**The Effect on Mental Health of Retiring During the Economic Crisis**



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# The effect on mental health of retiring during the economic crisis

Michele Belloni[1,2,3], Elena Meschi[1], Giacomo Pasini[1,2,*]

[1] Ca' Foscari University of Venice, Department of Economics, Venice, Italy
[2] Network for Studies on Pensions, Aging and Retirement, Tilburg, Netherlands
[3] CeRP – Collegio Carlo Alberto, Turin, Italy

**Abstract**
This paper investigates the causal impact of retirement on late-life mental health, a growing concern for public health, since major depressive disorders are the second leading cause of disability. We shed light on the role of economic conditions in shaping the effect of retirement on mental health by exploiting time and regional variation in the severity of the economic crisis across 10 European countries during 2004–2013. We use data from four waves of the Survey of Health, Ageing and Retirement in Europe and address the potential endogeneity of the retirement decision to mental health by applying a fixed effects instrumental variables approach. The results indicate that retirement improves the mental health of men but not that of women. This effect is stronger for blue collar men working in regions that have been severely hit by the economic crisis. These findings may be explained by the worsening of working conditions and the rise in job insecurity stemming from the economic downturn: Under these circumstances, exit from the labour force is perceived as a relief.

**Keywords**: Depression, stress, working conditions, retirement, IV fixed effect, recession

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[*] Corresponding author: Economics Department, Ca' Foscari University of Venice, Cannaregio 873, 30121 Venice, Italy; email giacomo.pasini@unive.it; tel. 041 234 9171; fax 041 234 9176.

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## 1. INTRODUCTION

The present study examines whether retiring during the recent economic crisis has led to benefits or losses in terms of mental health. Late-life depression is a serious and growing public health problem: The economic cost of depression is estimated at €118 billion in Europe (Sobocki *et al.*, 2006) and $83.1 billion in the United States (Greenberg *et al.*, 2003).

There is a growing literature on the effects of retiring on mental health but evidence collected before the recent crisis is not conclusive. Early works in social epidemiology find that retirement is associated with an improvement in mental health, while more recent contributions report either no association or a negative one (for a review of this literature, see Avendano and Berkman, 2015). Papers addressing the potential endogeneity of the retirement decision to mental health using panel data methods report mixed results. Dave *et al.* (2008) use US data from 2002–2005 and find a negative effect of retirement on mental health, while an opposite result is obtained by Jokela *et al.* (2010) and Mein *et al.* (2003) exploiting the Whitehall II study. Lindeboom *et al.* (2002) use a Dutch cohort study and find no effect of retiring on mental health (while other life events matter).

A potential reason for the inconclusiveness of these panel studies is that they do not account for the simultaneity issue that may arise from unobserved shocks in the individual environment affecting both mental health and the retirement decision. Recent studies exploit changes in pension eligibility to instrument retirement choices: Behncke (2012) uses three waves of the English Longitudinal Study of Ageing (ELSA) and finds that, while retirement increases the risk of being diagnosed with a chronic condition, it has no effect on mental health. De Grip *et al.* (2012) use Dutch data and find strong evidence that postponing the statutory retirement age leads to a worsening of mental health among those affected by the policy change who are still at work. Gorry *et al.* (2015), applying the same method to the US Health and Retirement Study, find that retirement improves mental health (as well as life satisfaction). Charles (2004) combines three sources of data for the US and exploits discontinuous retirement incentives in the Social Security System and changes in laws affecting mandatory retirement and Social Security benefits as a source of identification of the effect of retirement on subjective wellbeing. The author finds that retirement reduces the prevalence of depression and loneliness feelings. Coe and Zamarro (2011) rely on cross-country – rather than time – variation in the statutory retirement age, exploiting data from the first wave of the Survey on Health, Ageing and Retirement in Europe (SHARE); they conclude that retiring leads to lower depression scores. Eibich (2015) and Johnston and Lee (2009) apply Fixed Effects (FE) and a regression discontinuity design to German and English panel data, respectively, and find that retirement has a positive effect  a on mental health.

To the best of our knowledge, no previous study has analysed the effect on mental health of retiring during a period characterised by a severe economic crisis. This paper relies on the panel component of SHARE comprising waves from 2004 until 2013. During this period, the recession hit different European countries and regions at different moments and with different intensities: exploiting geographical and time heterogeneity in the severity of the economic crisis, we are able to shed light on the underlying mechanism of how and why mental health can improve at retirement. In regions in which economic conditions worsened substantially, employers retrenched reducing monetary and non-monetary rewards from working, and workers' perception of the risk of being laid off increased (Benitez-Silva *et al*, 2011; Eurofund, 2013). We argue that harsher labour market conditions lead to severe job-related stress, a key risk factor for old age depression (Lupien *et al.*, 2009). Retiring during the crisis can therefore be perceived as a relief and reducing depression scores.

Our identification strategy is based on an Instrumental Variables Fixed Effects (IV-FE) model whereby we control for unobserved time-invariant individual heterogeneity by including individual fixed effects, and address the possible issue of retirement endogeneity by exploiting the exogenous variation in retirement behaviour induced by country-specific early and standard retirement pension rules. Our results show that retirement improves mental health of men working in regions severely hit by the crisis and highlight that this effect is stronger for (ex-) workers in blue collar occupations, more intensively affected by the worsening of working conditions after the crisis. This evidence supports the stress-related explanation of the relation between retirement and mental health.

The remainder of the paper is organised as follows. Section 2 describes the data and the estimation method. Section 3 presents the results of the analysis. Section 4 reports a number of robustness checks, while Section 5 discusses the policy implications of our findings and draws some conclusions.

## 2. DATA AND METHODS

This study exploits data from the Survey of Health, Ageing and Retirement in Europe (SHARE). SHARE is a cross-national panel survey designed to provide comparable information on the health, employment, and social conditions of a representative sample of the non-institutionalised European population aged 50+. We use four waves of the survey: wave 1 (interview years 2004–2005), wave 2 (2006–2007), wave 4 (2011–2012) and wave 5 (2013).[1] The most recent wave is particularly useful for our scope, since it provides fresh evidence on the economic crisis. We select respondents from 10 countries included in all four abovementioned waves (i.e. Austria, Belgium, Denmark, France, Germany, Italy, the Netherlands, Spain, Sweden, and Switzerland).

Two key variables included in SHARE are exploited in our study: the EURO-D depression score (Prince *et al.*, 1999) and the self-reported job situation status. The EURO-D depression

[1] We exclude the third wave (a retrospective life history interview) because it is not directly comparable to the other waves.

score is a standardised scale of depressive symptoms designed to enhance cross-national comparability. The EURO-D consists of 12 items: feel depressed, pessimism, death wish, guilt, sleep, interest, irritability, appetite, fatigue, concentration, enjoyment, and tearfulness.

As in many other studies (e.g. Bonsang *et al.* 2012), the self-reported job situation status is exploited to define the variable *retired*. The latter is a dummy variable that takes the value zero if the individual reports being in the labour force at the time of the interview and one if the individual reports being retired.[2] The *retired* variable is adjusted using the information on the self-reported year of retirement: If the retirement status is missing but an individual reports the retirement year in any wave, the retirement status is then filled in accordingly (about 1.4 percent of observations are thus adjusted). Inconsistencies in the self-reported retirement status between waves are resolved by assuming that retirement is an absorbing state as e.g. Jimenez-Martin *et al.*, (1999): Once an individual reports being retired in one wave, that individual is considered retired in all subsequent waves. It turns out that only a very small fraction of observations (less than 0.3 percent) are adjusted this way.

We select individuals aged 55 to 70 (as in Eibich, 2015) who were working the first time they were observed in the panel (as in e.g. Behncke, 2012). Table 1 reports the pattern of individual participation in the panel: About 15 percent of the individuals were observed for four waves and around 25 percent for three waves. Attrition in SHARE is non-negligible and any analysis regarding mental health could suffer from sample selection: if the probability of remaining in the survey were inversely related to the mental health of the respondent, then the retention rate of the panel component of the sample is affected by the variable that is the object of interest of our analysis. Following the strategy proposed by Lindeboom *et al.* (2002), we conduct two

[2] Respondents who were disabled or self-reported as being a homemaker or in the residual ‘other’ category (about 14 percent of valid observations) were not considered in the analysis. This point distinguishes our study from that of Bonsang et al. (2012), who consider these categories the same as retired (not working for pay).

informal tests for the severity of selective attrition. First, we run a regression of the depression score on an attrition dummy (i.e. a dummy that takes the value one if the *i*th observation is not in the panel at time $t + 1$), its interaction with the retirement dummy, and the full set of controls including individual fixed effects: the attrition dummy and the interaction term are neither individually nor jointly statistically different from zero. Second, we regress the attrition dummy on the depression score plus controls and individual fixed effects and find no significant effect for the depression score, which means that individuals becoming more depressed are not more likely to leave the panel.[3]

Table 2 reports summary statistics of the variables included in the analysis. Women exhibit higher depression scores, are more often single or widowed and their household income distribution is shifted to the left. Additionally, about 42 percent (3,648) of the sampled individuals retired throughout the analysed period. Some preliminary evidence on the relation between depression and retirement status can be found in Figures 1 and 2. They show, for males and females, respectively, linear predictions from FE models where the EURO-D score is explained by a full set of age dummies interacted with the *retired* dummy. Although the standard errors are large, these figures indicate that being retired is associated with a lower level of depression (we limit this observation to the age range 57–65 years, which is when most retirement occurs in our data).

The first model we estimate is the following:

$$y_{ijt} = \beta_{RET} D_{ijt}^{RET} + \boldsymbol{X'}_{ijt} \boldsymbol{\beta_X} + \boldsymbol{W'}_{jt} \boldsymbol{\beta_W} + \vartheta_t + d_i + u_{ijt} \tag{1}$$

[3] The results of these tests are reported in the online appendix in Table A1.

The dependent variable $y_{ijt}$ is the EURO-D depression score for individual *i* living in region *j* in wave *t*; the main explanatory variable is the retirement status $D_{ijt}^{RET}$,a dummy variable taking the value one if the individual reports being retired at the time of the interview and zero if in the labour force.

A panel dataset allows us to control for unobservable time-invariant characteristics affecting both the individual retirement decision and mental health by including individual FE $d_i$ in all our specifications. The results obtained using a within estimator will, however, be unbiased only if $D_{ijt}^{RET}$ is orthogonal to the remaining time-varying residual term $u_{ijt}$ and its leads and lags. It is therefore important to control for potential time-varying observable confounders. In all specifications, $\boldsymbol{X}_{ijt}$ includes a second-order polynomial in age, that is, the effect of retirement is identified by changes in EURO-D at the specific retirement age, conditional on a long-term quadratic relation between mental health and age[4]. The term $\boldsymbol{X}_{ijt}$ also includes detailed marital status dummies (widow, divorced, not married until the year of the interview, married or cohabiting), the number of grandchildren, quintiles of household income,[5] and an aggregated physical health index constructed following Poterba *et al.* (2011).[6] Finally, $\boldsymbol{W}_{jt}$ includes GDP and unemployment rate of region *j* (NUTS 1 level) at time *t*: time-varying local labour market conditions may explain variation in depression scores at aggregate level.Finally, following Lindeboom *et al.*, (2002), we include in the model a full set of wave dummies $\vartheta_t$ to account for any possible time-variant shock to mental health, which is common to all individuals.

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[4] We test the sensitivity of our estimates to the particular age polynomial included in the regressions. We experimented with both a cubic and a quartic specification in age. Coefficient estimates of being retired are almost unaffected, while coefficients for higher than second order polynomials are never significant (see table A4 in Online appendix).

[5] Note that SHARE respondents were asked about their household's gross incomes in the first wave and about net incomes later. Therefore, we have to resort to a relative measure of income rather than an absolute one to use all the available waves (e.g. Kalwij *et al.* 2014).

[6] Details about the construction of the health index can be found in the Online Appendix of this paper.

Still, it is possible that unobservable transitory shocks in the idiosyncratic time-varying residual $u_{ijt}$ affect both the decision to retire and mental health. This type of endogeneity can be accounted for with an IV approach: We need a set of variables $\boldsymbol{Z}_{ijt}$ that affect mental health only through their effect on the retirement decision $D_{ijt}^{RET}$. We follow Coe and Zamarro (2011) and use cross-country variation in the rules determining eligibility to Social Security benefits as an instrument. More specifically, we use the statutory and early retirement ages from the Mutual Information System on Social Protection (MISSOC) database[7] as instruments for the retirement decision. The first (second) instrument for the retirement status takes the value one if the respondent has reached the standard (early) retirement age at the time of the interview and zero otherwise. The identifying assumption is that being age eligible for statutory and early retirement pension benefits does not affect mental health directly[8] but provides incentives to retire, as can be seen in Figures 3 and 4 (for males and females, respectively). These figures report, for each country and by gender, the retirement age distribution for our estimation sample, together with statutory old age and early retirement ages. Retirement age is computed exploiting information on the self-reported year of retirement. There is considerable variation in retirement age; however, most of the countries have clear spikes at legal retirement ages. Our identification strategy does not only rely on changes in retirement rules over time (see Figures 3 and 4) but, rather, on the discontinuity in the number of retired individuals at the legal retirement ages, conditional on a smooth function of age.

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[7] MISSOC collects information on social protection for the Member States of the European Union and other countries, including Switzerland (see http://www.missoc.org/MISSOC/INFORMATIONBASE/COMPARATIVETABLES/MISSOCDATABASE/comparativeTableSearch.jsp).

[8] It should be noted that the exclusion restrictions could potentially be violated because of justification bias and the use of subjective questions to evaluate mental health. In fact, individuals who are retired before the eligibility age of retirement may want to justify their non-working status by pretending that they are in poor health. Once they reach the age of eligibility for retirement, they may have less pressure to justify their non-working status. As a result, being eligible for retirement benefits might affect the dependent variable, conditional on retirement status. However, since EURO-D is a standardized, validated and widely used measure of depressive symptoms (based on a battery of 12 questions rather than on a single self-reported question), the bias due to subjective differences in self-assessment should be limited.

An open question is, nonetheless, what the underlying mechanism is that governs the relation between the retirement decision and mental health. The availability of a panel dataset covering periods before and after the peak of the Great Recession in 2008 and regions differently hit by the crisis helps shed light on this mechanism. A FE estimation allows us to control for any unobservable time-invariant individual characteristic. Therefore, the type of time-varying confounder inducing the unobservable transitory shock we deal with in the IV approach is a sudden change in the environment of the respondent, which affects both labour market participation decisions and mental health. The economic crisis and the corresponding changes in the working environment faced by individuals are clearly potential sources of this type of shock. Worsened economic prospects for a region are likely to increase work-related stress through at least two channels: First, working in a region where economic prospects are bad increases the risk of being laid off (Benitez-Silva *et al.*, 2011; Eurofund, 2013). Second, most employers retrenched substantially as a consequence of the crisis, reducing monetary and non-monetary rewards associated with working activities. An increased level of stress at work can induce individuals to anticipate retirement: leaving the labour force is likely to be perceived as a relief. Lupien *et al.* (2009), in a review of the literature on the physiological link between stress and brain disorders, report that even a single period of severe stress can lead to memory dysfunction, depression, and post-traumatic stress disorder. The data allow us to partially test whether the reduced stress channel is responsible for the mental health improvement at retirement: we first construct a binary indicator $D_{jt}^{HIT}$ that takes value one if a given European region *j* in a given year *t* was severely hit by the economic crisis and zero otherwise, and then we estimate the following model:

$$y_{ijt} = \beta_{RET} D_{ijt}^{RET} + \gamma D_{ijt}^{RET} D_{jt}^{HIT} + \beta_{HIT} D_{jt}^{HIT} + \boldsymbol{X'}_{ijt} \boldsymbol{\beta_X} + \boldsymbol{W'}_{jt} \boldsymbol{\beta_W} + \vartheta_t + d_i + u_{ijt} \qquad (2)$$

where $\beta_{HIT}$ captures the effect on mental health of living in a region-year severely hit by the economic crisis, and $\gamma$ accounts for the differential effect of retiring in such economic circumstances.

The crisis indicator $D_{jt}^{HIT}$ is obtained starting from the regional unemployment rate time series (Eurostat, 2015). We apply the Hodrick–Prescott (1997) filter with a smoothing parameter of 100 to each of these time series and split the unemployment rate into trend and cycle components. Figure 5 shows the unemployment rate cyclical component from the Hodrick–Prescott filter in our sample period (2004–2013). Our binary indicator $D_{jt}^{HIT}$ identifies severe shocks in economic conditions by taking on the value one if the cycle component is negative and greater than one standard deviation. As stated, we assume that the economic crisis affects the relationship between retirement and depression through its effect on working conditions close to retirement. Therefore, we attribute the value of our indicator $D_{jt}^{HIT}$ to each individual according to the individual's self-reported retirement year. In Section 4, we experiment with different indicators of the severity of the crisis and find similar results.

As for equation (1), also equation (2) is estimated both with FE and with IV-FE estimators. In the latter case, we use four instruments: the already described statutory and early retirement age eligibility dummies, and their interaction with the crisis indicator $D_{jt}^{HIT}$. Both equations are estimated separately by gender to account for potential differences due to different reservation wages and labour supply elasticities between men and women.

## 3. RESULTS

Table 3 reports estimation results for equation (1), by gender. Column (1) shows FE estimates for males: Moving into retirement leads to a significant reduction of 0.203 in depression scores, a 13 percent reduction compared to its mean value (equal to 1.55, see Table 2). Column (2) reports IV-FE estimates for males. Instruments are informative and valid: The F-test of joint significance of the excluded instruments (i.e. having reached statutory and early retirement age) in the first stage is highly significant and the Hansen J-test shows that the over-identification restriction is valid. The coefficient of the *retired* dummy is no longer statistically significant. Despite the Durbin–Wu–Hausman test not providing evidence of endogeneity of the retirement dummy, we do not think the IV-FE estimates can simply be dismissed as inefficient. First, there are good economic reasons to think that reverse causality could be an issue. Second, the lack of significance may hide heterogeneous effects.

Magnitude and significance of control variables are comparable between columns (1) and (2): being widowed increases depression scores, while the health index is negative and highly significant. We were concerned about the possible endogeneity of the health index. We therefore estimate a model without this variable among the explanatory variables and find that the coefficients of the variables of interest are essentially unchanged. Results for females, reported respectively in columns (3) and (4), are in line with those for males. An exception is that, for them, being in the highest household income quintile is associated with higher depression scores.

The IV-FE estimates account for unobservable transitory shocks affecting both retirement decisions and mental health. Between 2004 and 2013 a likely transitory unobservable shock to European men and women close to retirement has been a worsening of working conditions due to the economic crisis. As we explained in the previous section, the increased stress associated with worse working conditions can induce those who retire to feel relieved and experience a reduction in depression score; in other words, job-related stress can be the driving force behind the relation between retirement and depression. If this is true, IV-FE estimates in Table 3 do not capture an important source of heterogeneity: the effect of retirement should be stronger in regions that experienced a more pronounced slowdown during recent years.

Table 4 reports estimation results for equation (2) for males (top panel), and females (bottom panel). In comparison with equation (1), we additionally include a dummy variable that takes value one if an individual experienced a particularly bad economic slowdown in a given year and zero otherwise (*Hit by the crisis*) and its interaction with the retirement dummy (*Retired x hit by the crisis*).[9] Column (1) reports the IV-FE results for the full sample: living in a region and year particularly hit by the crisis significantly increases depression scores for males. As we anticipated, the non-significant effect of retiring we found in Table 3 hides important heterogeneity: the coefficient of the interaction term is negative and significant. This means that those who retire in a region and year not considerably hit by the crisis do not experience any significant reduction on mental health, while those who retire in a year and region severely hit by the crisis do experience such a reduction. The magnitude of the differential effect for those retiring when severely hit is a reduction in depression scores of 0.43 EURO-D points, which accounts for an approximately 27 percent reduction with respect to the mean value. For females, the crisis play no role: results of Table 3 are confirmed and retirement does not significantly affect their mental health.

Distinguishing between blue collar and white collar workers (columns (2) and (3) of Table 4), we find that both the negative effect of experiencing an economic crisis and the positive effect of retirement on males' mental health in the region-year hit by the economic crisis outlined in column (1) of Table 4 are driven by blue collar workers. For female blue collar workers, experiencing a severe crisis increases depression score, but retiring has no relieving effect. Not significant results are found for female white collar workers, nor for females overall (column 1; note that more than 80 percent of sampled female work(ed) in white collar occupations).

[9] The dummy $D_{jt}^{HIT}$ varies by year of retirement and region. We do not have enough heterogeneity in the data to estimate specifications of equation (2) that include also regional unemployment rate and GDP per capita.

The diagnostics reported in Table 4 again point to the validity of the chosen instruments for both men and women, whereas the *retired* regressor no longer passes the exogeneity test for women. [10]

These results are in line with the evidence showing that job insecurity increased considerably during the crisis (Eurofund, 2013) and that the extent of such increase varied by occupations. In fact, looking at aggregate figures from the 2005 and 2010 waves of the European Working Conditions Surveys (EWCS),[11] we find that perceived chance of losing one's job within the next six months, is very heterogeneous across occupations, increasing by 5 percent for blue collar workers and by 2 percent for white collar workers. Moreover, Eurofund (2013) reports that the recent economic downturn affected other dimensions of working life, leading to less choice for workers, wage freezes and wage cuts, greater work intensity, deterioration of work–life balance, and greater risk of harassment/bullying. All these factors affect workers' level of stress, and are likely to be particularly important for less-qualified workers, consistently with the results of our estimates.

## 4. ROBUSTNESS CHECKS

A critical element of our empirical strategy is the type of transition to retirement analysed and the corresponding definition of the dummy *retired*. As explained in Section 2, we defined *retired* as a dummy taking the value zero if the individual reports being in the labour force at the time of the interview and one if reports being retired. We have therefore considered two alternative dummies, based on individuals' status when in the labour force. In particular, we first excluded those who reported being unemployed and defined the variable *retired* equal to

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[10] Since, in this case, we have multiple endogenous variables, the F-statistic is computed according to Sanderson and Windmeijer (2013).

[11] The EWCS are cross-country surveys that assess and quantify the working conditions of both employees and the self-employed across Europe on a harmonized basis. More details about the EWCS and our elaborations can be found in the online appendix (see Figure A1).

one if the individual had retired and zero if the individual reported *working*. Second, we defined a corresponding dummy *retired* that considers only those who were unemployed among those in the labour force. Table 5 reports the results of this analysis. Our main results hold in the first case (*retirement from work*) but no longer do in the second case (*retirement from unemployment*). This evidence reinforces our hypothesis that the channel through which retirement reduces depression scores is by alleviating job-related stress.

A second key element of our analysis is the definition of the economic crisis, based on the *binary* indicator $D_{jt}^{HIT}$ obtained by detrending the time series of the regional unemployment rate. In Table 6 we report the results of a sensitivity analysis with respect to the crisis definition (only for males, IV-FE estimates). We first define a *continuous* indicator of the severity of the crisis based on the same detrended macroeconomic variable (see column 1 for the whole sample and column 2 for blue-collar workers). Figure 6 shows the deriving marginal effect of retirement on the depression score for male blue-collars as a function of the cyclical component of the unemployment rate: Retirement significantly reduces depression scores when economic conditions are very bad, namely, above the 80th percentile of the distribution of the indicator. Different values of the Hodrick–Prescott filter yield very similar results. We then define the crisis by applying the same de-trending method to another macroeconomic indicator: the regional real GDP per capita.[12] We define the crisis by means of a binary indicator, equal to 1 if the cycle component is negative. Results are reported in columns 3 and 4 for the whole sample and for blue-collars respectively. The coefficients of interest are essentially unchanged, thus confirming the robustness of our findings to alternative definitions of the crisis. Our preferred macroeconomic indicator of the crisis remains, however, the unemployment rate: the unemployment rate peaks only after the worst point of the business cycle (Stock and Watson,

[12] Eurostat provides data on nominal GDP per capita by NUTS 1 region. We compute the real GDP per capita by dividing the nominal regional data by the national price index (GDP deflator with the base year 2010).

1999) because individuals and firms do not adjust immediately to changes in GDP and the effects on workers' environment might occur with a lag. As a result, unemployment rate is likely to better reflect working conditions.

Finally, we tested whether the results are sensitive to the estimation method. The IV-FE method requires the instruments to be strictly exogenous, that is, independence of the contemporaneous error term as well as all its lags and leads. We re-estimated the model with an IV first differences (IV-FD) estimator that requires only weak exogeneity (independence of the contemporaneous error term and its leads but not its lags). The results are reported in columns (1) and (2) of Table 7 for males and females respectively and are very similar compared to the IV-FE estimates (although less precisely estimated due to sample reduction). Until now we always ran the IV-FE model under the assumption that some of the explanatory variables can be correlated with unobservable time-invariant individual characteristics. However, if we are willing to assume conditional mean independence between the time-invariant component of the error term and the included regressors, the IV-FE and the IV-FD estimators are less efficient than an IV Random Effects (IV-RE) estimator. Moreover, under the conditional mean independence assumption, the IV least squares estimator is also consistent and requires only contemporaneous exogeneity for the instruments rather than strict exogeneity, as the IV-RE estimator does. Columns (3) and (4) of Table 7 report IV-RE estimates and columns (5) and (6) IV least squares estimates. The interaction term of interest in the IV-RE estimates is in line with that of the IV-FE and IV-FD, while it loses statistical significance in the linear regression. In the results of Table 7, the coefficient of the retirement dummy becomes significant in both the IV-RE and linear IV regressions. This positive association disappears with the IV-FE or IV-FD, thus suggesting that there are unobserved time-invariant characteristics that are correlated with both the instruments and the dependent variables. The use of fixed effects model strengthens the assumption of independence between the instruments and the error term of the main equation.

## 5. DISCUSSION AND CONCLUSION

This paper studied the causal effect on mental health of retiring during the economic crisis. We addressed the potential endogeneity of the retirement decision by applying a Fixed Effects Instrumental Variables approach using country- and gender- specific statutory and early retirement ages as instruments for retirement behaviour. We found that retirement per se does not have a significant impact on depression scores, once endogeneity is taken into account. However, by exploiting time and geographical heterogeneity in the intensity of the economic crisis, we showed that retirement does improve mental health in periods and regions severely hit by the economic crisis. Moreover, our results indicated that this effect is entirely due to blue-collar (ex) workers and does not apply to white collars.

Our results help shedding light on the mechanisms behind the relationship between retirement and mental health. We suggest that retirement may affect mental health through its effect on stress. If retirement is perceived as a relief from job-related stress, our finding of a differential effect of retirement according to the type of occupation is entirely in line with the worsening of working conditions experienced in particular by blue collar workers as a consequence of the economic downturn.

The natural implication of our result is that policy makers willing to postpone the retirement age should account for the fact that this may increase health inequalities, since the implied costs in terms of worsening mental health would disproportionately hit low-skilled workers.

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## TABLES AND FIGURES

**Table 1.** Individual panel participation

| Wave 1 | Wave 2 | Wave 4 | Wave 5 | N individuals | Percent |
|---|---|---|---|---|---|
| X | X | X | X | 1281 | 14.88 |
| | X | X | X | 1315 | 15.27 |
| X | X | X | | 467 | 5.42 |
| X | X | | X | 176 | 2.04 |
| X | | X | X | 218 | 2.53 |
| | | X | X | 3628 | 42.13 |
| X | X | | | 821 | 9.53 |
| | X | X | | 316 | 3.67 |
| | X | | X | 208 | 2.42 |
| X | | | X | 100 | 1.16 |
| X | | X | | 81 | 0.94 |

**Table 2.** Summary statistics, by gender

| | Females | | | | | Males | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | *Obs* | *Mean* | *SD* | *Min* | *Max* | *Obs* | *Mean* | *SD* | *Min* | *Max* |
| EURO-D depression score | 10583 | 2.25 | 2.06 | 0 | 12 | 11377 | 1.55 | 1.72 | 0 | 12 |
| retired | 10583 | 0.27 | 0.44 | 0 | 1 | 11377 | 0.27 | 0.44 | 0 | 1 |
| Age | 10583 | 60.53 | 3.78 | 55 | 70 | 11377 | 60.73 | 3.82 | 55 | 70 |
| ***Marital status*** | | | | | | | | | | |
| Married or in a couple | 10583 | 0.72 | 0.45 | 0 | 1 | 11377 | 0.81 | 0.39 | 0 | 1 |
| Divorced | 10583 | 0.14 | 0.35 | 0 | 1 | 11377 | 0.10 | 0.30 | 0 | 1 |
| Widowed | 10583 | 0.07 | 0.26 | 0 | 1 | 11377 | 0.03 | 0.16 | 0 | 1 |
| Not married | 10583 | 0.07 | 0.25 | 0 | 1 | 11377 | 0.07 | 0.25 | 0 | 1 |
| Number of grandchildren | 10583 | 2.05 | 2.41 | 0 | 20 | 11377 | 1.63 | 2.23 | 0 | 19 |
| ***Income quintiles*** | | | | | | | | | | |
| 1 | 10583 | 0.15 | 0.35 | 0 | 1 | 11377 | 0.14 | 0.34 | 0 | 1 |
| 2 | 10583 | 0.16 | 0.37 | 0 | 1 | 11377 | 0.13 | 0.34 | 0 | 1 |
| 3 | 10583 | 0.20 | 0.40 | 0 | 1 | 11377 | 0.18 | 0.38 | 0 | 1 |
| 4 | 10583 | 0.23 | 0.42 | 0 | 1 | 11377 | 0.25 | 0.43 | 0 | 1 |
| 5 | 10583 | 0.25 | 0.44 | 0 | 1 | 11377 | 0.30 | 0.46 | 0 | 1 |
| Health index | 10583 | 59.63 | 26.44 | 1 | 100 | 11377 | 63.79 | 24.10 | 1 | 100 |

**Table 3. Number of depression symptoms and retirement**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | FE | IV-FE | FE | IV-FE |
| | Males | | Females | |
| Retired | -0.203*** | -0.181 | -0.240*** | 0.136 |
| | (0.062) | (0.312) | (0.082) | (0.264) |
| Age | -0.109 | -0.099 | -0.377** | -0.189 |
| | (0.126) | (0.201) | (0.164) | (0.207) |
| Age2 | 0.015 | 0.014 | 0.027** | 0.010 |
| | (0.009) | (0.016) | (0.011) | (0.017) |
| Marital status: Divorced | 0.107 | 0.108 | 0.224 | 0.248 |
| | (0.221) | (0.218) | (0.245) | (0.245) |
| Marital status: Widowed | 1.249*** | 1.249*** | 0.838*** | 0.835*** |
| | (0.229) | (0.227) | (0.232) | (0.231) |
| Marital status: Not married | -0.298 | -0.300 | -0.003 | 0.037 |
| | (0.354) | (0.354) | (0.433) | (0.378) |
| No. of grandchildren | -0.020 | -0.019 | 0.020 | 0.020 |
| | (0.016) | (0.016) | (0.020) | (0.020) |
| Household income quintile: 2 | -0.110* | -0.110* | -0.015 | -0.015 |
| | (0.064) | (0.063) | (0.087) | (0.086) |
| Household income quintile: 3 | -0.049 | -0.049 | 0.108 | 0.111 |
| | (0.064) | (0.063) | (0.082) | (0.082) |
| Household income quintile: 4 | 0.012 | 0.012 | 0.111 | 0.122 |
| | (0.050) | (0.050) | (0.083) | (0.081) |
| Household income quintile: 5 | 0.042 | 0.043 | 0.181** | 0.181** |
| | (0.064) | (0.064) | (0.080) | (0.078) |
| Health index | -0.016*** | -0.016*** | -0.014*** | -0.014*** |
| | (0.001) | (0.001) | (0.001) | (0.001) |
| Regional unemployment rate | -0.002 | -0.002 | -0.026 | -0.012 |
| | (0.012) | (0.013) | (0.022) | (0.023) |
| Log(regional real GDP per capita) | 1.077* | 1.122 | -0.796 | 0.329 |
| | (0.641) | (0.859) | (1.200) | (1.384) |
| Wave 2 | -0.233 | -0.239 | 0.109 | 0.010 |
| | (0.178) | (0.188) | (0.176) | (0.180) |
| Wave 3 | -0.400 | -0.413 | 0.471 | 0.325 |
| | (0.426) | (0.445) | (0.451) | (0.458) |
| Wave 4 | -0.608 | -0.622 | 0.375 | 0.216 |
| | (0.545) | (0.560) | (0.572) | (0.581) |
| First stage | | | | |
| Normal retirement age | | 0.234*** | | 0.259*** |
| | | (0.026) | | (0.031) |
| Early retirement age | | 0.023 | | 0.017 |
| | | (0.030) | | (0.026) |
| Weak identification: F-test of excluded instruments | | *153.54**** | | *255.93**** |
| Observations | 11,377 | 11,377 | 10,583 | 10,583 |
| Number of IDs | 4,461 | 4,461 | 4,150 | 4,150 |
| Hansen J-test statistic | | *1.809* | | *1.809* |
| p-Value | | *0.1786* | | *0.1786* |
| Durbin–Wu–Hausman test: p-value | | *0.883* | | *0.225* |

**Notes***:* This table shows FE (columns (1) and (3)) and IV-FE estimates (columns (2) and (4)) for the number of depression symptoms. The set of instruments includes the statutory normal and early retirement ages. The omitted categories are ‘married or in a couple’, being interviewed in the first wave, and the first quintile of the household income. Standard errors (clustered by regions – NUTS 1) are in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

**Table 4.** Number of depression symptoms and retirement in economic crisis specifications

| | (1) | (2) | (3) |
|---|---|---|---|
| Variables | **All sample** | **Blue collar** | **White collar** |
| | | **Males** | |
| Retired | -0.113 | -0.534 | 0.0892 |
| | (0.294) | (0.413) | (0.373) |
| Retired x hit by crisis | -0.426** | -0.558** | -0.191 |
| | (0.184) | (0.241) | (0.471) |
| Hit by the crisis | 0.224*** | 0.419*** | 0.138 |
| | (0.0781) | (0.0923) | (0.131) |
| Observations | 11,377 | 3,399 | 7,212 |
| Number of IDs | 4,461 | 1,371 | 2,805 |
| Sargan–Hansen statistic (p-value) | 4.708 (0.095) | 0.693 (0.707) | 2.021 (0.364) |
| Weak identification: Sanderson-Windmeijer F statistic: | | | |
| Retired | 58.11*** | 33.47*** | 26.06*** |
| Retired x hit by the crisis | 104.88*** | 221.06*** | 31.00*** |
| Durbin–Wu–Hausman test: p-value | 0.6901 | 0.1393 | 0.5705 |
| | | **Females** | |
| Retired | 0.175 | -0.0530 | 0.158 |
| | (0.229) | (0.784) | (0.285) |
| Retired x hit by the crisis | -0.338 | -1.094 | -0.279 |
| | (0.382) | (0.768) | (0.468) |
| Hit by the crisis | 0.0617 | 0.558*** | -0.0498 |
| | (0.0973) | (0.198) | (0.0856) |
| Observations | 10,583 | 1,691 | 8,064 |
| Number of IDs | 4,150 | 667 | 3,150 |
| Sargan–Hansen statistic (p-value) | 1.166 (0.558) | 2.578 (0.275) | 0.007 (0.996) |
| Weak identification: Sanderson-Windmeijer F statistic: | | | |
| Retired | 32.23*** | 17.86*** | 32.28*** |
| Retired x hit by the crisis | 53.45*** | 29.00*** | 66.60*** |
| Durbin–Wu–Hausman test: p-value | 0.2096 | 0.0209 | 0.3327 |

**Notes***:* This table reports the IV-FE estimates in an economic crisis specification for the overall sample (column (1)) and for the blue collar sub-sample (column (2)) and white-collar sub-sample (column (3)) for males (top panel) and females (bottom panel). The dummy *hit by the crisis* takes the value one if the cyclical component of the Hodrick–Prescott filter applied to the unemployment rate series of the period 2004–2013 is at least one standard deviation above zero and zero otherwise. The set of instruments includes the statutory normal and early retirement ages and their interaction with the dummy *hit by the crisis*. All specifications include the following additional covariates: second degree polynomial of age; dummy variables for marital status, income quintile, and wave; the number of grandchildren; and the health index. Standard errors (clustered by regions – NUTS 1) are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

**Table 5.** Number of depression symptoms and retirement in economic crisis specifications: Robustness checks for alternative definitions of retirement dummy

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | **Retirement from work** | | **Retirement from unemployment** | |
| | ***Males*** | ***Females*** | ***Males*** | ***Females*** |
| Retired | 0.150 | 0.293 | 0.114 | 0.535 |
| | (0.324) | (0.236) | (0.680) | (0.474) |
| Retired x hit by crisis | -0.483** | -0.126 | 1.146 | 0.035 |
| | (0.206) | (0.472) | (0.851) | (1.402) |
| Hit by the crisis | 0.178*** | 0.035 | -0.291 | 0.273 |
| | (0.069) | (0.095) | (0.385) | (0.452) |
| Observations | 10,344 | 9,613 | 3,027 | 3,299 |
| Number of id | 4,070 | 3,795 | 1,324 | 1,412 |
| Sargan-Hansen statistic (p-value) | 4.370 (0.112) | 0.731 (0.694) | 1.925 (0.382) | 3.459 (0.177) |
| Weak identification: Sanderson-Windmeijer F statistic: | | | | |
| Retired | 52.97*** | 33.22*** | 10.80*** | 18.95*** |
| Retired x hit by the crisis | 100.75*** | 45.39*** | 25.33*** | 4.94*** |

**Notes***:* This table reports the IV-FE estimates in an economic crisis specification for the overall sample for males (column (1) and column (3)) females (column (2) and column (4)). The dummy *hit by the crisis* takes the value one if the cyclical component of the Hodrick–Prescott filter applied to the unemployment rates of the period 2004–2013 is at least one standard deviation above zero and zero otherwise. The set of instruments includes the statutory normal and early retirement ages and their interaction with the dummy *hit by the crisis*. All specifications include the following additional covariates: second degree polynomial of age; dummy variables for marital status, income quintile, and wave; the number of grandchildren; and the health index. Standard errors (clustered by regions – NUTS 1) are in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.

**Table 6:** Number of depression symptoms and retirement in economic crisis specifications - Alternative crisis indicator – IV-FE estimates, males

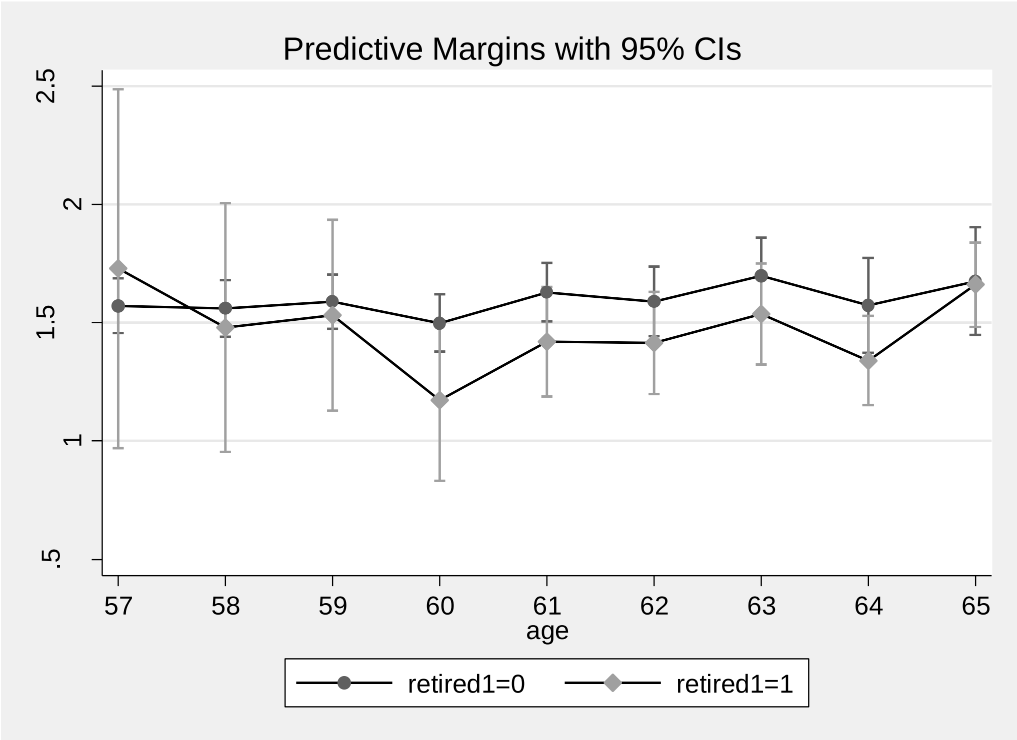
| Crisis indicator: | **Cycle component of unemployment rate** | | **real GDP per capita (dummy =1 if hit)** | |
|---|---|---|---|---|
| | **Whole sample** | **Blue collar** | **Whole sample** | **Blue collar** |
| Retired | -0.131 | -0.553 | -0.0162 | -0.260 |
| | (0.297) | (0.407) | (0.298) | (0.445) |
| Retired x crisis indicator | -0.141** | -0.224* | -0.324*** | -0.624** |
| | (0.0621) | (0.119) | (0.0945) | (0.253) |
| Hit by the crisis | 0.0357* | 0.0356 | 0.00558 | 0.0406 |
| | (0.0185) | (0.0289) | (0.0745) | (0.142) |
| Observations | 11,377 | 3,399 | 11,377 | 3,399 |
| Number of id | 4,461 | 1,371 | 4,461 | 1,371 |

Note: Robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1

**Table 7.** Number of depression symptoms and retirement in economic crisis specifications: Robustness checks for alternative estimation methods

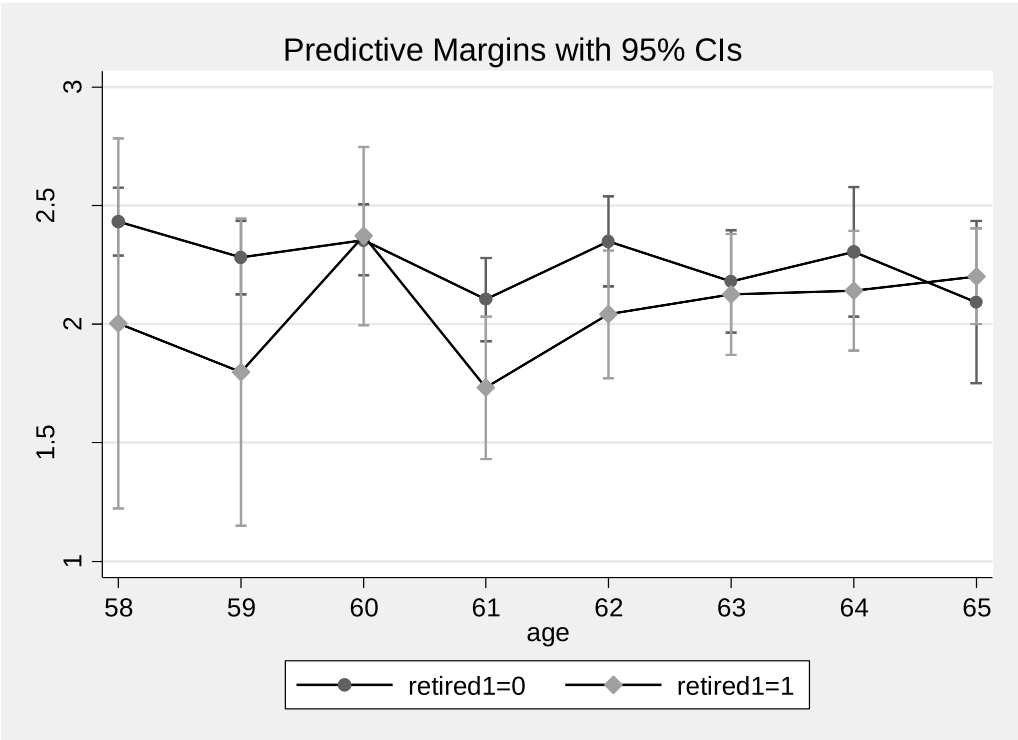
| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | **IV-FD** | | **IV-RE** | | **IV least squares** | |
| | ***Males*** | ***Females*** | ***Males*** | ***Females*** | ***Males*** | ***Females*** |
| Retired | -0.214 | 0.171 | 0.542*** | 0.475*** | 0.970*** | 0.615*** |
| | (0.314) | (0.263) | (0.139) | (0.113) | (0.135) | (0.107) |
| Retired x hit by crisis | -0.369 | 0.617 | -0.245* | 0.227 | -0.255* | 0.265 |
| | (0.244) | (0.517) | (0.141) | (0.205) | (0.134) | (0.190) |
| Hit by the crisis | 0.200* | -0.072 | 0.164*** | -0.016 | 0.192*** | -0.033 |
| | (0.103) | (0.122) | (0.055) | (0.070) | (0.061) | (0.076) |
| Observations | 6,488 | 6,078 | 20,380 | 19,343 | 20,380 | 19,343 |
| Number of IDs | 4,240 | 3,982 | 11,841 | 11,166 | 11,841 | 11,166 |

**Figure 1. EURO-D predictive margins with 95 percent confidence intervals, males**



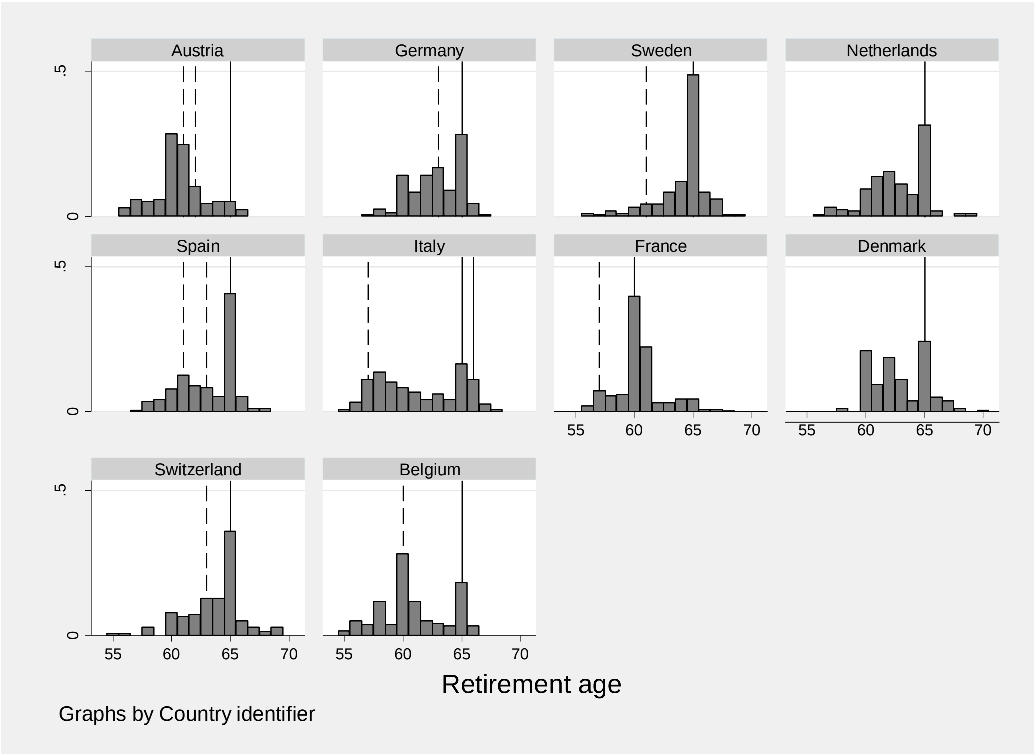

**Notes**: This figure shows the linear predictions from the FE models where the EURO-D score is explained by a full set of age dummies interacted with the dummy *retired*. Standard errors are computed using the delta method.

**Figure 2. EURO-D predictive margins with 95 percent confidence intervals, females**



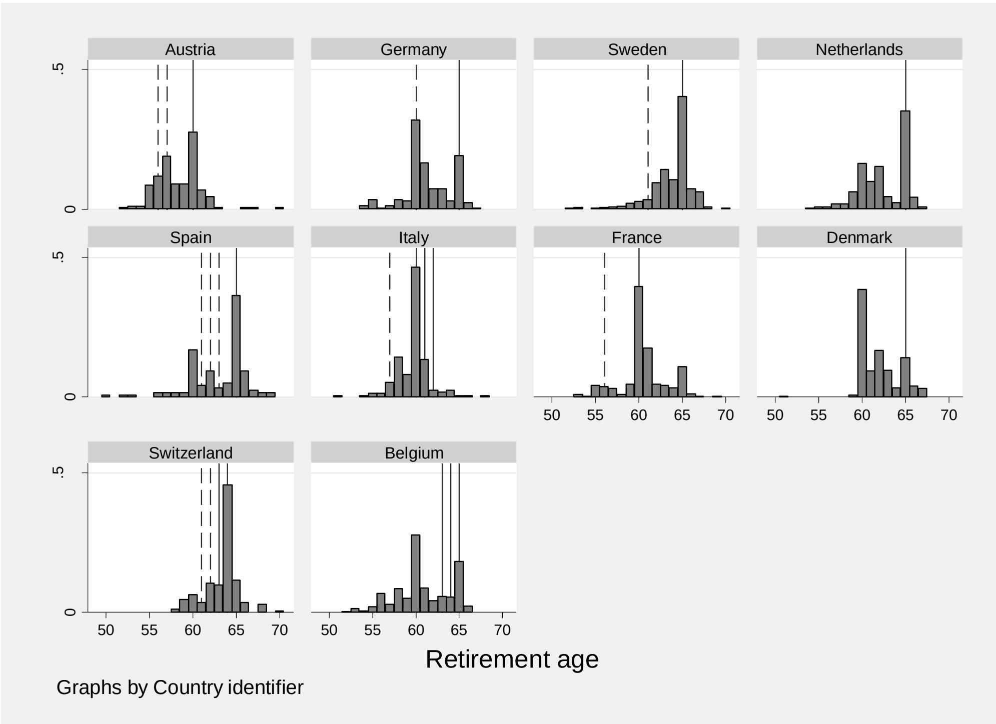

**Notes**: This figure shows the linear predictions from the FE models where EURO-D score is explained by a full set of age dummies interacted with the dummy *retired*. Standard errors are computed using the delta method.

**Figure 3. Retirement age distribution, old age, and early retirement eligibility rules: 2004–2013, males**



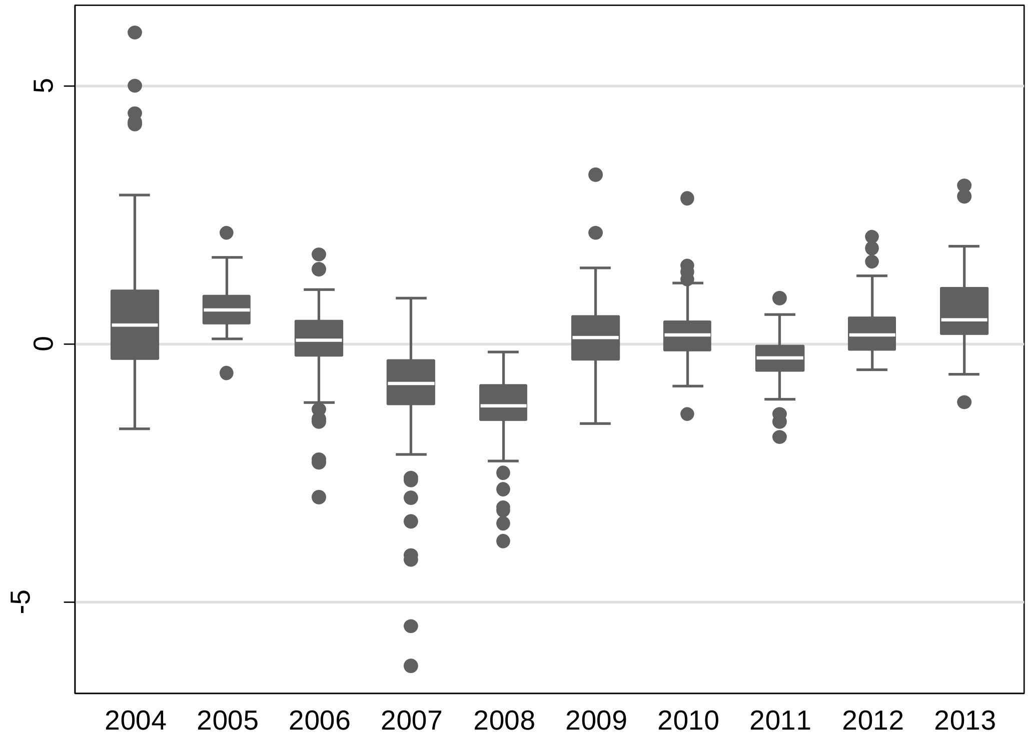

**Notes**: In this figure, a solid line represents old age retirement age(s) and a dashed line early retirement age(s). Early retirement conditions have been tightened dramatically in Italy, for which we only report the earliest retirement age.

**Figure 4. Retirement age distribution, old age, and early retirement eligibility rules: 2004–2013, females**



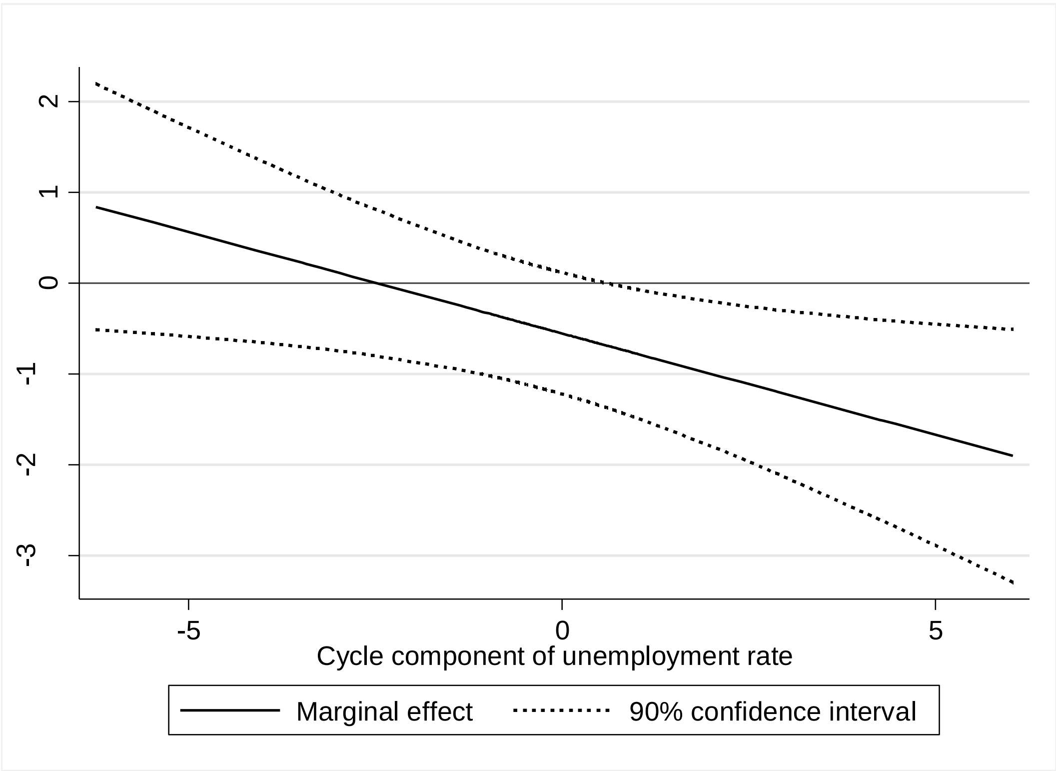

**Notes**: In this figure, a solid line represents old age retirement age(s) and a dashed line early retirement age(s). Early retirement conditions have been tightened dramatically in Italy, for which we only report the earliest retirement age.

**Figure 5. Unemployment rate cyclical component of the Hodrick–Prescott filter, 2004–2013**




**Source**: Our computations on Eurostat (2015) regional data on unemployment rates.

**Figure 6. Marginal effect of retirement on the depression score as a function of the cyclical component of the regional unemployment rate**




Note: Specification in Table 6 column (1).