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Old European Couples' Retirement Decisions: the Role of Love and Money*

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Abstract

This study investigates old European couples' retirement choices in order to bridge the gap between the European and the American literature. The typical European family approaching retirement is a dual-earner family: the dataset used in this paper reveals that 78 percent of working males is married, and amongst these 24 percent has a working wife. This results from dramatic changes in the labor force behaviour of both older men and older women after World War II. These trends signal a need of investigating retirement choices at a household level. Using an absolutely new international micro data (SHARE, Survey of Health, Ageing and Retirement in Europe - Release 2), we adopt a duration analysis approach and estimate both single and competing risks models by allowing for a flexible specification with and without unobserved heterogeneity. Our findings show that joint retirement is significantly correlated with education, age, and health status, together with partner's employment status, partner's education and partner's health status. We also perform a sensitivity analysis in order to check whether the results on the correlation of health status are robust to two alternative measures of health which possibly correct for subjectivity and cross-country incomparability of self-reported health. We find that allowing for different exit routes and taking into account partner's characteristics is critical to fully understand joint retirement.

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1 Introduction

Encouraging older people to stay in workforce longer would have a number of positive effects. From a macroeconomic point of view, it would boost labour force growth and help offset the impact of population ageing on economic growth; it would improve public finances via reduced public expenditures associated to early retirement while increasing tax revenues; and it would also help employers by smoothing the pace at which they will have to replace retiring workers with new entrants (OECD, 2006). A microeconomic perspective suggests that if Social Security as well as private pension systems are appropriately designed and individuals work longer, the amount of pension benefit could increase and the risk of poverty after retirement could be reduced.

Nowadays, the typical European family approaching retirement is a dual-earner family: the dataset used in this paper reveals that 78 percent of working males is married, and amongst these 24 percent has a working wife (in addition to some 36 percent of married working males with a retired wife). This results from dramatic changes in the labor force behaviour of both older men and older women after World War II.

Currently, labour market participation after the age of 50 varies substantially across OECD countries, ranging from less than 50% to more than 70%. Underlying the changes over time in participation rates of people aged 50-64, there are very different trends by gender. Participation rates for older men have decreased substantially since 1970 in almost all OECD countries with important differences across countries. For the OECD area, fewer than 1 in 6 older men were not in the labour force in 1970, this number has increased to 1 in 4 in 2004. On the other side, there has been an increasing trend in the participation of older women in every OECD country, which accounts for the rise over time in the total participation rate. This is the by-product of a constant increase in participation rates of successive cohorts of young women which has translated in a rise of participation rate in the older age group. The change has not been even across countries: there is a wide gap between countries like Belgium, Greece, Italy, Spain, etc. and those at the top of the scale where the participation rate of women aged 50-64 is 65% or higher (Denmark, Finland, Norway, Sweden, etc.).

The reasons why older people do not work may be very different. There are often employer barriers due to negative perceptions about the capacities of older workers to keep the pace of technological and organizational change; wages and non-wage labour costs that rise more steeply with age than productivity and the difficulties firms may face in adjusting employment because of employment protection legislation. On the supply side, older workers might experience a depreciation of their human capital; they may face difficulties to upgrade their skills and may not be very motivated to take up available opportunities for training. Bad health and poor working conditions might play a relevant role: many studies show that blue-collar and less qualified workers retire earlier than white-collar and highly qualified workers. Moreover, older workers who lose their job often face considerable difficulties finding a new job and large potential wage losses. On average across OECD countries, the hiring rate of workers aged 50 and over was less than half the rate of workers aged 25-49 (OECD, 2006).

This paper is aimed at helping to bridge the existing gap between European and US literature on joint retirement and at providing new evidence on couples' joint retirement. Historical participation rates trends have changed and gender differences in labor market attachment across ages signal a need of investigating retirement choices at a household level. This might also be informative for poverty reducing policies: correct incentives need to be provided to workers such that husbands and wives can take appropriate retirement decisions by looking at differences in the amount of benefits deriving from retirement close to or far from each other.

We adopt a duration analysis approach and estimate both single and competing risks models by allowing for a flexible specification with and without unobserved heterogeneity. Our findings show that joint retirement is significantly correlated with education, age, and health status, together with partner's employment status, partner's education and partner's health status. We also perform a sensitivity analysis in order to check whether the results on the impact of health status are robust to two alternative measures of health which possibly correct for subjectivity and cross-country incomparability of self-reported health. In the first case, we compute an objective and internationally comparable measure of health by combining subjective and objective health information in SHARE into a single index (Jurges, 2007). In the second case, we remedy for cross-country incomparability of self-assessed health by correcting reporting heterogeneity with anchoring vignettes (King et al., 2004).

The structure of the paper is as follows: section 2 briefly reviews the relevant economic literature, section 3 describes the data set, section 4 provides details on the empirical strategy, section 5 illustrates our results, section 6 shows the sensitivity analysis, and section 7 concludes.

2 Literature

The last decade has shown an increasing interest in couples' joint retirement decisions, mainly on US data. For example, Baker and Benjamin (1999) examine the effects of the introduction of the Spouse's Allowance to the Canadian Income Security system on couples' retirement behaviour, and they find that the introduction of such a scheme is associated with a relative increase in the labor force rates by 6-7 percentage points amongst males in eligible couples. Blau (1998) proposes a model where household utility function depends on both husband's and wife's labor supply status, so that households are assumed to behave as single decision makers. The revealed associations between labour force transition probabilities of one spouse and labour force status of the other spouse are not explained by financial incentives (like in Baker and Benjamin,

1999) but they seem to be the result of preferences for shared leisure. One main problem with this unitary approach is that there is a lot of empirical evidence that does not fit the data very well. Slutsky symmetry and negativity are regularly rejected when confronted with consumption or labor supply data (Browning and Chiappori, 1998; Michaud and Vermeulen, 2004).

Gustman and Steinmeier (2000, 2002, 2004) consider that older couples are composed of two individuals with own preferences. Household decisions result from a non-cooperative Nash bargaining approach. They find that a measure of how much each spouse values being able to spend time in retirement with the other accounts for a good portion of retirement decisions interdependence. Furthermore, they find an asymmetry in which husbands prefer their wife to be retired before themselves and a distaste of many husbands to retire if their wife is in poor health, whereas wives are willing to stay at home with sick husbands. Maestas (2001) bases her model in the cooperative game theory adapting the collective framework of Browning and Chiappori (1998) to a life-cycle utility maximization model. The model ignores the complexities that can arise because of highly non-linear pension incentives that may lead to multiple solutions to the first-order conditions. Maestas (2001) specifies two mechanisms through which husbands and wives interact in choosing their optimal retirement dates: leisure complementarity and decision-making power. Results suggest that couples with high leisure complementarity retire within a shorter time interval than other couples do, and this effect is enhanced when the wife has greater decision-making control. Coile (2004) focuses on the budget constraint in a family framework: husbands' and wives' retirement behaviour is affected by spillover effects coming from the spouse's financial incentives provided by pensions and Social Security. He estimates a reduced form model where the dependent variable, retirement probability, depends on labour force status of the spouse and on the head's and spouse's financial incentives. An et al. (2004) estimate a multivariate mixed proportional hazards model and find strong evidence of complementarity in leisure time of Danish husbands and wives.

The purpose of the paper is to explore the existence of cross-spouse effects by estimating reduced-form models of individual and joint retirement decisions. We do not model financial incentives for a number of reasons: the sample at hand contains only one cross-section and does not provide information about earnings before retirement for retirees, and legislative details are very different across the sample of countries available in SHARE.

3 Data

We use data from Release 2.0.1 of the Survey of Health, Ageing and Retirement in Europe (SHARE). It is a multi-disciplinary and cross-national data set which has information on the individual life circumstances of all eligible members of about 18,000 households. A household is eligible for participation in SHARE if at least one household member was born in or before 1954. An individual member of the household is eligible for an interview if she or he, or his/her partner, was born in or before 1954. The SHARE data was gathered in 2004 and is a random sample of the target population[1]. The survey contains information on a wide range of health indicators and socioeconomic variables of more than 30,000 individuals. SHARE covers 11 countries: Austria, Belgium, Denmark, France, Germany, Greece, Italy, the Netherlands, Spain, Sweden and Switzerland. The data set is designed after the Health and Retirement Study (HRS) and the English Longitudinal Study of Ageing (ELSA). Its cross-sectional dimension makes it a unique and particularly interesting data set in comparison to other micro data on the elderly.

The concept of retirement has a number of meanings. According to OECD (1995), three broad definitions of retirement are distinguished. Retirement can be understood

[1] Data for Belgium and France were collected in 2004/2005.

as:

1) being a recipient of a public or private old-age pension, regardless of the current employment;

2) being out of the labor force, regardless of the reasons to cease working and no matter whether an old-age pension is being drawn;

3) a self-declaration of people describing their own status, regardless of employment status and receipt of a pension.

We combine the last two definitions and we define retirees those who (1) self-assess their employment status as retiree and (2) are actually out of the labor force. This latter implies that an individual (2a) reports having done paid work at some time in her life and (2b) she can indicate the year in which the last job was terminated. In the data set about 10,000 individuals retired according to this definition in the data. However, this definition is not based on current pension benefits claiming and does not include self-reported unemployed, housewives, or disabled people. For our analysis, a sample of 31,115 records is reduced to 23,000 records by dropping individuals who declare they are "unemployed", "permanently sick or disabled", "housewives", or in "another current job situation". To enable investigation of partner effects, we focus on a sub-sample of married and non-married couples living together, where both household members are aged 50 or older (about 9,000 observations). We also drop observations for which calibrated individual weight is missing and we keep only respondents who worked at the age of 50 or over [2], and thus we end up with a sample of 6,300 individuals. About 56 percent of the individuals (approximately 3,400) are retirees and, as Table 1

[2] We keep only individuals who did not retire before age 50, because a model relating for example current financial and health status makes no sense for people who stopped working at a very young age.

shows, the majority of them (about 85 percent) stopped working at age 55 or beyond.[3] Lastly, as illustrated in Table 2, year of retirement goes from 1979 to 2004, with the average being 1996 and the median 1997, which indicates a distribution skewed to the right. The object of interest is elapsed years from when a worker is 50 years old to the date of retirement or to the survey year for right censored spells. Tables 3 and 4 report the distributions of individuals by observed duration up to retirement for all observations and non right-censored observations, respectively. Results show that most of the workers retire at the age of 60, and women tend to retire earlier than men.

The definitions of covariates used in the analysis are reported in Table 5 along with their sample means, but some clarifications are needed. First, there are no time-varying covariates; second, information about labour income before retirement is scarce. Unfortunately, retired respondents were not asked about their labour income before retirement. This omission is unfortunate since labour income is an important determinant of retirement decisions through a substitution effect (Dorn and Sousa-Poza, 2005). We include educational attainments in order to have a control for labour income.

Life-cycle theory suggests that the timing of retirement is affected by workers' wealth. The SHARE data set provides the amount of assets at household level: the variable is the sum of real and net financial assets. Real assets are defined as the sum of the value of primary residence net of mortgage, the value of other real estate, the owned share of own business and the owned cars. Net financial assets are equal to the sum of the values of bank accounts, government and corporate bonds, stocks, mutual funds, individual retirement accounts, contractual savings for housing and life insurance policies owned by the household minus financial liabilities. All these values are expressed in Euro and are purchasing power parity adjusted, to correct for price level

[3]We take account of the complex survey design. The potentially bias effects on descriptive statistics are accounted for by using sampling weights provided in the data set: these weights are approximately equal to the inverse of the probability of selection of each household into the sample. We use calibrated individual weight for the main and vignettes samples together to compensate to some extent for unit nonresponse.

differences across countries. We choose to categorize asset income into country-specific quartiles. The values of household assets were imputed where missing.[4] Five values were generated for each missing observation by running the same program five times and by using a different seed to perform the hot-deck imputation in each run of the program (Brugiavini et al., 2005). We always take into account that we have five different data sets (see Appendix A for technical details). We also have educational attainments: SHARE originally classified individuals into national education schemes based on the highest level of education reported. We reclassified it into three equivalent categories: levels 0-1 (none, pre-primary, primary education), 2-3 (lower and upper secondary education), 4-6 (post-secondary and tertiary education) of the ISCED (International Standard Classification of Education) (UNESCO, 1997).

As we mentioned in the introduction, health is an important factor which can affect retirement decisions. The main issue with health in retirement models is the lack of availability of a good and objective measure of individuals' health. Such an issue is exacerbated when the analysis includes more than one country because self-reported health status may not be comparable across countries. Arguably, these measures are affected by cultural/social norms, differences in thresholds for medical diagnoses, and access to health care resources so that cross populations health comparisons could be difficult (Groot, 2000; Sen, 2002; Lindeboom and Van Doorslaer, 2004). In other words, international comparisons of health cannot separate out the observed variation of subjective health responses into the variation of genuine health and into differences related to cultural or social norms. SHARE is a unique data source which provides a wide range of measures of physical health. Along with the single-item 5-point scales question on self-perceived health, also a dichotomous variable indicating whether an individual suffers from chronic or long-term health problems is included. This is a

[4]A breakout by item of the percentage of observations requiring imputation shows that bank accounts have the highest average rate of non-response (28 percent), followed by mutual funds and stocks (8 percent).

quasi-objective measure of health status, i.e. a subjective information of a factual matter (Jurges, 2007). Thus, as a proxy of true health, we prefer to include this quasi-objective indicator of health rather than the self-reported general health status given that the former should be less affected by the mentioned above issues. In the sensitivity analysis we will further investigate the reliability of results based on this proxy of health status.

We include household size to investigate whether the presence of dependents living in the same household affects retirement choices, and a dummy variable for individuals living in a metropolitan area to capture the cost of living and services level which is usually higher in urban areas.

Since people might want to align their status to their partner's, the variable "active spouse" indicates whether an individual lives with an employed partner at the date of the survey. We include age difference between spouses, computed as age of the individual minus age of the spouse, as a possible reason for joint retirement; spouse's education level and spouse's health status are intended to capture an added worker effect and correlation in observables. We control for cross-country and institutional changes by adding country-fixed effects and age dummies amongst the covariates.

4 Empirical Strategy

Since the duration variable of interest - time to retirement - is available in interval-censored form (yearly intervals), the appropriate approach to modeling retirement duration is a discrete-time hazard model. Following Meyer (1990), the discrete time hazard of exiting the state of employment can be modeled using a discrete-time proportional hazards model. In particular, the hazard of retirement in the $j-th$ year, $h(t_j)$, for individual i with a vector of covariates, x, who spent t years in employment and given that retirement has not occurred before t_{j-1}, is the following:

$$h_{ij} = 1 - \exp\left(-\exp\left(\gamma_j\left(t\right) + \left(x_i\beta\right)\right)\right),\ where\ \gamma_j\left(t\right) = \int_{-\infty}^{\infty} h_o(u)du \tag{1}$$

$\gamma_j\left(t\right)$ represents the baseline hazard. We assume a full non-parametric specification, i.e. we include a dummy for each year up to retirement. Rearranging (1), we get what is known as complementary log-log transformation of the conditional probability of exiting the state of employment at time t_j :

$$ln\left(-\ln\left(1 - h_{ij}\left(t_j|x_i\right)\right)\right) = x_i'\beta + \gamma_j\left(t\right) \tag{2}$$

The parameter β is interpreted as the effect of the covariates on the hazard rate of retirement in interval j, assuming a constant hazard rate over the $j - th$ interval.

In the duration literature it is consolidated that not accounting for unobserved heterogeneity might lead to biased estimates of the baseline hazard as well as of covariate effects (Heckman and Singer, 1984a and 1984b; Lancaster, 1990).[5] Taking this into account, we try to control for unobserved heterogeneity. The standard practice in the literature is to introduce a positive-valued random variable (mixture), v, into the hazard specification. In the context of the proportional hazard approach, the augmented hazard function, which incorporates a multiplicative mixture term, is given by:

$$ln\left(-\ln\left(1 - h_{ij}\left(t_j|x_i\right)\right)\right) = x_i'\beta + \gamma_j\left(t\right) + u_i \tag{3}$$

where $u_i = log(v_i)$. It is not possible to estimate the values of v themselves since, by construction, they are unobserved. Or equivalently, there are as many individual effects as individuals in the data set, and there are not enough degrees of freedom left to fit

[5] In our case, the effects of unobserved heterogeneity on coefficients should be mitigated, given that we choose a flexible specification for the baseline hazard (Meyer, 1990).

these parameters. However, if we suppose that the distribution of v has a shape whose functional form is summarized in terms of a few key parameters, then it is possible to estimate those parameters with the available data. So after specifying a distribution for the random variable v, we derive the "frailty" survivor function corresponding to this mixture distribution, and we write the likelihood function so that it refers to the initial and mixing distributional parameters rather than to each v. In the discrete-time case, the individual likelihood contribution incorporating unobserved heterogeneity is:

$$L_i(\beta, \gamma, \Delta) = \int_{-\infty}^{+\infty} \left(h_j(t, x_i | u_i)^{y_i} (1 - h_j(t, x_i | u_i))^{1-y_i} \right) g_u(u_i) du \tag{4}$$

where Δ is the vector of unknown parameters in $g_u(u_i)$. The unobserved heterogeneity term is assumed to be independent of observed covariates (x_i) and the random duration variable (T), and have density $g_u(u_i)$. We assume that u has a normal distribution with zero mean. In this case, there is no convenient closed form expression for the survivor function and likelihood contributions: the "integrating out" must be done numerically.

Finally, we estimate an alternative model with two routes out of employment (joint retirement with partner and retirement alone) by means of an independent competing risks model. Hence, we define the cause-specific hazard function to destination $r1$ (joint retirement with partner) and to destination $r2$ (retirement before or after partner) as:

$$h_{ij}^{r1} = 1 - \exp\left[- \int_{t_{j-1}}^{t_j} \theta_{fc}(t)\, dt \right]$$

$$h_{ij}^{r2} = 1 - \exp\left[- \int_{t_{j-1}}^{t_j} \theta_{oc}(t)\, dt \right]$$

where θ_{r1} and θ_{r2} are underlying destination-specific continuous time hazards. The overall discrete hazard and the survivor function for exit to any destination for t_j are given by:

$$h_{ij} = 1 - \left\{ \left[1 - h_{ij}^{r1}\right] \left[1 - h_{ij}^{r2}\right] \right\}$$

$$S_{ij} = S_{ij}^{r1} S_{ij}^{r2}$$

To further proceed, we assume that transitions can only occur at the boundaries of the intervals.[6] Then the overall likelihood contribution for a person with a spell length t_j is given by:

$$L_{ij} = \left(L_{ij}^{r1}\right)^{\delta_{r1}} \left(L_{ij}^{r2}\right)^{\delta_{r2}} \left(L_{ij}\right)^{1-\delta_{r1}-\delta_{r2}} = \left[\frac{h_{ij}^{r1}}{1-h_{ij}^{r1}}\right]^{\delta_{r1}} S_{ij}^{r1} \left[\frac{h_{ij}^{r2}}{1-h_{ij}^{r2}}\right]^{\delta_{r1}} S_{ij}^{r2} \tag{5}$$

where δ_{r1} and δ_{r2} are the destination-specific censoring indicators. Thus, the likelihood contribution (4) partitions into a product of terms, each of which is a function of a single destination-specific hazard only. Consequently, it is possible to estimate the overall independent competing risk model by estimating separate destination-specific models having defined suitable destination-specific censoring variables. In the competing risks model, as well as in the previous model, we have accommodated the presence of observed individual heterogeneity assuming a multiplicative error term associated with each specific hazard function.

[6]The assumption may not be appropriate in practice. So we have also estimated a multinomial logit model, originally developed for intrinsically discrete data. If the interval hazard rate was relatively small, this model would provide estimates that are a close approximation to a model for grouped data with the assumption that the (continuous) hazard is constant within intervals.

5 Retirement Patterns

As documented in the introduction, women are important actors in the labor market: participation rates for older men have decreased substantially, and there has been an increasing trend in the participation of older women, as by-product of a constant increase in participation rates of successive cohorts of young women, in every OECD country, which accounts for the rise over time in the total participation rate.

Another indicator of the new role of women in the labor market comes from the estimates of the Life-table's hazard functions.[7] Figure 1 illustrates gender disaggregated hazard functions. The hazard function shows the proportion of people exiting employment as time proceeds. The hazards for men are almost identical to those for women through age 65, where we observe the big spike which follows the first but smaller spike at age 60. The hazard for men is higher than for women after age 65. Thus, women's retirement decisions might be influenced by the same factors that affect those of men, namely Social Security, and that women do not simply follow their partners' choices.

Indeed, joint retirement is a common but not a universal phenomenon: of the couples in which both spouses are between 50 and 75 years of age and retired by the survey year, 27 percent retire within one year of each other. Figure 2 plots the fraction of retired couples in which both spouses are between 50 and 75 years of age by difference in age and in year of retirement. Age difference is computed as age of husband minus age of wife, and year of retirement difference is given by year of husband minus year of retirement of wife. Therefore, a negative age difference means that wife is older than her husband, and a negative year difference implies that husband retired first. The series labeled "All Years" and "All Ages" show the marginal distribution of the age difference and retirement dates difference between husbands and wives. Moving along the x-axis, each series indicates the distribution of age differences for couples in each

[7] A non-parametric approach to duration analysis: they are a generalization of the Kaplan-Meier hazard functions for interval-censored data. The estimates are available on request from the authors.

year-difference group. The joint retirement, defined as retirement within one year of each other, has spikes across the age-difference groups: it is interesting to note that very small is the fraction of couples with negative age differences, or in other words with wife older than husband. If we account for the low proportion of these couples, we can say that joint retirement, a quite common but far from universal event, is not simply explained by difference in age. If joint retirement could be explained by couples being of the same age, husband who are two years older than their wives should retire two years earlier. Starting from the left-hand corner and moving down along the diagonal, we do not observe high percentages of couples retiring for all groups except for the last: wives who are four years or more younger are likely to retire four or more years later. Overall, the proportion of joint retirements remains significantly high as age difference increases, therefore complementarity in leisure could be an important explanation of older couples' retirement decisions.

Figures 3-6 provide plots of the disaggregated hazard functions by partner's education. For example, men with less than college education are more likely to retire if their partner has a higher educational level. This could be related to an added worker effect, if we consider education a good proxy of earnings. This is true also for women with a college education. Figures 7 and 8 seem to support the hypothesis of complementarity in leisure time: the hazard of retirement for individuals living with an employed partner at the date of the survey is generally lower than for individuals with a partner retired. This holds both for men and women.

6 Results

The first set of results comes from the single risk duration model, and the second set is from an independent competing risks model. Finally, we show a sensitivity analysis of our main results.

6.1 Results from the Single Risk Models

As mentioned earlier in the description of the data set, we take account of the complex sample design in all estimations.[8]

The estimated coefficients of duration dependence with and without unobserved heterogeneity[9] reveal that the baseline hazard increases as time proceeds (see Figure 9). Hence the results indicate that, though there are some differences in the magnitude of the estimated hazards between the homogeneous and the mixing model, there is a general evidence of true positive duration dependence. This supports Meyer's (1990) suggestion that using a flexible specification for the baseline hazard attenuates the sensitivity of estimated parameters to the unobserved heterogeneity. It could be explained by the fact that as time passes individuals are more likely to retire since they accrue years of service. It seems also likely that people with a long tenure could take advantage of pension funds rules providing early retirement benefits only to workers above a certain tenure.

The main effects of covariates on the hazard of exit from employment are reported in the first column of Table 7. Starting with the effect of personal characteristics on the hazard of retirement, women do not have a different hazard of retirement with respect to men. Highly educated individuals, i.e. with post-secondary or tertiary education, are less likely to retire than those with low educational attainment (primary education or no schooling at all). Such a negative correlation between schooling and likelihood of retirement could be explained by the fact that the level of education is a proxy of earnings level. Individuals with higher education may decide to postpone retirement

[8]The sampling weight used is wgtaci, that is a calibrated individual weight for the main and the vignettes samples together.

[9]The baseline hazard functions are calculated by setting all covariate values equal to zero.

because they value future opportunities on the labor market (in terms of higher earnings) more than leisure and pension benefits. With regard to assets, the life-cycle model postulates that higher wealth increases the probability of retirement. The assets measures correspond to the survey time, which for most of the non-censored observations[10] takes place after retirement. We include country-specific household assets quartiles: the coefficients of the second and third quartile are not statistically significant. The last quartile has a negative and statistically significant coefficient: individuals who do not retire or retire later end up being the wealthiest after retirement. Household size affects the likelihood of retirement negatively: individuals living only with a spouse have a 16 percent higher hazard than those living in a household with more members. Our data supports the hypothesis that retirement behavior differs between urban and rural areas: living in a big city lowers the retirement hazard, perhaps reflecting the high costs of living in urban areas. Turning to health status, a problem similar to assets measures affects the interpretation of its coefficient, though it is a bit attenuated by how the question has been phrased: individuals are asked whether they suffer from a long-term health problem and by long-term the question means it has troubled the individual over a period of time or is likely to affect the individual over a period of time. In this case, for most of the non-censored observations the measure refers to a time after retirement, but the onset of the long-term disease could be traced back to a time before retirement and thus could have affected the retirees for a long time. Long-term illness positively influences the hazard of retirement, consistently with earlier studies (Gordon and Blinder, 1980; An et al., 2004; Dorn and Sousa-Poza, 2004). Hence bad health increases the probability of retirement. There are two possible explanations of it. First, poor health leads to decreases in productivity, decreases in wages, and more generally to less attractive employment opportunities. The second is that poor

[10] Excluding individuals who retired at the time of the survey, but they are a very small fraction of the working sample.

health might shift individual preferences toward leisure because work becomes more burdensome.

Turning to the effects of partner's characteristics, the coefficient of partner's high education is positive and statistically significant. A partner with a high education can be expected to strongly correlate with higher earnings, thus the former result could indicate an added worker effect, whereby individuals decrease their labor supply when spouse's earnings increase. On the other hand, we also find some evidence of complementarity in leisure because in our sample people tend to align their labor force status to their partner's status. Individuals with a partner in the labor force have a lower hazard of retirement with respect to those with an inactive partner. However, contrary to what we found in descriptive statistics and unlike An et al. (2004), we do not observe individuals having a higher hazard of retirement if the partner is closer to them in age. The fact that spouses close in age are more likely to retire together is due to Social Security incentives being more similar than for couples who differ in ages[11]. Partner's bad health has a different effect by gender: it reduces males' likelihood of retirement if the spouse has a long-term illness, whereas it increases the probability of retiring for women. So females could be more prone to take care of their partner and thus quit their job.

As to the effects of employment-related characteristics, self-employed have a 40 percent lower hazard of retirement since self-employment might allow for a more gradual transition to retirement through a gradual and flexible reduction of working hours. A positive effect on the hazard of retirement is also found for large firms. This result is consistent with the strong positive impact on retirement decisions for workers in firms above a specific size found by three micro studies for Europe (Røed and Haugen, 2003, for Norway; Dorn and Sousa-Poza, 2005, for Switzerland; Wübbeke, 1999 for

[11](If the pensionable age in country A is 65, two spouses who are the same age will reach 65 at the same time, and hence we will be more likely to observe them retiring jointly that if the husband become 65 five years before the wife)

Germany). We also test the effect of being a supervisor: supervising any number of workers, as opposed to having no supervisory power at all, does positively affect the probability of retirement.

In the second column of Table 7, we report the estimated coefficients when we introduce unobserved heterogeneity. The results are fairly similar to the ones reported in the first column[12], with the main difference being the significance of the gender dummy.

6.2 Results from the Independent Competing Risks Model

Now we turn to the multiple destination states model. Sample means of duration until retirement and of number of exits are reported in Table 6. Comparing individuals retiring jointly with their partner with individuals retiring alone shows that elapsed duration is slightly longer for the formers. The most common form of transition is retirement before or after partner: only about 8 percent of individuals in the sample retire jointly with the spouse, indicating that there is not a very high propensity to retire together.

The disaggregated version of the piecewise constant hazard regressions is reported in Table 8. It is immediately evident that some regression coefficients vary by destination state. An individual's educational level decreases the probability of retirement alone but does not affect the hazard of joint retirement. Assuming that education is a proxy of earnings level, it lowers the probability of retirement before or after the spouse but not that of joint retirement, i.e. the substitution effect is significant only in case of retirement alone.[13] Living in urban areas is significantly and negatively associated with the probability of retirement before or after spouse but not with the hazard of joint

[12]Standard errors are however not precisely estimated given that we could not clusterize by country and take account of the complex sample design.

[13]A t-test indicates that we can reject the null hypothesis of equality of the coefficients of a high education dummy.

retirement. Individuals who do not retire or retire later end up being the wealthiest after retirement.[14] Difference in age between spouses turns significant: the greater the difference the lower the probability of joint retirement, the opposite is true if we consider the hazard of retirement before or after the spouse. These findings warn against uncritical aggregation by destination state. Finally, poor health significantly increases the probability of retirement before or after the spouse but does not have any impact on the hazard of joint retirement[15]. Partner's education and health status do not significantly affect the decision of either retiring alone or jointly. Self-employed have a lower hazard rate of retirement before or after spouse. If we consider the hazard of joint retirement the effect of self-employment is opposite. Firm size positively influences the probability of retirement before or after spouse but it has no direct effect on joint retirement. The third and the fourth columns of Table 8 show the estimated coefficients taking account of unobserved heterogeneity: they are fairly similar to the coefficients in the homogeneous model.[16]

Very similar results are obtained estimating a multinomial logit model, under the assumptions that the interval-hazard is small and that the continuous hazard is constant within intervals.[17]

Baseline hazard functions, corresponding to the non-parametric specification, are illustrated in Figures 10 and 11. The baseline hazard function for joint retirement is stable over years. This is not the case for retirement before or after spouse which is characterized by decreasing exit rates. The hazard function of joint retirement is generally lower than that of retirement before or after her spouse. Very similar results are obtained when we take account of unobserved heterogeneity.

[14]A t-test shows that we can reject the null hypothesis of equality of the coefficients of a high assets dummy.

[15]We remind the reader that health measures suffers from the same problem as household assets, though it could be mitigated by the fact that the variable refers to a long-term disease rather than overall health status at the time of the survey

[16]See note 8.

[17]These results are available on request from the authors.

7 Sensitivity Analysis

In the main section, we estimate the effect of health status on the hazard of retirement. As we mentioned previously, we prefer the quasi-objective measure of health rather than the general subjective health assessment because the latter could be severely affected by measurement error. The "long-term illness" variable could not be completely reliable given that it is self-reported information. As we argued in section 3, given that our analysis is a cross-cultural study, respondents from different countries and cultures might not have the same reference levels of health. Furthermore, it is often stated that a justification bias could be encountered with self-reported measures of health: retirees could rate their health status worse than it actually is in order to justify their exit from the labor force. In this section, we examine the sensitivity of our estimates of health status to two different definitions.

First, we compute, as in Jurges (2007), an objective and internationally comparable measure of health by combining in a single index every subjective and objective information on health contained in SHARE.[18] We get a health index ranging from 0 (the worst observed health state) to 1 (very good health). The computed health index will be used as a proxy for true health.

Assuming that true health is a continuous and unobservable variable, the health index is computed as the linear prediction from an ordered probit regression, normalized to 0 for the worst observed health state and 1 for the best observed health state. Self-rated health ranging from excellent (1) to poor (5) is regressed on 20 different diagnosed physical conditions (as reported by respondents), whether ever treated for depression, body mass index, grip strength, age and gender. The absence of any condition implies

[18]SHARE contains a broad range of different health measures, both physical and mental. These include self-reported general health, self-reported diagnosed chronic conditions, medication, functional limitations, ADL and IADL limitations, symptoms, mental health as measured by two alternative depression scales (CES-D and Euro-D), physical measurements (hand grip strength and gait speed), self-reported height and weight (to compute BMI) and detailed information on health services utilisation (doctor visits, hospital stays, etc.).

perfect health, i.e. an index value of 1. The presence of a condition reduces the health index by some given amount or percentage, the so-called disability weight. Since the dependent variable is potentially subject to cross-cultural bias, as in Jurges (2007) we account for country-specific reporting styles by modeling ordered probit thresholds as a function of the country of residence. The identification of this model relies on the availability of an objective measure of health, i.e. grip strength, which allows to arrive at comparable scales.[19]

Formally, let GH_i^* be the latent (unobserved) perceived level of general health of individual i, which is modeled as an ordered probit model:

$$GH_i^* = X_i'\beta + \varepsilon_i \tag{6}$$

where X_i includes the previously mentioned covariates, β includes parameters and ε_i is an individual residual error term, assumed to be standard normal distributed, $\varepsilon_i \sim N(0; \sigma^2)$.

Respondent i turns the continuous perceived level of health into a reported category, gh_i,where:

$$gh_i = k \;\; if \;\; \tau_k^{m-1} \leq GH_i^* \leq \tau_k^m$$

and where -$\infty = \tau_k^1 < \tau_k^2 < \ldots\ldots < \tau_k^M = 5$. The thresholds are allowed to vary over countries as follows:

[19]Grip strength is a core physical measure of health that potentially enables cross-national comparability of health estimates and avoids some of the endogeneity problems inherent in more subjective health measures like self-rated health. It also helps to overcome the measurement issues related to biases that arise from subjectivity of self-reported health and health conditions due to cultural differences across and within countries, differential physician contacts or cross-national differences in the criteria for thresholds of medical diagnosis. At the same time, predictive validity of grip strength for assessing health was established in studies that found grip strength to be a better predictor of future medical problems than self-reported health (Rantanen et al., 1999; Al-Snih et al. 2002).

$$\tau_k^1 = \gamma^1 D_k$$

$$\tau_k^m = \tau_k^{m-1} + \exp(\gamma^m D_k)$$

where τ_k^m is the $m-th$ threshold for country k, γ^m is a vector of parameters in the $m-th$ threshold equation and D_k is a vector of country dummies. M is the number of categories of the dependent variable (five in our case). Note that while thresholds are allowed to vary across countries, disability weights are constrained to be the same in each country. The last assumption, that is the cross-national identity of disability weights, is crucial to distinguish true health from reporting effects.

Estimated results of the generalised ordered probit are reported in Table 9. Regression parameters have very small standard errors and are statistically significant. For the sake of simplicity, the specification of the health index equation does not account for co-morbidity. The second column reports the implied disability weights computed as the regression parameters from the generalized ordered probit model divided by the range of its linear prediction. As in Jurges (2007), the highest disability weight is found for Parkinson's disease and heart attack while the lowest is found for high blood cholesterol and hearing problems.

Third column of Tables 7 and 8 report estimates respectively of the single and the competing risks models when the health index, instead of self-reported long-term illness, is added as covariate.[20] Overall, the results with the single index of health confirm our previous findings: better health status is generally associated with a lower hazard of retirement and partner's health has a different effect according to gender.[21]

[20]Given that some variables used to estimate the health indicator are also explanatory variables of the hazard model, we perform several tests for multicollinearity. They reject the hypothesis that explanatory variables are collinear.

[21]As always when using a predicted variable as explanatory variable, the hazard regression model has to take generated regressor bias into account, since standard errors would be underestimated

In this case, however, both the hazard of joint retirement and the hazard of retirement alone are negatively associated with better health.

SHARE also offers a second way to purge the data from cross-country reporting bias, as information on anchoring vignettes was collected for a subset of respondents. In this case, we use the correction of response category cut points by means of anchoring vignettes. These are short descriptions of people in different states of health which respondents are asked to rate on a common scale. Vignette questions have been applied successfully in recent work on international comparisons of health, job satisfaction and political efficacy (King et al., 2004; Salomon et al., 2004; Kristensen and Johansson, 2008). The estimated model in this case explains both respondents' self-reports on health and their reports on several domains of health of hypothetical vignette people.[22] Self-report health is modeled with an ordered probit as in equation 3.[23] In this case, X_i includes gender, education, age and country-specific effects[24] and the thresholds are allowed to vary across observations as a function of covariates Z_{it} which may overlap X_i :

$$\tau_i^1 = \gamma^1 Z_i$$

$$\tau_i^m = \tau_i^{m-1} + \exp(\gamma^m Z_i) \tag{7}$$

The fact that different respondents can use a different response scale τ_i^m is known as "differential item functioning" (DIF). By using only self-reported own health, the parameters β and γ_1 cannot be separately identified.

otherwise. We bootstrap standard errors using 500 replications.

[22]The vignettes we use in the analysis are presented in Appendix B.

[23]The variance in this case is normalized to one to identify the variance of the latent variable in the vignette model.

[24]The vignette subsample does not include Austria, Denmark and Switzerland.

A subsample L answered vignette questions related to three domains of health (pain, mobility and work disability) of hypothetical people. Respondent l perceives the true unobserved actual level of health as θ_j with normal random error, so that:[25]

$$V_{l,j}^{*} \sim N\left(\theta_j,\ \sigma^2\right)$$

becomes respondent l's continuous unobserved perception of the actual level of health for vignette j. The evaluations of vignettes are modeled by using similar ordered response equations. Thus, the respondent turns the continuous $V_{l,j}^{*}$ into a categorical answer to the survey question v_{lj} via this observation mechanism:

$$v_{lj} = k \quad if \quad \tau_l^{m-1} \leq V_{l,j}^{*} \leq \tau_l^{m}$$

with thresholds determined by the same γ^1 coefficients as in (4) for gh_i and the same explanatory variables but with values measured for units l, V_l :

$$\tau_l^1 = \gamma^1 Z_l$$

$$\tau_l^m = \tau_l^{m-1} + \exp(\gamma^m Z_l) \tag{8}$$

By applying the same thresholds in the vignette model as in the self-assessment model, we assume that individuals use the response categories for the health question in

[25]Here there are two implicit assumptions. The first is that the domain levels represented in each vignette are understood in the same way by every respondent, irrespective of their country of residence and of other sociodemographic variables. The second assumption is that response scale differences are the same in all domains of health.

the same way when they evaluate hypothetical scenarios as they do when they provide self-reported assessments of their own current health, we adopt a response consistency assumption. It is clear how the vignette evaluations can be used to separately identify β and γ (=γ_1,...γ_5). From the vignette evaluations alone, γ, $\theta_1, \ldots \theta_3$ can be identified (up to the usual normalization of scale and location). From self-reported health , β can then be identified in addition. Thus, the vignettes can be used to solve the identification problem due to DIF. The two-step procedure is sketched only to make intuitively clear why the model is identified. In practice, all parameters will be estimated simultaneously by maximum likelihood. The model outlined in this section is called chopit model (compound hierarchical ordinal probit, see King et al; 2004). Results of chopit and of standard ordered probit are reported in Table 10. As far as the chopit model is concerned, we report only the estimated effects of respondent characteristics and of country indicators on the first response threshold, i.e. the threshold for determining whether someone claims to be in "excellent health". The other columns of table 10 are the coefficients in the general health equation not correcting or correcting for differential response scales. In the ordered probit model, i.e. without the DIF correction, the probability of reporting excellent health falls significantly with age. Health is positively associated with the highest educational level and women seem to be more likely to report worse health than their male counterparts. The DIF corrected results imply that the negative impact of age on health is even higher compared to the non corrected one since a higher initial threshold for self reported general health is used by older individuals, i.e. older individuals tend to have a "milder" view of their own health status compared to their younger counterparts. There is also evidence of correlation between the response scale and education level. Thus, for example, individuals with this educational level have a higher initial threshold, i.e. they are more likely to rate their health as excellent. This implies that the negative impact of the lowest educational level on general health is underestimated without correcting for DIF. As far as the country dummies are concerned, though they are generally significant in the threshold equation for the chopit model, the results reveal that the country ranking only differs very little between the chopit and the ordered probit models. The healthiest respondents live in Germany while the least healthy ones live in Sweden. Testing for rank correlation (Kendall's tau) we cannot reject that they have the same order.

The last columns of Tables 7 and 8 report estimates of single and competing risks model when the DIF-corrected health index is added as covariate. As in the previous cases, good health status is generally associated with a lower hazard of retirement but partner's health has not a different effect according to gender. For both men and women, we find that spouse's health is correlated negatively with the hazard of retirement. The estimated effect, however, is significant only for women. In the competing risks model, we find that the decision of either retiring jointly or alone are negatively associated with both individual and partner's good health.

8 Conclusions

Exploiting the retrospective questions in the employment section of SHARE, we estimate duration models to investigate the determinants of couples' retirement decisions. We allow for single risk (retirement) and competing risks (retirement alone vs. joint retirement) and test specifications with individual and partner's variables. Our findings suggest that leisure complementarity, correlation in observables, captured by partner's labor market status and education level, and age difference between spouses, are important determinants of couples' retirement decisions in addition to standard controls like household assets, individual education as a proxy for earnings, age, gender, worker type, etc. Partner's health status has a different effect according to gender: females are more likely to take care of their sick partners, and in that case they retire earlier, whereas husbands do not. When we allow for different exit routes, difference in age between spouses turns out to be significant and has a different effect: the larger the difference, the less likely joint retirement, whereas it has a positive impact on the likelihood of retirement alone.

To our knowledge, this is the first time subjectivity of an independent variable is corrected for since a justification bias and a cross-cultural comparability issue could emerge. We correct for such issues by constructing an index variable and by including a self-reported health variable adjusted for cross-cultural differences. Our findings about health status are confirmed with both the index variable and the multiple cultures corrected variable.

Allowing for different exit routes is of critical importance to avoid getting misleading results where coefficients attached to variables might capture the effect on the probability of retirement without taking into account the very different nature of joint vs. alone retirement from the labor force.

European couples' retirement choices are determined by individual and job characteristics as well as by partner's and couple's characteristics which points to the existence of complementarity in leisure and correlation in observables. To the extent that these effects capture the role of love, we can say that money and love do matter for the employment vs. retirement decision.

A Appendix: multiple imputations

The values of some of the covariates used to estimate both the single and the competing risks model were imputed where missing. Five values were generated for each missing observation by running the same program five times and using a different seed to perform the hot-deck imputation in each run of the program (Brugiavini et al., 2005). In estimating the duration models reported above, we always take into account that we have five different data sets. We perform the analysis on each data set separately and then combine results from all five data sets by using the results by Rubin (1987). Let m=1,...,M index the imputation run (with M in our case equal to 5) and let $\widehat{\beta}_m$ be our estimate of interest (e.g. sample median, regression coefficient, etc.) from the m_{th} implicate data set. The estimate from all M data sets is just an average of M separate estimates, i.e.

$$\overline{\beta}_M = \frac{1}{M}\sum_{m=1}^{M}\widehat{\beta}_m$$

The variance of this estimate consists of two parts. Let $\widehat{V}_m$ be the variance estimated from the m_{th} implicate data set. Then the first magnitude one needs to compute is an average of all M variances, which constitutes the within-imputation variance, i.e.

$$WV_m = \frac{1}{M}\sum_{m=1}^{M}\widehat{V}_m$$

The second component to compute is the between-imputation variance,which is given by:

$$BV_M = \frac{1}{M-1}\sum_{m=1}^{M}\left(\widehat{\beta}_m - \overline{\beta}_M\right)^2$$

Then the total variance of the estimate is equal to:

$$V_M = WV_M + \frac{M+1}{M} BV_M$$

As Little and Rubin (2002) point out, the second term in the above equation represents the share of total variance due to missing values. One can perform a t-test of significance by employing the following formula to compute the degrees of freedom n:

$$n = (M-1)(1 + \frac{1}{M+1}\frac{WV_m}{BV_m})^2$$

B Appendix: vignette description for several domains of health. Response categories: none (1); mild (2); moderate (3); severe (4); extreme (5)

1. Vignette for pain: "Charles has pain in his knees, elbows, wrists and fingers, and the pain is present almost all the time. Although medication helps, he feels uncomfortable when moving around, holding and lifting things. Overall in the last 30 days, how much of bodily aches or pains did Charles have?"

2. Vignette for mobility: "Rob is able to walk distances of up to 200 metres without any problems but feels tired after walking one kilometer or climbing more than one flight of stairs. He has no problems with day-to-day activities, such as carrying food from the market. Overall in the last 30 days, how much of a problem did Rob have with moving around?"

3. Vignette for work disability: "Mark has been diagnosed with high blood pressure. His blood pressure goes up quickly if he feels under stress. Mark does not exercise much and is overweight. How much is Mark limited in the kind or amount of work he could do?"

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Table 1: Distribution of retirees by age group and gender.

	Gender		
Retirement Age	Males	Females	Total
Age 50-54	52.77	47.23	15.75
Age 55-59	59.39	40.61	36.64
Age 60-64	63.76	36.24	38.69
Age 65 and over	68.53	31.47	8.91
Total	60.85	39.15	100

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 2: Year of retirement.

	Percentiles	**Smallest**		
1%	1984	1979		
5%	1987	1979		
10%	1990	1979	Obs	3377
25%	1993	1980	Sum of Wgt.	11602454.8
50%	1997		Mean	1996.23
		Largest	Std. Dev.	5.053829
75%	2000	2004		
90%	2003	2005	Variance	25.54119
95%	2003	2005	Skewness	-0.4543619
99%	2004	2005	Kurtosis	2.624729

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 3: Observed durations for complete spells by gender.

	Gender		
Observed durations	Males	Females	Total
0	57.18	42.82	2.28
1	57.04	42.96	2.77
2	45.09	54.91	2.46
3	54.07	45.93	4.22
4	50.68	49.32	4.03
5	48.59	51.41	7.58
6	58.98	41.02	7.04
7	60.35	39.65	7.43
8	62.61	37.39	7.72
9	67.04	32.96	6.88
10	57.91	42.09	18.96
11	64.9	35.1	5.49
12	66.58	33.42	5.7
13	74.56	25.44	5.41
14	73.34	26.66	3.14
greater than 15	68.53	31.47	8.91

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 4: Observed durations for right-censored spells by gender.

	Gender		
Observed durations	Males	Females	Total
0	26.39	73.61	5
1	34.01	65.99	9.21
2	37.56	62.44	7.78
3	37.76	62.24	9.05
4	48.5	51.5	10.73
5	51.31	48.69	9.09
6	52.93	47.07	8.75
7	48.4	51.6	9.77
8	58.36	41.64	7.12
9	48.23	51.77	5.94
10	49.67	50.33	4.85
11	46.81	53.19	3.14
12	53.64	46.36	2.6
13	55.05	44.95	1.87
14	65.35	34.65	1.35
greater than 15	55.4	44.6	3.75

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 5: Descriptive statistics.

Variable	**Mean**	**Linearized s.d.**
Retired (1, retired)	0.562	0.009
Gender (1, female)	0.458	0.009
Supervisor (1, if person have managerial power over others)	0.298	0.008
Size (1, if firm size is over 200 persons)	0.513	0.009
Self-employed (1, if person is self employed)	0.141	0.006
Low education (1, if person with none or primary edu)	0.186	0.006
Medium education (1, if person with secondary edu)	0.530	0.006
High education (1, if person with postsecondary or tertiary edu)	0.267	0.009
Sector1 (1, if employed in the manufacturing sector)	0.137	0.002
Sector2 (1, if employed in the construction sector)	0.227	0.008
Sector3 (1, if employed in the retail trade sector)	0.080	0.006
Sector4 (1, if employed in the hotels sector)	0.108	0.008
Sector5 (1, if employed in the transport sector)	0.021	0.005
Sector6 (1, if employed in the financial sector)	0.075	0.006
Sector7 (1, if employed in other business activities)	0.034	0.003
Sector8 (1, if employed in the public administration)	0.044	0.005
Sector9 (1, if employed in the education sector)	0.078	0.003
Sector10 (1, if employed in the health sector)	0.111	0.004
Sector11 (1, if employed in other sector)	0.085	0.005
Agediff (difference in spouses' age)	3.099	0.006
Health status (1, if person with long-term illness)	0.346	0.005
Low assets income at household level	0.252	0.046
Medium-low assets income at household level	0.244	0.009
Medium assets income at household level	0.250	0.007
High assets income at household level	0.254	0.007
Urban area (1, if living in a metropolitan area)	0.269	0.007
Household size (1, if household size equal to two)	0.549	0.008
Low education (1, if person with none or primary edu)	0.200	0.003
Medium education (1, if person with secondary edu)	0.557	0.008
High education (1, if person with postsecondary or tertiary edu)	0.400	0.003
Tertiary I or II spouse (1, if spouse with tertiary edu)	0.254	0.009
Active spouse (1, if spouse is active on the labor market)	0.426	0.002
Health status spouse (1, if spouse with long-term illness)	0.357	0.008

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 6: Distribution by destination state and mean durations.

	Retirement			
Status	no	yes	Total	**Mean duration**
Right Censored	100	0	43.92	5.94(4.13)
Joint retirement	0	13.3	7.46	8.78(3.84)
Retirement before or after spouse	0	86.7	48.62	8.16(4.03)
Total	43.92	56.08	100	

Notes: Weighted results. Source: SHARE release 2.0.1.

Table 7: Results of the single risk model.

	Model (1)	**Model (2)**	**Model (3)**	**Model (4)**
Female	0.054	0.349	0.272	0.461
	0.524	*0.001*	*0.412*	*0.010*
Supervisor	0.159	0.175	0.174	0.124
	0.045	*0.029*	*0.020*	*0.248*
Size	0.215	0.402	0.203	0.259
	0.000	*0.000*	*0.000*	*0.000*
Self	-0.527	-0.706	-0.532	-0.514
	0.000	*0.000*	*0.000*	*0.000*
Medium education	-0.097	-0.194	-0.074	-0.069
	0.048	*0.071*	*0.066*	*0.242*
High education	-0.315	-0.637	-0.281	-0.007
	0.006	*0.000*	*0.007*	*0.968*
Agediff	0.003	0.001	0.005	0.010
	0.648	*0.897*	*0.406*	*0.217*
Low-medium income	0.015	-0.131	0.014	0.034
	0.800	*0.154*	*0.815*	*0.579*
Medium income	-0.076	-0.302	-0.070	-0.024
	0.388	*0.002*	*0.462*	*0.852*
High income	-0.482	-0.499	-0.441	-0.528
	0.000	*0.000*	*0.000*	*0.002*
Urban	-0.215	-0.190	-0.197	-0.319
	0.072	*0.013*	*0.103*	*0.017*
Health status	0.160	0.340	-1.251	-1.825
	0.003	*0.000*	*0.000*	*0.041*
Household size	0.147	0.273	0.132	0.187
	0.142	*0.006*	*0.140*	*0.006*
Active spouse	-0.702	-0.876	-0.651	-0.650
	0.000	*0.000*	*0.000*	*0.000*
Medium education (spouse)	0.073	0.123	0.037	0.189
	0.388	*0.243*	*0.660*	*0.217*
High education (spouse)	0.142	0.196	0.108	0.301
	0.023	*0.120*	*0.154*	*0.174*
Health status spouse	-0.173	-0.101	0.446	-0.137
	0.042	*0.270*	*0.471*	*0.891*
Health status spouse*female	0.150	0.102	-0.202	-1.246
	0.112	*0.455*	*0.535*	*0.000*
LogL	-30079170	-8623.153	-29154348	-28720678
N	41589	41589	40381	26170

Notes: Weighted results. P-value reported below respective coefficient. All regressions include country-fixed effects, age dummies and sector dummies. Model(1): Main specification; Model(2): Specification with unobserved heterogeneity; Model(3): Specification with the Health index variable; Model(4): Specification with the self-reported health variable corrected for differential item functioning. Source: SHARE release 2.0.1.

Table 8: Results of the independent competing risks model.

	Model (1)		Model (2)		Model (3)		Model (4)	
	Joint ret	**Ret alone**	**Joint ret**	**Ret alone**	**Joint ret**	**Ret alone**	**Joint ret**	**Ret alone**
Female	0.523	0.072	0.711	0.315	0.799	0.422	0.526	0.431
	0.002	*0.113*	*0.365*	*0.269*	*0.003*	*0.001*	*0.000*	*0.000*
Supervisor	0.195	0.150	0.203	0.085	0.229	0.148	0.214	0.160
	0.000	*0.000*	*0.138*	*0.070*	*0.000*	*0.000*	*0.001*	*0.000*
Size	-0.054	0.159	-0.182	0.162	-0.097	0.151	-0.048	0.162
	0.706	*0.000*	*0.160*	*0.000*	*0.492*	*0.000*	*0.771*	*0.000*
Self	0.519	-0.354	0.543	-0.276	0.495	-0.356	0.531	-0.368
	0.000	*0.000*	*0.001*	*0.000*	*0.000*	*0.000*	*0.000*	*0.000*
Medium education	-0.165	-0.063	-0.012	-0.084	-0.149	-0.058	0.080	0.030
	0.216	*0.061*	*0.942*	*0.179*	*0.259*	*0.075*	*0.397*	*0.545*
High education	-0.226	-0.231	-0.245	-0.238	-0.219	-0.215	0.641	0.171
	0.078	*0.000*	*0.267*	*0.002*	*0.101*	*0.000*	*0.000*	*0.023*
Agediff	-0.098	0.015	-0.088	0.015	-0.101	0.016	-0.099	0.021
	0.000	*0.002*	*0.000*	*0.023*	*0.001*	*0.001*	*0.002*	*0.000*
Low-medium income	0.051	-0.053	-0.067	-0.032	0.082	-0.053	0.140	-0.040
	0.834	*0.208*	*0.650*	*0.554*	*0.768*	*0.201*	*0.629*	*0.337*
Medium income	0.096	-0.040	-0.211	-0.045	0.118	-0.035	0.187	-0.013
	0.770	*0.201*	*0.185*	*0.421*	*0.753*	*0.169*	*0.647*	*0.761*
High income	-0.247	-0.285	-0.010	-0.123	-0.144	-0.261	-0.192	-0.276
	0.366	*0.000*	*0.946*	*0.040*	*0.665*	*0.000*	*0.575*	*0.000*
Urban	0.072	-0.193	0.065	-0.076	0.056	-0.176	0.072	-0.200
	0.451	*0.032*	*0.590*	*0.091*	*0.469*	*0.062*	*0.508*	*0.059*
Health status	-0.084	0.098	-0.550	-0.680	-0.706	-0.529	-5.631	-2.725
	0.446	*0.039*	*0.333*	*0.001*	*0.006*	*0.000*	*0.000*	*0.000*
Household size	0.220	0.145	0.282	0.152	0.209	0.133	0.123	0.116
	0.600	*0.060*	*0.130*	*0.011*	*0.641*	*0.066*	*0.796*	*0.196*
Medium education (spouse)	-0.120	0.016	0.050	0.020	-0.117	0.006	0.137	0.123
	0.357	*0.657*	*0.778*	*0.749*	*0.403*	*0.882*	*0.317*	*0.014*
High education (spouse)	-0.131	0.042	-0.038	0.058	-0.131	0.040	0.626	0.408
	0.242	*0.320*	*0.857*	*0.442*	*0.269*	*0.452*	*0.000*	*0.000*
Health status spouse	-0.219	-0.049	0.105	-0.001	0.005	0.106	-4.732	-1.799
	0.282	*0.330*	*0.885*	*0.997*	*0.990*	*0.771*	*0.000*	*0.000*
Health status spouse*female	0.025	0.083	-0.278	-0.229	-0.358	-0.434	-0.418	-1.470
	0.838	*0.343*	*0.786*	*0.540*	*0.482*	*0.009*	*0.077*	*0.000*
LogL	-6614864.2	-32192476	-2419.161	-9095.505	-6570426.4	-31290898	-6571422.9	-31265249
N	41589	41589	40381	40381	40381	40381	40381	40381

Notes: Weighted results. P-value reported below respective coefficient. All regressions include country-fixed effects, age dummies and sector dummies. Model(1): Main specification; Model(2): Specification with unobserved heterogeneity; Model(3): Specification with the Health index variable; Model(4): Specification with the self-reported health variable corrected for differential item functioning. Source: SHARE release 2.0.1.

Table 9: Generalized ordered probit regression of self-assessed health on health indicators.

Variables	**Generalised ordered probit**	**Implied disability weight**
Heart attack	0.588	0.081
	0.035	
High blood pressure	0.228	0.031
	0.022	
High blood cholesterol	0.117	0.016
	0.025	
Stroke or celebral vascular desease	0.559	0.077
	0.063	
Diabetes	0.478	0.066
	0.035	
Chronic lung disease	0.459	0.063
	0.053	
Asthma	0.193	0.027
	0.053	
Arthritis or rheumatism	0.463	0.064
	0.026	
Osteoporosis	0.380	0.052
	0.041	
Cancer or malignant tumor	0.587	0.081
	0.052	
Stomach, duodenal or peptic	0.214	0.029
	0.048	
Parkison disease	0.819	0.113
	0.162	
Hip or femoral fracture	0.231	0.032
	0.094	
Other condition	*0.531*	0.073
	0.030	
Ever treated for depression	0.356	0.049
	0.022	
Hearing	*0.175*	0.024
	0.011	
Bite on hard foods	0.045	0.006
	0.006	
Eyesight	*0.265*	0.036
	0.012	
Orientiring	-0.108	-0.015
	0.019	
Grip strenght	*-0.119*	-0.016
	0.009	
Body mass index	0.028	0.004
	0.003	
LogL	–30607.40	
N	26572	

Notes: Weighted results. P-value reported below respective coefficient. Additional controls are gender and age. Thresholds parameter estimates not shown. Source: SHARE release 2.0.1.

Table 10: Results from ordered probit and chopit model with covariates in threshold equation.

Variables	**Ordered Probit**	**Chopit**	
	Beta	*Beta*	*Gamma (Threshold1)*
Sweden	-0.876	-1.580	-0.384
	0.032	0.074	0.079
Netherlands	-0.450	-0.789	-0.147
	0.032	0.065	0.073
Spain	-0.144	-0.450	-0.409
	0.035	0.073	0.087
Italy	-0.169	-0.436	-0.131
	0.035	0.076	0.087
France	-0.230	-0.395	-0.120
	0.030	0.060	0.072
Greece	-0.425	-0.573	-0.063
	0.031	0.062	0.070
Belgium	-0.446	-0.652	-0.162
	0.029	0.061	0.070
Low education	0.502	0.751	0.289
	0.030	0.058	0.063
Medium education	0.281	0.444	0.093
	0.030	0.060	0.062
Female	0.103	0.106	-0.087
	0.020	0.040	0.045
Age	0.056	0.108	0.028
	0.013	0.025	0.030
$(\text{Age})^2$	-1.948	-5.010	-2.005
	0.950	1.826	2.282
Vignettes			
Theta1		1.794	
		0.041	
Theta2		0.640	
		0.031	
ln(sigma)		-0.127	
s.e.		0.007	
LogL		-74982.540	
N		27122	

Notes: Weighted results. P-value reported below respective coefficient. Source: SHARE release 2.0.1.

Figure 1: Empirical hazard function by gender.

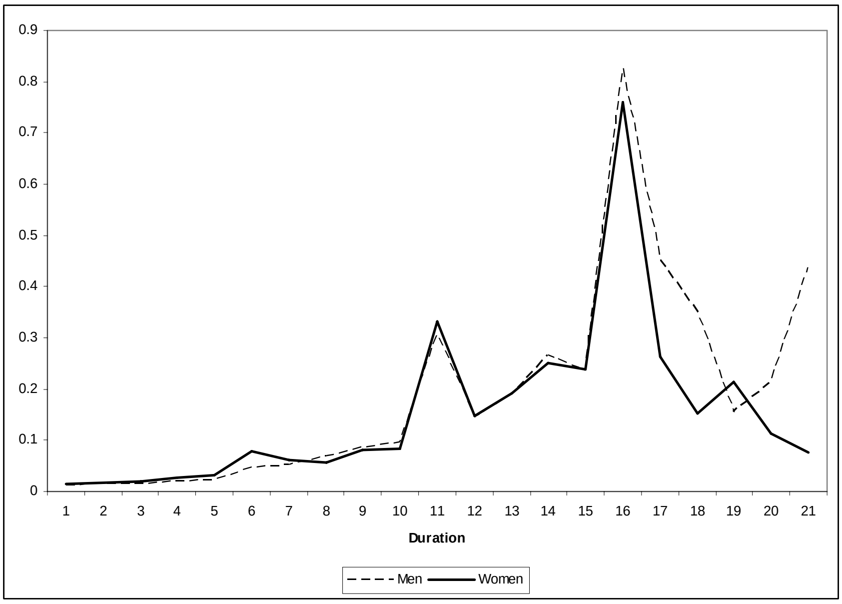


Figure 2: Fraction of couple retiring by age and year of retirement difference - (50-75 year of age).

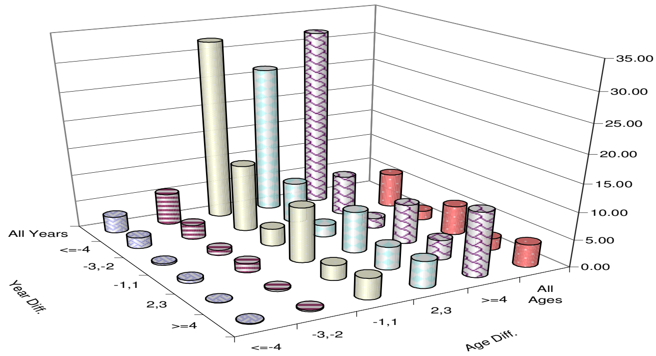


Figure 3: Empirical hazard function by partners education; women with less than college education.

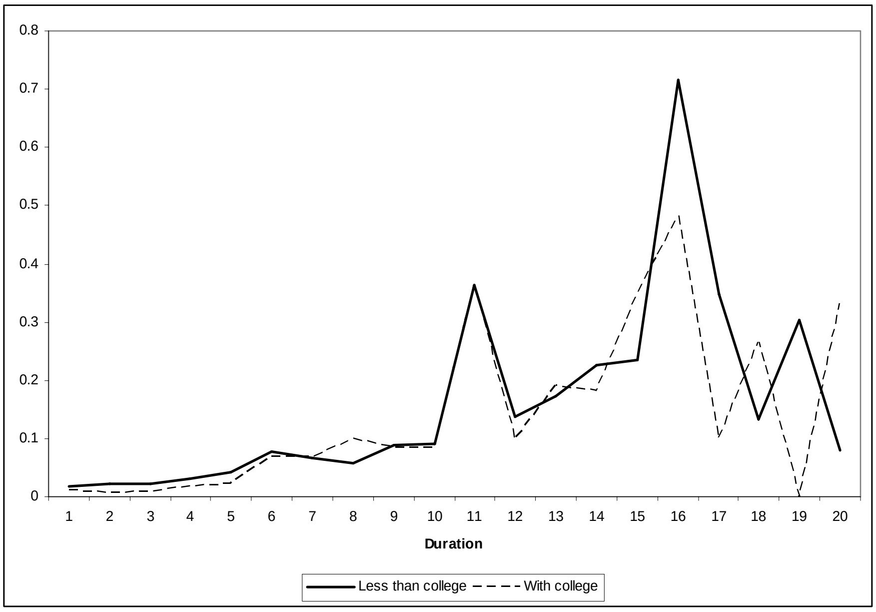


Figure 4: Empirical hazard function by partners education; women with college education.

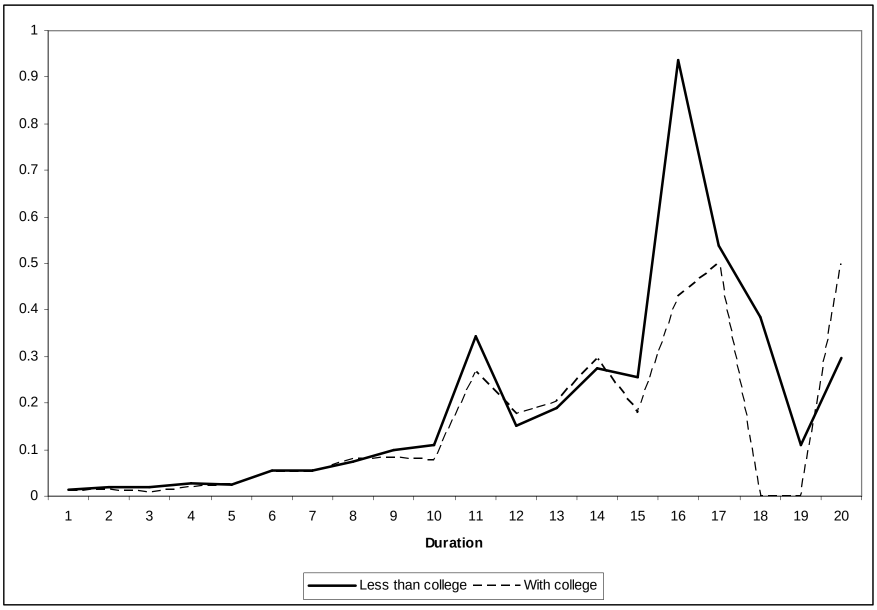


Figure 5: Empirical hazard function by partners education; men with less than college education.

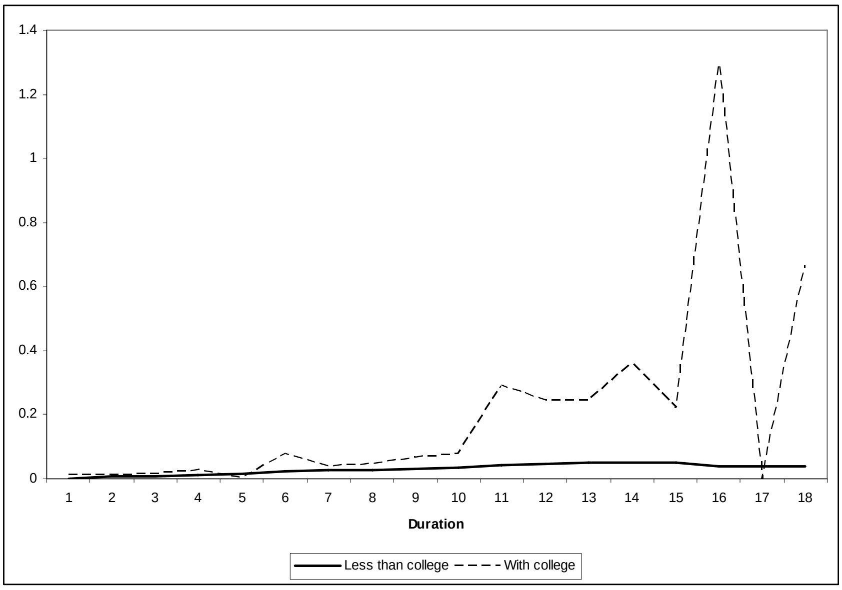


Figure 6: Empirical hazard function by partners education; men with college education.

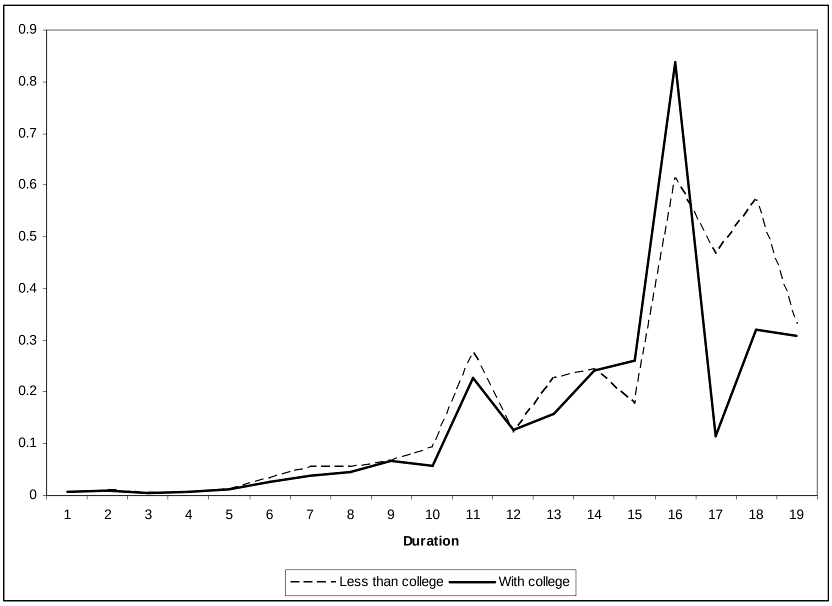


Figure 7: Empirical hazard function by partners employment condition; women.

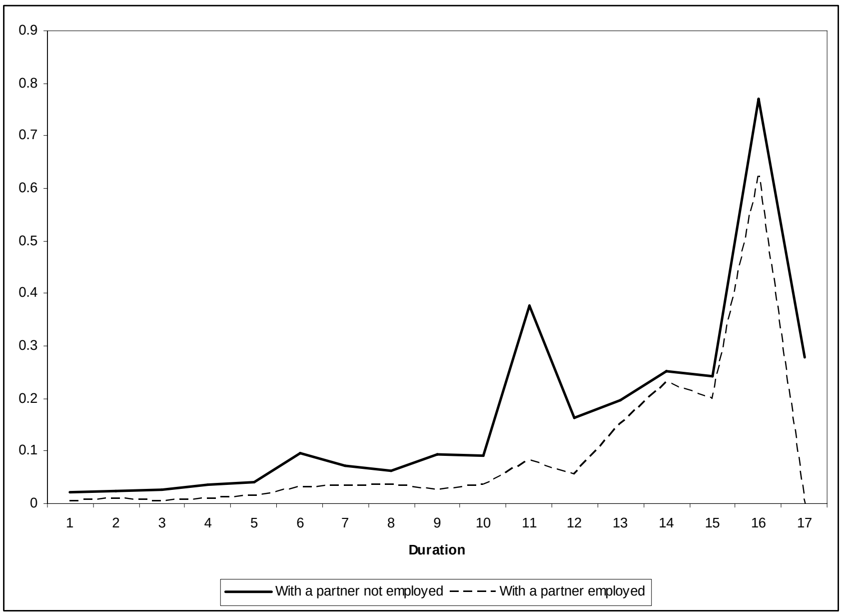
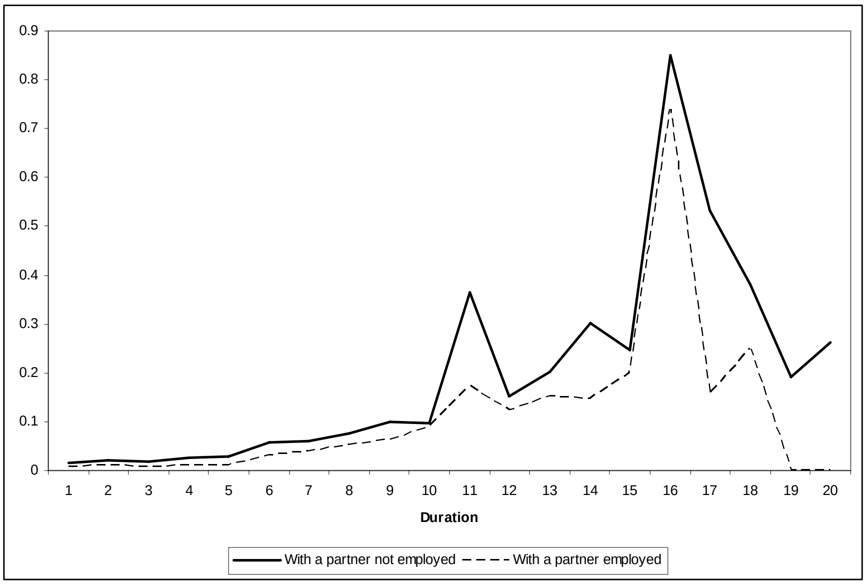


Figure 8: Empirical hazard function by partners employment condition; men.



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