The effect on mental health of retiring during the economic crisis

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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 variable 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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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 mixed. 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 that retirement has no effect 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 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. Finally, applying the same method to the US Health and Retirement Study, Gorry *et al.* (2015) find that retirement improves mental health (as well as life satisfaction). 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: Exploiting time and geographical heterogeneity of the intensity of the economic crisis, we are able to shed light on the underlying mechanism of how and why mental health can improve at retirement. Our identification strategy is based on an instrumental variable (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 highlight at least three important sources of heterogeneity in the effect of retirement on mental health. We find that retirement improves the mental health of men, but not of women, and only for (ex-) workers in blue collar occupations. Moreover, the effect is much stronger for blue collar men working in regions that were severely hit by the crisis than for those working in less affected regions. We argue that retirement may affect mental health through its effect on stress: Avendano and Berkman (2015) suggest that retirement may release low-skilled workers from job-related stress. This effect is more likely to be important in regions in which economic conditions worsened substantially, increasing workers' perception of the risk of

being laid off and reducing the rewards from working (Eurofund, 2013). According to the life cycle model of stress (Lupien *et al.*, 2009), relief from a period of severe stress induced by retiring during the crisis is likely to reduce the onset of depression.

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 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 the just released 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 score is a standardised scale of depressive symptoms designed to enhance cross-national

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

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, in the context of an analysis of the determinants of cognitive functioning using the Health and Retirement Study survey), 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 (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 selective attrition if the probability of remaining in the survey were inversely related to the mental health of the respondent. This would imply that 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),

[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).

we conduct two 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: Neither the attrition dummy nor its interaction is statistically significant. Second, we regress the attrition dummy on the depression score plus controls and find no significant effect for the depression score. Behncke (2012), facing very similar evidence in ELSA, concludes that the scope of health-related attrition bias cannot be large. A possible explanation for the results of our test is that SHARE includes newly sampled individuals in each wave to maintain population representativeness, a survey design feature that, by construction, alleviates any type of selective attrition.

Table 2 reports summary statistics of the variables included in the analysis, by gender. Women exhibit higher depression scores and are more often single or widowed and their household income distribution is shifted to the left. Additionally, about 42 percent (3,843) 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 model we estimate is the following:

$$y_{i,t} = \boldsymbol{X'}_{i,t}\boldsymbol{\beta_X} + \beta_{RET} D_{i,t}^{RET} + \vartheta_t + d_i + u_{i,t} \qquad (1)$$

The dependent variable $y_{i,t}$ is the EURO-D depression score for individual *i* in wave *t*; the main explanatory variable is the retirement status $D_{i,t}^{RET}$ 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. Following Lindeboom *et al.*, (2002), we include in the model a full set of wave dummies ϑ_t to account for any possible time-variant shock to mental health, which is common to all individuals.

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 $D_{i,t}^{RET}$ is orthogonal to the remaining time-varying residual term $u_{i,t}$ and its leads and lags. It is therefore important to control for potential time-varying observable confounders. In all specifications, $\boldsymbol{X}_{i,t}$ 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. The term $\boldsymbol{X}_{i,t}$ also includes marital status dummies, the number of grandchildren, quintiles of household income,[3] and an aggregated physical health index constructed following Poterba *et al.* (2011).[4]

Still, it is possible that unobservable transitory shocks in the idiosyncratic time-varying residual $u_{i,t}$ 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 $\boldsymbol{Z}_{i,t}$ that affect mental health only through their effect on the retirement decision $D_{i,t}^{RET}$. We follow Coe and Zamarro (2011) and use cross-country variation in the rules determining eligibility to Social Security benefits

[3] 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).

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

as an instrument. More specifically, we use the statutory and early retirement ages from the Mutual Information System on Social Protection (MISSOC) database[5] 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 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 sample, together with statutory old age and early retirement ages. 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.

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. An 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

[5] 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).

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 compare the effects on mental health of retiring in a region severely hit by the crisis and retiring in regions that were not so affected by the economic downturn.

We construct a binary indicator that takes on the value one if a given European region (NUTS 1) in a given year was hit by the economic crisis and zero otherwise. We then use the time series for the real gross domestic product (GDP) per capita from Eurostat regional data (Eurostat, 2015).[6] We apply the Hodrick–Prescott (1997) filter with a smoothing parameter of 6.25 to each of these time series and split the log of the real GDP per capita into trend and cycle components. Our binary indicator *hit by the crisis* takes on the value one if the cycle component is negative. Figure 5 shows the log of the GDP per capita cyclical component from the Hodrick–Prescott filter in our sample period (2004–2013). Regions below the horizontal line are defined as having been hit by the crisis: 3 percent of the regions in 2007 are hit, the percentage goes up to 100 percent in 2009. As stated, we assume that the economic crisis affects

[6] 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).

the relationship between retirement and depression through its effect on working conditions close to retirement. Therefore, we attribute the value of our indicator *hit by the crisis* 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.

We run all our estimations separately by gender to account for different reservation wages and labour supply elasticity between men and women. Finally, to explore the heterogeneity of the effects across different types of occupations, we estimate all equations separately for blue and white collar (ex-) workers.

3. RESULTS

Tables 3 and 4, for males and females, respectively, look at the retirement–mental health relationship without exploiting time and regional heterogeneity in the effect of the crisis, as standard in the literature. Column (1) of Table 3 reports the FE estimates: Moving into retirement leads to a significant reduction of 0.205 in depression scores. Column (2) reports the results from IV estimates with the exclusion restrictions of having reached statutory and early retirement age, respectively. Instruments are informative and valid: The F-test of joint significance of the excluded instruments in the first stage is highly significant and the Hansen J-test shows that the overidentification restriction is valid (p-value 0.52). 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.

Columns (3)–(6) of Table 3 report the FE and IV FE estimates of the same model of columns (1) and (2) but splitting the male sample between blue collar (columns (3) and (4)) and white

collar (columns (5) and (6)) (ex-) workers.[7] Whereas FE estimates of the coefficient of *retired* are not significantly different between occupational groups, the IV FE estimates are. For male blue collar workers, the coefficient of *retired* is -0.58 – that is, equivalent to a 35 percent reduction in the EURO-D score compared to its mean value (equal to 1.69) – and significant at the 10 percent level. For white collar workers, the estimated coefficient of *retired* is equal to 0.03 and is not statistically significant (while it is statistically different from blue collar workers).

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.

Table 4 reports the results of the same set of specifications for women. FE models would suggest that the effect of retirement on mental health for females was significant and not different from that observed for males. However, once we control for endogenous unobservable transitory shocks with IVs, the results for both white and blue collar workers are no longer statistically significant. Diagnostics confirm that the instruments are valid and informative. The results for the control variables are similar to those for males, with one exception: For female white collar workers, higher wages lead to higher depression scores.

The heterogeneous results shown in Tables 3 and 4 could be driven by the economic crisis. The IV estimates account for unobservable transitory shocks affecting both retirement decisions and stress prevalence. 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. Eurofund (2013) shows that job insecurity increased considerably

[7] Following Eurofund's classification, respondents with International Standard Classification of Occupations occupational codes 1 to 5 are defined as white collar workers and those with codes 6 to 9 are defined as blue collars (see http://www.eurofound.europa.eu/surveys/ewcs/2005/classification).

during the crisis. We back up this result with our computations from the last two (2005 and 2010) waves of the European Working Conditions Surveys (EWCS).[8] We find that the increase in job insecurity, as measured by one's perceived chance of losing one's job within the next six months, is very heterogeneous across occupations, increasing by 33 percent for blue collar workers and by 16 percent for white collar workers. The crisis can also affect other dimensions of working life. Eurofund (2013) reports that the recent economic downturn has led to less choice for workers, wage freezes and wage cuts, greater work intensity, deterioration of work–life balance, and greater risk of harassment/bullying. These are all factors that may well affect workers' level of stress and are more likely to affect less-qualified workers. As mentioned in the previous section, the increased stress associated with worse working conditions can induce workers to anticipate retirement and those who retire feel relieved and experience a reduction in depression score.

Stress-related motivation can also explain why we observe a positive effect of retirement on mental health only for males. Men have lower labour supply elasticity and higher labour market attachment than women, who are more likely not to work full time and have a higher marginal utility from leisure (Alesina *et al.*, 2005). Therefore, work-related stress is more likely to affect men than women and the relief associated with retirement is stronger for men.

If the driving force is increased work stress associated with the economic crisis, the results should be stronger in regions that experienced a more pronounced slowdown during recent years. SHARE reports the region of residence of the respondents: After linking individual data to real GDP time series at the regional level, we define a dummy indicator that takes the value one if an individual experienced a particularly bad economic slowdown in a given year and

[8] The EWCS are cross-country surveys that assess and quantify the working conditions of both employees and the self-employed across Europe on a harmonised basis.

zero otherwise. Column (1) of Table 5 reports the estimates of the same set of specifications as column (2) in Tables 4 and 5, but with the dummy *hit by the crisis* and its interaction with the retirement dummy added to the regressors. The top panel reports the results for males: Working in a region and period particularly hit by the crisis has no statistically significant effect on mental health. Moreover, retiring per se has no significant effect. What drives the mental health improvement we observe in Table 3 for males is the positive effect of retiring in the regions and periods severely hit by the crisis: The magnitude of the effect is a reduction in depression scores of 0.32 EURO-D point, which accounts for an approximately 20 percent reduction with respect to the mean value. Looking at the results for women, retiring per se as well as living in a region–year hit by the crisis does not affect mental health. However, retiring during a crisis increases depression scores. This finding can be explained by the fact that, after retirement, women tend to work harder in informal care, especially during periods of economic downturn (Costa-i-Font *et al.*, 2015) and this may be associated with an increase in depressive symptoms (e.g. Coe and van Houtven, 2009).

The diagnostics reported in Table 5 again point to the validity of the chosen instruments, for both men and women,[9] whereas the *retired* regressor no longer passes the exogeneity test for women.

Distinguishing between blue collar and white collar workers (columns (2) and (3) of Table 5), we find that the positive effect of retirement on males' mental health in the region-year hit by the economic crisis outlined in column (1) in Table 5 is entirely driven by blue collar workers. On the contrary, we that find the negative effect of retiring during the crisis for females is

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

driven by white collar workers, with more than 80 percent of all female occupations being white collar.

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 the variable *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 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 6 reports the results of this analysis. Our main results hold – and the coefficients of interest become even larger – 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. From Table 5 onwards, we used the binary indicator *hit by the crisis*, obtained by detrending the time series of the regional real GDP. We alternatively define a continuous indicator based on the same detrended macro variable as a measure of the severity of the crisis. Figure 6 shows the marginal effect of retirement on the depression score as a function of the cyclical component of the real GDP per capita: Retirement significantly (at the 10 percent level) reduces depression scores when economic conditions are bad, namely, up to the 20th percentile of the distribution of the indicator, corresponding to a cyclical component of the GDP equal to –0.22 percent. Different

values of the Hodrick–Prescott filter yield very similar results. We then define the crisis by applying the same method to another macroeconomic indicator: the regional unemployment rate. We define the crisis by means of both a binary and a continuous specification of the unemployment rate's cyclical component. The coefficients of interest are essentially unchanged, thus confirming the robustness of our findings to alternative definitions of the crisis.[10] Our preferred macroeconomic indicator of the crisis remains, however, the real GDP per capita, since the unemployment rate peaks only after the worst point of the business cycle (Stock and Watson, 1999).

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, independent of the contemporaneous error term as well as all its lags and leads. We re-estimated the model with an IV first differences (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 and are virtually unchanged compared to the FE estimates. Until now we always ran the FE model under the implicit consideration that some of the included variables can be correlated with the unobservable time-invariant characteristic. However, if we are willing to assume conditional mean independence between the time-invariant component of the error term and the included regressors, the FE and the FD estimators are less efficient than a random effects (RE) estimator. Moreover, under the conditional independence assumption, the IV least squares estimator is also consistent and requires only contemporaneous exogeneity for the instruments rather than strict exogeneity, as the 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 RE estimates is in line with that of the FE and FD, while it

[10] All the detailed results of this set of robustness checks are available in the Appendix.

loses statistical significance in the linear regression. A potential advantage of linear IV and RE estimates is the possibility of directly controlling for regional dummies, to separately identify regional time-invariant differences from severity of the crisis. The key interaction effect remains significant, even after regional dummies are added to the controls.[11] In the results of Table 7, the coefficient of the retirement dummy becomes significant in both the RE and linear IV regressions: This positive association disappears with the FE or FD, thus signalling the importance of controlling for unobservable characteristics driving both the retirement decision and depression.

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 an FE IV approach using country- and gender- specific statutory and early retirement ages as instruments for retirement behaviour. We find that retirement reduced the depression scores of male ex-blue collar workers but did not affect those of male ex-white collar workers. By exploiting time and geographical heterogeneity in the intensity of the economic crisis, we note that it is not retirement per se that has a protective effect on mental health but, rather, doing so in a period and region severely hit by the economic crisis.

Our results help shed 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

[11] The results are available upon request.

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	1291	14.47
	X	X	X	1316	14.76
X	X	X		458	5.14
X	X		X	177	1.98
X		X	X	217	2.43
		X	X	3683	41.29
X	X			821	9.21
	X	X		315	3.53
	X		X	210	2.35
X			X	100	1.12
X		X		82	0.92

Table 2. Summary statistics, by gender

	Female					Males				
	Obs	***Mean***	***SD***	***Min***	***Max***	***Obs***	***Mean***	***SD***	***Min***	***Max***
Eurod depression score	10781	2.25	2.06	0	12	11558	1.56	1.73	0	12
retired	10781	0.27	0.44	0	1	11558	0.27	0.44	0	1
Age	10781	60.51	3.77	55	70	11558	60.73	3.81	55	70
Marital status										
Married or in a couple	10781	0.72	0.45	0	1	11558	0.81	0.39	0	1
Divorced	10781	0.14	0.35	0	1	11558	0.10	0.30	0	1
Widowed	10781	0.07	0.26	0	1	11558	0.03	0.16	0	1
Never married	10781	0.07	0.25	0	1	11558	0.07	0.25	0	1
Number of grandchildren	10781	2.05	2.41	0	20	11558	1.63	2.23	0	19
Income quintiles										
1	10781	0.15	0.35	0	1	11558	0.14	0.34	0	1
2	10781	0.16	0.37	0	1	11558	0.13	0.34	0	1
3	10781	0.20	0.40	0	1	11558	0.18	0.38	0	1
4	10781	0.23	0.42	0	1	11558	0.25	0.43	0	1
5	10781	0.26	0.44	0	1	11558	0.30	0.46	0	1
Health index	10781	59.56	26.48	1	100	11558	63.76	24.13	1	100

Table 3. Number of depression symptoms and retirement, males

	(1) FE	(2) IV FE	(3) FE	(4) IV FE	(5) FE	(6) IV FE
	Whole sample		Blue collar		White collar	
Retired	-0.205***	-0.174	-0.171*	-0.597*	-0.254***	0.028
	(0.0519)	(0.252)	(0.100)	(0.355)	(0.0622)	(0.373)
Age	-0.0819	-0.066	0.122	-0.091	-0.109	0.056
	(0.130)	(0.190)	(0.261)	(0.313)	(0.155)	(0.265)
Age2	0.0145	0.013	-0.00380	0.013	0.0204*	0.006
	(0.00927)	(0.015)	(0.0185)	(0.023)	(0.0110)	(0.021)
Marital status: Divorced	0.112	0.112	0.312	0.279	0.00255	0.007
	(0.234)	(0.236)	(0.479)	(0.493)	(0.278)	(0.279)
Marital status: Widowed	1.110***	1.246***	0.715*	1.134***	1.032***	1.046***
	(0.228)	(0.228)	(0.404)	(0.411)	(0.291)	(0.293)
Marital status: Never married	-0.321	-0.302	-0.401	-0.258	-0.0159	-0.027
	(0.361)	(0.361)	(0.637)	(0.640)	(0.457)	(0.458)
No. of grandchildren	-0.0207	-0.022	-0.00259	-0.007	-0.0289	-0.027
	(0.0156)	(0.016)	(0.0294)	(0.029)	(0.0189)	(0.019)
Household income quintile: 2	-0.109*	-0.107*	-0.132	-0.115	-0.0724	-0.069
	(0.0644)	(0.064)	(0.105)	(0.105)	(0.0891)	(0.089)
Household income quintile: 3	-0.0477	-0.047	-0.0626	-0.078	-0.00890	-0.005
	(0.0636)	(0.064)	(0.110)	(0.110)	(0.0843)	(0.085)
Household income quintile: 4	0.0113	0.015	-0.00815	0.003	0.0281	0.037
	(0.0642)	(0.064)	(0.116)	(0.115)	(0.0832)	(0.084)
Household income quintile: 5	0.0446	0.049	0.0466	0.070	0.0477	0.062
	(0.0657)	(0.066)	(0.127)	(0.128)	(0.0836)	(0.086)
Health index	-0.0158***	-0.016***	-0.0181***	-0.018***	-0.0147***	-0.015***
	(0.000968)	(0.001)	(0.00177)	(0.002)	(0.00120)	(0.001)
Wave 2	-0.228	-0.235	-0.229	-0.071	-0.320*	-0.386*
	(0.156)	(0.170)	(0.323)	(0.345)	(0.182)	(0.202)
Wave 3	-0.496	-0.497	-0.385	-0.022	-0.750	-0.886*
	(0.402)	(0.424)	(0.818)	(0.857)	(0.470)	(0.504)
Wave 4	-0.756	-0.756	-0.529	-0.072	-1.133*	-1.293**
	(0.518)	(0.543)	(1.055)	(1.102)	(0.607)	(0.644)
First stage						
Normal retirement age		0.246***		0.334***		0.199***
		(0.014)		(0.026)		(0.018)
Early retirement age		0.020*		0.038*		0.019
		(0.012)		(0.022)		(0.015)
Weak identification: F-test of excluded instruments		153.54***		87.13***		63.37***
Observations	11,558	11,547	3,595	3,586	7,380	7,380
Number of IDs	4,606	4,604	1,548	1,546	2,959	2,959
Hansen J-test statistic		1.809		0.715		0.374
p-Value		0.1786		0.3977		0.541
Durbin–Wu–Hausman test: p-value		0.883		0.225		0.441

Notes: This table reports the FE estimates (columns (1) and (2)) and FE IV estimates (columns (3) 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 household income. Standard errors are in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1.

Table 4. Number of depression symptoms and retirement, females

	(1) FE	(2) IV FE	(3) FE	(4) IV FE	(5) FE	(6) IV FE
	Whole sample		**Blue collar**		**White collar**	
Retired	-0.222***	0.164	-0.171	0.045	-0.229***	0.168
	(0.064)	(0.235)	(0.165)	(0.661)	(0.072)	(0.261)
Age	-0.389**	-0.165	-0.536	-0.441	-0.265	-0.016
	(0.161)	(0.209)	(0.441)	(0.516)	(0.180)	(0.240)
Age2	0.028**	0.009	0.035	0.026	0.023*	0.002
	(0.012)	(0.017)	(0.032)	(0.041)	(0.013)	(0.019)
Marital status: Divorced	0.218	0.250	-0.046	-0.055	0.311	0.360
	(0.292)	(0.294)	(0.682)	(0.683)	(0.323)	(0.326)
Marital status: Widowed	0.840***	0.841***	0.365	0.374	1.054***	1.057***
	(0.182)	(0.182)	(0.415)	(0.416)	(0.217)	(0.218)
Marital status: Never married	-0.045	0.043			-0.103	-0.015
	(0.938)	(0.943)			(0.915)	(0.920)
No. of grandchildren	0.019	0.019	-0.065	-0.067	0.029	0.030
	(0.019)	(0.019)	(0.051)	(0.051)	(0.021)	(0.021)
Household income quintile: 2	-0.005	-0.008	-0.171	-0.178	0.097	0.099
	(0.073)	(0.073)	(0.164)	(0.165)	(0.087)	(0.088)
Household income quintile: 3	0.115	0.116	0.097	0.100	0.196**	0.200**
	(0.075)	(0.075)	(0.177)	(0.177)	(0.089)	(0.089)
Household income quintile: 4	0.113	0.122	-0.111	-0.116	0.220**	0.239**
	(0.078)	(0.079)	(0.199)	(0.201)	(0.092)	(0.093)
Household income quintile: 5	0.185**	0.184**	0.110	0.107	0.271***	0.276***
	(0.081)	(0.082)	(0.212)	(0.212)	(0.094)	(0.095)
Health index	-0.014***	-0.014***	-0.021***	-0.021***	-0.013***	-0.013***
	(0.001)	(0.001)	(0.003)	(0.003)	(0.001)	(0.001)
Wave 2	0.080	0.005	-0.023	-0.055	-0.014	-0.100
	(0.185)	(0.191)	(0.459)	(0.470)	(0.210)	(0.217)
Wave 3	0.400	0.261	0.639	0.600	0.039	-0.141
	(0.480)	(0.489)	(1.214)	(1.224)	(0.544)	(0.558)
Wave 4	0.280	0.117	0.674	0.634	-0.179	-0.398
	(0.620)	(0.629)	(1.574)	(1.583)	(0.701)	(0.717)
First stage						
Normal retirement age		0.291***		0.266***		0.300***
		(0.013)		(0.033)		(0.015)
Early retirement age		-0.004		0.050		-0.021
		(0.012)		(0.032)		(0.013)
Weak identification: F-test of excluded instruments		255.96***		33.65***		202.51***
Observations	10,781	10,776	1,775	1,773	8,275	8,272
Number of IDs	4,313	4,313	747	747	3,331	3,331
Hansen J-test statistic		1.809		2.048		0.210
p-Value		0.1786		0.1524		0.6466
Durbin–Wu–Hausman test: p-value		0.089		0.738		0.113

Notes*:* This table shows the FE estimates (columns (1) and (2)) and IV FE estimates (columns (3) 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 are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 5. Number of depression symptoms and retirement in economic crisis specifications

	(1)	(2)	(3)
Variables	**All sample**	**Blue collar**	**White collar**
		Males	
Retired	-0.0162	-0.260	0.120
	(0.298)	(0.445)	(0.356)
Retired x hit by crisis	-0.324***	-0.624**	-0.224
	(0.0945)	(0.253)	(0.155)
Hit by the crisis	0.00558	0.0406	-0.0236
	(0.0745)	(0.142)	(0.107)
Observations	11,377	3,399	7,212
Number of IDs	4,461	1,371	2,805
Sargan–Hansen statistic (p-value)	2.406(0.300)	1.432(0.489)	0.581(0.748)
Weak identification: Sanderson-Windmeijer F statistic:			
Retired	54.68***	25.42***	38.26***
Retired x hit by the crisis	145.64***	108.91***	129.55***
Durbin–Wu–Hausman test: p-value	0.2631	0.1767	0.6374
		Females	
Retired	0.00869	0.127	-0.00455
	(0.248)	(0.802)	(0.298)
Retired x hit by the crisis	0.381*	-0.0846	0.416*
	(0.203)	(0.491)	(0.216)
Hit by the crisis	-0.144	0.0627	-0.158
	(0.114)	(0.187)	(0.135)
Observations	10,583	1,691	8,064
Number of IDs	4,150	667	3,150
Sargan–Hansen statistic (p-value)	0.474(0.789)	1.342(0.511)	0.210(0.900)
Weak identification: Sanderson-Windmeijer F statistic:			
Retired	52.77***	26.68***	41.22***
Retired x hit by the crisis	114.00***	51.78***	102.88***
Durbin–Wu–Hausman test: p-value	0.0705	0.8616	0.1002

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 log of the real GDP per capita for the period 2004–2013 is negative 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. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

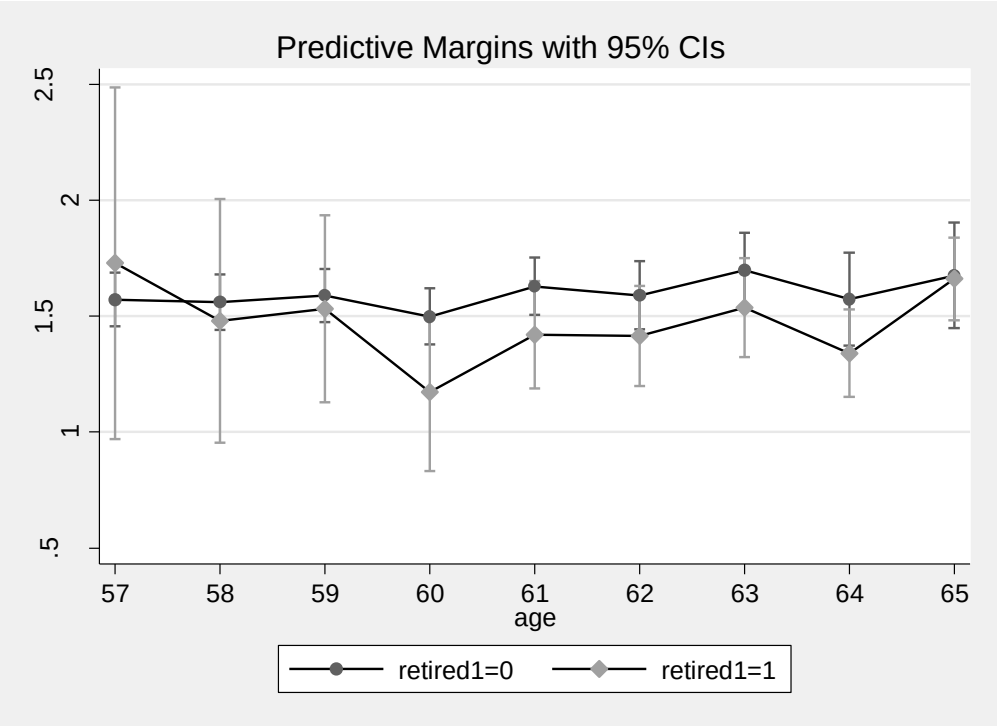
Table 6. Number of depression symptoms and retirement in economic crisis specifications: Robustness checks for alternative definitions of retirement dummy

	(1)	(2)	(3)	(4)
	Retirement from work		**Retirement from unemployment**	
	Males	***Females***	***Males***	***Females***
Retired	0.294	0.0765	0.0384	1.045
	(0.315)	(0.247)	(1.032)	(1.204)
Retired x hit by crisis	-0.400***	0.530**	0.476	-0.279
	(0.117)	(0.217)	(0.654)	(0.924)
Hit by the crisis	0.0107	-0.224**	-0.397	-0.0996
	(0.0883)	(0.112)	(0.407)	(0.542)
Observations	10,344	9,613	2,547	2,346
Number of id	4,070	3,795	1,137	1,038
Sargan-Hansen statistic (p-value)	2.447 (0.2941)	0.284 (0.8676)	2.488 (0.2882)	1.912 (0.3844)
Weak identification: Sanderson-Windmeijer F statistic:				
Retired	56.57***	53.33***	6.92***	8.31***
Retired x hit by the crisis	128.16***	108.53***	13.39***	7.02***

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

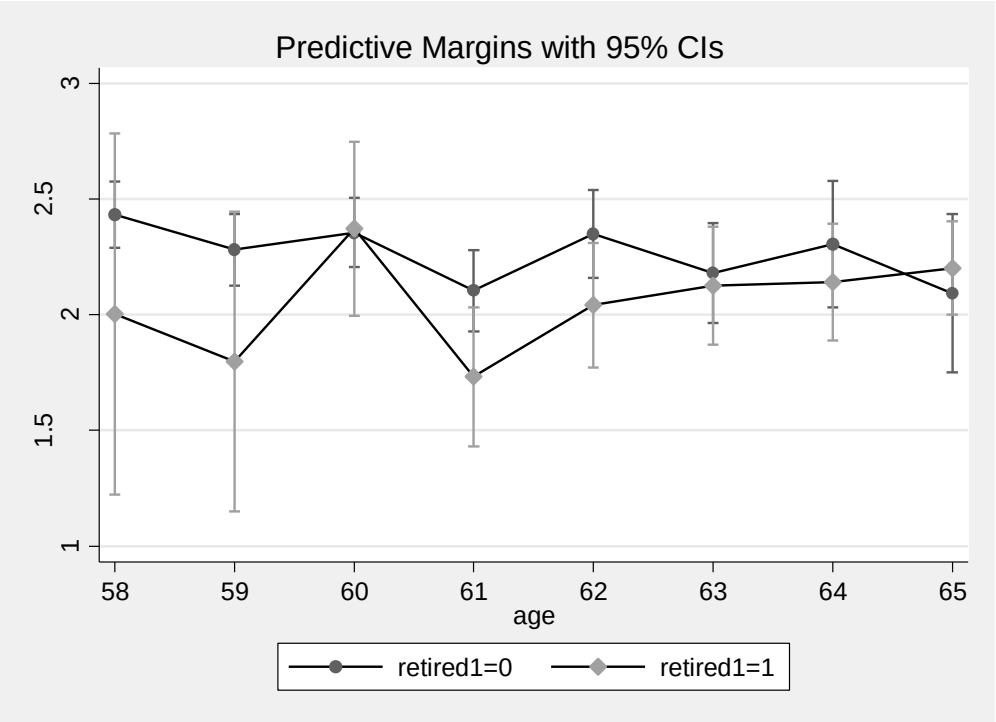
	(1)	(2)	(3)	(4)	(5)	(6)
	FD		**IV RE**		**IV least squares**	
	Males	***Females***	***Males***	***Females***	***Males***	***Females***
Retired	-0.0428	-0.0117	0.597***	0.647***	1.264***	1.171***
	(0.367)	(0.322)	(0.221)	(0.211)	(0.284)	(0.439)
Retired x hit by crisis	-0.413**	0.452*	-0.285**	0.270*	-0.182	0.186
	(0.167)	(0.253)	(0.134)	(0.163)	(0.165)	(0.174)
Hit by the crisis	-0.00417	-0.157	-0.0544	-0.218***	-0.166*	-0.297*
	(0.0898)	(0.130)	(0.0639)	(0.0788)	(0.0962)	(0.154)
Observations	6,488	6,078	11,491	10,712	11,491	10,712
Number of IDs			4,575	4,279		

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

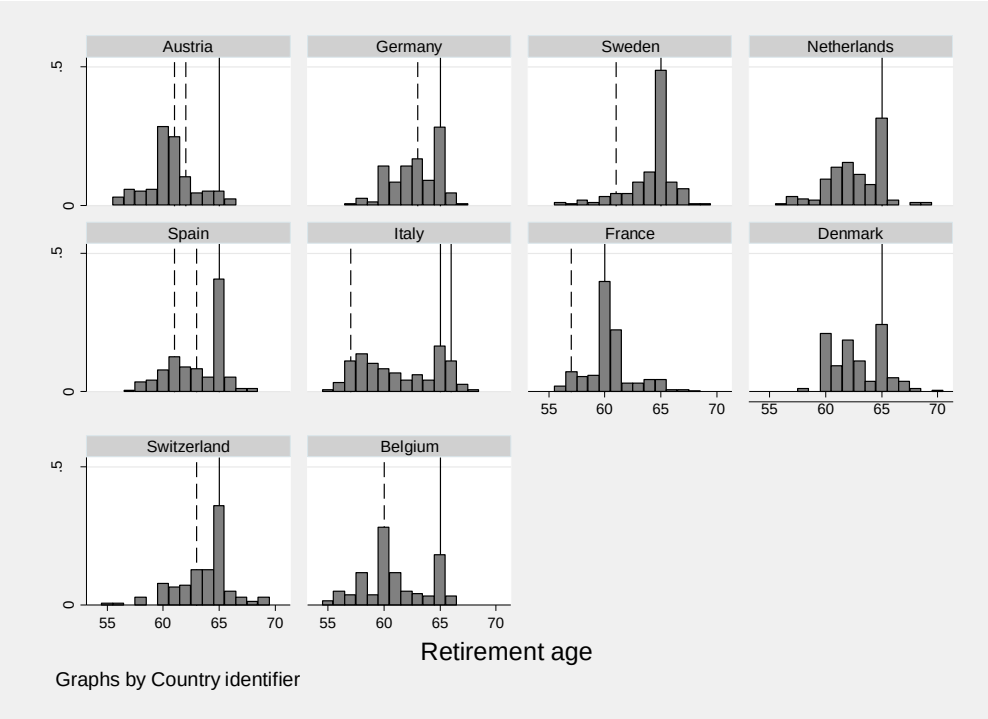
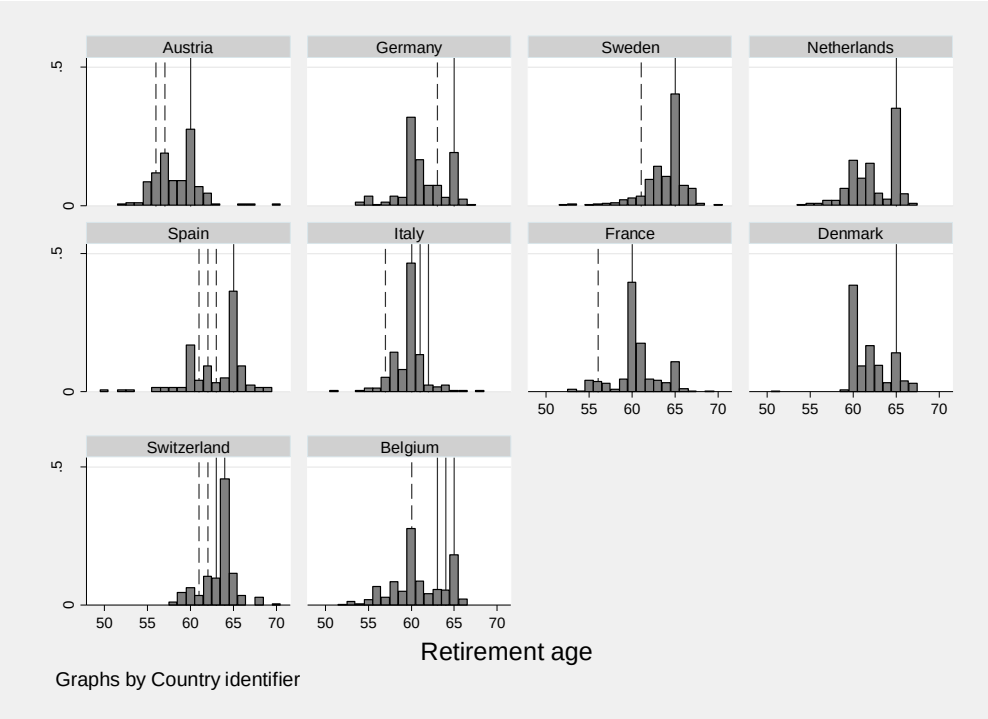
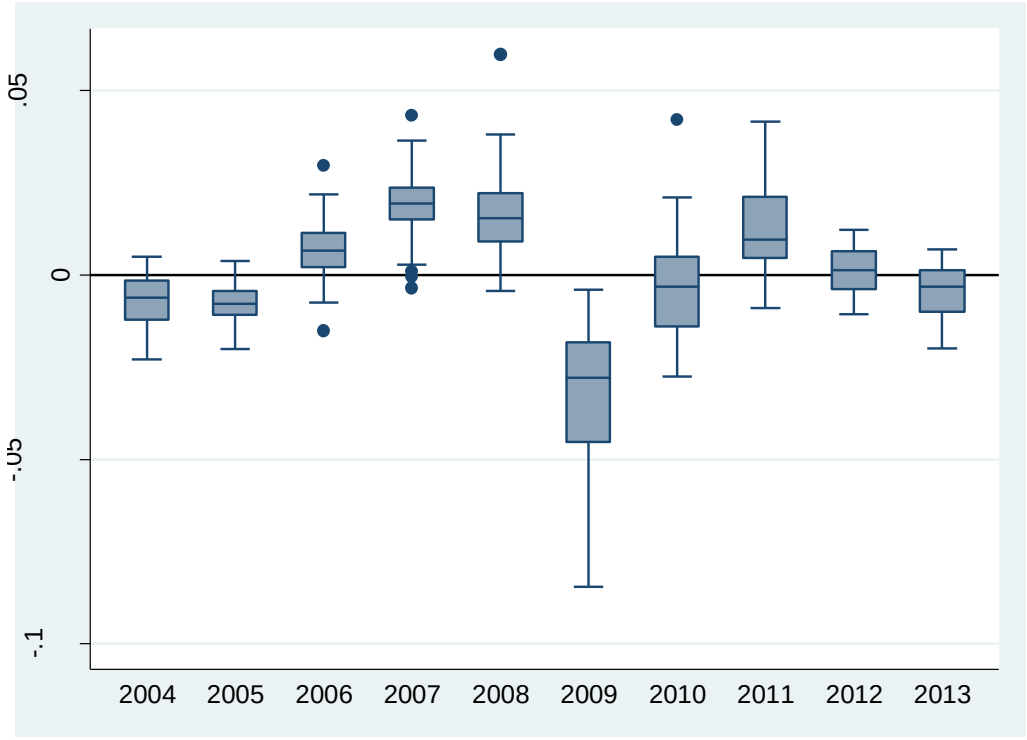


Figure 4. Retirement age distribution, old age, and early retirement eligibility rules: 2004–2013, females



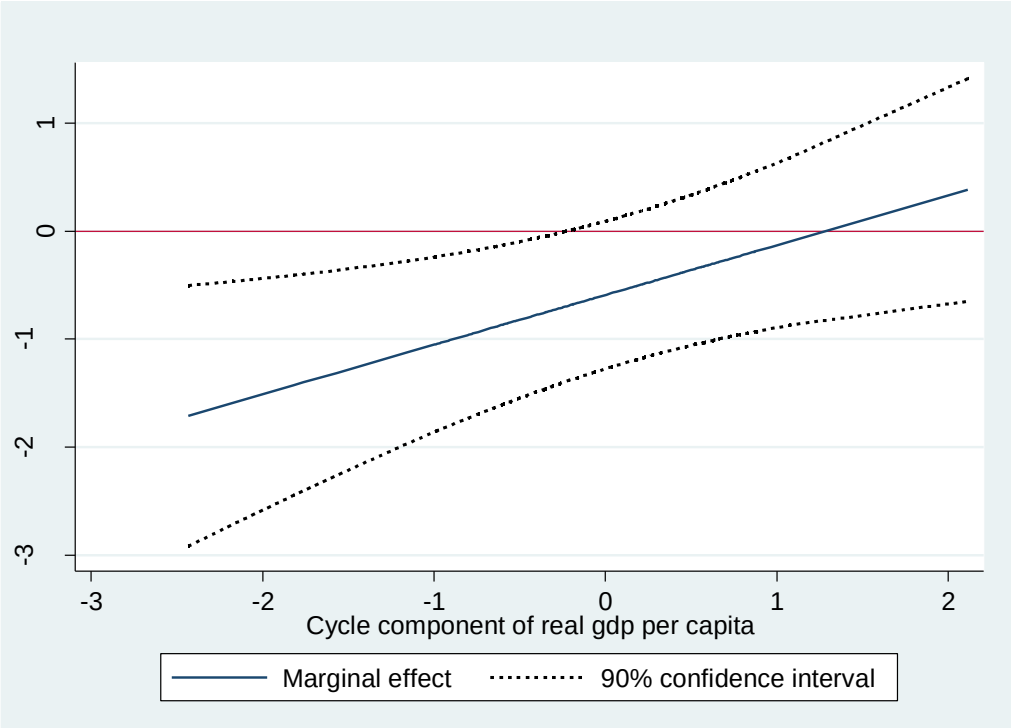
Notes: In this figure, a solid line represents old age retirement age(s) and a dashed line early retirement age(s). Early retirement conditions have been tightened dramatically in Italy, for which we only report the earliest retirement age.

Figure 5. log GDP per capita cyclical component of the Hodrick–Prescott filter, 2004–2013



Source: Our computations on Eurostat (2015) GDP regional data.

Figure 6. Marginal effect of retirement on the depression score as a function of the cyclical component of the log of the real GDP per capita



APPENDIX

6. HEALTH INDEX

The aggregated physical health index used in the analysis is constructed following Poterba et al. (2011). In particular, for each wave we run a Principal Component Analysis (PCA) on a rich set of variables related to objective health conditions (i.e.: presence of mobility limitations, such as walking 100 meters or climbing several flights of stairs; having at least one limitation in the activities of daily limits, having back problems, heart disease, stroke, hypertension, diabetes or cancer; having had a doctor visit, overnight stay in the hospital or in the nursing home and BMI). We did not to include in the PCA self-reported measures, such as self-reported health status or having a mental illness diagnosed, since these are variables that are likely to be affected by mental health status. The first component is then extracted and the obtained index is translated into percentile scores of objective health index. Table A1 reports the results of the Principal Component Analysis on the set of health measures. The first five components for which eigenvalues are greater than 1 are shown. The first component explains 20 per cent of the overall variance, almost as much as the four following components.

Table A1: First five components of Principal component analysis (PCA) for the aggregate health index.

Variable	Components				
	1st	2nd	3rd	4th	5th
Diff. walking	0.301	-0.107	0.140	-0.140	-0.074
Diff. lifting weights	0.316	-0.160	0.021	0.046	-0.082
Diff. pulling objects	0.312	-0.207	0.062	-0.021	-0.081
Limited ADL	0.296	-0.149	0.097	-0.129	0.049
Diff. climbing stairs	0.318	0.072	-0.453	-0.029	-0.143
Diff. stooping	0.318	0.064	-0.152	-0.021	0.004
Diff. getting up from a chair	0.304	-0.037	-0.159	-0.056	0.018
Diff. reaching arms	0.251	-0.167	-0.020	-0.084	-0.013
Arthritis	0.158	0.093	-0.453	0.307	0.181
Diff. sitting for 2h	0.224	-0.107	-0.166	-0.019	-0.018
Diff. picking a coin	0.193	-0.194	0.090	-0.180	0.064
Back problem	0.198	0.138	-0.474	0.245	0.170
Heart disease	0.137	0.204	0.291	0.154	-0.110
Stayed in the hosp.	0.139	0.119	0.343	0.386	0.054
Visited a doctor	0.101	0.313	0.063	0.361	0.136
Stroke	0.117	0.005	0.324	-0.046	0.302
Hypertension	0.114	0.485	0.055	-0.140	0.107
Lung disease	0.106	0.079	0.083	0.178	-0.565
Diabetes	0.107	0.390	0.149	-0.282	-0.012
BMI	0.097	0.473	-0.086	-0.365	-0.016
Stayed in a nursing home	0.068	-0.112	0.222	0.028	0.652
Cancer	0.052	0.015	0.228	0.439	-0.099
Eigen value	4.562	1.427	1.287	1.130	1.000
Explained variance	0.207	0.065	0.059	0.051	0.046

7. ROBUSTNESS

We test the sensitivity of our estimates to the particular age polynomial included in the regressions. This is a crucial issue in our empirical setting despite the lack of significance of age in most specifications: the identification power of the causal effect of interest comes from a discontinuity at specific retirement ages, conditional on an underlying smooth ageing process of mental health. Column (2) and (6) of table A2 report the same results with a quadratic polynomial in age as in table 5 for convenience. We experimented with a linear, a cubic and a quartic specification in age. Coefficient estimates of being retired are almost unaffected.

Table A2: Number of depression symptoms and retirement, in specifications with economic crisis. Alternative age polynomial

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Males				***Females***			
Retired	0.112	-0.016	-0.013	0.015	0.053	0.009	0.051	0.043
	(0.175)	(0.298)	(0.297)	(0.333)	(0.195)	(0.248)	(0.285)	(0.290)
Retired x hit by crisis	-0.321***	-0.324***	-0.324***	-0.325***	0.389*	0.381*	0.383*	0.381*
	(0.095)	(0.095)	(0.095)	(0.095)	(0.199)	(0.203)	(0.203)	(0.203)
Hit by the crisis	-0.013	0.006	0.005	0.002	-0.153	-0.144	-0.149	-0.148
	(0.063)	(0.075)	(0.075)	(0.076)	(0.106)	(0.114)	(0.118)	(0.119)
Age	0.098*	-0.055	0.723	0.744	-0.061	-0.144	2.021	12.936*
	(0.057)	(0.209)	(0.616)	(0.649)	(0.071)	(0.220)	(4.098)	(6.706)
Age 2		0.012	-0.114	-0.105		0.007	-0.342	-3.000*
		(0.017)	(0.092)	(0.203)		(0.018)	(0.654)	(1.704)
Age 3			0.683	0.459			1.870	30.566
			(0.452)	(4.793)			(3.454)	(19.817)
Age 4				0.138				-11.592
				(2.998)				(8.736)
Observations	11,377	11,377	11,377	11,377	10,583	10,583	10,583	10,583
Number of id	4,461	4,461	4,461	4,461	4,150	4,150	4,150	4,150
Sargan stat p-value	0.295	0.300	0.325	0.387	0.780	0.789		

Table A3: Alternative definitions of crisis indicator

	(1) **IV- FE males BC**	(2) **IV- FE males BC**	(3) **IV- FE males BC**
Retired	-0.534	-0.592	-0.553
	(0.413)	(0.414)	(0.407)
Retired x hit by crisis (unemployment rate)	-0.558**		
	(0.241)		
Hit by the crisis (unemployment rate)	0.419***		
	(0.0923)		
Retired x Cycle component of real GDP per capita		0.460*	
		(0.236)	
Cycle component of real GDP per capita		-0.0999	
		(0.137)	
Retired x Cycle component of unemployment rate			-0.224*
			(0.119)
Cycle component of unemployment rate			0.0356
			(0.0289)
Observations	3,399	3,399	3,399
Number of id	1,371	1,371	1,371

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