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Keywords
depression, stress, retirement, crisis

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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 depression, a growing concern for public health as 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 the time and regional variation in the severity of the economic crisis across ten European countries over the 2004-2013 period. We use data from four waves of the Survey of Health, Ageing and Retirement in Europe (SHARE) and address the potential endogeneity of retirement decision to mental health by applying a fixed-effect instrumental variable approach. Results indicate that retirement improves mental health of men, but not of women. This effect is stronger for those men working in regions that are severely hit by the economic crisis and in blue-collar jobs. These findings may be explained by the worsening of working conditions and the rise in job insecurity stemming from the economic downturn: In these circumstances, the exit from the labor force is perceived as a relief.

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1. INTRODUCTION

The present study examines whether retiring during the recent economic crisis leads to benefits or losses in terms of mental health. We focus on late life depression, a serious and growing public health problem: it is estimated that the economic cost of depression is $83.1 in the United States (Greenberg et al., 2003) and €118 billion in Europe (Sobocki et al, 2006).

We are not the first to address this issue: evidence regarding both the United States and Europe before the economic crisis is mixed. Early works in social epidemiology find a positive association, while more recent contribution report either no association or a negative correlation (see Avendano, Berkman, 2015 for a review of the relevant social epidemiology literature). Papers addressing the potential endogeneity of retirement decision to mental health induced by unobserved individual preferences towards time and risk still lead to mixed results. Dave et al (2008) use US panel data from 2002 until 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 longitudinal data on a different cohort study and find no effect of retirement on mental health (while they find evidence of an effect of other life events).

A potential reason for the inconclusiveness of these panel studies is that the simultaneity issue may also be induced by unobserved shocks in the individual environment that affect mental health and retirement decision. If this is the case, fixed effect methods do not solve the endogeneity problem. Behncke (2012) addresses the issue using an instrumental variables approach, which exploits changes in pension eligibility in England. The author uses three waves of the English Longitudinal Study of Ageing and a matching method coupled with an IV estimator. While finding that retirement increases the risk of being diagnosed a chronic condition, she finds no statistically significant relation with mental health. A first piece of evidence that partially contradicts Behncke (2012), but relies on a similar identification

strategy, is De Grip et al (2012). The authors exploit the same source of variation, namely changes in eligibility for state pension, but look at Dutch data and find strong evidence that postponing the statutory retirement age leads to a worsening of mental health of those affected by the policy change that are still at work. Still, a negative effect of a policy change on those who are not retired may not imply that retirement is beneficial for mental health. Coe and Zamarro (2011) look at statutory retirement age as well, but exploit a different source of exogenous variation: they use data on individuals observed across different European countries interviewed in the first wave of the Survey on Health Ageing and Retirement in Europe (SHARE), and conclude that retiring actually leads to lower depression scores. To our knowledge, there is no previous study analysing the effect on depression of retiring during a period characterised by a severe economic crisis. This paper relies on the panel component of SHARE, comprising waves from 2004 and including the just released 2013 wave. During this period, European countries and regions were differently affected by the crisis and we exploit this heterogeneity to identify the role of economic conditions in shaping the effect of retirement on mental health.

Our paper improves on the existing literature along two dimensions. First, we combine the identification strategies previously proposed in the literature and apply a fixed-effect instrumental-variable approach, where unobserved individual heterogeneity is controlled by including fixed effects and endogeneity is addressed exploiting the exogenous variation in retirement behavior induced by country-specific rules on early and full retirement. Second, by exploiting the above-mentioned time and geographical heterogeneity in the intensity of the economic crisis, we are able to shed some light on the underlying mechanism for how and why mental health can improve at retirement. Our results highlight at least two important sources of heterogeneity in the effect of retirement on depression. We find that retirement improves mental health of men, but not of women. Moreover, the effect is much stronger for men

working in regions that are severely hit by the economic crisis than for those working in less affected regions and this effect is concentrated in blue-collar occupations. We argue that retirement may affect mental health through its effect on stress: Avendano and Berkman, (2015), suggest that retirement may release in particular low-skilled workers from job-related stress. This effect is more likely to be important in regions in which the economic conditions worsened substantially, increasing workers' perception of the risk of being laid off and reducing the rewards from working (Eurofund, 2013). According to stress process theory (Pearlin et al., 2005), the relief from a period of augmented stress induced by retiring during the crisis is likely to reduce the onset of depression.

The remainder of the paper is organized as follows. The next section describes the data and the estimation method. Section 3 presents the results of the analysis, while in the last section we discuss the policy implications and draw the 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-institutionalized 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). The most recent wave is particularly useful for our scopes since it provides fresh evidence on the economic crisis. We exclude the third wave since it is not directly comparable to the other waves (it was a retrospective life history interview). We select respondents from ten countries present in all four abovementioned waves (i.e. Austria, Germany, Sweden, the Netherlands, Spain, Italy, France, Denmark, Switzerland and Belgium).

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

Alike in many other studies (see, 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 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.[1] The retired variable is adjusted using the information on self-reported year of retirement: if retirement status is missing but an individual reports the retirement year in any wave, the retirement status is then filled in accordingly. The inconsistencies in the self-reported retirement status between waves are solved assuming that retirement is an absorbing state (see, e.g. Jimenez-Martin et. al., 1999; Mastrogiacomo et. al., 2004): once an individual reports to be retired in one wave, she is considered as retired in all the subsequent waves. It turns out that only a very small fraction of observations (less than 1 %) are adjusted in this way.

Table 1 reports the pattern of individual participation to the panel. Attrition in SHARE is non-negligible, and any analysis regarding health or mental health may suffer of selective attrition. Therefore, it is reasonable to expect the probability of participating in a SHARE wave to be inversely related to mental health of the respondent: this means that the retention rate of the

[1] Respondents who are observed in disability or self-report themselves as being homemaker or in the residual "other" category are not considered in the analysis. On this point, we distinguish from Bonsang et.al. 2012, who consider these categories as "retired" (not working for pay). Generally speaking, most econometric issues we deal with in this study are common to Bonsang et.al. 2012.

panel component of the sample may be affected by the variable that is the object of interest of our analysis. Following the strategy proposed by Lindeboom et al. (2002), we run an informal test for the severity of the selective attrition issue. We run a regression of our depression scale on the probability of dropping from the sample (i.e., a dummy which takes value 1 if the *i*-th observation is not in the panel at time t+1) and on the key variables of our analysis. Controlling for a full set of wave dummies, the attrition dummy is not statistically significant.

Table 2 reports summary statistics of the variables included in the analysis, by gender. Additionally, it is worth mentioning that about 6 % (more than 5,000) of sampled individuals retired throughout the analysed period. Some preliminary evidence on the relationship between depression and the retirement status can be found in Figures 1 and 2. They show, for males and females respectively, linear predictions from Fixed Effects models where EURO-D score is explained by a full set of age dummies interacted with the dummy retired. Although standard errors are large, these figures point out that being retired is associated with a lower level of depression at every age (we limit this observation to the age range 57 to 65, which is when retirement occurs in our data).

The model we estimate on our panel dataset is the following:

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

Our dependent variable $y_{i,t}$ is the EURO-D depression score measure for individual *i* in wave *t*; the main explanatory variable is the retirement status $D_{i,t}^{RET}$ which is, as described earlier, a dummy variable taking value one if the individual reports being retired at the time of the interview and zero in the labour force. Following Lindeboom et al, (2002), we include in the model a full set of wave dummies ϑ_t. This accounts for any possible time variant shock to mental health, which is common to all individuals.

The use of a panel dataset allows us to control for unobservable time invariant characteristics affecting both individual retirement decision and mental health, by including individual fixed effects d_i in all our specifications. Fixed effect estimates are based on the within observation variation over time, and therefore account for any observable time invariant individual characteristic, such as education or country of residence. Still, results obtained using a within estimator will be unbiased only if $D_{i,t}^{RET}$ is orthogonal to the remaining time-varying residual term $u_{i,t}$. It is therefore important to control for potential time-varying observable confounders. All the specifications include in $\boldsymbol{X}_{i,t}$ a second order polynomial in age, while a set of richer specifications includes marital status differentiating between "married or in couple", "divorced or separated", "widowed" and "never married"; number of grandchildren, quintile of imputed household net (wave 1) or gross (subsequent waves) income and an aggregated physical health index (Table 2 reports some descriptive statistics of all variables included in the analysis). The health index is constructed following the methodology of 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 decided not to include in the PCA self-reported measures, such as self-reported health being fair or poor or having a mental illness diagnosed, since these are variables that are likely to be affected by mental health status. Then the first component is extracted and the obtained index is translated into percentile scores of objective health index. Table 3 reports the results of the Principal Component Analysis on the set of health measures. The first five components for which eigen-values are greater than 1 are shown. The first component explains 20% of the overall variance, almost as much as the four following components.

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 Instrumental Variables approach: we need a set of variables $\boldsymbol{Z}_{i,t}$ that affect mental health only through their effect on the retirement decision $D_{i,t}^{RET}$, in other words we need instruments $\boldsymbol{Z}_{i,t}$ to be informative (i.e., the instruments must correlate with the decision to retire) and exogenous ($\boldsymbol{Z}_{i,t}$must be orthogonal to $u_{i,t}$). We follow Coe and Zamarro (2011), Charles (2004) and Dave et al (2008) and use the rules determining eligibility to Social Security benefits as instruments. Being more specific, we use the statutory and early retirement ages from the *Mutual Information System on Social Protection* (MISSOC) database[2] as instruments for retirement decision. The first (second) instrument for retirement status takes value 1 if respondent has reached the standard (early) retirement age at the time of the interview and 0 otherwise. The identifying assumption that justifies the use of these instruments is the same as in Coe and Zamarro (2011): being age eligible to statutory and early retirement pension benefits does not affect mental health directly, but it provides incentives to retire, as it emerges from 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 a lot of variation in retirement age; however, spikes at legal retirement ages clearly emerge for most of the countries. Note that our identification strategy relies not so much on the change in retirement rules over time (see Figures 3 and 4 and footnote thereby), but rather on the discontinuity in the number of retired individuals at the legal retirement ages, conditional on a smooth function of age.

2 The 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

An open question is nonetheless what is the underlying mechanism that governs the relation between labour market participation decision and depression. We think that 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 can help shading light on this mechanism. Fixed Effect estimation allows to control for any unobservable time invariant individual characteristic, therefore the type of time varying confounder inducing the unobservable transitory shock we deal with the IV is a sudden change in the environment of the respondent, which affects both labour market participation decision 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 shocks. Worsened economic prospects for a region are likely to increase work related stress at least through two channels. First, working in a region where economic prospects are bad increases the risk of being laid off (see Benitez-Silva et al., 2011 and Eurofund, 2013). Second, most employers retrenched substantially as a consequence of the crisis reducing monetary and non-monetary rewards associated to working activities. An increased level of stress at work can induce individuals to anticipate retirement, given the same financial incentives: leaving the labour force is likely to be perceived as a relief. At the same time, according to stress process theory (Pearlin et al., 2005), a continuous exposure to higher levels of stress can have a cumulative effect on the allostatic load, which refers to the burden placed on the organism and its biological functions in response to hardship and demand (Seeman et al., 2001). The greater the allostatic load, the more difficult it becomes for bodily systems to function optimally, a mechanism through which a stressful working life can lead to mental health deterioration. Anticipating retirement can therefore block the stress proliferation mechanism and thus reduce depression scores at the time of retirement. The data at hand allows to partially test the stress channel as responsible for the mental health improvement: we look

at the differential effect on mental health of retiring in a region severely hit by the crisis, compared to retiring in regions that were not so hardly affected by the economic downturn.

We build a binary indicator which takes on value 1 if a given European region (NUTS 1) in a given year was hit by the economic crisis and zero vice versa. Time series on GDP per capita from Eurostat regional data (EUROSTAT, 2015) are used for this purpose.[3] We apply the Hodrick-Prescott filter (Hodrick and Prescott, 1997) to each of these time series and split log of GDP per capita into trend and cycle components. Our binary indicator *hit by the crisis* takes on value 1 if the cycle component is negative. Figure 5 shows the log of GDP per capita cyclical component from Hodrick-Prescott filter - years 2004-2013. Regions below the horizontal line are defined as hit by the crisis: their percentage by year ranges from 0 % (year 2007) to 100 % (2009). As stated earlier, we assume that the economic crisis affects the relationship between retirement and depression through its effect on working conditions in proximity of retirement. Therefore, we attribute the value of our indicator *hit by the crisis* to each individual in the sample retiring in the considered period according to her self-reported retirement year.

We run all our estimations separately by gender, in order to account for the different reservation wages and labour supply elasticity. Finally, to explore the heterogeneity of the effects across different types of occupation, we estimate equation (1) separately for blue and white-collar (ex) workers.

3. RESULTS

Table 4 reports the regression results for males in our baseline specification. Columns (1) and (2) report the Fixed Effect estimates without controlling for the potential endogeneity due to

[3] Data on GDP per capita by NUTS 2 regions from EUROSTAT (EUROSTAT, 2015) are currently (15/6/2015) available up to year 2011. We impute GDP for the last two years included in the data (i.e. 2012-2013) to each region using information on national GDP per capita (EUROSTAT, 2015b) and rescaling it according to the GDP growth experienced in 2011 by each NUTS relative to that experienced in the same year by the country to which the region belongs.

transitory unobservable shocks affecting both mental health and retirement decision. Retiring reduces depression scores: moving into retirement leads to a significant 0.2 reduction in depression scores. Difference in the parameter of interest between the specification with only a quadratic in age and the specification with the full set of controls is not statistically significant. As we discussed in the previous section, there are good reasons to expect that controlling for time invariant unobservables with fixed effects only, is not enough to solve the endogeneity issue: it may well be that sudden unobservable changes in individual environment affect both retirement decision and mental health. Columns (3) and (4) report results from IV estimates where we include the two instruments discussed above, namely a dummy for having reached statutory retirement age, and a dummy for being age eligible to early retirement, given the country and cohort of each individual. Looking at first stage results both instruments are informative and the F test is highly significant for both men and women. The Hansen J test shows that the over-identification restrictions are valid. Instruments are therefore valid and informative, while Hausman tests rejects the null of FE against the IV specification.

The IV estimates are larger than the non-instrumented ones, pointing to an attenuation bias due to endogeneity. Nevertheless, differences between comparable specifications (e.g. between column (2) and column (4)) are not statistically different. The overall picture is that retirement significantly reduces the depression score: moving into retirement leads to a 0.24 reduction in EURO-D. This is not a small effect: EURO-D is measured on a scale from 0 to 12 but the mean in the estimation sample is 1.8 which means the marginal effect of retirement is equivalent to about a 13 percent reduction in the EURO-D score compared to its mean value.

As regards other regressors, results are in line with expectations: age induces a significant increase in depressions scores, and the marginal effect is increasing with age: it is 1.17 EURO-D points at age 50, and rises to 1.52 at age 65. Being widowed increases depression scores, while having grandchildren reduces depression. The Poterba Venti and Wise health index is

significant and negative. Interpreting results regarding income is not straightforward. Unfortunately, SHARE respondents were asked about gross incomes in the first wave and about net incomes later on. Therefore, we have to resort to a relative measure of income rather than an absolute in order to use all the available waves, as it was done e.g. in Kalwij et al (2014). The fixed effect already accounts for differences in permanent income, therefore income quintiles highlight the effect of transitory changes in relative income within a country. Wave dummies are never significant, reducing the threat to validity due to potential non-random attrition.

Table 5 reports the results of the same set of specifications for women. Here the picture is quite different: in fixed effect models, where we control for time invariant unobservables, the effect of retirement on mental health is significant and not different from what we observe for males. However, significance is lost once we control for endogenous unobservable transitory shocks with instrumental variables. Our interpretation of these differences is the following: given the period of observations, the most likely transitory unobservable shock to men and women close to retirement has been a worsening of working conditions due to the economic crisis. Working during a profound downturn as many European regions experienced after 2008 means facing a substantially higher unemployment risk, and thus being exposed to higher job insecurity. The crisis can also affect other dimensions of working life. For example, a recent report by Eurofund (2013) showed that the recent economic downturn affected working conditions leading to less choice for workers, wage freezes and wage cuts, greater work intensity, deterioration of work–life balance increasing stress at work, greater risk of harassment/bullying. These are all factors that may well affect workers' level of stress. The link between stress and depression is well documented in the medical literature. As we mention in the previous section, the increased stress associated to worse working conditions can induce workers to retire as soon as possible, and those who retire look back to their last years of

working feeling a relief and thus experiencing a reduction in depression scores. A stress related motivation for the positive effect of retirement on mental health is coherent with the gender difference we observe: men have lower labour supply elasticity and a higher labour market attachment than women, who are more likely not to work full time and have a higher marginal utility from leisure (See Alesina et al, 2005 for a review of the literature on labour supply elasticity estimates). Therefore, work related stress is more likely to affect men than women and the relief associated to retirement is stronger for males.

If the driving force is the increased work stress associated with the economic crisis, 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 GDP time series at regional level, we define a dummy indicator, which takes value 1 if an individual experienced a particularly bad economic slowdown in a given year, 0 otherwise. Table 6 reports estimates of the same set of specifications, where we added among the regressors the crisis dummy, and the interaction of this dummy with the retirement dummy[4]. The top panel reports the results for males: from column (1), we can see that working in a particularly bad economic environment has no statistically significant effect on mental health. On the other hand, also retiring per se has no significant effect. What drives the mental health improvement we observed in table 4 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 of depression scores of 0.29 EURO-D points, which accounts for about 16 percentage reduction with respect to the mean value. Looking at results for women, again the picture is completely different: retiring has a positive effect only if transitory unobservable effects are not accounted for, i.e.

[4] In this set of estimates, the sample size reduces significantly because we kept in the sample only individuals who were working the first time they are observed in the panel. This choice is due to the fact that we link the macroeconomic variables to SHARE individual data, based on the retirement year and therefore we need to consider only individuals whose retirement year lies in our sample period.

in the FE specifications, and in these specifications, there is no differential effect of being severely hit by the crisis. Once the transitory unobservable shocks are accounted for with instrumental variables, the effect of retiring during the crisis on mental health disappears completely.

When we divide the sample according individuals' type of occupation, distinguishing between white (WC) and blue collars (BC)[5] (see columns 2 and 3), we find that the positive effect of retirement on depression in the regions hit by the economic crisis only holds for blue-collars. Again, as the lower panel of the Table 6 suggests, there is no significant effect of retirement during the economic crisis for women.

Overall, our results indicate that the positive effect of retirement on mental health is stronger in the regions most affected by the crisis and is particularly visible for individuals employed in blue-collar occupations. This finding may be explained by the worsening of working conditions experienced in particular by blue-collar unskilled workers as a consequence of the economic downturn. For instance, job insecurity has considerably increased after the crisis (see e.g. Eurofound, 2013). Data from the 2005 and 2010 waves of the European Working Conditions Surveys (EWCS)[6] also show that such an increase has been heterogeneous across occupations. Job insecurity, as measured by perceived chance of losing one's job within the next six months, increased by 33 percent for blue-collar while it increased by 16 percent for white collars. Given that job insecurity can have important implications for psychosocial health and well-being, this evidence may help explaining why retiring in bad economic conditions improves individuals' mental health, particularly in the types of occupations that are more exposed to the crisis.

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[5] Following to the classification of Eurofund, the respondents with ISCO occupational codes from 1 to 5 are defined as white collars and those with ISCO from 6 to 9 as blue collars (see http://www.eurofound.europa.eu/surveys/ewcs/2005/classification)

[6] The EWCS is a cross-country survey that assesses and quantifies working conditions of both employees and the self- employed across Europe on a harmonised basis.

4. DISCUSSION AND CONCLUSION

This paper studied the causal effect on depression of retiring during the great recession. We addressed the potential endogeneity of retirement decision to mental health by applying a fixed-effect instrumental-variable approach, where we used country-specific early and full retirement ages as instruments for retirement behavior.

We find convincing evidence that for men retirement reduces the symptoms of depression. By exploiting time and geographical heterogeneity in the intensity of the economic crisis, we showed that this effect is mainly due to individuals who retired in regions and periods most affected by the crisis. This effect is not homogenous across all types of occupations, but it is rather concentrated in blue collar occupations.

We argue that our results help shedding some light on the mechanisms behind the relationship between retirement and depression. We suggest that retirement may affect mental health through its effect on stress. In fact, if retirement is perceived as a relief from job-related stress, our finding of a differential effect of retirement according to the intensity of the crisis and to the type of occupation, is completely 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 retirement age should account for the implied costs in terms of mental health, in particular in periods of economic downturn.

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TABLES AND FIGURES

Table 1: Pattern of individual panel participation

Wave 1	Wave 2	Wave 4	Wave 5	N indviduals	Percent
X	X	X	X	6928	23.48
	X	X	X	3306	10.42
X	X	X		1743	5.49
X	X		X	878	2.77
X		X	X	779	2.45
		X	X	10866	34.23
X	X			3773	11.89
	X	X		982	3.09
	X		X	479	1.51
X			X	395	1.24
X		X		357	1.12

Table 2 – summary statistics of the variables included in the analysis - by gender

	Female					Males				
	Obs	*Mean*	*sd*	*Min*	*Max*	*Obs*	*Mean*	*sd*	*Min*	*Max*
Eurod	40692	2.55	2.21	0	12	40383	1.79	1.91	0	12
retired	40692	0.65	0.48	0	1	40383	0.66	0.47	0	1
Age	40692	66	10	50	102	40383	66	9	50	101
Marital status										
Married or in couple	40692	0.62	0.48	0	1	40383	0.80	0.40	0	1
Divorced	40692	0.11	0.32	0	1	40383	0.08	0.26	0	1
Widowed	40692	0.20	0.40	0	1	40383	0.06	0.24	0	1
Never married	40692	0.06	0.24	0	1	40383	0.06	0.23	0	1
Number of grandchildren	40692	2.78	3.12	0	22	40383	2.49	2.96	0	20
Income quinitiles										
1	40692	0.24	0.42	0	1	40383	0.16	0.36	0	1
2	40692	0.21	0.41	0	1	40383	0.20	0.40	0	1
3	40692	0.19	0.39	0	1	40383	0.21	0.41	0	1
4	40692	0.18	0.39	0	1	40383	0.22	0.41	0	1
5	40692	0.18	0.38	0	1	40383	0.21	0.41	0	1
Health index	40692	50.17	28.79	1	100	40383	56.35	26.62	1	100

Table 3: 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

Table 4: Number of depression symptoms and retirement, males

	(1) FE	(2) FE	(3) IV-FE	(4) IV-FE
Retired	-0.203***	-0.188***	-0.356***	-0.241**
	(0.0410)	(0.0402)	(0.115)	(0.113)
Age	-0.231***	-0.167***	-0.189***	-0.140***
	(0.0376)	(0.0371)	(0.041)	(0.041)
Age2	0.0197***	0.0146***	0.017***	0.013***
	(0.00155)	(0.00157)	(0.002)	(0.002)
Mar st: Divorced		0.207		0.177
		(0.140)		(0.142)
Mar st: Widowed		0.887***		0.869***
		(0.0913)		(0.095)
Mar st: Never married		-0.396		-0.399
		(0.267)		(0.265)
N° of grandchildren		-0.0186**		-0.022***
		(0.00782)		(0.008)
H income quint: 2		-0.0845**		-0.089***
		(0.0333)		(0.034)
H income quint: 3		-0.0121		-0.014
		(0.0350)		(0.036)
H income quint: 4		0.0568		0.053
		(0.0367)		(0.037)
H income quint: 5		0.0179		0.023
		(0.0390)		(0.040)
Health index		-0.0165***		-0.016***
		(0.000530)		(0.001)
Wave 2	0.0197	-0.00401	0.001	-0.017
	(0.0807)	(0.0789)	(0.082)	(0.080)
Wave 3	0.126	0.0873	0.104	0.065
	(0.211)	(0.207)	(0.215)	(0.211)
Wave 4	0.0791	-0.0357	0.046	-0.066
	(0.273)	(0.267)	(0.278)	(0.272)
First stage				
Normal retirement age			0.251***	0.251***
			(0.006)	(0.006)
Early retirement age			0.151***	0.150***
			(0.006)	(0.006)
Observations	41419	40383	40255	39284
Number of id	15929	15671	15766	15473
Hansen J			0.126	0.472
P-value			0.7225	0.4919
Weak identification Cragg-Donald Wald F statistic)			1777.04***	1714.06***

Notes*:* 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 1st wave and the 1st quintile of the household income. Standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1

Table 5: Number of depression symptoms and retirement, females

VARIABLES	(1) FE	(2) FE	(3) IV-FE	(4) IV-FE
Retired	-0.222***	-0.203***	-0.187*	-0.097
	(0.045)	(0.045)	(0.104)	(0.105)
Age	-0.221***	-0.179***	-0.237***	-0.221***
	(0.041)	(0.041)	(0.045)	(0.045)
Age2	0.018***	0.015***	0.018***	0.016***
	(0.002)	(0.002)	(0.002)	(0.002)
Mar st: Divorced		0.287*		0.326*
		(0.169)		(0.170)
Mar st: Widowed		0.704***		0.681***
		(0.078)		(0.081)
Mar st: Never married		0.176		0.030
		(0.391)		(0.413)
N° of grandchildren		0.011		0.012
		(0.009)		(0.009)
H income quint: 2		-0.068**		-0.073**
		(0.034)		(0.036)
H income quint: 3		-0.011		-0.004
		(0.038)		(0.039)
H income quint: 4		-0.009		-0.028
		(0.041)		(0.042)
H income quint: 5		0.016		0.008
		(0.044)		(0.045)
Health index		-0.015***		-0.015***
		(0.001)		(0.001)
Wave 2	-0.083	-0.090	-0.043	-0.037
	(0.088)	(0.087)	(0.091)	(0.090)
Wave 3	0.053	0.004	0.181	0.169
	(0.232)	(0.231)	(0.240)	(0.239)
Wave 4	-0.018	-0.140	0.149	0.080
	(0.300)	(0.298)	(0.310)	(0.309)
First stage				
Normal retirement age			0.370***	0.367***
			(0.006)	(0.006)
Early retirement age			0.091***	0.094***
			(0.006)	(0.006)
Observations	42203	40692	40439	39028
Number of id	16360	16069	16073	15730
Hansen J			4.415	4.696
P-value			0.0356	0.0302
Weak identification Cragg-Donald Wald F statistic)			2852.85***	2679.94***

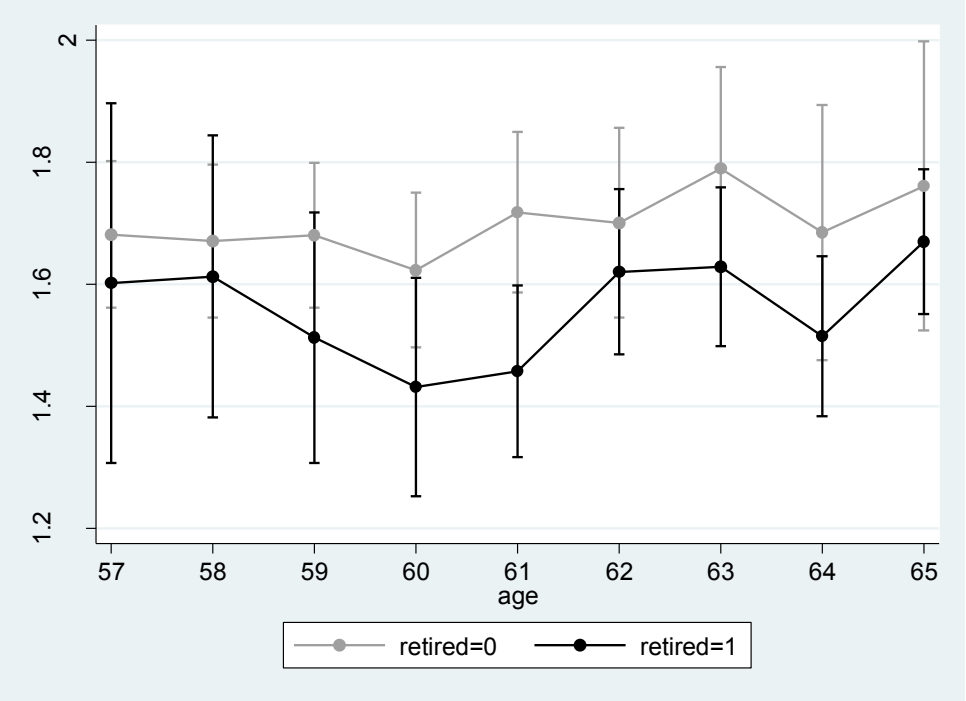
Notes*:* FE estimates (columns (1) and (2)) and IV-FE estimates (columns (3) and (4)) for the number of depression symptoms. The set of instruments includes the statutory normal and early retirement ages. The omitted categories are "married or in a couple", being interviewed in the 1st wave and the 1st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

Table 6: Number of depression symptoms and retirement, in specifications with economic crisis

	(1)	(2)	(3)
VARIABLES	**All sample**	**Blue collars**	**White collars**
		Males	
Retired	0.0229	-0.0130	0.0371
	(0.196)	(0.304)	(0.264)
Retired x hit by crisis	-0.290*	-0.627**	-0.204
	(0.153)	(0.281)	(0.188)
Hit by the crisis	0.0169	0.0459	0.0329
	(0.0613)	(0.112)	(0.0744)
Observations	14741	5052	9348
Number of id	5659	2143	3619
Sargan J stat.	2.303	0.975	0.493
P-value	0.316	0.614	0.781
Weak identification Cragg-Donald Wald F statistic)	166.278***	81.043***	82.916***
		Females	
Retired	-0.0632	0.341	-0.185
	(0.192)	(0.496)	(0.218)
Retired x hit by the crisis	0.198	0.00155	0.318
	(0.182)	(0.471)	(0.204)
Hit by the crisis	-0.00629	0.127	-0.0430
	(0.0718)	(0.190)	(0.0792)
Observations	14344	2573	11352
Number of id	5506	1094	4398
Sargan J stat.	5.786	7.162	0.914
P-value	0.055	0.027	0.633
Weak identification Cragg-Donald Wald F statistic)	268.114***	40.273***	207.496***

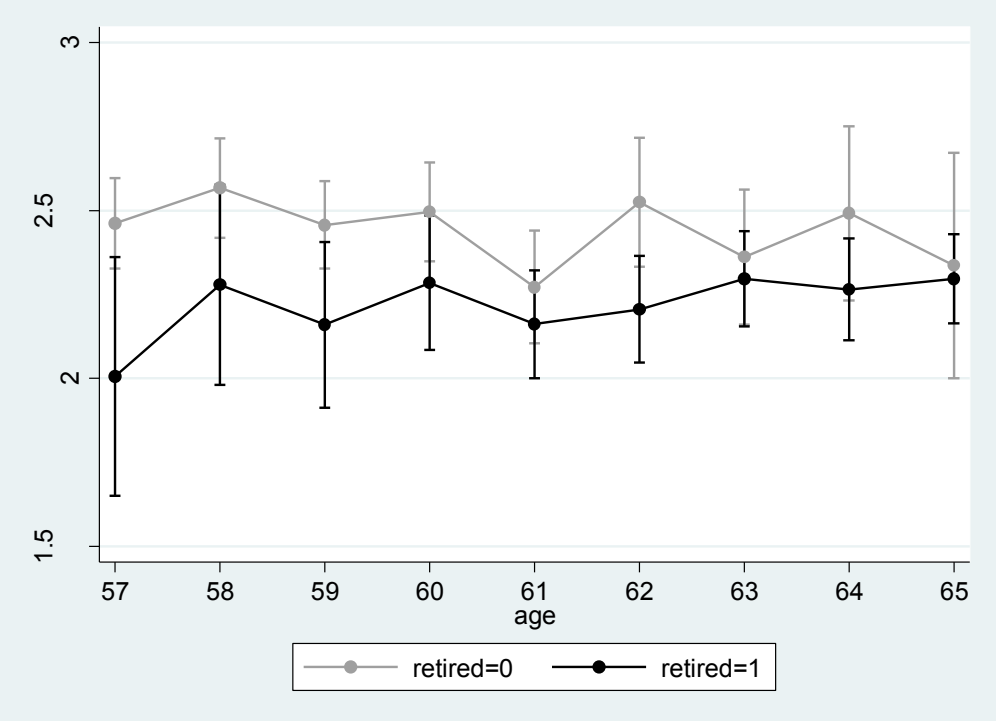
Notes*:* IV-FE estimates in a specification with the economic crisis for the overall sample (column (1)) and for blue-collar (column (2)) and white-collar (column (3)) sub-samples, for males (upper panel) and females (lower panel). The dummy hit by the crisis takes value 1 if the cyclical component of the Hodrick-Prescott filter applied to the log of GDP per capita for the period 2004-2013 is negative and value 0 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 quintiles, wave; number of grandchildren, health index. Standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1

Figure 1 – EURO-D predictive margins with 95% Cis: Males



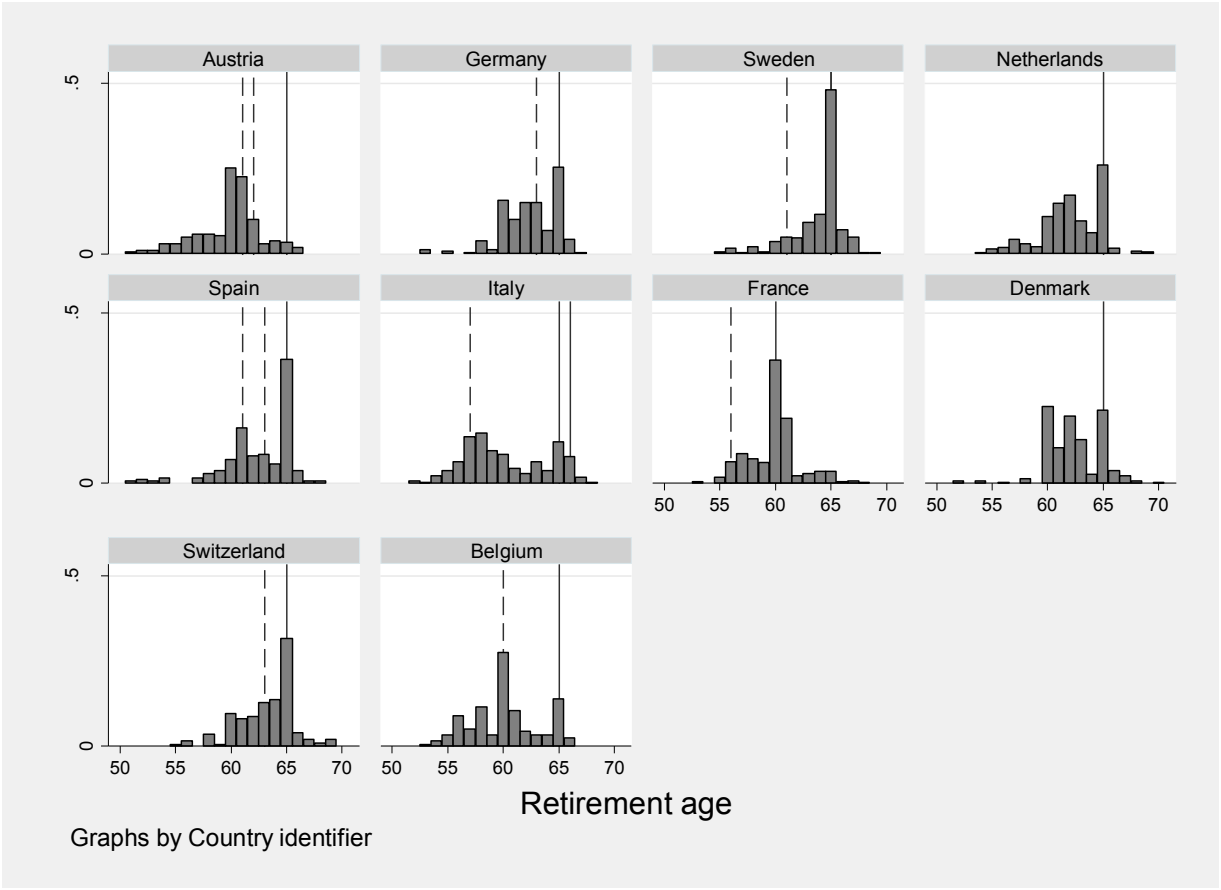
Notes*:* Linear predictions from Fixed Effects models where EURO-D score is explained by a full set of age dummies interacted with the dummy retired. Standard Errors are computed using the delta method

Figure 2 - EURO-D predictive margins with 95% Cis: Females



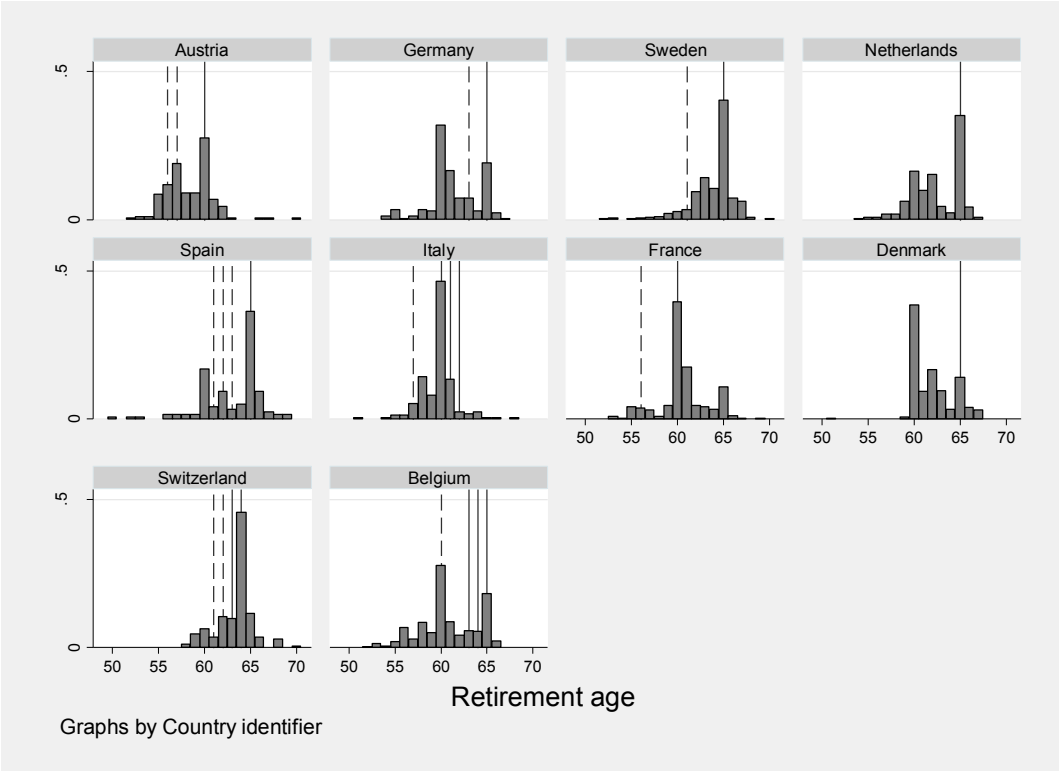
Notes*:* Linear predictions from Fixed Effects models where EURO-D score is explained by a full set of age dummies interacted with the dummy retired. Standard Errors are computed using the delta method

Figure 3 – Retirement age distribution, old age and early retirement eligibility rules: 2004-2013 - Males



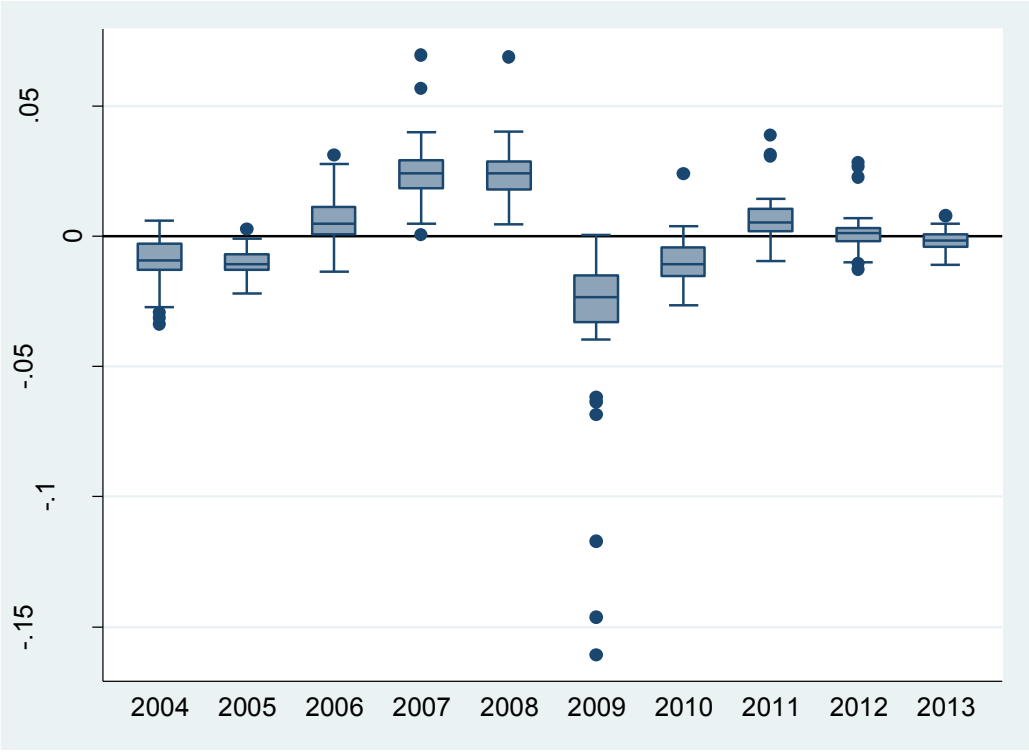
Notes: solid lines = old age retirement age(s); dashed line = early retirement age(s). Early retirement conditions have been tightened dramatically in Italy: for this country we only report the earliest retirement age

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



Notes: solid lines = old age retirement age(s); dashed line = early retirement age(s). Early retirement conditions have been tightened dramatically in Italy: for this country we only report the earliest retirement age

Figure 5 – log GDP per capita cyclical component from Hodrick-Prescott filter: years 2004-2013



Source: our computations on EUROSTAT (2015) GDP regional data