement in the Netherlands was possible from the age of 60 until 1995, 61 until 2004, and in 2005 the Dutch early-retirement scheme was abolished.
Spain: early retirement in Spain was possible from age 60 until 2001; in 2002 a reform shifted the minimum age to 61 and significantly increased reductions for workers voluntarily leaving the labor market before age 65 (so that those who would stop working at 61 would have incurred a reduction of 24-32%). Taking into account minimum age, years of contribution and related reductions, we set ERA at 63 for individuals retiring from 2002 onward.
Italy: under the system in force before 2008, workers could retire at age 57 if they had contributed to the system for 35 years, while the reform approved in 2008 introduced a quota system based on a combination of age and seniority that increases the minimum age gradually from 58 to 61 by 2013. Based on the combination of age and years of contributions we set ERA to 58 for men and 59 for women (who are more likely to have interrupted their careers) who retired until 2008, and then to 60 between 2009 and 2011 and 61 until 2012. The 2011 reform abolished the quota system, making retirement possible without penalty from age 62, with at least 42 years of contribution (in force from 2013).
France: early retirement in France is possible from the age of 60.
Denmark: early retirement in Denmark is being phased out but it is still available to workers born before January 1st, 1959 (our whole sample) and to workers aged 60 to 65 who continue working for 12 to 30 hours a week.
Switzerland: early retirement was possible two years before the standard retirement age, i.e. from 63 for men and 62 for women since 2000. Until 2000 for women and 1997 for men, retirement was possible only at the standard age (65 for men, 63 for women), and for men retiring between 1997 and 2000 early retirement eligibility age was 64.
Belgium: early retirement in Belgium was possible from the age of 60 for men and 55 for women until 1987; after 1987 the minimum statutory age for early retirement for both men and women has been set at 60.

Table A2 Description and means of control variables

Variable	***Description***	***Mean***		
		Whole Sample	*Employed*	*Retired*
Occupations 1-digit ISCO-88				
ISCO 1 - Managers		0.125	0.132	0.116
ISCO 2 and 3 - Professionals and Technicians		0.408	0.422	0.393
ISCO 4 - Clerks		0.139	0.139	0.139
ISCO 5 - Service workers		0.111	0.131	0.088
ISCO 6 and 7 - Crafts & Trade workers		0.096	0.076	0.12
ISCO 8 - Plant & Machine operators		0.049	0.034	0.065
ISCO 9 - Elementary occupations		0.072	0.066	0.079
Industries 1-digit NACE				
Agriculture and mining		0.024	0.021	0.028
Manufacturing		0.189	0.141	0.244
Construction		0.062	0.064	0.059
Wholesale, Hotels and Transports		0.148	0.134	0.165
Financial and Real Estate services		0.111	0.109	0.112
Public Administration		0.104	0.111	0.096
Education, Health and other Services		0.361	0.419	0.296
Weekly hours worked				
$\geq$ 48hours	dummy = 1 if respondent works at least 48 hours a week (when employed)	0.149	0.142	0.123
5^{th}*quintile*	dummy = 1 if respondent's hours worked are in the upper quintile of the weekly distribution	0.159	0.153	0.135
Income				
difficulty	dummy = 1 if Household makes ends meet with difficulty or great difficulty	0.187	0.154	0.226
Household's income deciles				
1		0.046	0.044	0.047
2		0.044	0.026	0.065
3		0.062	0.045	0.081
4		0.08	0.064	0.098
5		0.093	0.072	0.118
6		0.114	0.099	0.131
7		0.13	0.132	0.127
8		0.142	0.16	0.121
9		0.149	0.183	0.11
10		0.141	0.176	0.102
N		9,266	4,953	4,313

Table A3 Doctor's visits around retirement - No health controls - FE-IV

	Number of visits			***More than 4 visits***		
	(1)	(2)	(3)	(4)	(5)	(6)
retired	0.531*	1.071**	0.679**	0.0510	0.133**	0.0735*
	(0.310)	(0.443)	(0.283)	(0.0432)	(0.0613)	(0.0402)
age	-0.000793	-0.0170	-0.00525	0.00630	0.00383	0.00562
	(0.142)	(0.143)	(0.142)	(0.0207)	(0.0208)	(0.0207)
Demographics	✓	✓	✓	✓	✓	✓
Industry and Occupation	✓	✓	✓	✓	✓	✓
Income	✓	✓	✓	✓	✓	✓
Wave and Country dummies	✓	✓	✓	✓	✓	✓
Individual fixed-effects	✓	✓	✓	✓	✓	✓
First stage results						
Above ERA	0.368***		0.314***	0.368***		0.314***
	(0.0146)		(0.0153)	(0.0146)		(0.0153)
Above ORA		0.264***	0.160***		0.264***	0.160***
		(0.0148)	(0.0146)		(0.0148)	(0.0146)
R^2	0.4275	0.3802	0.4451	0.4275	0.3802	0.4451
F-stat of excluded instruments	634.89	315.51	410.92	634.89	315.51	410.92
N	9,266	9,266	9,266	9,266	9,266	9,266

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the set of controls as in columns 2 and 5 in Table 2.
Hansen J statistic for overidentification is 1.341 (p = 0.2468) for column 3 and 1.741 (p = 0.1871) for column 6.

Table A4 Alternative definitions of retirement FE-IV

	Number of visits			***More than 4 visits***		
	(1)	(2)	(3)	(4)	(5)	(6)
			Panel A - Year/month of retirement			
retired[a]	0.613**	0.522*	0.521*	0.0726*	0.0598	0.0625
	(0.259)	(0.270)	(0.272)	(0.0375)	(0.0394)	(0.0397)
$\geq$ 48hours		-0.0824			-0.0353	
		(0.173)			(0.0250)	
retired$\times \geq$ 48hours		0.866**			0.146**	
		(0.412)			(0.0629)	
5^{th}quintile			-0.0895			-0.0326
			(0.169)			(0.0246)
retired$\times 5^{th}$quintile			0.786**			0.117**
			(0.387)			(0.0598)
First stage results[b]						
Above ERA	0.330***	0.333***	0.334***	0.330***	0.333***	0.334***
	(0.0161)	(0.0169)	(0.0169)	(0.0161)	(0.0169)	(0.0169)
Above ORA	0.191***	0.170***	0.166***	0.191***	0.170***	0.166***
	(0.0155)	(0.0159)	(0.0159)	(0.0155)	(0.0159)	(0.0159)
F-stat of excluded instruments	461.18	243.75	248.87	461.18	243.75	248.87
N	8,730	8,730	8,730	8,730	8,730	8,730
			Panel B - Self-reported retirement and no work performed			
retired[c]	1.238***	1.124***	1.104**	0.159***	0.134**	0.133**
	(0.415)	(0.431)	(0.433)	(0.0592)	(0.0617)	(0.0620)
$\geq$ 48hours		-0.161			-0.0360	
		(0.179)			(0.0246)	
retired$\times \geq$ 48hours		1.106			0.259**	
		(0.793)			(0.121)	
5^{th}quintile			-0.169			-0.0350
			(0.178)			(0.0245)
retired$\times 5^{th}$quintile			1.139			0.246**
			(0.749)			(0.115)
First stage results[b]						
Above ERA	0.226***	0.229***	0.229***	0.226***	0.229***	0.229***
	(0.0154)	(0.0160)	(0.0160	(0.0154)	(0.0160)	(0.0160)
Above ORA	0.161***	0.145***	0.143***	0.161***	0.145***	0.143***
	(0.0155)	(0.0156)	(0.0155)	(0.0155)	(0.0156)	(0.0155)
F-stat of excluded instruments	196.02	102.31	103.46	196.02	102.31	103.46
N	7,821	7,821	7,821	7,821	7,821	7,821

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.

[a] Year/month of retirement is used to define retirement status. Individuals are considered to be retired if the year/month of the interview is larger or equal to the year/month of retirement - i.e. the individual has been retired for at least one month.

[b] Hansen-J statistic for overidentification rejects the null hypothesis in all estimates.

[c] We consider an individual as retired only if he reports to be retired and did not do any paid work during the 2 weeks preceding the interview.

Table A5 Alternative sample and specifications - coefficient on the *retired* variable - FE-IV[a]

	Number of visits	***More than 4 visits***	***Obs.***
1. Drop countries:range [min;max][b]	[0.571**-0.943***]	[0.0730*- 0.110**]	[7,990-9,333]
2. Country-specific time trends	0.766***	0.0898**	9,266
	(0.274)	(0.0392)	
3.Country-specific age trends	0.761***	0.0892**	9,266
	(0.274)	(0.0392)	

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01.

[a] Retirement is instrumented with both ERA and ORA.

[b] The range of estimates is obtained excluding one country at a time from our preferred specification.

Table A6 Dynamic IV

	Number of visits (t+1)	***More than 4 visits (t+1)***
retired$(t, t+1)$	0.710*	0.0927*
	(0.413)	(0.0560)
number of visits(t)	0.311***	
	(0.0212)	
>4 visits (t)		0.239***
		(0.0187)
age	0.00554	0.000520
	(0.0218)	(0.00293)
female	0.212**	0.0172
	(0.108)	(0.0144)
poor health(t)	0.292*	0.0347
	(0.168)	(0.0217)
diagnosed conditions(t)	0.419***	0.0454***
	(0.0553)	(0.00740)
mental health index(t)	0.0434	0.00723*
	(0.0282)	(0.00396)
R^2	0.199	0.128
N	4,747	4,747

Robust standard errors in parentheses, clustered at the individual level. Significance: * p<.1, ** p<.05, *** p<.01. All results are obtained using the full set of controls of columns 3 and 6 in Table 2.
First-stage statistics for both models confirm the relevance of the instruments. However, the overidentification test for the equation of the number of visits is rejecting the null with a p-value of 0.083.

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