Gaining weight through retirement? Results from the SHARE survey

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Abstract

This paper estimates the causal impact of retirement on the Body Mass Index (BMI) of adults aged 50–69 years old, on the probability of being either overweight or obese and on the probability of being obese. Based on the 2004, 2006 and 2010–11 waves of the Survey of Health, Ageing and Retirement in Europe (SHARE), our identification strategy exploits variation in European Early Retirement Ages (ERAs) and stepwise increases in ERAs in Austria and Italy between 2004 and 2011 to examine an exogenous shock to retirement behavior. Our results show that retirement induced by discontinuous incentives in early retirement schemes causes a 12 percentage point increase in the probability of being obese among men within a two- to four-year period. We find that the impact of retirement is highly non-linear and mostly affects the right-hand side of the male BMI distribution. Additional results show that this pattern is driven by men retiring from strenuous jobs and by those who were already at risk of obesity. In contrast, no significant results are found among women.

Keywords: Body Mass Index; Obesity; Retirement; Instrumental Variables
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1 Introduction

In its 1998 report, the World Health Organization (WHO) ranked the obesity epidemic among the ten leading global public health issues. Obesity rates in the world have more than doubled over the last 30 years (WHO (2012)). In the 27 European Union member states, approximately 60% of the adult population – 260 millions of adults – is either overweight (Body Mass Index (BMI) from 25 to 29.9 kg/m^2) or obese (BMI 30 kg/m^2 and above) (International Obesity Task Force (IASO/IOTF (2010)). Obesity has become a pan-European epidemic (IASO/IOTF (2002)), and prevalence rates in the EU-27 range from 7.9% in Romania to 24.5% in the United Kingdom (OECD (2010)).

Obesity is a risk factor for numerous highly prevalent and costly chronic diseases (cardiovascular diseases, type II diabetes, hypertension and certain types of cancer) and for disability. It reduces quality of life, shortens life expectancy and lowers levels of labor productivity (Must et al. (1999); Rosin (2008)). Moreover, it places a heavy financial burden on both individuals and society – particularly on public transfer programs and private health plans (Finkelstein et al. (2003)). At the individual level, Emery et al. (2007) find that healthcare costs for French obese individuals are, on average, twice those for normal-weight individuals. At the aggregate level, obesity-related healthcare expenditures account for 1.5 to 4.6% of total health expenditures in some European countries (see Schmid et al. (2005) and Emery et al. (2007) for evidence on France and Switzerland, respectively).

In most European countries, obesity rates reach their peak around age 60.5[1] (Sanz-de Galdeano (2005)). Recent studies have highlighted the particularly strong impact of overweight, obesity and increased BMI on morbidity and disability among adults aged 50 and older (Jenkins (2004); Andreyeva et al. (2007); Peytremann-Bridevaux and Santos-Eggimann (2008)), thereby attracting policymakers' attention to the substantial burden that obesity places on the general health and autonomy of adults over 50.

Understanding the causes of obesity among the elderly is therefore a key issue. Unlike other age groups – such as children or adolescents – the elderly have not yet received much attention. As the elderly are characterized by low labor participation and high job exit rates, one might wonder whether transitions out of employment have an impact on the weight trajectories of individuals over 50 years old. In this paper, we focus on the most common transition out of employment, i.e., retirement.

There are some reasons to believe that retirement might trigger weight changes. The Grossman model of the demand for health (Grossman (1972)) is consistent with the interpretation that individuals are likely to adopt health-producing activities after retirement.

[1]This figure does not allow us to disentangle age and cohort effects. Using the 2004, 2006 and 2010 waves of the Survey of Health, Ageing and Retirement in Europe (SHARE), we find that obesity rates among those 50–70 years old reach their peak between age 55 and 65 for all cohorts born between 1940 and 1954.

Although retirees have tighter budget constraints, they have more time to allocate to leisure: they may engage in physical activity or healthier diets, for instance, which are time consuming but not money consuming. Empirical findings seem to corroborate this view. In a three-year follow-up of French middle-aged adults, Touvier et al. (2010) find that retirement is associated with an increase in leisure-time physical activities of moderate intensity, such as walking. For food intake, the findings are more mixed. Using US longitudinal data, Chung et al. (2007) find that households spend less on eating out ($10 per month, on average) following retirement, while their monthly spending on food at home does not change. In a recent review of the literature, however, Hurst (2008) argues that due to an increase in home food production, overall food intake does not decline following retirement. Overall, these results suggest that retirement would affect weight through changes in physical activity rather than via food consumption.

At the same time, new retirees may lose some incentive to invest in their health, as their income (pension benefits) is no longer health dependent. This could lead to lower health investments, and to a lower health stock in the long-run. Moreover, retirement might also increase the risk of social isolation and depression (Friedmann and Havighurst (1954); Bradford (1979)), potentially leading individuals to reduce the efforts devoted to health-producing activities or to develop addictive behaviors (e.g., alcohol or tobacco consumption). Finally, the loss of a structured use of time may also encourage snacking in-between meal times and sedentary habits (television watching). In the study mentioned before, Touvier et al. (2010) find that retirement is also associated with an increase in time spent watching TV.

Overall, the direction of the effect is not clear. Besides, the effect of retirement is likely to be highly heterogeneous, particularly across job types. As retirement induces a direct reduction in job-related exercise, individuals who retire from strenuous jobs are at a higher risk of weight gain if they do not compensate by increasing their leisure-time physical activity or decreasing their food intake. Conversely, retirees from sedentary jobs may lose weight if their leisure-time activities after retirement are more physically demanding than their time at work.

The purpose of the present paper is to estimate the causal impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese. Identifying such a causal impact is problematic in the presence of confounding factors and reverse causality. Indeed, retirement is often a choice, and it is often based on unobservable characteristics, which may be correlated with weight (e.g., time preference[2], deteriorating health or psychological status). Reverse causality may also be a concern. As overweight and obese individuals are, on average, paid less, promoted less (Cawley (2004); Morris (2006); Brunello and d'Hombres (2007); Schulte et al. (2007)) and experience worse health, their incentives to retire might be higher than normal-weight individuals. Burkhauser and Cawley (2006) show that fatness and obesity are indeed strong predictors of early receipts of old-age

[2]See Smith et al. (2005), Anderson and Mellor (2008) and Ikeda et al. (2010) for empirical evidence of the positive relationship between time preference and BMI.

benefits in the USA.

To address this endogeneity issue, we use an instrumental variable approach. Our identification strategy exploits the fact that as individuals reach the Early Retirement Age (ERA) at which they are entitled to receive either reduced pensions or full pensions – conditional on a sufficient number of years of social security contributions – the probability that they retire strongly increases. In other words, this discontinuous incentive in retirement schemes provides a strong exogenous shock on retirement behavior. We exploit variation in ERAs across European countries and over time (in countries that implemented stepwise increases in ERAs over the study period) to solve the major identification problems related to confounding factors and reverse causality. We implement a fixed effects instrumental variable model in order to control for both time-invariant factors (such as genetics) and time-varying omitted variables and/or reverse causality. This allows us to estimate the short-term causal effect of a transition to retirement on weight. We use the 2004, 2006 and 2010 waves of the SHARE (the Survey of Health, Ageing and Retirement in Europe). Our results show that retirement induced by early retirement rules causes a 12 percentage point increase in the probability of being obese within a two- to four-year period[3] among men. We find that the impact of retirement is highly non-linear and mostly affects the right-hand side of the male BMI distribution. Additional results show that this effect is driven by men retiring from strenuous jobs and by those who were already at risk of obesity. In contrast, no significant results are found among women. Overall, our results show that most individuals do not gain weight through retirement. The impact of retirement is never significant among men who retire from sedentary jobs or among women. However, a small subgroup of retirees – i.e., men who retire from strenuous jobs and already at risk of obesity – may lose potential health benefits in retirement by risking obesity.

This paper relates to several strands of the literature. First and foremost, it contributes to the literature on the effects of retirement on weight. Most papers in this body of literature estimate mere correlations, disregarding the possibility that retirement is endogenous. The results have been quite consistent so far. Nooyens et al. (2005) find that the effects of retirement on changes in weight and waist circumference depend on one's former occupation: weight gain is higher among men who retire from an active job. Forman-Hoffman et al. (2008) find no significant relation for men, but they do observe weight gain for women retiring from blue-collar jobs. Gueorguieva et al. (2010) find a significant increase in the slopes of BMI trajectories only for individuals retiring from blue-collar occupations. To the best of our knowledge, Chung et al. (2009) and Goldman et al. (2008) are the only studies addressing the endogeneity issue. Both studies use longitudinal data from the Health and Retirement Study (HRS) – the USA's equivalent of the European SHARE – and estimate fixed effect models with instrumental variables. They use Social Security and Medicare eligibility ages (62 and

[3] There is a two-year period between the 2004 and 2006 SHARE waves and a four-year period between the 2006 and 2010 waves.

65, respectively) as instruments for retirement.[4] Chung et al. (2009) conclude that people who are already overweight and those with lower wealth retiring from physically demanding occupations gain weight (an increase of 0.294 and 1.013 BMI units, respectively). Goldman et al. (2008) find that males retiring from strenuous jobs gain weight (0.5 BMI units) during the first six years of retirement, while sedentary job leavers experience virtually no weight change (0.1 BMI units over a 10-year period). Our study improves this literature in three respects: first, we identify a causal effect of retirement on weight, while most papers (with the notable exceptions of Chung et al. (2009) and Goldman et al. (2008)) document mere correlations. Second, variation in ERAs within Europe and over time allows us to explore the effect of retirement on weight at different ages rather than just between the ages of 62 and 65, as in Chung et al. (2009) and Goldman et al. (2008). Weaker assumptions in terms of weight trajectories by cohort and age are needed in our empirical setup. Finally, our paper is the first to exploit European data. Most of the above-mentioned studies – except Nooyens et al. (2005) – use US data from the HRS. Given their differences in terms of labor markets, social security schemes and social policies, it is not clear whether the results obtained for the USA should hold for Europe.

This paper also relates to a substantial, recent body of literature that explores the effects of retirement on general health and health-related outcomes – e.g., mental health, cognitive functioning and well-being. The results in this literature are quite ambiguous, and whether retirement has a detrimental effect on health is still an open question. Most papers, however, support the idea that retirement improves general health (Charles (2004); Neuman (2008); Coe and Lindeboom (2008); Coe and Zamarro (2011)[5]; Blake and Garrouste (2012); Eibich (2015)). But not all findings in the literature are consistent with this interpretation (see, for instance, Dave et al. (2006) and Behncke (2011)[6]). Besides, there is strong evidence that retirement leads to a significant negative effect of retirement on cognitive functioning (Rohwedder et al. (2010); Bonsang et al. (2012)). An interesting way to solve this retirement puzzle is to look, as we do, at behavioral outcomes following retirement. These behavioral outcomes can be rapidly modified over the short run and precede longer term health outcomes (such as chronic diseases and mortality). Rather than looking at the long-term health consequences of retirement – which may not easily be disentangled from the effects of age –

[4] Chung et al. (2009) use spousal pension eligibility as an additional instrument. However, recent work highlights asymmetries in spouses' retirement strategies (Gustman and Steinmeier (2009); Stancanelli (2012)). Using spousal pension eligibility as an instrument might thus be a questionable strategy.

[5] Our identification strategy is in the spirit of Coe and Zamarro (2011), who use the 2004 SHARE wave and country-specific early and full retirement ages as instruments for retirement behavior. However, we offer an improvement in two respects. First, we take advantage of the panel structure of the SHARE data, which allows us to control for individual time-invariant, unobservable characteristics. Second, reforms implemented in Austria and Italy over the 2004–2011 period provide additional variation in ERAs. Finally, rather than investigating the effects of retirement on health, we consider the effects of retirement on under-studied dimensions of health and major risk factors for numerous diseases among the elderly: weight change and obesity.

[6] Dave et al. (2006) find negative effects of retirement on both objective and subjective measures of health; Behncke (2011) shows that retirement significantly increases the risk of being diagnosed with a chronic condition (cardiovascular disease and cancer) and worsens self-assessed health.

we analyze how weight change is modified following retirement over the short run. We also investigate the extent to which this effect is heterogeneous across several dimensions, such as gender, occupational strenuousness, and baseline weight category. As weight change is likely to be an important mechanism by which retirement affects health, this paper contributes to a recent and growing body of literature by exploring one of the potential mediating channels between retirement and health.[7]

Finally, this paper contributes to a growing body of literature that investigates the impacts of various dimensions of professional activity on body weight and obesity, such unemployment (Marcus (2012)), working conditions (Lallukka et al. (2008b)), occupational mobility (Ribet et al. (2003)), job insecurity (Muenster et al. (2011)), physical strenuousness of work (Böckerman et al. (2008)), working overtime (Lallukka et al. (2008a)), and income (Cawley et al. (2010), Schmeiser (2009), Colchero et al. (2008)).

This paper proceeds as follows. Section 2 presents our empirical approach, and Section 3 describes the data (the 2004, 2006 and 2010 waves of the SHARE). Section 4 presents the results, and Section 5 concludes.

2 Empirical approach

We investigate the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese. As a first step, we pool the observations from the 2004, 2006 and 2010 waves of the SHARE and estimate the following standard Pooled Ordinary Least Squares (POLS) model:

$$Y_{it} = \alpha + \gamma R_{it} + X_{it}\beta + D_i + D_t + u_{it}, \tag{1}$$

where Y_{it} denotes the weight outcome of individual i at time t.[8] R_{it} is a binary variable indicating whether individual i is retired at time t, X_{it} a vector of individual characteristics either time-varying or time-invariant, D_i a country dummy, D_t a time dummy, and u_{it} the error term.

However, retirement status R_{it} may be correlated with the error term u_{it}, in which case the POLS estimate of γ is inconsistent. Endogeneity may arise from several sources. Omitted variables, such as unobservable time preference or latent health status, may have an impact on both retirement status and weight. Similarly, reverse causality may be a concern: obese individuals are more likely to seek early retirement benefits (Burkhauser and Cawley (2006)).

[7] See, for instance, Eibich (2015) on the respective roles of exercise, sleep and work-related strain in the retirement-health relationship.

[8] Y_{it} can be either continuous (BMI) or binary (being either overweight or obese/being neither overweight nor obese; being obese/not being obese). The POLS (presented) and pooled probit models (not presented but available upon request) yield very similar results when the dependent variable is binary.

Faced with these endogeneity problems, we consider the following fixed effects (FE) model:

$$Y_{it} = \alpha + \gamma R_{it} + K_{it}\beta + D_t + \alpha_i + v_{it}, \tag{2}$$

where Y_{it} denotes the weight outcome, R_{it} the individual retirement status, K_{it} a vector of time-varying individual characteristics, D_t a time dummy, α_i an individual fixed effect – including the country fixed effect – and v_{it} the error term.

The FE model allows regressors to be endogenous, provided that they are correlated only with α_i, the time-invariant component of the error, but not with the idiosyncratic error v_{it}. If some unobservable time-varying characteristics are correlated with R_{it}, however, $\hat{\gamma}$ continues to be biased. Moreover, reverse causality is still a concern.

To address the endogeneity problem, we estimate a Fixed Effects Instrumental Variable (FEIV) model. This model allows us to control for both time-invariant factors (such as genetics, food preferences over the life-course or time preferences) and time-varying omitted variables (such as health deterioration) as well as for reverse causality. Our identification strategy exploits the fact that as individuals reach the ERA for their country, the probability that they retire strongly increases.[9] This exogenous shock to retirement behavior allows us to estimate the causal impact of a transition to retirement on weight over the short run – within a two- to four-year period.[10]

Retirement decisions in industrialized countries depend on a number of institutional features. In particular, the earliest age at which individuals are entitled to pension benefits has been shown to exert a powerful influence on their retirement behaviors (Gruber and Wise (1999)). This ERA is defined as the earliest age at which individuals are entitled to either reduced pensions or full pensions – conditional on a sufficient number of years of social security contributions. The ORA is the age at which workers are entitled to either minimum guaranteed pensions or full old-age pensions irrespective of their contributions or work histories. The ORA appears to be less important in predicting retirement behavior than the ERA (Gruber and Wise (1999)). Few individuals actually work until the ORA. As a consequence, there is a gap between the ORA and the average effective age at which older workers withdraw from the labor force in almost all industrialized countries.

The earliest, official and effective retirement ages in Europe are presented in Table 1. As

[9] One could use a health shock between two subsequent waves of the survey as an alternative instrument for retirement behavior. However, in practice, the exclusion restriction – according to which this health shock does not affect weight except through the increased probability of retiring – is not likely to hold in the data.

[10] There is a two-year period between the 2004 and 2006 SHARE waves and a four-year period between the 2006 and 2010 waves. In this setup, we assume that the effect of retirement on weight over a two-year period is the same as over a four-year period. Another option is to consider the two- and four-year periods separately and estimate the regressions using two different samples. However, we chose to consider a single sample (i) to obtain a larger sample, which is important when investigating the heterogeneous impact of retirement (ii) to obtain three observations per individual over the period.

evidenced in columns (1) and (2), the ORA varies very little across countries and genders. In contrast, the ERA varies quite a lot across countries and genders (columns (3) and (4)). There is no early retirement option in Denmark, and early retirement was abolished in 2005 in the Netherlands (OECD, 2011). Effective retirement ages are lower than the ORAs in almost every country (see columns (5) and (6) for men and women, respectively), suggesting that ORAs are less binding than are ERAs in most cases. A number of countries in our sample implemented substantial reforms to their ERAs over the study period. In Austria, for instance, the 2004 pension reform introduced a gradual increase in the ERAs for men and women. Immediately before the reform, workers in Austria could retire at ages 61.5 (men) and 56.5 (women). After the reform, the ERA increased by two months for each quarter of birth for men (women) born in the first two quarters of 1943 (1948). Following these increases, the ERAs were increased by one month for each quarter of birth for men born in the third quarter of 1943 and later and for women born in the third quarter of 1948 and later. Furthermore, the 2004 pension reform created special corridor pensions for men born in the last quarter of 1943 and later, thereby making the ERA beyond age 62 non-binding in many cases (Manoli and Weber (2012)). Italy also introduced a stepwise increase in the minimum age to request early retirement, from age 57 in 2004 to age 60 in 2011. Finally, Belgium implemented a substantial reform to the ORA for women, which increased from age 63 in 2004 to 64 on January 1, 2006 and to age 65 on January 1, 2009 (OECD, 2011). More information on the Austrian, Italian and Belgian reforms can be found in the notes for Table 1.

Our main analysis takes advantage of ERA variation across countries and over time to explore the causal effect of retirement on weight at different ages. An extended analysis uses ERAs and ORAs as joint instruments for retirement behavior. The latter specification is not our preferred one, however, as ORAs appear to be less important in predicting retirement behavior than ERAs (see Section 4). In our main analysis, we thus instrument retirement status R_{it} using a dummy variable indicating whether individual i's age at time t is above or below the ERA in force at time t in his country c.[11] Let age_{it} be individual i's age at time t and ERA_{ct} be the ERA in i's country c at time t. Our instrument is defined as:

$$Z_{ict} = \mathbb{1}_{\{age_{it} > ERA_{ct}\}}. \tag{3}$$

A good instrument should be strongly correlated with actual retirement behavior but should not directly affect weight outcomes.

As shown in Table 1, Z appears to be well correlated with retirement status. Suggestive evidence is provided by columns (7) and (8): in each country, there is a large gap in the fraction of individuals retired before and after the ERA cutoff. For example, only 17% of individuals in the pooled sample are retired before age 60 in France – which is when they are first entitled to pension benefits – but this proportion increases to 88% after age 60. Taking

[11] For countries where there were no increases in ERAs over the study period – i.e., all countries, except Italy and Austria – our identification exploits a non-linear version of age, therefore identifying using the functional form of age.

advantage of the panel structure of our data, for each country, we then compute the proportion of individuals retiring when reaching their country's ERA between two subsequent waves of the survey (see column (9)). This proportion is high in most countries. For instance, in Belgium, 34.3% of individuals reaching age 60 between two waves of the survey actually retire between these two waves.

At the same time, after controlling for age, reaching the ERA cutoff is highly unlikely to be correlated with weight outcomes, except through the increased probability of retiring. This exclusion restriction holds if we assume there are no discontinuities in the weight trajectories at the ERAs, except for the effect of retirement at these ages. In our setup, the variation in ERA is large across countries, between genders and (for Austria and Italy) over time. That the ERA cutoff may systematically be related to a discontinuity in population health is not likely to hold in our sample.[12] We show in the robustness section that it is not likely to be the case.

Equation (2) is then estimated by fixed effects two-stage least squares, where Z_{ict} is an instrument for R_{it}. In the first stage, retirement status R_{it} is regressed on Z_{ict} and other covariates, and individual fixed effects are included. In the second stage, equation (2) is estimated using a FE regression/FE linear probability model, where R_{it} is replaced with its predicted value from the first stage. The covariance matrix of $\hat{\gamma}$ is corrected accordingly.

Our FEIV estimate $\hat{\gamma}$ can be given a causal interpretation as a Local Average Treatment Effect (LATE) without requiring a constant treatment assumption. In our case, the "treatment" is defined as retiring between two subsequent waves of the survey. Specifically, $\hat{\gamma}$ is identified for the subset of individuals whose behavior is shifted by our instrument, i.e., the compliers. In this setup, compliers are (i) individuals who became eligible for early retirement schemes between two subsequent waves of the survey and *did* retire but would not have retired had they not become eligible and (ii) individuals whose eligibility for early retirement schemes did not change between two subsequent waves of the survey and *did not* retire but would have retired had they become eligible. As the ERA is probably more binding for individuals with long careers, we expect compliers to be less educated.

Overall, our estimation strategy allows to us to measure the causal effect of a transition to retirement on weight within a two- to four-year period among this subpopulation of compliers. In an extended analysis, we take ERAs and ORAs as joint instruments for retirement behaviors to produce a more general representation of the population of retirees (i.e., includ-

[12] Positions on increasing retirement age are salient variables when people choose a candidate or political party to support. If people take a position based on their health (or weight) status, the ERA may be correlated with the general health level of the population, and the instrument might be correlated with the error term. This criticism applies to all studies using country-level regulations as an exogeneous shock to retirement behavior, and particularly to studies in a single-country setting. This issue is potentially less of a concern in a cross-country analysis where the variation in ERA is large across countries, between genders and over cohorts.

ing compliers at later ages).

Our empirical setup allows us to explore the effect of retirement on a wide range of ages, not just the ages of 62 and 65 as in studies of the USA. Moreover, weaker assumptions in terms of weight trajectories by cohort and age are needed for this setup.

3 Data

3.1 Presentation of the sample

We use data from the SHARE, a multidisciplinary and cross-national panel database containing individual information on health, socio-economic status and social and family networks. Approximately 85,000 individuals over 50 years old and their spouses/partners (independent of their age) from 19 European countries (including Israel) have been interviewed so far. Four waves have been conducted, and further waves are being planned to take place on a biennial basis. We use the 2004, 2006 and 2010 SHARE waves.[13] To produce a balanced panel, our sample includes the ten European countries that took part in the 2004 SHARE baseline survey and participated in the 2006 and 2010 waves, i.e., Austria, Germany, Sweden, the Netherlands, Spain, Italy, France, Denmark, Switzerland and Belgium.

Our sample contains all individuals interviewed in waves 2004, 2006 and 2010[14], aged 50 to 69 years old[15], who declared in each wave that they were either employed or retired. In other words, we only consider the traditional and most frequent pattern of retirement wherein individuals transition directly from work to retirement. Transitions from employment to unemployment, disability or inactivity are thus excluded. We also exclude transitions from retirement to employment, unemployment, disability or inactivity. We further eliminate individuals who have not worked since age 50, regardless of whether due to individual choice or physical or mental limitations. Our main specification exploits the ERA as an instrument for retirement behavior. As there is no early retirement option in Denmark and early retirement was abolished in 2005 in the Netherlands, both countries are excluded from the main sample.[16] Finally, we exclude individuals reporting a height shorter than 1.20 meters as well as individuals reporting weights below 30 kilograms or above 200 kilograms. Overall, our

[13]The 2008–2009 wave, SHARELIFE, is a retrospective survey that focuses on people's life histories. Although it can be linked to the existing SHARE data, it is not of direct use here and is not included.

[14]We thus consider a balanced panel. Attrition rates are rather high in SHARE – 30% between the 2004 and 2006 waves of the survey. High attrition rates are a concern if non-response is systematically related to weight. We show in the robustness section that this is not likely to be the case. Additional robustness checks show that our results do not significantly vary when re-estimating our regressions using an unbalanced panel.

[15]The 50–69 age window broadly corresponds to the ages at which individuals reach the ERA in their country and become entitled to pension benefits.

[16]Our results are virtually unchanged (although slightly less significant) when including Denmark and the Netherlands in the sample and taking full retirement ages as the earliest ages at which people are entitled to retirement benefits.

dataset contains 2599 individuals[17] from eight countries (Austria, Germany, Sweden, Spain, Italy, France, Switzerland and Belgium) across the three waves.

3.2 Variables

We use a question about self-declared current job situation to determine whether an individual is retired. According to this definition, anyone who declares herself as retired, regardless of whether she has been in a paid job during the month preceding the interview – even for a few hours – is considered retired. Conversely, anyone who declares herself to be employed or self-employed is considered to be currently working.[18] Self-declared retirement status seems to be reliable in the SHARE: it is strongly associated with the eligibility for either public or private pensions in the dataset.[19] We also consider alternative definitions of retirement as robustness checks (see Section 4.6).

Tables 2 and 3 provide summary statistics for the full sample – pooled over the 2004–2010 period – for men and women, respectively. Each table also presents characteristics for the individuals continuously employed across waves (column (2)), continuously retired across waves (column (3)), or who retired across waves (column (4)). According to Tables 2 and 3, 57% of men and 60% of women in the full sample were employed or self-employed, with the rest being retired. Eight hundred and sixteen individuals (23% of the individuals working in 2004) retired between 2004 and 2010.

BMI is calculated for each wave as the self-declared weight in kilograms divided by the square of the self-declared height in meters (kg/m^2). We derive clinical weight categories from the BMI: underweight (BMI less than 18.5 kg/m^2), normal (BMI from 18.5 to 24.9 kg/m^2), overweight (BMI from 25 to 29.9 kg/m^2) and obese (BMI 30 kg/m^2 and above). We also compute the individual weight change (in kg) as well as a dummy variable indicating whether the individual experienced a weight change of at least 10% between two subsequent waves of the survey. BMI is a rather crude measure of body composition, as it does not distinguish fat from lean mass (Prentice and Jebb (2001); Burkhauser and Cawley (2008)). However, it has been shown to be highly correlated with more precise measures of adiposity. When self-reported – as is the case here – BMI may also suffer from measurement error (Niedhammer et al. (2000); Burkhauser and Cawley (2008)). Following Brunello et al. (2013) and Sanz de Galdeano (2007), we note that the rank correlation between country-level self-reported and objective measures of weight, however, is very high in Europe.

[17]Once conditioning on having no missing value on weight, height and any covariate included in the model, our sample goes down to 2406 individuals across the three waves (1089 men and 1317 women), i.e., 7218 observations in the pooled sample (3951 men and 3267 women).

[18]The SHARE also includes information about the year and month of retirement. This measure is not reliable in our data, so we do not use it. Hence, we know whether a given individual retires between two waves of the survey, but we do not have any information about the exact month or year of retirement.

[19]Of the 3020 retired individuals in the pooled sample, 84% declared that they had received income from either a public or occupational old age pension during the year preceding the interview.

The average BMI of the full sample was 26.93 kg/m^2 for men and 25.72 kg/m^2 for women, slightly above the overweight threshold in both cases. Eighteen percent of men in the full sample were obese, 50% were overweight, 32% were normal weight, and less than 1% were underweight. As for women, 16% were obese, but less than 32% were overweight and 50% were normal weight. Interestingly, while only 15% of men employed in all waves were obese, 20% of men retired in all waves were obese. The same pattern was found for women (the corresponding figures are 14% and 23%, respectively). This large gap is probably best explained by the fact that individuals employed in all waves are, on average, younger than individuals retired in all waves. However, it suggests that those who are 50–69 years old undergo serious weight changes around the retirement age.

Additional descriptive statistics seem to support the claim that those 50–69 years old undergo serious weight changes: in the pooled sample – irrespective of retirement status – 10.4% of individuals experienced a weight change (gain or loss) of at least 10% between two subsequent waves of the survey. Seventeen percent switched from the underweight or normal weight categories to overweight or obesity between two subsequent waves of the survey, while 8% of overweight or obese individuals switched back to the underweight ot normal weight categories during the same period. These figures provide evidence of high within-individual weight variation in our sample, suggesting that weight change among the elderly can be rapid.

Interestingly, Figure 1 suggests that weight changes are even more important among individuals who retired between waves. Figure 1 plots the distribution of weight change for individuals who retired across waves against the distribution of weight change for individuals continuously employed in all waves for men and women. A simple look at each graph suggests that the distribution is flatter for individuals who retired across waves: the peak around zero – indicating no weight change – is indeed less clear-cut in both graphs (especially for men). Although the distributions are not significantly different – for either men or women – it suggests that individuals who retire experience weight change to a greater extent than do individuals who are continuously employed.

Different sets of covariates are used, depending on the model specification used (POLS, FE or FEIV). We introduce age and age squared into all specifications to properly control for the age trend and to account for potential non-linear effects of age on weight. Each specification includes marital status (lives with a spouse/partner;does not live with a spouse/partner) and time dummies for 2006 and 2010. The average age of men and women in the pooled sample was 59.8 and 59.7 years old, respectively. On average, men and women who retired during the study period were 60 and 60.4 years old, respectively. Eighty-eight percent of men in the full sample lived with a spouse or partner, while only 72% of women did so. Gender, educational level[20] (primary education/lower secondary/upper secondary/post secondary), occupation[21] (blue collar/white collar/technician/manager and professional) and

[20]Based on the 1997 International Standard Classification of Education (ISCED 97)

[21]Based on the 1988 International Standard Classification of Occupations (ISCO 88). Occupation is mea-

country dummies are only included in the POLS specification, as FE and FEIV models do not permit the identification of the effects of time-invariant variables. Summary statistics for gender, educational level, occupation and country can be found in Tables 2 and 3 for men and women, respectively. Seventeen percent of men in the full sample had completed primary education, 17% lower secondary education, 33% upper secondary education and 33% post-secondary education. The corresponding figures for women were 15%, 18%, 30% and 37%, respectively. Thirty-two percent of males in the pooled sample were in blue-collar occupations, 13% were in white-collar occupations, 20% were technicians and 35% were managers or professionals. Similarly, 18% of women in the full sample were in blue-collar occupations, 32% in white-collar occupations, 20% were technicians and 30% were managers or professionals. Men and women who retired across waves exhibited the same patterns of education and occupation as individuals in the full sample. Belgium, Sweden, France, Italy and Germany were the most represented countries in the male and female pooled samples.

In extended specifications of POLS and FE models, we control for health characteristics. As health status is co-determined with retirement as well as weight, including it in the model is likely to generate some endogeneity. To minimize endogeneity issues, we introduce lagged measures of health into our POLS and FE models.[22] Whenever introduced in our regressions, these health measures are self-assessed health (measured on a five-point scale as excellent/very good/good/fair/poor) and the Euro-D depression index (measured on a twelve-point scale, where twelve is highly depressed). Descriptive statistics for self-rated health and the Euro-D depression index can be found in Tables 2 and 3 for men and women, respectively. Men and women in the full sample exhibited similar patterns of self-assessed health. Conversely, women were, on average, more depressed than men in all samples. Individuals employed in all waves – men and women – were in better health than individuals who were retired in all waves and individuals who retired between waves.

The impact of retirement on weight is likely to be highly heterogeneous by the previous job's strenuousness. To assess the physical strenuousness of work, we use a question asking workers their opinions about the following statement: "My job is physically demanding". Four answers are available, ranging from "strongly agree" to "strongly disagree". We dichotomize the responses into strenuous work (strongly agree/agree) and sedentary work (disagree/strongly disagree). This information is only available in the SHARE for individuals who were working at baseline, i.e., 938 men and 810 women across the three waves. Among individuals who were working at baseline, 43.2% (44.1%) of men (women) declared that they worked a strenuous job (see columns (1) of Tables 2 and 3, respectively).[23] Forty-four percent

sured at baseline and is not time-varying in our data. Given that we focus on elderly workers, this seems to be a plausible assumption.

[22]By construction, POLS and FE models including lagged values of health are estimated on the smaller sample of individuals interviewed in waves 2006 and 2010 (2634 men and 2178 women).

[23]A similar proportion of men and women declared working a strenuous job. Let us underline, however, that physical demands of a "strenuous job" (as declared by the respondents) are likely to be far more intense for men than women. Focusing on individuals 45–79 years old in the EPIC-Norfolk cohort, Barnett et al.

of men who retired across waves declared that they were retiring from a strenuous job, while 41.8% of women declared so (see columns (4) of Tables 2 and 3, respectively).

Finally, we supplement our dataset with the ERA in force in each country at the time of the survey (see Table 1). We build a dummy variable indicating whether an individual's age at time t was above or below the ERA in force at time t in his country.

4 Results

4.1 Determinants of retirement

Almost 23% (816 individuals) of those working at baseline retired between 2004 and 2010. Among them, 45% (365 individuals) had reached the national ERA during the same period. This suggests that actual retirement behavior is well correlated with the ERA.

The first-stage results are reported in Table 4 for both men and women. The coefficients displayed in columns (1)–(2) are estimated by fixed effects probability models. As expected, they indicate that the ERA is an important predictor of retirement. Reaching the ERA increases the probability of retiring by 21 and 28 percentage points for men and women, respectively (both effects are significant at the 1% level). These coefficients can also be interpreted as the proportions of compliers in our sample[24] (21% among men and 28% among women), which are high. These results, combined with the F-statistics of the excluded instruments of 121.3 and 162.8 for men and women, respectively, show that reaching the ERA provides a strong exogenous shock to retirement behavior.

After controlling for these country-specific age breaks, the probability of retiring decreases with age up to a certain point after which it increases again – probably when reaching the ORA. Finally, neither time dummies for 2006 and 2010 nor marital status appear to be statistically significant for retirement behavior.

4.2 The impact of retirement on BMI, overweight and obesity

Given the differences in terms of both biological constitutions and labor market histories, we estimate separate models for men and women. Tables 5 and 6 report the POLS estimates for BMI (columns (1) and (2)), probability of being either overweight or obese (columns (3) and (4)) and probability of being obese (columns (5) and (6)) for men and women, respectively.[25] All specifications include age, age squared, time dummies for 2006 and 2010, marital status and time-invariant variables such as education, occupation and country dummies. The first

(2014) show that occupational physical activity (measured in Metabolic Equivalent of Task (MET) (h/wk)) of individuals employed in manual jobs was twice as high for men as for women (112.5 versus 56.5 MET (h/wk)).

[24] This is true because both R_{it} and Z_{ict} are dummy variables and because our model is estimated by fixed effects two-stage least squares.

[25] Standard errors are clustered at the individual level in all POLS regressions.

column of each pair presents the results of our baseline specification, while the second column presents the results after controlling for lagged measures of self-rated health and depression.

Most of the control variables are statistically significant and of the expected sign. A steep education gradient in BMI, overweight and obesity is found for women and, to a lesser extent, for men. Compared with primary education, post-secondary education is indeed associated with a lower BMI and a lower probability of being either overweight or obese as well as of being obese for both men and women. Controlling for education, occupation is not significantly associated with BMI, overweight or obesity, except for women: females in managerial or professional occupations have a lower probability of being overweight than blue-collar females. Living with a spouse or partner does not seem to be correlated with BMI or the probability of being obese but is associated with a higher risk of being either overweight or obese among men. Most country indicators are significant.[26] Surprisingly, once we control for retirement behavior, age has a small and insignificant impact on BMI, overweight and obesity.

Our baseline specification reveals a positive and significant association between retirement and weight outcomes for men and women. Retirement is positively correlated with BMI, increasing BMI by 0.36 and 0.49 units for men and women, respectively (both effects are significant at the 10% level).[27] Retirement also increases men's probability of being either overweight or obese by 4.9 percentage points (the effect is significant at the 5% level). This coefficient corresponds to a 10% increase in the probability of being overweight or obese for men (compared with the sample average). Retirement also increases women's probability of being obese by 3.3 percentage points – although the effect is marginally significant at the 15% level. This represents a 21% increase in the probability of being obese for women. These results are fairly consistent after controlling for lagged health variables, although the magnitude of the effects declines with most becoming only marginally significant. After controlling for past health, retirement among men is no longer significantly associated with BMI. However, it is still positively and significantly associated with men's probability of being either overweight or obese (a 5.2 percentage point increase, significant at the 10% level). For women, retirement is no longer statistically associated with BMI or the risk of obesity.

These correlations, however, are difficult to interpret because they potentially reflect the effects of unobserved characteristics that may affect both weight outcomes and retirement behavior. The importance of confounding factors is apparent when we look at the coefficients on retirement after estimating fixed effect regressions (see Table 7 and Table 8 for men and women, respectively). Once taking into account the potential endogeneity arising from the correlation between retirement and time-invariant unobserved characteristics, retirement is no longer significantly associated with weight outcomes among men. The sign of the coeffi-

[26] These results not shown but available upon request. Note that our results are virtually unchanged when including country*time fixed effects, suggesting that the time trend in weight among the elderly is fairly consistent across countries.

[27] For an average man measuring 1.75m and weighing 82 kg, this corresponds to a 1.1 kg weight gain. For an average woman measuring 1.63m and weighing 69 kg, this corresponds to a 1.3 kg weight gain.

cient associated with retirement becomes negative or close to zero for each weight outcome. In contrast, retirement leads to weight gain (by 0.25 BMI units, significant at the 5% level) among women, although the magnitude of the estimates declines compared to the POLS results. Not controlling for time-invariant factors – such as time preferences, which have a positive effect both on the probability of retiring and on weight gain – may indeed generate an upward bias and explain the larger effect of retirement on weight in the POLS models. In an extended specification, we control for past health deterioration by including a lagged measure of self-rated health and depression. The results are presented for men and women in columns (2), (4) and (6) of Tables 7 and 8), respectively. In contrast to POLS models, controlling for lagged health variables in the fixed effects models only marginally affects the estimates, suggesting that time-invariant health characteristics affect weight outcomes to a larger extent than past health changes.

However, the fixed effects estimates cannot be interpreted as causal: a number of omitted time-varying factors may easily generate biases in the results. Contemporaneous health or psychological deterioration, for instance, may trigger both retirement and weight change. Hence, we need to take into account the remaining endogeneity in the model by instrumenting retirement behavior. Let us underline first that whether fixed effects estimates should over or underestimate the true effect of retirement is not entirely clear. Contemporaneous health or psychological deterioration is likely to trigger retirement, but may or may not lead to weight gain. It is a well-established fact, for instance, that the onset of acute and chronic diseases as well as of psychiatric disorders is a leading cause of involuntary weight loss among the elderly (Fischer and Johnson, 1990). Ignoring the correlation between some unobservable time-varying characteristics and retirement may thus generate a downward bias in our fixed effects estimates. The results when instrumenting retirement behavior are presented in Table 9 for men (columns (1) to (3)) and women (columns (4) to (6)). For each sex, Table 9 displays the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese. Our FEIV results suggest that our previous fixed effects estimates strongly underestimate the true impact of retirement. At this point, it should also be noted that our FEIV estimates identify a LATE among a subpopulation of compliers, i.e., the effect of retirement for those who effectively retire at the country-specific ERA. As the ERA is probably more binding for individuals with long careers, we expect compliers to be less educated. By contrast, the fixed effects model estimates the average effect of retirement for all those who retire during the study period.

Under the hypothesis that reaching the ERA is a valid instrument, our preferred IV estimates show that retirement induced by discontinuous incentives in early retirement schemes does not significantly affect men's BMI or their probability of being either overweight or obese, although both coefficients are positive. Retirement, however, causes a 12 percentage point increase in the probability of being obese (at the 5% level) within a two- to four-year period among men. It corresponds to a 60% increase in the probability of being obese within a two- to four-year period.[28][29] Note that we find a significant impact of retirement on men's BMI if we restrict our sample to men who had a BMI between 26 and 30 at baseline. The coefficient associated with retirement is equal to 1.07 (standard error: 0.64) and significant at the 10% level. This result is highly consistent with the significant impact of retirement on men's probability of being obese. In contrast, Table 9 shows that women do not experience weight changes following retirement. The coefficients associated with retirement (although positive) are never significant, whatever the outcome.

Overall, our results suggest a non-linear impact of retirement on men's BMI, and no significant impact for women. Retirement would mostly affect the right-hand side of the male BMI distribution, thus increasing the risk of obesity. To examine this further, we estimate the male BMI distributions under different treatments for the subpopulation of compliers, following Imbens and Rubin (1997b).[30] Figure 2 plots the estimated distributions of BMI standardized by age for winning and losing compliers. In our setup, winning compliers are individuals who became eligible for early retirement schemes between two subsequent waves of the survey and *did* retire but who would not have retired had they not become eligible. Losing compliers are individuals whose eligibility for early retirement schemes did not change between two subsequent waves of the survey and *did not* retire but who would have retired had they become eligible. According to Figure 2, the density function of winning compliers is shifted to the right compared to losing compliers. Winning compliers also seem to be more dispersed than losing compliers. Interestingly, the right tail of the winning compliers' density is fatter after threshold 1 – broadly corresponding to a BMI around 30 for all ages.[31] This is evidence that obese individuals are more frequent among winning compliers. This graphical evidence is consistent with the FEIV results discussed above and with the idea that retirement has a non-linear impact on men's BMI.

[28] When choosing an alternative threshold for obesity, e.g., a BMI of 31, we find that the impact of retirement on the probability of being obese is marginally significant (at the 15% level) and of a similar magnitude (coefficient: 0.08, standard error: 0.05)

[29] The coefficients associated with the effect of retirement on BMI in FEIV models are very close to those obtained for the USA using similar FEIV strategies. We find that retirement causes an increase of 0.39 and 0.31 BMI units within a two- to four-year period among men and women, respectively (although both coefficients are insignificant at conventional levels). These estimates are comparable to Chung et al. (2009) and Goldman et al. (2008). Chung et al. (2009) finds that retirement causes an increase of 0.24 BMI units within a two-year period (at the 10% significance level). Goldman et al. (2008) shows that retirees from strenuous jobs gain approximately 0.5 units of BMI during the first six years of retirement. Unfortunately, the above-mentioned studies did not explore the causal impact of retirement on the probability of being either overweight or obese or on the probability of being obese, other comparisons based on the magnitude of the coefficients cannot be made.

[30] We provide a brief explanation of this method in Appendix B.

[31] Looking at Figure 2, one may wonder why the distribution of the BMI standardized by age for losing compliers takes negative values in the [1.8; 3.2] range. This point is discussed by Imbens and Rubin (1997b). According to the authors, this negativity can be due either to sampling variation or to violation of the assumptions. In our case, as the density takes negative values when it is very close to zero and for a limited range of values, this negativity is most likely due to sampling variation.

4.3 Heterogeneous effects of retirement

The impact of retirement on weight outcomes is likely to be highly heterogeneous across job types. In particular, individuals who retire from physically demanding jobs are likely to gain weight if they do not compensate for the direct reduction in job-related exercise by increasing their leisure-time physical activity or by decreasing their food intake. To test for this, we re-estimate our FEIV models by adding an interaction term between retirement status and a measure of the previous job's physical strenuousness.[32] Table 10 shows the results when interacting retirement status with our indicator of job strenuousness for men (columns (1) to (3)) and women (columns (4) to (6)). As shown in column (3), the retirement effect on obesity seems to be mainly driven by men having retired from strenuous jobs. The coefficient associated with retirement is equal to 0.16 and insignificant at conventional levels (p-value: 0.16), but the interaction term is equal to 0.10 and significant at the 5% level. Both coefficients are jointly significant at the 5% level. Overall, retirement causes a 26 percentage point increase in the probability of being obese among men who retired from strenuous jobs within a two- to four-year period (at the 5% significance level). In contrast, retiring from a sedentary job does not seem to affect men's weight outcomes. Irrespective of occupation strenuousness, we do not find that retirement has a significant impact on men's BMI or their probability of being either overweight or obese. For women, columns (4), (5) and (6) show no weight changes following retirement, regardless of whether they retired from strenuous or sedentary jobs. The coefficients associated with the retirement indicator and the interaction term are never significant, whatever the outcome.

The impact of retirement on weight outcomes is also likely to be highly heterogeneous across weight at baseline. Additional results show that the causal impact of retirement on the probability of being obese is only significant for men who already had a BMI higher than 24 at baseline – regardless of whether we estimate the model with the interaction term *retirement*job strenuousness*.[33] The "marginal" individual – the individual likely to become obese

[32]We derive an additional instrument for this interaction term by interacting our instrument Z_{ict} with our indicator of the job's physical strenuousness. Occupation strenuousness, however, is potentially endogeneous to weight. People who are overweight may opt out of physically-demanding occupations (Lakdawalla and Philipson, 2007). While the use of fixed effects models should partly offset this risk, dynamic selction into occupations may still be a concern. It might be the case that individuals gaining weight over the study period switch to less demanding occupations prior to retirement. Given that employment perspectives and career mobility are low among the elderly, this might not happen very often.

[33]When considering men who already had a BMI higher than 24 at baseline, our sample decreases to 1026 men across the three waves. We re-estimate our FEIV models on this subsample to estimate the effect of retirement on the probability of being obese. The coefficient associated with retirement is equal to 0.13 (standard error: 0.07) and significant the 10% level. The coefficient associated with retirement is insignificant for the subsample of men who had a BMI lower than 24 at baseline (291 men across the three waves). We also re-estimate our FEIV models after including the interaction term *retirement status*job strenuousness*. When considering men who had a BMI higher than 24 and who were working at baseline, our sample decreases to 721 men across the three waves. When estimating the effect of retirement on obesity, the coefficient associated with retirement is equal to 0.17 (standard error: 0.14) and insignificant at conventional levels. The coefficient associated with the interaction term is equal to 0.11 (standard error: 0.06) and significant at the 10% level.

through retirement – is thus a man who is already at risk of obesity, i.e., already overweight or not far from the overweight threshold before retirement.

Overall, our results show that retirement effects can be highly heterogeneous: in particular, retiring from a strenuous job while being at risk of obesity (having a BMI higher than 24 at baseline) is likely to trigger obesity for men.

4.4 Underlying mechanisms

In this section, we further investigate the heterogeneous response to retirement by gender. According to our results, retirement increases the probability of being obese among men but has no effect on women's weight outcomes. As retirement is likely to operate on weight through physical activity and food intake, we try to assess whether changes in food intake or physical activity following retirement are gender specific. Due to data limitations on food intake[34], we focus on changes in leisure-time physical activity after retirement. Leisure-time physical activity is captured in the SHARE by the following question: "How often do you engage in activities that require a moderate level of energy, such as gardening, cleaning the car, or going for walk?". Four answers are available, ranging from "more than once a week" to "hardly ever or never". We first dichotomize the responses into high (more than once a week/once a week) and low moderate leisure-time physical activity (one to three times a month/hardly ever or never). When using this dichotomization, our FEIV models show that women tend to increase their leisure-time physical activity following retirement while men do not. Retirement causes a 14 percentage point increase in the probability of performing a moderate physical activity at least once a week (at the 5% significance level) among women. The corresponding figure for men is equal to 4 percentage points and insignificant at conventional levels. This would be evidence that the heterogeneous impact of retirement across genders is partly explained by women's higher propensity to engage in leisure-time physical activities following retirement. However, the results do not seem robust to alternative dichotomizations of leisure-time physical activity. When redefining the leisure-time physical activity variable (hardly ever or never versus more than once a week/once a week/one to three times a month), we find that retirement causes a 7 (11) percentage point increase in the probability of performing a moderate physical activity at least one to three times a month among men (women); these coefficients are significant at the 15% (5%) level.

Overall, our data lead to inconclusive results with regard to gender-specific patterns of leisure-time physical activity following retirement. This is not surprising, as only very precise

Both coefficients are insignificant for the subsample of men who had a BMI lower than 24 and who were working at baseline (213 men across the three waves).

[34]The SHARE contains only two measures of food consumption: monthly household expenditures on food consumed away from home and monthly household expenditures on food consumed at home. These two measures, however, are difficult to interpret: they are likely to reflect a household joint decision concerning food consumption. Thus, they do not necessarily reflect an individual's change in food consumption – and even less an individual's change in food intake.

measures of physical activity – both at work and during leisure time – would have allowed us to investigate whether women and men compensate for the direct reduction in job-related exercise to different extents. Thus, data limitations make it difficult to explore the underlying mechanisms through which retirement affects weight.

4.5 Extending the analysis: ERA and ORA as joint instruments for retirement behavior

Until now, we have used eligibility for early retirement pensions as an instrument for retirement behavior. As mentioned before, the ERA is more binding for individuals with long careers, so that compliers are likely to be less educated. As less-educated individuals are typically in worse health than the higher educated, this may affect the extent to which our results generalize to the whole population of retirees. In this paragraph, we extend our analysis and use ERA and ORA as joint instruments for retirement behavior. By including compliers at later ages, our results will produce a better representation of the population of retirees.

We re-estimate our FEIV models using ERA and ORA as joint instruments for retirement behavior. Denmark and the Netherlands are now included in the empirical analysis. The results are presented in Table A1, where Panel A displays the first stage results and Panel B the second stage results. As can be seen in Panel A, both instruments significantly predict retirement behavior for men (see column (1)) and women (see column (2)). Second-stage results are presented in Panel B. For both men and women, the Sargan-Hansen test cannot reject that the over-identifying restrictions are valid at the 5% confidence level.[35] Overall, the estimates presented in Table A1 provide evidence that using ERA and ORA as joint instruments for retirement behavior hardly change the results. This suggests that the results obtained with the main specification can be extended to a more general population of retirees. This points to a homogeneous effect of retirement on weight (especially for men) – and possibly to a moderate age gradient of the retirement effect.

4.6 Robustness checks

Our estimation strategy is likely to yield unbiased results if properly controlling for the age trend. As one may worry that our results are driven by the inadequate estimation of the age effect, we have tested linear, quadratic (presented) and quartic age terms as robustness checks. The results are qualitatively similar.[36]

[35] The Sargan-Hansen test rejects that the over-identifying restrictions are valid at the 10% significance level when considering women's BMI. This may not necessarily point to an identification failure, but rather towards heterogeneous treatment effects (Angrist and Pischke, 2008). It is not very surprising, as women compliers for official retirement ages are likely to have very specific characteristics, hence differential responses to retirement.

[36] When introducing age as a linear term, the point estimate associated with the effect of retirement on the probability of being obese in FEIV models for men is very similar to that obtained when introducing age as a quadratic term (presented). The coefficient associated with retirement is equal to 0.11 (standard error: 0.06) and significant at the 5% level. The corresponding figure when introducing age as a quartic term is 0.12 (standard error: 0.08), which is significant at the 15% level. Finally, both specifications lead to insignificant results for men's BMI and probability of being either overweight or obese. No significant results are found for women.

As underweight status is associated with a higher risk of morbidity and mortality for the elderly (Corrada et al. (2006)), one may fear that underweight individuals exhibit a different response to retirement. It might be the case that underweight individuals lose weight because of retirement, thus leading to an overall insignificant impact of retirement on BMI. We verify that our results are robust to the exclusion of underweight individuals by re-estimating our FEIV models for normal, overweight and obese individuals at baseline. The results are virtually unchanged.[37]

Until now, retirement has been defined by the respondents self-declared current job situation (see Section 3). According to this definition, anyone who declares herself to be retired is considered retired. This definition may be problematic. Hospido and Zamarro (2014) note that a non-negligible proportion of women in Europe describe themselves as homemakers rather than retired, even though they report having worked at age 50. Our results are robust to reintroducing homemakers into our sample and redefining retirement as reporting to be either a retiree or a homemaker. In particular, when considering the probability of being obese as an outcome, the point estimate associated with retirement in the FEIV model for men is equal to 0.13 (standard error: 0.06) and significant at the 5% level. Another concern with our retirement definition could be that individuals declare themselves retired despite working full- or part-time simply because they have left their "career" job. We use an alternative definition according to which anyone who declares herself to be "retired" and who did not do any paid work during the month preceding the interview is considered retired. The point estimates obtained for the retirement indicator using this alternative definition do not significantly vary from to those presented in Table 9. In particular, when considering the probability of being obese as an outcome, the coefficient associated with retirement in the FEIV model for men is equal to 0.11 (standard error: 0.08) and significant at the 15% level. Given that only 395 individuals retire between 2004 and 2010 according to this alternative definition, this result is likely to be due to a power problem.

An additional concern is that our model does not control for country-specific time trends (e.g., differential trends in food supplies, health policies or early life conditions). If these country-specific trends cause a non-linear relationship between weight and age at the country-specific ERAs, our model may not estimate the true effect of retirement on weight. Given that we consider country and time-varying ERAs as well as several cohorts, this seems highly unlikely. However, an imperfect way to test for this problem is to introduce age*country and age^2*country interaction terms to our FEIV models. By doing so, we test whether age has a differential impact on weight across countries. All coefficients associated with these

[37] When considering the probability of being obese as the outcome in the FEIV model for men, the coefficient associated with retirement is equal to 0.12 (standard error: 0.06) and significant at the 5% level. When considering either the BMI or the probability of being obese as the outcome, the coefficient associated with the retirement indicator in FEIV models for men is still insignificant. No significant results are found for women.

additional terms are insignificant in our FEIV models. The point estimates obtained for the retirement indicators do not significantly vary from those presented in Table 9. In particular, when considering the probability of being obese as an outcome, the coefficient associated with retirement in the FEIV model for men is equal to 0.12 (standard error: 0.07) and significant at the 10% level.

One may worry that our results are driven by the particularly strong effect of retirement on weight in a specific country. Our results, however, are virtually unchanged when countries are excluded one at a time from our sample.[38] Similarly, one may worry that the reforms undertaken in Italy and Austria – i.e., stepwise increases in ERAs between 2004 and 2011 – may lead to anticipatory behavior, which may bias our results. Our results are unchanged by excluding Italy and Austria from our sample.

Finally, we conduct a placebo test to confirm the reliability of our results. We evaluate the impact of retirement in a fictive state of the world wherein ERAs are interchanged across countries.[39] We re-estimate our FEIV regressions using this fictive instrument. As expected, the coefficient associated with this fictive instrument in the first stage is close to 0 and nonsignificant at conventional levels. The F-statistic of the excluded instrument is equal to 2.40 – below the standard requirement of 10 (Bound et al. (1995)) – thus pointing to a weak instrument problem. As expected, we do not find significant effects of retirement on weight outcomes in the second stage.

We conduct additional robustness checks. In particular, we check that our FEIV results are robust to the presence of serial correlation in the error terms. We also consider an unbalanced panel and alternative estimation strategies.

If the error terms in the FEIV model were serially correlated, the usual standard errors obtained from it would be very misleading. We re-estimate our FEIV models, allowing for clusters at the individual level. The results are virtually unchanged.[40]

As previously mentioned, attrition rates are non-negligible in the SHARE. In our setup, high attrition rates are a concern if non-response is systematically related to weight. We address panel attrition using the approach developed by Becketti et al. (1988) and Fitzgerald

[38] When excluding Belgium or Sweden from our sample, the impact of retirement on men's probability of being obese becomes marginally significant at the 15% level. Given that Belgium and Sweden are two of the most represented countries in the sample of men (22% and 15% of men, respectively), this result is likely to be due to a power issue.

[39] The design of the placebo reform is as follows. We interchange ERAs across countries, e.g., we assign to each country an ERA in force in another country of the sample. France's ERA is set to 61. The corresponding ERAs for Germany, Sweden, Switzerland, Spain, Austria, Italy and Belgium are 57, 60, 60, 63, 59, 62 and 62, respectively.

[40] In particular, when considering the probability of being obese as the outcome in the FEIV model for men and clustering at the individual level, the coefficient associated with retirement is equal to 0.11 (standard error: 0.06) and significant at the 10% level (p-value: 0.057). No significant results are found for women.

et al. (1998). This approach is based on the assumption that all determinants of attrition can be controlled for (selection on observables). In the test, the value of BMI at the initial wave of the survey is regressed on future attrition A (i.e., whether the individual later attrits). The test for attrition selection is simply based upon the significance of A in that model. The results (available upon request) indicate that A is not significant in that model, suggesting that people with higher BMI are no less likely to participate in future SHARE waves. This is evidence of the absence of attrition bias due to weight. As an additional robustness check, we re-estimate our FEIV models on an unbalanced sample[41] to support the reliability of our results. The point estimate obtained on the retirement indicator when considering the probability of being obese as the outcome does not significantly vary from that presented in Table 9: it is equal to 0.07 (standard error: 0.05) and significant at the 15% level (p-value: 0.13). No significant results are found for women.

Finally, we check that our results are robust to alternative estimation strategies. We first consider a pooled IV model. If our instrument Z_{ict} is truly exogenous, the results obtained from the pooled IV model should not be markedly different from those of the FEIV model. We estimate a two-stage least squares (2SLS) model wherein retirement status is instrumented by a dummy indicator Z_{ict}. The covariates included in the model are, as usual, age, age squared, marital status, occupation, education, and time and country dummies. Standard errors are robust and clustered at the individual level. We find no significant results for men's BMI or probability of being either overweight or obese. As for the probability of being obese, the coefficient associated with the retirement indicator is of a similar magnitude to that obtained through the FEIV model: it is equal to 0.13 (standard error: 0.15) but not significant at conventional levels. No significant results are found for women. We further investigate the impact of retirement on the probability of being obese in the pooled IV setting by implementing a bivariate probit with Z_{ict} as an identifying variable. The 2SLS and the bivariate models are consistent, but only the bivariate model is efficient in our case, as both endogenous variables (the retirement and obesity indicators) are dichotomous. The marginal effect associated with the retirement indicator after implementing the endogenous bivariate probit model is equal to 0.09 and significant at the 10% level.

As a final check, we allow for state dependence in the outcome variable. Taking into account the dynamics of weight may be particularly relevant as weight typically shows strong persistence over time. Following Coe and Lindeboom (2008), we address this issue by estimating a dynamic IV model. We select individuals who were working at baseline and regress the weight outcome observed in survey wave $s+1$ on the (lagged) value of weight in survey wave s, a dummy variable indicating whether the individual has retired between waves s and $s+1$, and a vector of individual characteristics (age, age squared, education, occupation, marital status, time and country dummies). Retirement status is instrumented by the dummy

[41] After conditioning on having no missing values for weight, height and any covariate included in the model and having at least two observations per individual across the three waves, the unbalanced sample consists of 18,199 observations in eight countries (9,693 men and 8,506 women).

indicator Z (i.e., reaching the ERA between waves s and $s+1$). The model is estimated by 2SLS, and standard errors are clustered at the individual level to account for repeated observations of the same person. In this setup, the retirement coefficient will pick up the effect of retirement given previous weight status: it will reflect the effect of retirement on the onset of a weight outcome (e.g., obesity) rather than on its prevalence. These results are presented in Table A2. Overall, our results show that state dependence is important, as can be seen by looking at the always significant coefficient of the lagged weight status. However, taking into account state dependence in weight hardly changes the results. We find that retirement produces a bump in men's weight (a 21 percentage point in the probability of being obese), although this effect is marginally significant at the 15% significance level. This difficulty in deriving statistically significant results is likely to be due to a power issue, as the dynamic IV model is estimated on a smaller sample (1876 men and 1620 women). Additional results show that the retirement effect on obesity is driven by men who retire from strenuous jobs (the retirement indicator and the interaction term are jointly significant at the 10% level – results not shown but available upon request). Consistently with our previous findings, we do not find any significant effects of retirement on women's weight outcomes. Overall, these results make us confident that the bump after retirement among men is not merely a continuation of a trend that begins before retirement.
Our results are robust to alternative estimation strategies. They confirm the positive and significant impact of retirement on men's risk of obesity. The estimated impact of retirement on the probability of being obese is always of similar magnitude, ranging between 0.09 and 0.21. In contrast, these results confirm that retirement does not lead to weight changes among women.

5 Conclusion

This paper examines the effect of retirement on several weight outcomes using the 2004, 2006 and 2010 SHARE waves. It exploits European variation in ERAs and stepwise increases in ERAs implemented in Austria and Italy during the study period to produce an exogenous shock to retirement behavior. This allows us to estimate the short-term causal impact of retirement on weight among those aged 50–69 years old. Our results show that retirement induced by early retirement rules does not lead to significant weight change among women. In contrast, retirement causes a 12 percentage point increase in the probability of being obese among men within a two- to four-year period. Our findings suggest that retirement has a non-linear impact on men's BMI, mostly affecting the right-hand side of the BMI distribution. We provide evidence that this effect is highly heterogeneous and driven by men who retire from strenuous jobs who are already at risk of obesity. In an extended analysis, we use ERA and ORA as joint instruments for retirement behavior. We show that including compliers at later ages hardly changes the results, thus suggesting that the main results discussed in the paper can be extended to a more general population of retirees.

Our findings extend and support previous results reported in the literature on retirement and weight. Using HRS data, Goldman et al. (2008) find that males retiring from strenuous jobs gain weight (by 0.5 BMI units), while those retiring from sedentary jobs experience virtually no weight change (by 0.1 BMI units over a 10-year period). Using the same data, Chung et al. (2009) show that retirement leads to weight gain among individuals who are already overweight and people with lower wealth who retire from physically demanding occupations (by 0.294 and 1.013 BMI units, respectively). Combined with our own findings, this provides strong evidence that males retiring from physically demanding jobs are particularly at risk of weight gain in the USA and Europe. However, the extent to which our results for women are consistent with previous findings remains an open question: Goldman et al. (2008) considers a sample of males and Chung et al. (2009) do not conduct separate analyses by gender.

Overall, our results show that most individuals do not gain weight through retirement. The impact of retirement is never significant among men who retire from sedentary jobs or among women. However, a small subgroup of retirees – i.e., men who retire from strenuous jobs and already at risk of obesity – suffer from gain weight following retirement. A straightforward interpretation of our findings is that men who retire from strenuous jobs face an important reduction in job-related exercise for which they do not compensate. By contrast, women in strenuous jobs – who typically face less intense physical demands at work – may have to balance a smaller reduction in job-related exercise following retirement. Due to data limitations, we were not able to investigate this question in greater detail. Recent research in epidemiology, however, strongly supports this interpretation. Focusing on individuals aged 45–79 years old in the EPIC-Norfolk cohort, Barnett et al. (2014) show that retirement is associated with a decrease in occupational physical activity (measured in Metabolic Equivalent of Task (MET) (h/wk)), with the largest decrease for men and manual social classes (men, manual: -78.2 MET; men, non-manual: -51.3 MET; women, manual: -39.9 MET; women, non-manual: -38.2 MET). Interestingly, retirement was associated with a net decrease in overall physical activity[42], with again, the largest decrease for men and manual classes (men, manual: -49.6 MET; men, non-manual: -41 MET; women, manual: -31.6 MET; women, non-manual: -26.9 MET). These results strongly support the idea that physical activity is the first-order explanation of our results.

Another interpretation of our results might reflect reporting bias in BMI: men and women tend to underestimate their weights in surveys (Niedhammer et al. (2000); Gorber et al. (2007)). Yet, if retirement is associated with a higher propensity to see a doctor, new retirees are more likely to have their weight measured by a physician following retirement and may acquire more accurate knowledge of their "true" weight. When interviewed in subsequent waves of the survey, they may adjust their self-reported weight and thus declare a higher weight. In this context, the impact of retirement on self-reported BMI would result from a decline in the reporting bias in weight rather than from a true increase in BMI. However,

[42]Physical activity is measured in different domains (recreational, household, occupational and transport). Weekly total energy expenditure is based on the frequency, intensity and duration per episode of physical activity.

this interpretation relies on several assumptions that do not necessarily hold. First, there is no evidence that retirement increases the use of medical care.[43] Second, there is no evidence that misreporting of weight results from a lack of knowledge, nor is there evidence that individuals adjust their self-reported weights when obtaining more accurate knowledge (Niedhammer et al. (2000)).[44] Finally, this alternative explanation would not account for the heterogenous effects of retirement by gender and occupation type that we find.

This paper has several limitations. The FEIV method that we use typically yields quite imprecise second-stage results. Although we are able to show that retirement has a clear worsening effect among men who retire from strenuous jobs and are already at risk of obesity, we cannot rule out the damaging impact of retirement on men's BMI or probability of being either overweight or obese. In the same line, although we find a clear worsening effect for men who retire from strenuous jobs and already at risk of obesity, we cannot rule out a smaller (or weaker) impact for men who retire from sedentary jobs or for women. Another limitation of this study is that we only consider the traditional and more frequent pattern of retirement wherein individuals transit directly from work to retirement. We do not consider more complex pathways to retirement (via unemployment, disability or inactivity). This sample selection implies that our results do not necessarily generalize to other retirement transitions. Further research is required to obtain a fuller picture of the impacts of different patterns of retirement on subsequent weight and health changes.

Health is multidimensional, and its dimensions can be diversely affected by retirement. If retiring reduces the amount of stress and physical strain, it may improve both subjective and objective measures of health (self-rated health, physical health, mental health or well-being). In fact, most papers in the literature do support the idea that retirement improves general health. However, if retirement also reduces the amount of physical activity and mentally stimulating activities an individual experiences from work, specific dimensions of health may deteriorate, such as cognitive functioning or body weight. Our findings are highly consistent with this interpretation. They suggest that most individuals do not gain weight through retirement, which is consistent with the overall positive effect of retirement on weight. However, they also suggest that a specific subgroup of retirees (men retiring from strenuous jobs who are already at risk of obesity) may lose health benefits in retirement by risking obesity. From an inequality perspective, retirement may exacerbate weight and health disparities, as it seems to affect the most vulnerable individuals.

[43]On the contrary, Fe and Hollingsworth (2012) find that retirement decreases primary care use in the UK.

[44]In a study of the validity of self-reported weight and height in the French GAZEL cohort, Niedhammer et al. (2000) provide convincing evidence that "the reporting bias in BMI results more from inexact reporting of weight and height that are accurately known than from lack of knowledge".

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Figure 1: Distribution of weight change (in kg) for individuals who retire across waves and individuals continuously employed in all waves for men and women, respectively.

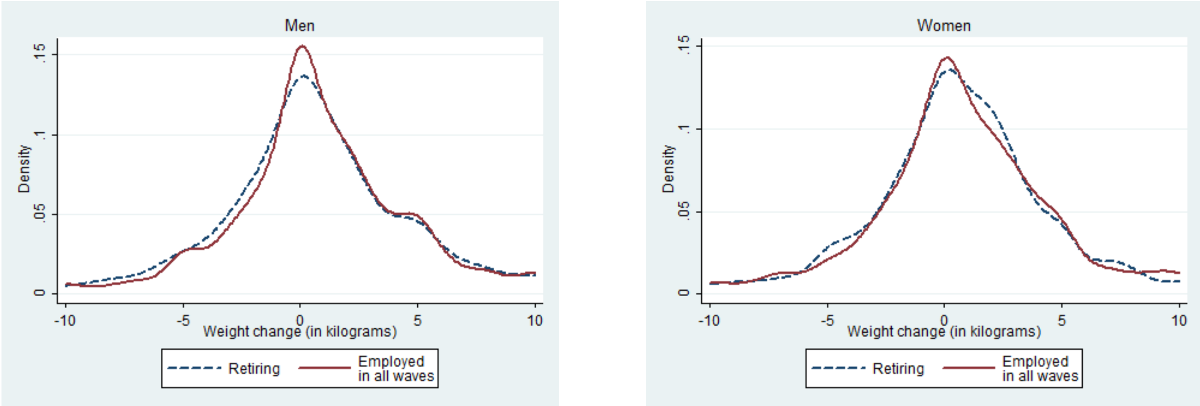

Figure 2: Counterfactual distributions of men's BMI (standardized by age) for the subpopulation of compliers.

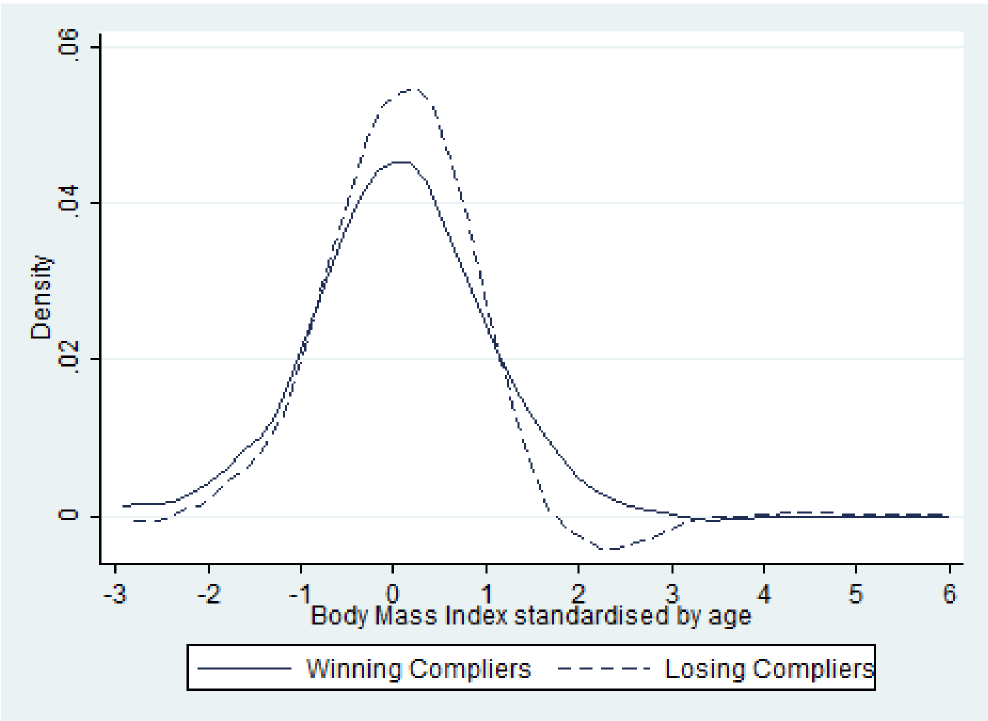

Table 1: Official (ORA), earliest (ERA) and effective retirement ages; Proportion of individuals retired below and above the ERA cutoff and proportion of individuals who retire when reaching the ERA between two subsequent waves of the survey.

Country	Official retirement age (ORA)[a]		Earliest retirement age (ERA)[a]		Effective retirement age[e,f]		% of retired below ERA[e]	% of retired above ERA[e]	% of individuals who retire when reaching ERA across waves[g]
	Men	Women	Men	Women	Men	Women			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Austria	65	60	61.5[b]	56.5[b]	61.3	59.5	31.4	94.2	28.6
Belgium	65	63[c]	60	60	61.4	61.4	21.1	81.6	34.3
France	65	65	60	60	60.5	62.1	17.0	88.0	44.5
Germany	65	65	63	60	63.6	62.6	11.5	79.9	38.6
Italy	65	60	57[d]	57[d]	61.4	61.4	18.6	81.6	27.2
Spain	65	65	61	61	65	63	8.9	68.9	21.6
Sweden	65	65	61	61	65.5	65.9	6.5	50.7	19.4
Switzerland	65	64	63	62	65	65	5.4	65.4	39.0
Netherlands	65	65	-	-	-	-	-	-	-
Denmark	65	65	-	-	-	-	-	-	-

[a] Official and earliest retirement ages are provided by Keese (2006) and OECD (2011). They concern workers retiring in 2005 under the main mandatory pension schemes and exclude special arrangements for public-sector workers and other workers, such as the long-term unemployed or disabled.

[b] In 2004, workers in Austria could retire at ages 61.5 (men) and 56.5 (women). The 2004 pension reform introduced a gradual increase in ERAs for both men and women. Specifically, the ERA was increased by two months for each quarter of birth for men (women) born in the first two quarters of 1943 (1948). Following these increases, the ERAs were increased by one month for each quarter of birth for men born in the third quarter of 1943 and later and for women born in the third quarter of 1948 and later. Moreover, the 2004 pension reform created special corridor pensions for men born in the last quarter of 1943 and later. The minimum entry age for these corridor pensions was 62, thereby making the ERA beyond age 62 non-binding in many cases (Manoli and Weber (2012)). Additional details about the reform can be found in Manoli and Weber (2012). We assign to each individual living in Austria the ERA corresponding to their quarter of birth and sex. We take age 62 as the binding age for men born in the last quarter of 1943 and later.

[c] Belgium implemented a substantial reform in women's ORAs: the full retirement age increased from age 63 in 2004 to 64 on January 1, 2006 and to age 65 on January 1, 2009 (OECD, 2011).

[d] Before 2008, workers in Italy could retire at age 57 if they had contributed to the system for 35 years. According to a reform approved as part of the 2008 budget process, the minimum age to request early retirement in Italy increased from 57 to 61 years old in 2013. The minimum age to request early retirement in Italy was 59 years old from July 1, 2009 to December 31, 2010 and 60 years old from January 1, 2011 to December 31, 2012 (OECD (2011)). We thus consider age 57 as the ERA in effect in Italy when waves 1 and 2 of the SHARE were fielded. As almost all the individuals of the 2010–2011 wave of the SHARE were surveyed in 2011, we take age 60 as the ERA in force when wave 3 of the SHARE was fielded.

[e] The figures in columns (5)–(8) are computed using the pooled sample, i.e., 7218 observations. We do not use the panel structure of the data to compute these estimates.

[f] We compute the effective retirement age as the average age of individuals who retired between 2004 and 2006 or between 2006 and 2010 in our data. As we do not have reliable information on the month and year in which the individuals retire, we cannot use the actual average age at which they retire. For this reason, the figures in columns (5)–(6) may be misleading because they systematically overestimate the effective retirement age, which is calculated in 2006 for individuals having retired between 2004 and 2006 and in 2010–11 for individuals having retired between 2006 and 2010–11.

[g] The panel structure of our data allows us to compute the proportion of individuals actually who retire between two subsequent waves of the survey when reaching the ERA in force in their country during the same period.

Table 2: Summary statistics for the pooled sample of men.

Characteristics		Whole sample Average (1)	Employed in all waves Average (2)	Retired in all waves Average (3)	Retiring between waves[a] Average (4)
Demographics					
Age		59.79	56.84	62.87	60.07
Marital status	Lives with spouse-partner	0.88	0.85	0.88	0.90
	Doesn't live with spouse-part.	0.12	0.15	0.12	0.10
Education	Post-secondary	0.33	0.41	0.24	0.32
	Upper secondary	0.33	0.29	0.33	0.37
	Lower secondary	0.17	0.18	0.16	0.17
	Primary education	0.17	0.11	0.26	0.15
Occupation	Managers and professionals	0.35	0.43	0.26	0.32
	Technicians	0.20	0.17	0.20	0.21
	White collars	0.13	0.12	0.14	0.13
	Blue collars	0.32	0.27	0.40	0.33
Employment					
Retirement status	Retired	0.43	0.00	1.00	-
	Employed or self-employed	0.57	1.00	0.00	-
Health-related measures					
Weight category	Underweight	0.01	0.00	0.01	0.01
	Normal	0.32	0.36	0.28	0.30
	Overweight	0.50	0.48	0.52	0.50
	Obese	0.18	0.15	0.20	0.19
Body Mass Index		26.93	26.49	27.33	27.08
Self-assessed health	Excellent	0.14	0.19	0.10	0.13
	Very good	0.26	0.30	0.19	0.27
	Good	0.42	0.39	0.47	0.43
	Fair	0.15	0.11	0.21	0.15
	Poor	0.03	0.01	0.03	0.02
Euro-D	Euro-D depression index (measured on a 1-12 scale)	1.44	1.50	1.44	1.38
Country	Austria	0.08	0.05	0.12	0.08
	Belgium	0.22	0.19	0.25	0.24
	France	0.15	0.13	0.20	0.15
	Germany	0.09	0.11	0.07	0.09
	Italy	0.15	0.09	0.25	0.12
	Spain	0.07	0.08	0.06	0.08
	Sweden	0.15	0.24	0.04	0.17
	Switzerland	0.07	0.11	0.02	0.07
Observations		3951	1497	1137	1317
Job strenuousness	Strenuous	43.2	42.8	-	43.6
	Sedentary	56.8	57.2	-	56.4
Observations[b]		2802	1494	-	1308

[a] An individual retiring between waves is defined as an individual who retire either between 2004 and 2006 or between 2004 and 2006. [b] Information on the physical strenuousness of work is only available for individuals who were working at baseline, i.e, 2802 men in the pooled sample.

Table 3: Summary statistics for the pooled sample of women.

Characteristics		Whole sample Average (1)	Employed in all waves Average (2)	Retired in all waves Average (3)	Retiring between waves[a] Average (4)
Demographics					
Age		59.56	56.45	63.23	60.41
Marital status	Lives with spouse-partner	0.72	0.73	0.68	0.75
	Doesn't live with spouse-part.	0.28	0.27	0.32	0.25
Education	Post-secondary	0.37	0.42	0.28	0.37
	Upper secondary	0.30	0.30	0.30	0.29
	Lower secondary	0.18	0.17	0.19	0.19
	Primary education	0.15	0.11	0.23	0.15
Occupation	Managers and professionals	0.30	0.34	0.27	0.28
	Technicians	0.20	0.19	0.15	0.23
	White collars	0.32	0.32	0.34	0.30
	Blue collars	0.18	0.15	0.23	0.19
Employment					
Retirement status	Retired	0.40	0.00	1.00	-
	Employed or self-employed	0.60	1.00	0.00	-
Health-related measures					
Weight category	Underweight	0.02	0.01	0.01	0.02
	Normal	0.50	0.53	0.38	0.53
	Overweight	0.32	0.30	0.38	0.31
	Obese	0.16	0.14	0.23	0.14
Body Mass Index		25.72	25.43	26.83	25.22
Self-assessed health	Excellent	0.14	0.18	0.06	0.15
	Very good	0.26	0.30	0.18	0.27
	Good	0.41	0.39	0.46	0.40
	Fair	0.16	0.11	0.25	0.16
	Poor	0.03	0.02	0.04	0.02
Euro-D	Euro-D depression index (measured on a 1-12 scale)	2.32	2.24	2.71	2.13
Country	Austria	0.07	0.02	0.17	0.06
	Belgium	0.20	0.19	0.25	0.17
	France	0.17	0.14	0.17	0.20
	Germany	0.13	0.14	0.10	0.13
	Italy	0.11	0.07	0.19	0.11
	Spain	0.04	0.06	0.01	0.03
	Sweden	0.21	0.27	0.09	0.24
	Switzerland	0.07	0.11	0.02	0.06
Observations		3267	1299	837	1131
Job strenuousness	Strenuous	44.1	46.1	-	41.8
	Sedentary	55.9	53.9	-	58.2
Observations[b]		2424	1296	-	1128

[a] An individual retiring between waves is defined as an individual who retire either between 2004 and 2006 or between 2004 and 2006. [b] Information on the physical strenuousness of work is only available for individuals who were working at baseline, i.e, 2424 women in the pooled sample.

Table 4: First-stage results : the impact of reaching the Earliest Retirement Age (ERA) on retirement behavior.

	Retired	
	Men	Women
	(1)	(2)
Above the ERA	.213***	.279***
	(.019)	(.022)
Age	-.057*	-.166***
	(.031)	(.032)
Age squared	.000**	.001***
	(.000)	(.000)
Time dummy for 2006	.049	.047
	(.045)	(.047)
Time dummy for 2010	.187	.166
	(.131)	(.137)
Lives with spouse/partner	.010	.030
	(.042)	(.037)
R-squared	0.31	0.36
F-Stat of excluded instruments	121.32	162.79
Observations	3951	3267

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level. (2) Standard errors in parentheses are robust. (3) Columns 1-2 are estimated by fixed-effect linear probability models.

Table 5: Pooled OLS results for men : the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Men					
	BMI		Overweight or Obese (BMI≥25)		Obese (BMI≥30)	
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.357*	.252	.049**	.052*	.024	.009
	(.215)	(.229)	(.025)	(.028)	(.021)	(.024)
Age	.110	.071	.051$^{\mu}$	.037	-.008	.001
	(.273)	(.387)	(.035)	(.048)	(.028)	(.039)
Age squared	-.001	-.001	-.001$^{\mu}$	.000	.000	.000
	(.002)	(.003)	(.000)	(.000)	(.000)	(.000)
Time dummy for 2006	.124$^{\mu}$	-.306	.019$^{\mu}$	-.045***	.029***	-.012
	(.080)	(.136)	(.012)	(.017)	(.009)	(.014)
Time dummy for 2010	.439**	-	.061***	-	.044**	-
	(.190)	-	(.023)	-	(.019)	-
Marital status						
(Ref : Does not live with a spouse/partner)						
Lives with spouse/partner	.186	.262	.072**	.091**	.004	.015
	(.330)	(.336)	(.035)	(.036)	(.028)	(.029)
Education (Ref : Primary education)						
Post secondary education	-1.435***	-1.327***	-.130***	-.108***	-.125***	-.122***
	(.361)	(.368)	(.041)	(.042)	(.036)	(.037)
Upper secondary education	-.635*	-.635*	-.063*	-.053$^{\mu}$	-.059*	-.066*
	(.352)	(.359)	(.036)	(.037)	(.034)	(.035)
Lower secondary education	-.715**	-.749**	-.030	-.017	-.061*	-.067*
	(.356)	(.367)	(.039)	(.039)	(.035)	(.036)
Occupation (Ref : Blue collars)						
Managers and professionals	.027	.105	.014	.018	-.012	-.002
	(.293)	(.296)	(.033)	(.034)	(.028)	(.029)
Technicians	.429	.474	.050$^{\mu}$	.047	.023	.034
	(.337)	(.344)	(.035)	(.035)	(.032)	(.032)
White collars	.313	.314	.034	.035	.047	.052$^{\mu}$
	(.329)	(.328)	(.036)	(.036)	(.034)	(.035)
Country fixed-effects	yes	yes	yes	yes	yes	yes
Lagged values of health	no	yes	no	yes	no	yes
R-squared	0.05	0.08	0.04	0.04	0.03	0.05
Observations	3951	2634	3951	2634	3951	2634

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are clustered at the individual level. (3) Columns 3-6 are estimated by linear probability models. (4) By construction, models including lagged values of health (columns 2, 4 and 6) are estimated on the smaller sample of men interviewed in waves 2006 and 2010 (2634 men).

Table 6: Pooled OLS results for women : the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Women					
	BMI		Overweight or Obese (BMI≥25)		Obese (BMI≥30)	
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.488*	-.047	.036	-.016	$.033^{\mu}$	-.002
	(.299)	(.305)	(.030)	(.033)	(.023)	(.024)
Age	-.194	-.591	-.027	-.070	-.009	-.015
	(.388)	(.529)	(.039)	(.053)	(.031)	(.043)
Age squared	.002	.005	.000	.001	.000	.000
	(.003)	(.004)	(.004)	(.000)	(.000)	(.000)
Time dummy for 2006	.236**	.188	$.020^{\mu}$	.033*	.018*	.014
	(.105)	(.171)	(.013)	(.021)	(.010)	(.015)
Time dummy for 2010	.032	-	-.016	-	.007	-
	(.238)	-	(.026)	-	(.019)	-
Marital status						
(Ref : Does not live with a spouse/partner)						
Lives with spouse/partner	.188	.283	.004	.019	-.007	-.004
	(.279)	(.281)	(.030)	(.030)	(.022)	(.023)
Education (Ref : Primary education)						
Post secondary education	-2.345***	-2.158***	-.170***	-.123**	-.142***	-.141***
	(.560)	(.571)	(.055)	(.056)	(.043)	(.046)
Upper secondary education	-1.578***	-1.588***	-.131***	-.108**	-.106**	-.125**
	(.535)	(.541)	(.051)	(.051)	(.042)	(.044)
Lower secondary education	-.911*	-.762	-.053	-.010	-.084**	-.089**
	(.537)	(.546)	(.049)	(.050)	(.043)	(.045)
Occupation (Ref : Blue collars)						
Managers and professionals	-.555	-.272	-.138***	-.131**	-.039	-.022
	(.508)	(.505)	(.053)	(.053)	(.040)	(.042)
Technicians	.283	.546	-.036	-.024	-.009	.011
	(.520)	(.453)	(.052)	(.052)	(.043)	(.044)
White collars	-.394	-.203	$-.066^{\mu}$	-.054	-.023	-.009
	(.453)	(.453)	(.044)	(.044)	(.037)	(.038)
Country fixed-effects	yes	yes	yes	yes	yes	yes
Lagged values of health	no	yes	no	yes	no	yes
R-squared	0.06	0.09	0.06	0.08	0.04	0.05
Observations	3267	2178	3267	2178	3267	2178

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are clustered at the individual level. (3) Columns 3-6 are estimated by linear probability models. (4) By construction, models including lagged values of health (columns 2, 4 and 6) are estimated on the smaller sample of women interviewed in waves 2006 and 2010 (2178 women).

Table 7: Fixed-effects results for men : the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Men					
	BMI		Overweight or Obese (BMI≥25)		Obese (BMI≥30)	
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	-.134	-.124	-.002	-.019	.019	.027
	(.131)	(.189)	(.023)	(.033)	(.018)	(.025)
Age	.355$^{\mu}$	.776$^{\mu}$	.095**	.139**	.008	.025
	(.235)	(.490)	(.039)	(.070)	(.026)	(.057)
Age squared	-.002	-.003	-.001***	-.001**	.000	.000
	(.002)	(.003)	(.000)	(.000)	(.000)	(.000)
Time dummy for 2006	-.220	1.560	-.018	.113	-.014	.087
	(.265)	(1.430)	(.051)	(.223)	(.033)	(.176)
Time dummy for 2010	-.533	-	-.039	-	-.083	-
	(.750)	-	(.146)	-	(.099)	-
Marital status						
(Ref : Does not live with a spouse/partner)						
Lives with spouse-partner	-.046	-.158	-.051	-.045	.003	.000
	(.245)	(.389)	(.054)	(.086)	(.034)	(.005)
Lagged values of health	no	yes	no	yes	no	yes
R-squared	0.91	0.94	0.80	0.86	0.83	0.87
Observations	3951	2634	3951	2634	3951	2634

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust in columns 1, 3 and 5 and clustered at the individual level in columns 2, 4 and 6. (3) Columns 3-6 are estimated by fixed-effect linear probability models. (4)By construction, models including lagged values of health (columns 2, 4 and 6) are estimated on the smaller sample of men interviewed in waves 2006 and 2010 (2634 men).

Table 8: Fixed-effects results for women : the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Women					
	BMI		Overweight or Obese (BMI≥25)		Obese (BMI≥30)	
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.248**	.336*	.010	.019	.023	.015
	(.123)	(.195)	(.021)	(.032)	(.017)	(.026)
Age	.237	.286	.027	.084	-.015	-.025
	(.235)	(.688)	(.040)	(.085)	(.027)	(.071)
Age squared	-.002	-.002	.000	.000	.000	.000
	(.002)	(.004)	(.000)	(.000)	(.000)	(.000)
Time dummy for 2006	.268	.384	.047	.168	.026	.138
	(.329)	(1.992)	(.054)	(.270)	(.040)	(.206)
Time dummy for 2010	.257	-	.084	-	.031	-
	(.961)	-	(.157)	-	(.117)	-
Marital status						
(Ref : Does not live with a spouse-partner)						
Lives with spouse/partner	.440$^{\mu}$	.190	.021	-.002	-.002	.035
	(.281)	(.415)	(.039)	(.060)	(.040)	(.059)
Lagged values of health	no	yes	no	yes	no	yes
R-squared	0.93	0.95	0.85	0.90	0.84	0.87
Observations	3267	2178	3267	2178	3267	2178

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust in columns 1, 3 and 5 and clustered at the individual level in columns 2, 4 and 6. (3) Columns 3-6 are estimated by fixed-effect linear probability models. (4) By construction, models including lagged values of health (columns 2, 4 and 6) are estimated on the smaller sample of women interviewed in waves 2006 and 2010 (2178 women).

Table 9: Second-stage results: the causal impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Men			Women		
	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.387	.043	.115**	.314	.037	.026
	(.445)	(.073)	(.059)	(.361)	(.058)	(.044)
Age	.401**	.995***	.016	.246	.030	-.014
	(.202)	(.032)	(.220)	(.200)	(.032)	(.024)
Age squared	-.002$^{\mu}$	-.002***	.000	.000	.000	.000
	(.001)	(.000)	(.000)	(.001)	(.000)	(.000)
Time dummy for 2006	-.259	-.021	-.021	.266	.046	.026
	(.240)	(.043)	(.029)	(.286)	(.043)	(.034)
Time dummy for 2010	-.658	-.050	-.106	.249	.080	.031
	(.685)	(.122)	(.086)	(.826)	(.125)	(.098)
Lives with spouse/partner	-.056	-.052	.001	.438	.020	-.002
	(.197)	(.041)	(.028)	(.215)	(.030)	(.030)
Observations	3951	3951	3951	3267	3267	3267

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust. (3) Columns 2-3 and 5-6 are estimated by FEIV linear probability models. (4) As the **xtivreg2** command in Stata (Schaffer (2010)) only computes the whithin R-squared, the overall R-squared is not reported here.

Table 10: Second-stage results : the impact of retirement by occupation type (strenuous/sedentary).

	Men			Women		
	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.419	.050	.162	.733	.085	.088
	(.894)	(.139)	(.116)	(.734)	(.126)	(.090)
Retirement*	.461	.011	.104**	.033	-.078	-.037
strenuous occupation	(.371)	(.058)	(.053)	(.036)	(.064)	(.045)
Age	.728	.112	.127$^{\mu}$	.552	.030	.024
	(.739)	(.102)	(.087)	(.620)	(.112)	(.076)
Age squared	-.005	-.001	-.001	-.005	.000	.000
	(.006)	(.001)	(.001)	(.005)	(.001)	(.001)
Time dummy for 2006	-.431	-.016	-.070	.360	.068	.017
	(.441)	(.065)	(.050)	(.389)	(.059)	(.042)
Time dummy for 2010	-.935	-.026	-.222$^{\mu}$	.492	.154	-.011
	(1.217)	(.181)	(.140)	(1.084)	(.168)	(.121)
Lives with spouse/partner	.009	-.037	.012	.524**	-.021	.003
	(.261)	(.054)	(.039)	(.236)	(.032)	(.036)
Observations	2802	2802	2802	2424	2424	2424

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust. (3) Columns 2-3 and 5-6 are estimated by FEIV linear probability models. (4) Information on the physical strenuousness of work before retirement is only available for individuals who were working at baseline, i.e, 2802 men and 2424 women in the pooled sample. (5) As the **xtivreg2** command in Stata (Schaffer (2010)) only computes the whithin R-squared, the overall R-squared is not reported here.

A Tables

Table A1: ERA and ORA as joint instruments for retirement behaviour: First-stage (Panel A) and second-stage results (Panel B).

Panel A : First-stage results (FEIV)

Outcome variable : Retired	Men (1)	Women (2)
Above the ERA	.196***	.270***
	(.019)	(.021)
Above the ORA	.058***	.211***
	(.028)	(.025)
R-squared	.30	.018
F-stat of excluded instruments	50.56	114.72
Observations	4980	4071

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust. (3) Models are estimated by fixed-effect linear probability models; they control for age, age squared, marital status as well as time dummies.

Panel B : Second-stage results (FEIV)

	Men			Women		
Outcome variable :	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)
	(1)	(2)	(3)	(4)	(5)	(6)
Retirement	.607	.102	.120*	.156	.041	-.024
	(.454)	(.078)	(.064)	(.289)	(.051)	(.038)
Sargan-Hansen test P-val (H0: overidentifying (restrictions are valid)	0.56	0.55	0.98	0.06	0.59	0.15
Observations	4980	4980	4980	4071	4071	4071

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are robust. (3) Models control for age, age squared, marital status as well as time dummies. (4) Columns 2-3 and 5-6 are estimated by FEIV linear probability models. (5) As the **xtivreg2** command in Stata (Schaffer (2010)) only computes the whithin R-squared, the overall R-squared is not reported here.

Table A2: Dynamic IV models: the impact of retirement on BMI, on the probability of being either overweight or obese and on the probability of being obese.

	Men			Women		
Weight outcome at *s+1*	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)	BMI	Overweight or obese (BMI≥25)	Obese (BMI≥30)
	(1)	(2)	(3)	(4)	(5)	(6)
Retired *(s;s+1)*	.500	.018	.210$^{\mu}$	.669	.109	.063
	(1.353)	(.190)	(.148)	(.967)	(.147)	(.103)
Weight outcome at *s*	.875***	.674***	.751***	.892***	.780***	.768***
	(.024)	(.022)	(.027)	(.023)	(.019)	(.029)
R-squared	0.72	0.49	0.51	0.77	0.61	0.57
Observations	1876	1876	1876	1620	1620	1620

Notes : (1) *** : significant at the 1% level, ** : significant at the 5% level, * : significant at the 10% level, $^{\mu}$: significant at the 15% level. (2) Standard errors in parentheses are clustered at the individual level. (3) Models are estimated by 2SLS on the sample of individuals who were working at baseline; they control for age, age squared, education, occupation, marital status as well as time and country dummies.

B Estimating Outcome Distributions for Compliers in Instrumental Variables (IV) Models.

Imbens and Rubin (1997b) extend the results of the IV literature (Imbens and Angrist (1994); Imbens and Rubin (1997a); Angrist et al. (1996)) by showing that under the usual assumptions, one can estimate the entire marginal distribution of the outcome under different treatments for the subpopulation of compliers. We briefly explain this method below.[45]

Let Z_i be a binary instrument. Let the pair $D_i(0)$ and $D_i(1)$ denote the values of the treatment for individual i that would be obtained given the instrument $Z_i = 0$ and $Z_i = 1$ respectively. If $D_i(0) = 0$ and $D_i(1) = 1$, unit i is called a *complier*. Let us denote $Y_i(0)$ the outcome that would be observed if the treatment were $D_i = 0$, and $Y_i(1)$ the outcome that would be observed if the treatment were $D_i = 1$.

The population is partitioned by the effect of the treatment assignment on treatment received; for never-takers (units with $D_i(0) = 0, D_i(1) = 0$), let $C_i = n$; for always-takers (units with $D_i(0) = 1, D_i(1) = 1$), let $C_i = a$; finally for compliers (units with $D_i(0) = 0, D_i(1) = 1$); let $C_i = c$. Assuming the monotonicity assumption (the "no defiers" assumption), these three types exhaustively partition the population. Let ϕ_n, ϕ_a and ϕ_c denote the populations frequencies of the three types of individuals. These proportions are known to the econometrician. Although we cannot identify the compliers from the observed data, we can identify some of the non-compliers. If $Z_{obs,i} = 0$ and $D_{obs,i} = 1$, then individual i must be an always-taker with $C_i = a$ and if $Z_{obs,i} = 1$ and $D_{obs,i} = 0$, then individual i must be a never-taker with $C_i = n$. Because of randomisation, the instrument is independent of C_i. Hence, in large samples, we

[45] This discussion heavily borrows from Imbens and Rubin (1997b).

know the distribution of $Y_i(1)$ for always-takers (denoted as $g_a(y)$) and the distribution of $Y_i(0)$ for never-takers (denoted as $g_n(y)$).

We are interested in the distributions of $Y_i(0)$ and $Y_i(1)$ among the compliers, denoted as $g_{c0}(y)$ and $g_{c1}(y)$. These distributions cannot be observed directly from the data because among those assigned to $Z_{obs,i} = 0$, both never-takers and compliers will be observed to have $D_{obs,i} = 0$. Analogously, among those assigned to $Z_{obs,i} = 1$, both always-takers and compliers will be observed to have $D_{obs,i} = 1$.
We write the directly estimable distributions of Y_i for the subsample defined by $Z_{obs,i} = z$ and $D_{obs,i} = d$ as $f_{z,d}(y)$. This implies that $g_a(y) = f_{01}(y)$ and $g_n(y) = f_{10}(y)$. Imbens and Rubin (1997b) show that the distributions for the winning and losing compliers can be expressed in terms of the directly estimable distributions in the following way:

$$\begin{aligned} \mathrm{g}_{c0}(y) &= \tfrac{\phi_n+\phi_c}{\phi_c} f_{00}(y) - \tfrac{\phi_n}{\phi_c} f_{10}(y) \\ \mathrm{g}_{c1}(y) &= \tfrac{\phi_a+\phi_c}{\phi_c} f_{11}(y) - \tfrac{\phi_a}{\phi_c} f_{01}(y) \end{aligned}$$