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The effect of work disability on the job involvement of older workers

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ABSTRACT

This paper analyzes the effect of work disability on the job involvement of workers aged 50–65 living in Europe. We elicit a measure of job involvement from a question asking respondents to think about their job and declare whether they would like to retire as early as they can. We exploit objective health indicators and anchoring vignettes to enhance the comparability across individuals of work disability self-assessments. Individuals' evaluations of their health-related work limitations are found to be mildly affected by justification bias but to depend on individual heterogeneity in reporting behaviour. Work disability significantly reduces the job involvement of workers. After controlling for individual fixed-effects and an extensive set of time-varying covariates, moving from the first to the third quartile of the work disability distribution is associated with a 8% increase (4 percentage points) in the probability of desiring to retire as soon as possible. The effect is larger for blue-collar workers. Justification bias and heterogeneity in reporting behaviour do not alter the magnitude of these effects.[1]

1. Introduction

The changes in the demographic structure of the population have triggered a season of reforms in many developed countries to guarantee the sustainability of pension systems. The main results of these reforms have been the lengthening

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of working life due to the progressive increase in the statutory retirement age and the penalization or removal of early retirement pathways (Gruber and Wise, 1999).

The real economic inclusion of older workers does not only depend on the extension of their working careers but also on their actual job attitudes. Individuals formally at work, but who are waiting for retirement as a relief from their current occupation, are unlikely to be fully involved in their jobs. This negative attitude might limit the productivity at the workplace and future investments in human capital accumulation. One of the main risks challenging the job involvement of an ageing workforce is related to health as the occurrence of work limiting health problems increases as people age (Currie and Madrian, 1999 and O'Donnell et al., 2015). In this paper, we quantify the effect of work disability on job involvement in a sample of working individuals aged 50–65 and living in eight European countries. Data are drawn from the Survey of Health, Ageing and Retirement in Europe (SHARE).

Individuals' attitudes towards their job are typically described by work quality and job satisfaction indicators, which have been proved to be determinants of mental health, actual labour market decisions and productivity (Clark, 2001; Dirlam and Zheng, 2017; Bellet et al., 2020 and Böckerman and Ilmakunnas, 2020). In our work, we instead focus on the broader notion of job involvement, which we define as the attractiveness of current job as compared with the retirement scenario.[2] Job involvement is measured in our data by exploiting the answers to the question "*Thinking about your present job, would you like to retire as early as you can from this job?*", which is asked to all respondents at work regardless of their actual eligibility to retirement routes. The retirement intentions elicited from this question result from a comparison between the values placed by individuals on their job and on a forecast of their well-being during retirement years. This question does not allow retrieving a subjective prediction of the timing of retirement, which is the focus of a related research strand briefly surveyed in the next section. Instead, we argue it complements the information about expected timing of retirement by providing a subjective assessment of workers' current job involvement, which is expected to be low for individuals desiring to retire as soon as possible from their job.

Job satisfaction and our measure of job involvement are two related but distinct concepts. Siegrist et al. (2007) and Dal Bianco et al. (2015) based on SHARE data show that individuals experiencing adverse work quality conditions are more likely to wish to retire as soon as possible. Nevertheless, the comparison between current job and the retirement scenario is also influenced by other characteristics, such as family ties, social networks, alternative uses of time, private and Social Security wealth. The attractiveness of the retirement scenario might counterbalance positive working conditions, increasing the desire to retire as soon as possible of workers satisfied with their job. If the retirement scenario is instead less appealing, individuals might have a milder desire to retire, even if they are facing adverse working conditions. In our sample, amongst workers who are satisfied with their job, 43% still would like to retire as early as they can. Vice versa, 28% of workers who are not satisfied with their job do not have a desire to retire as soon as possible. Sec. 5 will provide further empirical evidence showing that in our sample job satisfaction and our measure of job involvement are correlated but do not measure the same concept.

Analysing our measure of job involvement amongst older workers is also important to understand differences in their actual behaviour. Individuals considering retirement as more attractive than current job are likely to experience a working arrangement that does not motivate them to invest in training to acquire new skills and increase their productivity. They might also be more prone to retire earlier or, if they do not retire, they might have a stronger propensity to change their job in the future. Sec. 5 will document that these conjectures find an empirical support in our sample. All these situations might signal a discomfort of workers at the workplace. Despite at work, individuals do not feel actually involved in a job that is not stimulating their interests. They are likely to develop a passive attitude in carrying out their job duties that does not represent a real inclusion in the production process.

In this paper, we analyze to what extent job involvement is influenced by the presence of work disability problems. Health status is a dimension of paramount importance for mature workers as it clearly affects the amount and the type of work they can carry out and it typically deteriorates with age. In particular, instead of considering a general measure of health (McGarry, 2004) we will focus on the concept of work disability. Any given health problem translates in work disability only if it limits individuals' capacity in carrying out their job tasks. Work disability is then expected to be informative of actual individuals' capacity of work (Dwyer and Mitchell, 1999). The onset of work disability problems can limit job involvement via different channels. Health-impaired workers might perceive worse working conditions and an augmented disutility of work due to higher stress or physical effort required to manage job tasks effectively. Individuals experiencing poor health might also be less prone to undertake human capital investments to improve their skills and make them aligned with the technological and organizational change undergoing at their workplace (Lumsdaine and Mitchell, 1999). This is expected to lower workers' future earnings and to marginalize their role in the firm. Even if firms are able to adapt job contents to the health limitations of employees and preserve their work quality, work disability might make the retirement scenario more attractive. Indeed, work limiting health problems might reduce individuals' time horizon and increase the annualized consumption available from current wealth during retirement years (Disney et al., 2006). A reduced time horizon might also increase the current and future marginal utility of non-labour market activities and lead to a stronger preference for

[2] The psychological and sociological literature (Lorence and Mortimer, 1985, Reeve and Smith, 2001 and Lambert et al, 2016) has defined job involvement in terms of the importance of job among the interests and the activities of individuals' life. Although establishing an exact match between our definition of job involvement and the one adopted in this literature is not among our purposes, it is worth stressing that, consistently with our study, this literature points out that job involvement is a multi-faceted factor that is not solely determined by job attributes.

retirement in order to have more time for leisure, family ties and social networks. Although we expect that work disability reduces job involvement, assessing the magnitude of this effect remains an empirical issue.

The SHARE questionnaire is valuable for our research question since it asks respondents to self-assess the presence of health problems or impairments limiting the amount or kind of work they can do. The estimates of the effect of health on retirement behaviour have been proved sensitive to the measure of health used (Kerkhofs et al., 1999; French and Jones, 2017). Work disability self-assessments are widely used since they summarize in a single measure all the aspects individuals consider relevant to determine their health. However, there are several concerns challenging the interpretation of health self-assessment differentials as genuine differences in health status and questioning its exogeneity in our framework.

First, work disability is an inherently multidimensional concept that might arise as a result of a variety of health limitations and its severity depends on the type of job tasks carried out by workers. Different individuals might have different beliefs about the health dimensions to consider when assessing their own work disability and how to aggregate them to provide an overall evaluation. Second, health self-assessments are affected by individual heterogeneity in reporting styles (see for instance, Lindeboom and van Doorslaer, 2004, Kapteyn et al., 2007 and Angelini et al., 2011 and 2012). Individuals might provide diverse evaluations of the same underlying health level because they base their reporting behaviour on different benchmarks due to, for instance, heterogeneity in their age, socioeconomic condition or in the prevailing norms in their reference group. Finally, justification bias might be an issue, as pointed out in several papers analysing actual retirement decisions. Individuals might strategically overstate their health problems to rationalize their exit from the labour market and lead to an upward bias in the health effect on retirement (Kerkhofs and Lindeboom, 1995; Lindeboom and Kerkhofs, 2009; O'Donnell et al., 2015). Justification bias constitutes a state-dependent reporting behaviour that varies with the labour market status (Lindeboom and Kerkhofs, 2009). When looking at retirement expectations of individuals currently at work, the justification bias might be still a concern since individuals might exaggerate health limitations to explain their plan to retire earlier. Dwyer and Mitchell (1999), McGarry (2004) and Ilmakunnas and Ilmakunnas (2018) model retirement expectations and discuss how their data and econometric frameworks are suited to tackle this concern. In the specific context of this paper, justification bias remains a threat for the identification of the effect of health. Individuals might use health-related work limitations to rationalize their desire to retire as soon as possible.

We exploit SHARE data to adjust work disability self-evaluations and enhance their cross-individual comparability by accounting for measurement error and justification bias. Following Bound et al. (1999), we base our work disability measures on an extensive set of more objective health indicators less sensitive to measurement error and justification bias. Further, we extend this approach and implement an anchoring vignette methodology (see King et al., 2004 and Kapteyn et al., 2007) to control for the presence of heterogeneity in reporting styles in individuals' self-evaluations correlated with their observable characteristics. Building upon Lindeboom and Kerkhofs (2009) and Kapteyn et al. (2011), we control for the presence of state dependent reporting behaviour in our framework by allowing the way individuals report their health-related work limitations to vary with their desire to retire as soon possible. To the best of our knowledge, this is the first paper that uses objective health indicators and vignette information to produce work disability measures enhancing cross-individual comparability in a retirement intention analysis.

We find that work disability significantly reduces job involvement. An increase in the work disability level from the first to the third quartile of its empirical distribution is associated with a 8% increase in the probability of desiring to retire as soon as possible in our sample. Our findings propose mild evidence of justification bias in work disability self-assessments, which are nevertheless found to be significantly affected by individual heterogeneity in other sources of reporting behaviour. However, justification bias and measurement error do not alter the magnitude of the health effect in the retirement intention equations. The paper proceeds as follows. Sec. 2 offers a brief overview of related works on the effect of health on retirement intentions. In Sec. 3, we present the econometric specifications used to derive our measures of work disability and assess how it affects the probability of desiring to retire as soon as possible. Sec. 4 describes the data used in our analyzes and Sec. 5 presents the work disability measures and job involvement indicator we consider. Sec. 6 summarizes the main findings. Sec. 7 concludes.

2. Health and retirement intentions – A literature review

Our paper refers to the literature investigating how retirement intentions vary with health. The contributions in this literature define retirement intentions as the expected retirement age of workers or their subjective probability of working after some age threshold describing typical eligibility requirements for early or statutory retirement. These measures abstract from current job working conditions and provide a subjective forecast of the timing with which individuals plan to exit from the labour market. Our work instead considers a measure of retirement intentions that is elicited by explicitly asking respondents to think about their current job and declare whether they would like to retire as soon as possible from it. This question does not capture individuals' opinions about the number of years they will still spend at work but focuses on the attractiveness of current job as compared with the retirement scenario. Our paper will add to this literature by assessing whether health-related work limitations predict retirement intentions described by this alternative indicator.

A first research strand in the retirement intention literature models health by using individuals' self-evaluations. Chan and Stevens (2004) use the first four waves of the Health and Retirement Study (HRS) and measure the retirement intentions of employed individuals by considering their self-rated probabilities of working full-time after age 62 and 65.

They find that individuals who self-rate their current health as fair or poor are significantly less likely to plan to work full time after the two age thresholds considered. Mermin et al. (2007) exploit HRS and consistently find that these retirement intention indicators are also negatively affected by the presence of health-related work limitations. Cobb-Clark and Stillman (2009) analyze the determinants of expected retirement age of Australians based on the waves 1 and 3 of the survey Household, Income and Labour Dynamics in Australia. Using a first-difference estimator, they find cross-gender heterogeneity in the health effect. Whereas men retirement plans result to be unaffected by changes in health, women adjust retirement age according to variations in their health conditions and in those of their partners. Women expected retirement age reduces for those who experience an adverse health shock and for those with healthier partners. Coppola and Wilke (2014) analyze the impact on expected retirement age produced by a pension system reform implemented in Germany in 2007 that raises the statutory retirement age from 65 to 67. They draw data from the survey Saving and Old-Age Provision in Germany and implement a difference-in-differences approach that controls for time-varying covariates including self-reported health, which does not turn out to be a relevant predictor of retirement expectations for the individuals in the sample.

Other papers instead explicitly take into account measurement error and endogeneity concerns related with the use of self-reported health in an equation modelling retirement intentions. The “true” health of an individual is difficult to observe due to the complex nature of this concept. Health self-assessments might suffer from measurement error threatening their comparability across individuals and potentially bringing about a downward bias in the estimated effect. Health self-assessments might also be affected by “justification bias”: individuals who plan to retire earlier might misclassify strategically their self-evaluations and declare a poor health status to justify their intention to retire. Dwyer and Mitchell (1999) draw data from the first wave of HRS and estimate the effect of health on expected retirement age by linear regression models. They use subjective indicators, namely self-rated health and self-rated presence of work limitations, but also deal with the measurement error and justification bias potentially affecting these measures, which are then replaced with more objective health indicators derived from respondents’ assessments about the presence of specific health problems. Examples of these objective indicators are the presence of chronic and acute diseases, as well as problems with carrying out daily living activities. These indicators are less likely to be affected by measurement error and reporting behaviour as they focus on precise aspects of individuals’ health.[3] Their results show that individuals experiencing poor health conditions plan to retire earlier. These findings are also confirmed if both subjective and objective health indicators are allowed to be affected by measurement error and instrumented with exogenous health determinants arguably uncorrelated with reporting behaviour, such as parents’ health and mortality. McGarry (2004) uses the first two waves of HRS to assess the effect of health on individuals’ subjective probability of working after age 62. She shows that the effect of self-assessed health on retirement intention remains statistically significant when controlling for subjective survival probability of living to age 85, the presence of activity limitations and the presence of diagnosed health conditions. This pattern also holds when looking at the predictive power of health variations over time to explain the changes over time in retirement intentions, net of other time-varying factors like earnings and wealth. Ilmakunnas and Ilmakunnas (2018) investigate the effect of health on expected retirement age by using the waves 2003 and 2008 of the Quality of Life Surveys of Statistics Finland supplemented by register data.[4] In line with previous studies, they show that better self-assessed health is associated with delayed expected retirement plans. To tackle the potential measurement error and justification bias concerns of self-reported health, they exploit more objective indicators (pains and chronic illnesses). In their specifications, these indicators replace self-reported health within the set of right-hand-side variables or alternatively instrument it. The sign and the significance of the effect of health on expected retirement age are confirmed. The comparison between OLS and IV estimates suggests a downward bias produced by the measurement error in self-reported health.

In our paper, we will share the concerns discussed in Dwyer and Mitchell (1999), McGarry (2004) and Ilmakunnas and Ilmakunnas (2018) about the measurement error and justification bias potentially compromising the interpretation of the results produced when using health self-assessments. We will overcome these issues by exploiting objective health indicators and anchoring vignettes to develop work disability measures enhancing the comparability across individuals of work disability self-assessments. The next section will describe the framework in which our work disability measures are constructed.

3. Econometric specifications

3.1. A model of work disability measurement

The utilization of individuals’ self-assessments of work disability as measure of their actual work disability can be criticized by the fact that different individuals might consider different sets of health dimensions when asked to self-rate their own work disability. Moreover, individuals might have different beliefs about how to aggregate the dimensions considered when providing the overall self-assessment. Moreover, as discussed in the Introduction, self-evaluations are affected by justification bias if individuals use them to justify their desire to retire as soon as possible. Bound et al. (1999) specify a latent

[3] However, this kind of objective health indicators is not immune to biases (see for instance, Baker et al., 2004 and van Ooijen et al., 2015).

[4] Their analysis develops a comprehensive framework investigating the effect of health on actual retirement age, expected retirement age and the difference between them, which is interpreted as an expectation error.

variable model to develop a health index based on more objective self-assessed measures and individual-specific characteristics. These more objective health self-evaluations are expected to be less exposed to reporting heterogeneity and justification bias than overall self-assessments because they focus on precise health domains and presence of specific health problems. We will follow the approach by Bound et al. (1999) to derive an alternative work disability measure with enhanced comparability across individuals.

We assume that the unobserved true level of work disability of a generic individual $i = 1,\ldots,N$ at a given time period $t = 1,\ldots,T$ is

$$wd_{it}^{*} = \boldsymbol{Z}_{it}'\boldsymbol{\gamma}_1 + \boldsymbol{X}_{it}'\boldsymbol{\gamma}_2 + \varepsilon_{it} \tag{1}$$

where $\boldsymbol{Z}_{it}$ is a vector including objective health indicators, $\boldsymbol{X}_{it}$ is a vector including individual socioeconomic characteristics and ε_{it} is a random component. Including individual characteristics other than objective health indicators is in line with Bound et al. (1999) and Disney et al. (2006). This allows work disability to reflect individual heterogeneity not captured by the objective health measures considered but, for instance, related with age, job characteristics and economic resources.

Let swd_{it}^{*} be a continuous latent variable indicating how the generic individual i self-perceives her own work disability at time t and assume that $swd_{it}^{*} = wd_{it}^{*} + u_{it}$ or equivalently that

$$swd_{it}^{*} = \boldsymbol{Z}_{it}'\boldsymbol{\gamma}_1 + \boldsymbol{X}_{it}'\boldsymbol{\gamma}_2 + \varepsilon_{it} + u_{it} \tag{2}$$

where u_{it} is an idiosyncratic error component.

The observed work disability self-assessments of an individual i at time t, denoted by swd_{it}, is the discrete counterpart of swd_{it}^{*}. The variable swd_{it} is defined as a discrete ordered outcome taking on value $j = 1,\ldots,J$. More precisely, swd_{it}^{*} maps the observed self-reported health status swd_{it} as follows

$$swd_{it} = j \text{ if } \alpha_{j-1} < swd_{it}^{*} \leq \alpha_j \tag{3}$$

where $-\infty = \alpha_0 < \alpha_1 < \ldots < \alpha_J = \infty$.

Note that imposing invariant thresholds α_j implies that all individuals set the same criteria to report a given work disability level.

If $\varepsilon_{it} + u_{it}$ follows a standard normal distribution, we can obtain valid estimates for $\boldsymbol{\gamma}_1$ and $\boldsymbol{\gamma}_2$ by running a standard ordered probit regression of swd_{it} on $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$. The linear prediction produced by this regression is a measure of work disability based on individuals' achievements with respect to a battery of dimensions ($\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$) aggregated according to the $\boldsymbol{\gamma}_1$ and $\boldsymbol{\gamma}_2$ vectors of coefficients.

This approach develops a measure of work disability characterized by an improved cross-individual comparability as compared with raw self-assessments. Indeed, this measure is based on a set of dimensions invariant across individuals, which are the covariates in the ordered probit regressions, and aggregates them according to the same weighting scheme, which is described by the coefficients $\boldsymbol{\gamma}_1$and $\boldsymbol{\gamma}_2$. The linear prediction of the ordered probit regression is the first work disability measure we consider in our analysis.

To account for the fact that the coefficients $\boldsymbol{\gamma}_1$and $\boldsymbol{\gamma}_2$ are not observed a priori but estimated based on a sample, we follow Little and Rubin (1987) and produce a set of M imputations for the linear prediction of the ordered probit regression model. We estimate the ordered probit regression model by maximum likelihood to obtain the estimated vector of coefficients $\hat{\boldsymbol{\gamma}}_1$and $\hat{\boldsymbol{\gamma}}_2$ and their estimated variance and covariance matrix, $\hat{\boldsymbol{\Sigma}}$. Then, we repeat M times the following procedure. First, we draw a set of coefficients $\boldsymbol{\gamma}_1^m$ and $\boldsymbol{\gamma}_2^m$, $m = 1\ldots, M$, from a multivariate normal distribution having $\hat{\boldsymbol{\gamma}}_1$and $\hat{\boldsymbol{\gamma}}_2$ as mean vector and the matrix $\hat{\boldsymbol{\Sigma}}$ as variance and covariance matrix. Second, we impute our work disability measure by taking $\boldsymbol{Z}_{it}'\boldsymbol{\gamma}_1^m + \boldsymbol{X}_{it}'\boldsymbol{\gamma}_2^m$.

The traditional approach by Bound et al. (1999) previously described assumes that the cut-off points (thresholds) α_j, $j = 0\ldots J$, according to which respondents rate their perceived work disability levels, are invariant in the population. This statistical assumption neglects the presence of heterogeneity in the reporting styles in individuals' self-assessments. Let us consider two individuals who share identical work disability levels but have different beliefs about the severity of the health problems that can actually limit the amount of work they can do. These two individuals might provide different answers to the work disability self-evaluation question due to individual heterogeneity in the cut-off points applied to map their perceived work disability swd^{*} in the discrete ordered outcome swd. Reporting behaviour in work-disability self-assessments has been shown to vary with country of residence, health and socioeconomic characteristics (see for instance Kapteyn et al., 2007 and Angelini et al., 2011 and 2012). Self-assessments might reflect both genuine variability in health levels as well as variability in the reporting styles individuals use to rate the extent of their work limiting health problems.

If heterogeneity in reporting styles is systematically related with the individual health indicators and socioeconomic characteristics $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$, the work disability measure based on the standard ordered probit regression previously specified is not reliable. The coefficients on the variables in the vectors $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$ included in these specifications capture a combination of true work disability differentials and the correlation of these variables with individual reporting behaviour. We then extend the approach proposed by Bound et al. (1999) by replacing the standard ordered probit model with the Hopit model, which is a generalized ordered probit specification that formally relaxes the assumption of cut-off points invariant in the population.

The Eq. (2) is replaced with

$$swd_{it}^{*} = \boldsymbol{Z}_{it}'\tilde{\boldsymbol{\gamma}}_1 + \boldsymbol{X}_{it}'\tilde{\boldsymbol{\gamma}}_2 + \varepsilon_{it} + \tilde{u}_{it} \tag{4}$$

where $\tilde{u}_{it}$ is an idiosyncratic error component, such that $\varepsilon_{it} + \tilde{u}_{it}$ follows a standard normal distribution. The latent variable swd^{*}_{it} is mapped into the discrete observed outcome swd_{it} by individual-specific cut-off points,

$$swd_{it} = j \text{ if } \tilde{\alpha}_{ij-1} < swd^{*}_{it} \leq \tilde{\alpha}_{ij} \tag{5}$$

where $\tilde{\alpha}_{i0} = -\infty, \tilde{\alpha}_{iJ} = \infty,\ \tilde{\alpha}_{i1} = \boldsymbol{Z}'_{it}\boldsymbol{\delta}^{1}_{1} + \boldsymbol{X}'_{it}\boldsymbol{\delta}^{1}_{2} + itr_{it}\delta^{1}_{3},\ \tilde{\alpha}_{ij} = \tilde{\alpha}_{ij-1} + \exp(\boldsymbol{Z}'_{it}\boldsymbol{\delta}^{j}_{1} + \boldsymbol{X}'_{it}\boldsymbol{\delta}^{j}_{2} + itr_{it}\delta^{j}_{3}),\ j = 2,\ldots,J\text{-}1.$[5] The coefficients in the right-hand-side of Eq. (4) reflect how the true level of work disability depends on the covariates in the vectors $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$, net of their correlation with the reporting behaviour described by the corresponding coefficients in the threshold equations. In the Hopit model, the thresholds are allowed to correlate with the health and socioeconomic characteristics $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$ as well as with our job involvement indicator itr_{it} that takes on value 1 for respondents desiring to retire as soon as possible and 0 otherwise.

Following Lindeboom and Kerkhofs (2009) and Kapteyn et al. (2011), we control for state-dependent reporting and allow thresholds to vary with our job involvement indicator. If justification bias is an issue, conditional on a true level of work disability, respondents might decide to report different self-evaluations according to their desire to retire as soon as possible. Testing the significance of the coefficients δ^{j}_{3}, $j = 1,2,\ldots,J\text{-}1$ reveals whether reporting behaviour is correlated with job involvement, giving rise to justification bias in work disability assessments.

The estimation of the Hopit model requires the availability of anchoring vignettes, which are a survey instrument consisting in asking respondents to rate the work disability of hypothetical persons briefly described in vignettes kept constant across respondents. As long as the situations described in the vignettes are perceived by respondents in the same way (*vignette equivalence hypothesis*), differences in respondents' evaluations of vignettes reflect heterogeneity in their reporting behaviour when assessing work disability problems. If respondents use the same reporting styles to evaluate vignettes and their own conditions (*response consistency hypothesis*), work disability self-assessments and vignette evaluations provide the information required to separately identify the correlation of the vectors of covariates $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$ with the perceived work disability swd^{*}_{it} as well as the reporting styles described by the cut-off points $\tilde{\alpha}_{ij}$, $j = 0,\ldots,J$. The Hopit model is estimated by maximum likelihood.[6]

The linear prediction for swd^{*}_{it} produced by the Hopit specification is the second work disability measure used in our analysis. Following the same procedure described before, we produce a set of M imputations of this second work disability measure.

3.2. Modelling how job involvement varies with work disability

As argued in the Introduction, we interpret the intention to retire (*itr*) as soon as possible of workers as a signal of their low job involvement. We specify the following linear probability model

$$itr_{it} = \beta_0 + \beta_1 \widehat{wd}_{it} + \boldsymbol{\beta}_2 \boldsymbol{X}_{it} + c_i + e_{it} \tag{6}$$

The variable $\widehat{wd}_{it}$ is the measure of work disability obtained by standard ordered probit and Hopit regressions. To ease the comparability of the results produced by these two alternative measures, we rescale them to vary between zero and one. The vector $\boldsymbol{X}_{it}$ includes time-invariant and time-varying individual characteristics introduced in Sec. 3.1.[7] The error term in the equation is decomposed in a time-invariant component c_i and a time varying component e_{it}.

We are interested in β_1, the coefficient measuring how retirement intentions vary with work disability. Ordinary Least Squares (OLS) estimates of β_1 yield unbiased results only if the orthogonality condition between the explanatory variables and the error term is satisfied. This assumption might be quite restrictive as it rules out that retirement intentions and work disability might be jointly determined by unobserved common factors. Estimating Eq. (6) using fixed-effects methods for panel data allows relaxing this orthogonality condition. The fixed-effects estimation allows the time-invariant unobserved heterogeneity c_i to be arbitrarily correlated with the explanatory variables in the model. The theory offers several examples of individual-specific and time-invariant characteristics included in the c_i component. For instance, permanent income might influence work disability through health investments, but it might also affect retirement intentions via wealth accumulation, since higher levels of financial and real wealth can be used to anticipate the labour market exit and finance consumption during retirement years. In addition, higher time-discount rates of future periods might strengthen retirement intentions because agents assign a higher value to leisure but also affect how individuals evaluate the opportunity cost of health investments.

Finally, it is worth noting that the fixed-effects estimation supports the identification of work disability differentials in retirement intentions net of reporting behaviour heterogeneity. Indeed, typically unobserved factors, such as taste for work, might affect both individuals' desire to retire and their willingness to adopt strategic reporting behaviour in work disability

[5] In a given threshold equation $j = 2,3,4$ the generic coefficient δ^{j} is proportional to the effect of the corresponding variable on the distance between thresholds j and *j-1*. The overall effect of an explanatory variable on threshold $j = 2,3,4$ can be calculated by considering the nonlinear structure of the threshold equation. Full details of the calculation are reported in Section A1 of Online Appendix A.

[6] Section A1 in Online Appendix A offers a more detailed explanation of the Hopit model.

[7] The notion of job involvement is intrinsically subjective and this reduces the concerns for mismeasurement. However, if there is heterogeneity in social norms across countries and socioeconomic groups about the opportunity of reporting a desire to retire as soon as possible, this is captured by the control factors $\boldsymbol{X}_{it}$ used in our specifications.

self-evaluations to rationalize their intentions. As long as these unobserved factors fall into the c_i component, the fixed-effects estimation allows to properly account for this latter source of endogeneity.

Given the presence of multiple imputations for our work disability measure $\widehat{wd}_{it}$, the estimation of Eq. (6) will be carried out by using the multiple imputation techniques developed by Little and Rubin (1987), which exploit the variability within and between each set of imputations.

4. Data

Our sample has been drawn from the first two waves of SHARE collected in 2004/5 and 2006/7 respectively. It includes 11,194 observations referring to 8098 individuals aged 50–65 at work and living in Germany, Sweden, the Netherlands, Spain, Italy, France, Greece and Belgium.[8] In both waves, in addition to the standard CAPI (Computer Assisted Personal Interview) questionnaire, a subsample of respondents have filled-in a paper and pencil questionnaire that asks to self-assess their own work disability by answering to the following question "*Do you have any impairment or health problem that limits the kind or amount of work you can do?*" according to the following discrete ordered scale "*1. None*" (65.16%), "*2. Mild*" (22.85%), "*3. Moderate*" (9.38%), "*4. Severe*" (2.26%), "*5. Extreme*" (0.36%). The respondents in this subsample are also asked to rate the work disability of hypothetical individuals described in anchoring vignettes. Respondents are presented with short descriptions of hypothetical individuals including some key aspects of their health and socioeconomic conditions relevant to rate their work disability. For instance, one of the anchoring vignette administered in SHARE is the following "*Kevin suffers from back pain that causes stiffness in his back especially at work but is relieved with low doses of medication. He does not have any pains other than this generalized discomfort. How much is Kevin limited in the kind or amount of work he could do?*" ("*1. None, 2. Mild, 3. Moderate, 4. Severe, 5. Extreme*"). Notice that the scale used to rate individuals in anchoring vignettes is the same as the one used by respondents to self-assess their own work disability.[9] The vignette sample consists of 1685 observations referring to 1482 individuals.

5. Measuring job involvement and work disability

5.1. Job involvement

The standard CAPI interview of SHARE asks respondents at work about their retirement intentions by the question "*Thinking about your present job, would you like to retire as early as you can from this job?*" (Yes/No). This is our job involvement indicator. Individuals who would like to retire as soon as possible from their job (46% in our sample) arguably have a low level of job involvement. As outlined in the Introduction, our measures of job involvement and job satisfaction are two related but distinct concepts. The SHARE questionnaire asks respondents to provide a general job satisfaction assessment ("*All things considered I am satisfied with my job.*"). Answers are collected according to a 4-point Likert-scale ("*1. Strongly agree, 2. Agree, 3. Disagree, 4. Strongly disagree*"). The tabulation of these variables provides a clear support to the hypothesis that they capture two different concepts. In our sample 91% are satisfied with their job (they select the categories 1 or 2 in the Likert scale), this percentage about doubles that of respondents who would retire as soon as possible. We also run logit regressions to assess to what extent heterogeneity in job satisfaction is able to replicate the heterogeneity in our job involvement indicator. The set of covariates in this specification includes one dummy for each answering category in the job satisfaction question and controls for gender, age, country of residence and wave. Respondents who are more satisfied with their job have a significantly weaker desire to retire as soon as possible. However, only 64% of the predictions for the dependent variable match its observed value. This descriptive analysis shows that retirement intentions and job satisfaction are strongly intertwined but also signals that these two concepts cannot be considered as substitutes as a substantial part of the desire to retire is explained by other factors than job satisfaction. This evidence is consistent with the hypothesis that our job involvement indicator also captures determinants of the attractiveness of the retirement scenario unrelated with job satisfaction. These results hold when considering more detailed dimensions of job quality. Indeed, the SHARE questionnaire includes an extensive battery of quality of work questions that ask respondent to evaluate to what extent their job is physically demanding and stressful, whether they have freedom in managing their work duties and the possibility of developing new skills, whether they receive adequate support in difficult situations and the recognition they deserve for their efforts. Finally, respondents are asked to assess whether their wages are adequate, their job security is poor and whether they have prospects for job advancements and promotions. We re-run our logit specification replacing the job satisfaction evaluation with the set of indicators of job quality listed above.[10] Although experiencing better job quality conditions is generally associated with a lower desire to retire as soon as possible, this regression confirms that controlling for job quality indicators leaves unexplained a substantial part of the heterogeneity of job involvement in our sample.[11]

[8] Results are quantitatively the same if we select individuals aged 50-60. Results are available upon request.

[9] Section A1 in Online Appendix A reports the full text of all the vignettes used in our analysis. Section A3 shows summary statistics for respondents' work disability self-assessments and vignette evaluations.

[10] Standard errors are clustered at the individual level. Results of these logit regressions are available upon request.

[11] The predictions are consistent with observed values of the dependent variable only in the 65% of cases.

Table 1
Current and future probabilities of attending a training course, retiring and changing job and their variations (Δ) associated with the current desire to retire as soon as possible.

	(1) **Training**	(2) **Retirement**	(3) **Change job**
Current time	13.10	/	/
Δ	-5.25^{***}	/	/
After two years	14.75	12.35	6.84
Δ	-4.79^{***}	7.45^{***}	1.50^{**}
After four years	23.88	27.18	15.48
Δ	-6.27^{***}	11.83^{***}	0.6
After six years	23.83	43.57	26.83
Δ	-5.83^{***}	13.30^{***}	1.63
After eight years	20.69	56.65	29.9
Δ	-4.81^{***}	11.25^{***}	3.53^{*}

Note: Variations are calculated by estimating logit regressions controlling for gender, age, country of residence and wave. Future probabilities of training and changing job and corresponding variations are calculated conditional on being employed. Standard errors are clustered at the individual level.

We also want to assess to what extent current and future behaviour of workers are associated with our job involvement indicator. Individuals with poor job involvement might have developed a passive attitude in executing their job duties and do not find incentives in undertaking training decisions to improve their skills. We use logit models to check how the current and future probability of attending a training course varies with the current intention to retire.[12] We exploit the longitudinal dimension of SHARE and use up to the fifth wave of the survey to construct a time horizon covering eight years. The first column of Table 1 summarizes the results. The probability of attending a training course at the time of the interview for respondents in our sample is 13.10% and it reduces by 5.25 percentage points for those desiring to retire as soon as possible. Conditional on remaining employed, the probability of being involved in a training course after two years is 14.75% and it shrinks by about one third for those who currently would leave their job as soon as possible. Analogous results are found in the following points in time considered in our exercise. This evidence supports the hypothesis that our job involvement indicator captures a passive job attitude that leads to a lower propensity to invest in human capital accumulation. We also tested whether those who prefer to retire as soon as possible instead of carrying out their job actually have a higher probability of retiring earlier. Although they are forced to remain employed longer by pension reforms, these individuals might do not find any interest in prolonging their working career and prefer to opt out of the labour market at the earliest convenience. The results in the second column of Table 1 show that this is the case. A stronger desire to retire as soon as possible is associated with a higher probability of exiting the labour market. Finally, we assess whether the discomfort with respect to current job captured by our job involvement indicator correlates with the probability of changing job for those who remain in the labour force. The third column of Table 1 shows that the probability of changing job within the next two years is significantly higher (+22%) for those with a current desire to retire as soon as possible.

5.2. Generating work disability measures

This section explains the estimation of the work disability measures based on ordered probit and Hopit models. These specifications can be estimated only on the vignette sample as only in this sample we have work disability self-evaluations available. Following the approach proposed by Bound et al. (1999), the key information needed to enhance the comparability of work disability self-assessments consists of an extensive battery of objective health indicators expected to be less sensitive (or not sensitive at all) to individuals' reporting behaviour. Using the notation previously introduced, these health indicators fall in the vector $\boldsymbol{Z}_{it}$, which includes: the number of limitations with Activities of Daily Living (ADLs) and with Instrumental Activities of Daily Living (IADLs); Body Mass Index (BMI); the presence of reduced muscle strength (*sarcopenia*, see Bertoni et al., 2018) based on a hand-grip strength test; whether the respondent does not perform the grip strength test (*grip_miss*), which is expected to capture problems with the use of hands; the number of chronic diseases, mobility limitations and symptoms; the Global Activity Limitation Indicator (GALI); an indicator of mental health (based on the EURO-D depression scale) and a set of cognitive functioning indicators, namely orientation in time and numeracy. We also define the vector $\boldsymbol{X}_{it}$, collecting a set of covariates describing socioeconomic characteristics: country of residence, gender, age, having a cohabiting partner, number of children, number of grandchildren, education (ISCED levels), job characteristics (blue/white collar), type of occupation (self-employed/employees), sector (public/private), earnings quartiles, wealth quartiles and the number of years to and since the minimum retirement age in place in the country of residence at the time of the interview.[13] Finally, the vector $\boldsymbol{X}_{it}$ includes a dummy to discriminate between interviews conducted in the wave 1 and the wave 2 of SHARE. The sample averages of the explanatory variables used to derive our work disability measures are displayed in Table B1 in Online Appendix B.

[12] Likewise before, specifications control for gender, age, country of residence and a wave dummy. Standard errors are clustered at the individual level.

[13] These indicators are designed to capture penalizations for early retirees and incentives to prolong the working career set by the Social Security systems.

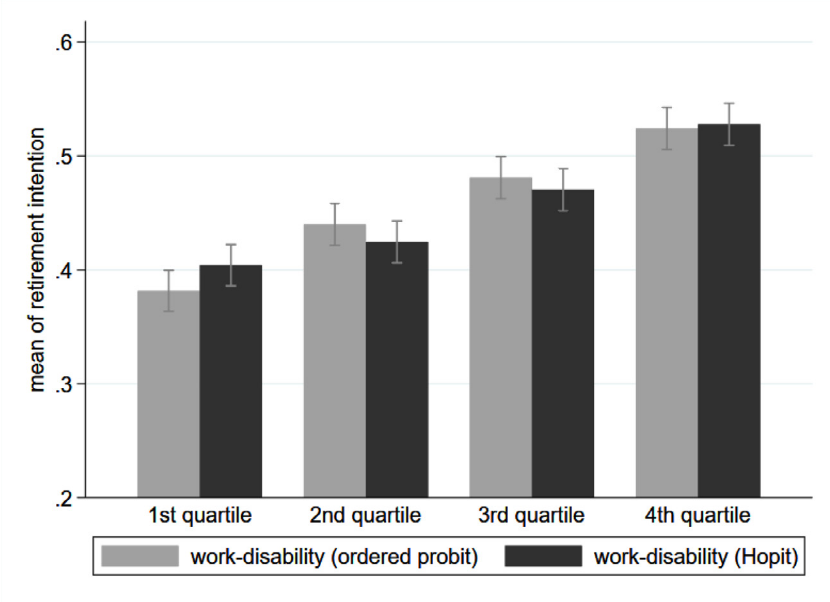

Fig. 1. Probability of desiring to retire as soon as possible by work disability quartiles.

The first column of Table 2 reports the results of the estimation of Eq. (2) by standard ordered probit techniques.[14] The linear prediction of the outcome of this equation is our first work disability measure. Work-limiting health problems are more prevalent when our battery of objective health indicators $\boldsymbol{Z}_{it}$ detects poor health episodes.

The following columns of Table 2 summarize the results of the Hopit regression, which formally relaxes the assumption of invariant cut-off points in the population and allows them to be dependent on the covariates in the vectors $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$ and on our job involvement indicator (*itr*). The second column collects the coefficient estimates of Eq. (4). Our second measure of work disability is based on the linear prediction of the outcome of this equation.[15] In order to account for the randomness embedded in any estimation exercise and produced by the fact that any estimator is a random variable following a distribution, we follow the procedure described in Sec. 3 to produce $M = 10$ multiple imputations of both our work disability measures. Even if the ordered probit and Hopit regressions can be run only on the vignette sample, the work disability measures based on the coefficient estimates they produce can be taken for the overall sample since all the explanatory variables in the models are defined based on the information collected by the standard CAPI interview administered to all SHARE respondents. As a result, we produce multiply imputed sets of work disability measures for all the 11,194 observations (8098 individuals) in our main sample.

Fig. 1 shows how the probability of desiring to retire as soon as possible varies across the quartiles of the empirical distribution of the work disability measures based on the ordered probit and the Hopit models. According to both measures, the proportion of individuals desiring to retire as soon as possible from work is positively correlated with work disability. Amongst those in the first quartile of the work disability distribution, about 40% want to retire as soon as possible. The percentage increases to 53% in the fourth quartile. Poor health episodes clearly propose as a powerful determinant of retirement intentions. Sec. 6 will assess whether work disability remains a predictor of retirement intentions once we control for household and individual heterogeneity.

The popularity of anchoring vignettes as well-established survey instrument to control for reporting styles heterogeneity is witnessed by their extensive utilization. Their advantages and disadvantages, including the discussion and tests of the validity of the vignette equivalence and response consistency hypotheses, are the focus of several contributions, which we review in the Sec. A2 of Online Appendix A. Nevertheless, for our purposes it is reassuring that, as documented in the Sec. A3 of this appendix, our findings are confirmed in a robustness analysis of the Hopit model estimation that varies the vignette choice to assess its sensitivity to deviations from the vignette equivalence hypothesis.

In our analysis, we consider only vignettes that are common in the first and second waves of SHARE. We check that our results are confirmed if we include also additional vignettes administered only in wave 1. We also test the robustness of

[14] The full set of estimation results is in Table B2 in Online Appendix B.

[15] The achievement of the convergence in the maximum likelihood estimation of the Hopit model is facilitated by defining explanatory variables ranging over comparable supports. To this end, we standardize the *BMI* by subtracting its sample average and dividing it by its standard deviation, we subtract 50 to *age* and we divide *nchild* and *ngrchild* by 10. For sake of comparability, these transformations are applied both in the ordered probit and the Hopit estimation.

Table 2
Estimation results of the ordered probit and Hopit models used to derive the work disability measures.

Variables	(1) Ordered probit	(2) Hopit	(3) Hopit–Threshold equations $\tilde{\alpha}_1$	(4) $\tilde{\alpha}_2$	(5) $\tilde{\alpha}_3$	(6) $\tilde{\alpha}_4$
ADL	0.308**	0.325**	0.137	−0.020	−0.158	0.011
	(0.137)	(0.149)	(0.099)	(0.110)	(0.116)	(0.185)
IADL	0.186	0.314**	0.036	0.098	0.093	−0.485**
	(0.143)	(0.160)	(0.105)	(0.086)	(0.089)	(0.196)
BMI	0.044	0.043	0.004	−0.001	−0.032*	−0.083**
	(0.031)	(0.036)	(0.022)	(0.020)	(0.020)	(0.036)
Sarcopenia	0.287*	0.300	0.087	−0.070	−0.033	0.065
	(0.172)	(0.192)	(0.119)	(0.120)	(0.111)	(0.192)
grip_miss	0.267	−0.028	−0.440***	0.325***	−0.061	−0.053
	(0.166)	(0.194)	(0.146)	(0.102)	(0.101)	(0.190)
Chronic	0.116***	0.113***	−0.000	−0.009	−0.014	0.014
	(0.034)	(0.038)	(0.023)	(0.021)	(0.021)	(0.040)
Mobility	0.134***	0.115***	−0.044*	0.023	0.044**	0.047
	(0.032)	(0.036)	(0.024)	(0.021)	(0.019)	(0.041)
Symptoms	0.133***	0.131***	−0.017	0.022	0.027	0.004
	(0.032)	(0.036)	(0.022)	(0.020)	(0.020)	(0.041)
GALI	0.741***	0.853***	0.171***	−0.118**	−0.086*	−0.018
	(0.077)	(0.088)	(0.054)	(0.049)	(0.050)	(0.095)
Depressed	0.405***	0.359***	−0.035	−0.023	0.023	0.040
	(0.089)	(0.100)	(0.061)	(0.057)	(0.055)	(0.113)
Orientation	−0.007	0.129	0.197**	−0.131*	−0.040	0.344**
	(0.125)	(0.144)	(0.096)	(0.076)	(0.073)	(0.148)
Numeracy	0.007	0.038	0.027	0.011	0.052	0.014
	(0.075)	(0.084)	(0.048)	(0.043)	(0.043)	(0.085)
Female	−0.012	0.127	0.153***	0.027	−0.107**	−0.067
	(0.077)	(0.087)	(0.051)	(0.047)	(0.046)	(0.087)
Age	0.054	0.059	−0.009	0.073***	−0.069***	0.040
	(0.042)	(0.046)	(0.024)	(0.024)	(0.022)	(0.051)
Couple	0.083	0.088	0.029	−0.050	0.024	0.014
	(0.091)	(0.102)	(0.057)	(0.052)	(0.054)	(0.098)
nchild	0.515*	0.057	−0.530**	0.365**	−0.112	0.173
	(0.294)	(0.340)	(0.211)	(0.183)	(0.182)	(0.361)
ngrchild	−0.563**	−0.682**	−0.115	−0.170	0.357**	0.062
	(0.248)	(0.283)	(0.170)	(0.156)	(0.144)	(0.238)
isced_34	−0.080	−0.031	0.032	0.051	−0.073	0.131
	(0.085)	(0.096)	(0.058)	(0.052)	(0.050)	(0.098)
isced_56	−0.067	0.053	0.134**	−0.061	−0.018	−0.021
	(0.096)	(0.108)	(0.063)	(0.058)	(0.056)	(0.106)
Public	−0.075	−0.146*	−0.072	0.021	−0.033	0.041
	(0.077)	(0.087)	(0.050)	(0.045)	(0.045)	(0.084)
Self employed	0.123	0.184*	0.051	−0.028	0.133**	0.012
	(0.093)	(0.105)	(0.061)	(0.055)	(0.054)	(0.122)
Blue collar	0.217***	0.246***	0.031	0.008	−0.040	0.109
	(0.080)	(0.091)	(0.054)	(0.049)	(0.047)	(0.095)
itr			−0.076*	−0.030	−0.027	0.094
			(0.040)	(0.039)	(0.039)	(0.077)
earnings_q2	−0.100	−0.136	−0.067	0.020	0.163***	−0.009
	(0.090)	(0.102)	(0.060)	(0.055)	(0.054)	(0.105)
earnings_q3	−0.137	−0.152	−0.029	0.014	0.110**	0.105
	(0.086)	(0.097)	(0.057)	(0.052)	(0.052)	(0.102)
earnings_q4	−0.272***	−0.333***	−0.077	0.065	0.048	−0.006
	(0.093)	(0.106)	(0.061)	(0.056)	(0.055)	(0.104)
totwealth_q2	0.097	0.107	−0.001	0.045	0.009	−0.100
	(0.095)	(0.107)	(0.061)	(0.058)	(0.057)	(0.103)
totwealth_q3	0.070	−0.005	−0.092	0.064	0.108*	0.067
	(0.100)	(0.113)	(0.065)	(0.060)	(0.058)	(0.108)
totwealth_q4	0.196*	0.166	−0.065	0.077	0.109*	0.002
	(0.102)	(0.115)	(0.066)	(0.061)	(0.059)	(0.115)
yearsto_mra	0.627	0.725	0.048	0.532**	−0.615***	0.619
	(0.416)	(0.464)	(0.244)	(0.239)	(0.223)	(0.510)
yearssince_mra	−0.497	−0.412	0.186	−0.558*	0.456	−0.455
	(0.521)	(0.581)	(0.303)	(0.292)	(0.292)	(0.606)
wave2	−0.025	0.186**	0.229***	−0.093**	0.145***	0.290***
	(0.072)	(0.083)	(0.051)	(0.045)	(0.043)	(0.080)
Constant			1.740***	−0.631*	0.719**	−1.319*
			(0.609)	(0.343)	(0.324)	(0.715)
Observations	1685	1685				

Note: Specifications control also for country dummies. Standard errors in parentheses are clustered at the individual level. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 3
Counterfactual distributions of work disability self-assessments according to alternative reporting behaviour settings.

%	None	Mild	Moderate	Severe	Extreme
Desire to retire as soon as possible	75.61	19.23	4.69	0.47	0.00
No desire to retire as soon as possible	77.51	18.40	3.68	0.42	0.00
Men	74.72	19.58	5.16	0.53	0.00
Women	78.75	17.03	3.86	0.36	0.00
51 years of age	77.98	15.01	6.71	0.18	0.12
61 years of age	75.43	22.08	2.20	0.30	0.00
ISCED 0–2	74.78	19.82	4.87	0.47	0.06
ISCED 5–6	78.69	16.56	4.39	0.30	0.06
Minimum retirement age reached	76.02	17.21	6.47	0.18	0.12
10 years to minimum retirement age	77.21	20.42	2.02	0.36	0.00

Note: Counterfactual simulations are based on the coefficient estimates of the Hopit model (the full set of results is reported in Table B2 in Online Appendix B). In each of the cases considered other covariates in the model are set to respondents' observed values.

our results to the inclusion of health history information from the third wave of SHARE (SHARELIFE). Again, all the results of our main analysis are confirmed. Details are provided in Sec. A3 of Online Appendix A.

5.3. Justification bias and measurement error in work disability self-assessments

As displayed in columns (3)-(6) of Table 2, reporting behaviour is found to be significantly correlated with individual characteristics. In each of the first three threshold equations ($\tilde{\alpha}_j,\ \ j = 1, 2, 3$) the Wald joint significance tests of the coefficients reject the null that the parameters are equal to zero at the 1% significance level. For the fourth threshold equation, the joint significance test instead does not reject (p-value of 0.16).[16] The Hopit specification allows netting out this heterogeneity in reporting behaviour from the computation of the work disability indicator.

As described in Sec. 3, our Hopit specification allows testing the presence of justification bias. Table 2 shows that the coefficient on the job involvement indicator (*itr*) in the first threshold is negative and statistically significant at the 10% level. This means that the first threshold, which discriminates between absent and mild work disability levels, is significantly lower for individuals with lower job involvement. Holding everything else constant, individuals who desire to retire as soon as possible set significantly looser criteria to report mild instead of absent work disability problems in the discrete scale used in the questionnaire. Given the same self-perceived work disability, individuals with low job involvement are less likely to declare absence of health-related work limitations than similar individuals with high job involvement. We also assessed how job involvement affects the level of the other thresholds in the model. We found that the second and third thresholds are significantly lower for individuals who desire to retire as soon as possible. Everything else constant, a given perceived level of work disability is more likely to be classified as moderate or severe by individuals with a low job involvement. The effect on the fourth threshold is still negative but not statistically significant.[17]

These findings support the hypothesis of the presence of justification bias in our data. We assess the magnitude of its consequences by a counterfactual exercise. Based on the estimated coefficients of the Hopit model and the observed values for the covariates $\boldsymbol{Z}_{it}$ and $\boldsymbol{X}_{it}$, we compute for each individual i at any time t the predicted value of the latent outcome swd^*_{it}. We then predict the discrete outcome swd_{it} by comparing the predictions of swd^*_{it} and the thresholds $\tilde{\alpha}_{ij}$, $i = 1,\ldots, N$ and $j = 0,\ldots, J$. Thresholds are computed under two alternative settings for reporting behaviour. In the former we impose for all individuals a strong desire to retire as soon as possible (itr=1 for all individuals in the sample). In the latter we do the opposite (itr=0 for all individuals in the sample). For all the other covariates in the threshold equations we use the respondents' actual values. By construction, the differences in the distribution of the discrete outcome swd_{it} between the two cases considered are solely related to the different reporting behaviour adopted. Overall, the differences between the two distributions are quite limited. The first two rows of Table 3 shows that when considering the reporting style of individuals desiring to retire as soon as possible, the probability of reporting absence of health-related work limitations drops from 77.51% to 75.61% and, conversely, the probability of having mild limitations increases from 18.40% to 19.23%. This is consistent with the negative and statistically significant coefficient on the dummy *itr* in the first threshold equation. When looking at the remaining part of the distribution, we notice that individuals with a weaker job involvement are

[16] The tests are asymptotically distributed as a χ^2 with 39 degrees of freedom (test statistic for $\tilde{\alpha}_1$ is 105.51, for $\tilde{\alpha}_2$ is 114.21, for $\tilde{\alpha}_3$ is 119.92 and for $\tilde{\alpha}_4$ is 47.88).

[17] As described in Section A1 of Online Appendix A, given the nonlinear component in the second, third and fourth thresholds, the partial effect of job involvement depends on the values taken on by the other covariates. We set them at their sample averages. When testing the significance of the partial effect of interest, we take into account the variance and covariance matrix estimator for the coefficients in the Hopit model.

more likely to declare moderate and severe work limitations. This pattern aligns with the sign of the partial effects of job involvement on the second and third threshold levels previously discussed. However, although we cannot reject the hypothesis of state-dependent reporting styles, justification bias is not found to alter the distribution of work disability self-assessments substantially. Regardless of the reporting behaviour considered, around 25% of individuals declare at least some work disability. Of them, about 80% and 18% declare mild and moderate limitations, respectively.

We carry out the same analysis for other covariates included in the threshold equations to provide a more extensive assessment of the measurement error in self-evaluations produced by individual heterogeneity in reporting behaviour neglected by standard ordered probit models. We first look at the heterogeneity in reporting styles across genders. Table 3 reports the distribution of declared work disability self-assessments when alternatively applying the response behaviour of men and women. The mechanics of the exercise is as in the previous case. Holding everything else constant, when the reporting styles of men is used, the percentage of respondents declaring absence of work disability problems reduces (from 78.75% to 74.72%) and the percentages of mild and moderate disability problems increases (from 17.03% to 19.58% and from 3.86% to 5.16%, respectively). Women response behaviour seems to underrate work disability problems. It is worth stressing that changes in the thresholds discriminating absent, mild and moderate work disability are particularly relevant because most of the respondents in our sample (97%) fall in these three categories. Further, as pointed out by Kapteyn et al. (2011), it is more likely to find differences in the evaluation of the work disability consequences of relatively weak health problems than in the evaluation of those associated with more serious health conditions. We also compute the distribution of self-assessed work disability according to the response behaviour of individuals aged 51 and 61 (10th and 90th percentiles of the age distribution in our sample). When using the reporting behaviour of those aged 61, the probability of declaring mild work disability problems rises from 15.01% to 22.08% and that of declaring moderate work disability reduces from 6.71% to 2.20%. These findings suggest that, as individuals age, self-image and self-esteem reasons might lead them to adopt stricter conditions to declare work disability. Table 3 also compares the reporting styles characterizing individuals with the lowest education level in our sample (ISCED 0–2) and the highest ones (ISCED 5–6). Everything else constant, the probability of declaring absence of work disability problems increases from 74.78% to 78.69% when using the thresholds of the highly educated individuals. This difference is mainly counterbalanced by the reduction of the percentage of individuals declaring mild work disability, which shrinks by about one fifth (from 19.82% to 16.56%) when applying the reporting styles of the highly educated. We finally look at the number of years to the attainment of the minimum retirement age in place in the country of residence at the time of the interview. We compare the distribution of the work disability self-assessments predicted under the hypotheses that all individuals in the sample use the reporting styles characterizing those who have already reached the minimum retirement age and those who need 10 more years to attain this requirement. Table 3 documents that the distribution of the extent of health-related work impairments shifts to the right if we use the response behaviour of individuals who are age eligible for retirement. The probability of reporting mild work disability problems reduces from 20.42% to 17.21% and the probability of declaring moderate problems increases to 2.02% to 6.47%. Everything else constant, individuals in our sample who have attained the minimum retirement age eligibility are more comfortable with declaring at least moderate health-related work limitations.[18]

6. The effect of work disability on job involvement

We analyze the effect of work disability on the probability of experiencing low job involvement by estimating linear probability models.[19] Table 4, columns 1 to 3, reports the results for our specifications based on the work disability measures produced by the ordered probit regression implemented according to the standard approach by Bound et al. (1999).[20] The first column shows the results produced by estimating the linear probability models by OLS. If we look at Eq. (6) in Sec. 3, this means that we are currently considering the time-invariant individual-specific component c_i as uncorrelated with the other explanatory variables in the model. Standard errors are clustered to allow for arbitrary correlation in the error term at the individual level. All the regression analyses are carried out by following the multiple imputation technique in Little and Rubin (1987) to formally account for the variability within and between the sets of imputations produced.

Work disability is an important determinant of job involvement. The coefficient is highly significant and shows that on average passing from the first to the third quartile (i.e. the interquartile range)[21] of the empirical distribution of our work disability measure is associated with a 8 percentage point increase in the probability of desiring to retire to as soon as possible. As reported in Table B4 in Online Appendix B, which contains the full set of results, this variation is comparable to the effect of moving from the highest level of educational attainments (*isced_56*=1) to the lowest level (*isced_02*=1).

Our full sample is longitudinally unbalanced as we include all individuals who have been interviewed in at least one wave. If we select only individuals appearing in both waves, the sample size reduces to 6192 observations (3096 individuals). The second column of Table 4 shows that this sample selection leaves our results quantitatively unaffected. Most importantly, defining a balanced sample is particularly relevant in order to run a fixed-effects (FE) analysis and fully exploit

[18] We did the same analysis considering the number of years past after the attainment of the minimum retirement age but we did not find substantial differences in the predicted distributions.

[19] Table B3 in Online Appendix B reports summary statistics for all the explanatory variables considered.

[20] The full set of estimation results is in Table B4 in Online Appendix B.

[21] The interquartile range of work disability is 0.127 when we use the ordered probit specification and 0.120 for the Hopit specification.

Table 4
Estimation results for the retirement intention model. Work disability measure from the ordered probit model (columns 1 to 3) and from the Hopit model (columns 4 to 6).

Variables	Work disability (ordered probit) (1) OLS unbalanced	(2) OLS balanced	(3) FE	Work disability (hopit) (4) OLS unbalanced	(5) OLS balanced	(6) FE
Work disability	0.636***	0.636***	0.287**	0.680***	0.688***	0.316**
	(0.092)	(0.110)	(0.125)	(0.103)	(0.125)	(0.133)
Observations	11,194	6192	6192	11,194	6192	6192
Individuals	8098	3096	3096	8098	3096	3096

Note: Specifications control for gender, age, presence of a cohabiting partner, education, job characteristics, earnings, wealth, number of children, number of grandchildren, country of residence, a wave dummy and the number of years to and since the minimum retirement age. Standard errors in parentheses are clustered at the individual level. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Estimation is based on 10 sets of multiply-imputed work disability measures combined according to Little and Rubin (1987).

the longitudinal dimension of SHARE. Using panel estimation controlling for individual fixed-effects is more suitable than OLS to account for time-invariant unobserved factors correlated with retirement intentions and the explanatory variables in our specifications, including work disability. The results are reported in the third column of Table 4. Although the point estimate shrinks, work disability remains a significant predictor of retirement intentions. Everything else constant, switching the work disability level from the first to the third quartile is associated with a 4 percentage point increase in the probability of desiring to retire as soon as possible. This effect is sizeable as it accounts for 8% of this probability in the sample. The comparison between OLS and FE estimates suggests that OLS overestimates the effect of interest supporting the hypothesis that the individual heterogeneity summarized by the c_i component includes factors that make individuals both more tempted to retire and more likely to suffer from work limiting health problems. Columns 4 to 6 of Table 4 summarize the results produced by replicating our analysis considering the work disability measure obtained by the Hopit model in order to account for individual heterogeneity in reporting behaviour. Our previous findings are overall confirmed across all the specifications considered and proved not to be driven by reporting bias in work disability self-assessments controlled by the anchoring vignette methodology.

Attrition between waves 1 and 2 of SHARE is a potential concern in our analysis as it might produce biased estimates. We follow Wooldridge (2010) and use inverse probability weighting (IPW) to assess whether this is the case. Specifically, we estimate a probit model for the probability of participating in the second wave of SHARE controlling for a wide set of individual and fieldwork characteristics observed in the first wave. The controls include the set of variables used in the Ordered probit/Hopit models, the outcome variable (*itr*) and specific interviewer's characteristics, such as indicators for rounding behaviour in measurements (Bristle et al., 2019). Results show that, amongst those working in the first wave, individuals in poor health, having a partner, having grandchildren, working in the private sector and interviewed by less diligent interviewers are less likely to participate in the second wave. The inverse of the fitted probabilities are then used as weights in our main specifications. Table B5 in Online Appendix B compares FE estimations with and without weights when using both Ordered probit and Hopit models to construct the work disability measure. Our findings are confirmed. If anything, the point estimates slightly increase.

6.1. Heterogeneous effects

From a policy point of view, it is particularly important to understand the profile of the workers who are exposed to the higher risk of compromising their job involvement as a result of a poor health episode. We allow the relationship of interest to vary along a wide spectrum of individual characteristics. Results are presented in Table 5. First, we look at gender differences. Women are typically more exposed to interrupted working careers and the attractiveness of the retirement scenario might be limited by lower amount of accumulated Social Security Wealth. The consequences of a negative health shock on the desire to retire as soon as possible might be smoother if current pension wealth does not allow maintaining an adequate standard of living during retirement years. Although the sign of the coefficient on the interaction term between work disability and the female dummy supports this conjecture, it is not statistically significant. We might also hypothesize that household and family composition affects the consequences of work disability on retirement intentions since individuals limited in their working career by health-related problems might put a higher value on the retirement option to spend more time with their family. However, we do not detect any heterogeneity with respect to presence of partners, children and grandchildren.

We further assess whether human capital can influence the effect of work disability on retirement intentions. Workers with higher levels of human capital might be more flexible in adapting their skills to offset work-limiting problems. Moreover, individuals with higher human capital levels might be more valuable to their firms, which might be more willing to create a job environment suitable to maintain their productivity, or at least to reduce the negative consequences of

Table 5
Heterogeneity analysis.

X	(1) Female	(2) Couple	(3) nchild	(4) ngrchild	(5) isced_56	(6) earnings_q4	(7) Blue collar	(8) Self-employed
				Work disability (ordered probit)				
Work disability	0.419**	0.527**	0.219	0.214	0.322*	0.389**	0.139	0.286**
	(0.177)	(0.253)	(0.155)	(0.133)	(0.175)	(0.186)	(0.133)	(0.132)
*Work disability*X*	−0.251	−0.299	0.032	0.067	−0.134	−0.340	0.513**	0.002
	(0.228)	(0.270)	(0.047)	(0.052)	(0.231)	(0.240)	(0.247)	(0.268)
X	–	0.097	−0.020	−0.246	–	0.086	−0.074	−0.063
		(0.093)	(0.268)	(0.163)		(0.061)	(0.096)	(0.088)
				Work disability (Hopit)				
Work disability	0.401**	0.589**	0.325*	0.263*	0.351*	0.491**	0.167	0.336**
	(0.185)	(0.264)	(0.175)	(0.141)	(0.186)	(0.200)	(0.142)	(0.139)
*Work disability*X*	−0.160	−0.342	−0.004	0.048	−0.112	−0.417	0.527**	−0.143
	(0.233)	(0.276)	(0.053)	(0.056)	(0.240)	(0.262)	(0.257)	(0.276)
X	–	0.095	0.014	−0.020	–	0.105	−0.062	−0.027
		(0.087)	(0.026)	(0.016)		(0.064)	(0.092)	(0.086)
Observations	6192	6192	6192	6192	6192	6192	6192	6192

Note: Specifications control for age, presence of a cohabiting partner, job characteristics, earnings, wealth, number of children, number of grandchildren, a wave dummy and the number of years to and since the minimum retirement age. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Estimation is based on 10 sets of multiply-imputed work disability measures combined according to Little and Rubin (1987).

their health conditions. We describe human capital by educational attainments and levels of earnings, but do not find any significant heterogeneity in this respect.

Further, we investigate whether work disability differentials in retirement intentions change with job characteristics. Everything else constant, the consequences of a poor health episode might be more severe if individuals carry out more physically demanding jobs (Bound et al., 1995). They might be forced to change their job or at least to change some of the tasks they were required to manage. Blue-collar workers face a higher risk of being more hampered by a health shock. Indeed, if we interact our work disability measures with the blue-collar dummy, we find that the same increase in work disability is associated with a higher increase in the propensity towards retirement for blue-collars. This suggests that firms are reluctant to adapt the job contents of older blue-collar workers to their worsened health conditions due to, for instance, organizational constraints or the short time horizon over which their investments to train older workers to carry out new job tasks can be recouped. Finally, the work disability effects on job involvement do not differ between employees and self-employed.

7. Conclusions

This paper investigates the consequences of work-limiting health problems on the job involvement of older workers. Job involvement is defined as the attractiveness of carrying out the current job as compared with the retirement scenario and it is elicited from a specific survey question in the SHARE questionnaire asking respondents to think about their job and reporting whether they would like to retire as soon as possible from it.

The Social Security reforms of the last twenty years strengthened the financial sustainability of pension systems by increasing retirement age and penalizing early withdrawals from the labour market. However, older individuals, albeit formally at work, are not necessarily either productive or actively involved in the productive process. Work limiting health problems are expected to be amongst the major risk factors challenging the job involvement of older workers because of the sharp increase in their incidence as individuals age. Quantifying the significance and the magnitude of the effect of work disability on the job involvement of older workers becomes of primary importance for policy makers to improve the actual economic inclusion of the older part of the workforce. We draw data from the first two waves of SHARE and base our analysis on a panel sample representative of the population of workers aged 50–65 and living in eight European countries.

Measuring work disability is a complex task since it cannot be directly observed. Our paper extends the approach by Bound et al. (1999) to enhance the comparability across individuals of work disability self-assessments and control for justification bias. We developed two alternative measures of work disability. First, we follow Bound et al. (1999) and use standard ordered probit regressions to generate a work disability measure based on a set of objective health indicators – arguably less exposed to reporting behaviour and justification bias concerns – and individual characteristics. Still, this approach imposes the assumption that reporting behaviour does not vary with individual characteristics, which has been questioned in several studies (Kapteyn et al., 2007 and 2011 and Angelini et al., 2011 and 2012). Our second measure of work disability exploits the anchoring vignettes collected in the first two waves of SHARE to control for reporting heterogeneity. Controlling for the measurement error in self-evaluations produced by individual heterogeneity in response styles is expected to improve the cross-individual comparability of our second measure of work limiting disabilities. Moreover, as long as the incentives to report work disability problems vary with individuals' job involvement, controlling for reporting styles is an additional way to limit the consequences of justification bias in our estimation exercise as job involvement can be included within the set of observable characteristics determining response behaviour (Lindeboom and Kerkhofs, 2009; Kapteyn et al., 2011). In line with previous studies, our findings confirm the presence of individual heterogeneity in the way individuals report their work disability problems. Reporting behaviour varies significantly with job involvement. Those who wish to retire as soon as possible from their job set significantly looser criteria to declare mild, moderate and severe work disability problems. However, the consequences of justification bias do not appear to be dramatic. Counterfactual simulations show that using reporting behaviour characterizing those with or without the desire to retire as soon as possible does not change substantially the distribution of the predicted work disability self-assessments. We also find that reporting behaviour is correlated with a number of other individual characteristics, such as gender, age, education and the number of years still needed to attain the minimum retirement age. Overall, our results suggest that part of the variability in work disability self-assessments does not reflect genuine differences in health-related problems but individual heterogeneity in reporting behaviour.

We then investigate the effect of work disability on job involvement. We estimate fixed-effects linear probability models to analyze how the desire to retire as soon as possible varies with our work disability measures. Our findings strongly indicate that the presence of work limiting health problems significantly increases the probability of older workers of desiring to retire as soon as possible. An increase from the first to the third quartile of the distribution of work disability limitations is associated with a 8% increase (4 percentage points) in the probability of desiring to retire as soon as possible. We also find that the positive effect of work disability on the propensity to retire as soon as possible is higher for blue-collars. Individuals who carry out more physically demanding jobs are also those more at risk of experiencing more dramatic job-related consequences from a negative health shock.

Our job involvement indicator is correlated with actual behaviour of workers signalling a negative attitude toward job and lack of incentives to undertake training investments. Although a causal estimation of the effect of our job involvement indicator on workers behaviour is beyond the scopes of this work, according to the descriptive evidence presented in our paper, moving from the lowest to the highest level of work disability in our sample increases the desire to retire as possible by 31.6 percentage points when work disability is estimated by the Hopit model; this is associated with a reduction of 10% in the probability of doing training activities, an increase of 19% in the probability of retirement and of 7% in the probability of changing job within the next 2 years.

Our findings are overall confirmed regardless of the work disability measure considered. Controlling for measurement error in work disability self-evaluations reflecting individual heterogeneity in reporting behaviour remains an issue to explain the heterogeneity in work disability self-assessments, but it does not turn out to substantially affect the relationship between work disability and job involvement in our specifications. This evidence aligns with the limited extent of justification bias found in our data and with the results presented in Dwyer and Mitchell (1999) and Ilmakunnas and Ilmakunnas (2018). Moreover, they support the argument by McGarry (2004) who points out that, when modelling retirement expectations of individuals currently at work, the scope of justification bias is smaller than when modelling actual retirement decisions.

Overall, our results indicate that public policies aimed at improving health conditions of individuals through prevention, access to health care services and promotion of healthier life styles are not just advisable from a social point of view but can have important economic consequences in preserving the actual economic inclusion of the older workforce. The mere postponement of the exit from the labour market is not enough to guarantee that older workers are actually involved in their job and in the production process. Moreover, firms using human resource management practices fostering employee involvement at the workplace (Cottini et al., 2011; Böckerman et al., 2012) should develop policies that adapt job tasks to the health limitations of older workers. The risk of health impairments increases with age and they might be particularly limiting for individuals involved in more physically demanding tasks, as outlined in our heterogeneity analysis. Health is found to play an important role in determining a positive attitude of workers in carrying out their job tasks and their willingness to work longer. This is a further dimension to consider when assessing the consequences of health and health care investments along the life cycle.

Declarations of interest

None

Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.jebo.2021.10.021.

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