THE RETIREMENT CONSUMPTION PUZZLE: EVIDENCE FROM A REGRESSION DISCONTINUITY APPROACH

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Abstract

In this paper we investigate the size of the consumption drop at retirement in Italy. We use micro data on food and total non-durable household spending covering the period 1993-2004, and evaluate the change in consumption that accompanies retirement by exploiting the exogenous variability in pension eligibility to correct for the endogenous nature of the retirement decision. We take a regression discontinuity approach, and make the identifying assumption that consumption would be the same around the threshold for pension eligibility if individuals would not retire. We check in our data that a non-negligible fraction of individuals retire as soon as they become eligible, and estimate at 9.8% the part of the non-durable consumption drop that is associated with retirement induced by eligibility. We show that such fall is not driven by liquidity problems for the less well off in the population, and can be accounted for by drops in goods that are work-related expenses or leisure substitutes. However, we also show that retirement induces a significant drop in the number of grown children living with their parents, and this can account for most of the retirement consumption drop.

Keywords: Consumption, Regression Discontinuity Design, Retirement
JEL Classification: D9; E2

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1 Introduction

In most developed countries, consumption accounts for over two-thirds of GDP. In these countries a rising fraction of the population is approaching or past retirement age. The way consumers respond to retirement and the way they spend in their old age is thus a topic of great interest in the analysis of aggregate economic fluctuations and in the economic policy debate.

The standard model to analyze the consumption-saving choice by the household sector has been Modigliani's life cycle model. The model has been extended to cover uncertainty, leisure choice and a bequest motive (Deaton, 1992, Browning and Lusardi, 1996), but its key prediction can still be described as follows: consumers form intertemporal plans aimed at smoothing their standard of living (or marginal utility of consumption) over their life-cycle.

However, recent micro evidence has emphasized that there is a one-off drop in consumption at the time of retirement that may be hard to reconcile with life-time optimizing behavior. This is documented for the UK (Banks, Blundell and Tanner, 1998), for the US (Bernheim, Skinner and Weinberg, 2001), and for Italy (Miniaci, Monfardini and Weber, 2003) and is known as the retirement consumption puzzle (or retirement savings puzzle). The literature mentions as possible reasons for this drop changes in preferences due to increased non-market time, unexpectedly low pensions or liquidity problems as well as myopic or perhaps time-inconsistent behavior.

In this paper we investigate the size of the consumption drop associated to retirement in Italy by exploiting the exogenous variability in pension eligibility to identify the causal effect of retirement on consumption expenditures in a regression discontinuity framework (see Hahn, Todd and van der Klaauw, 2001). To this end, we use data from the Bank of Italy Survey on Household Income and Wealth (SHIW), covering the 1993-2004 period, that has information on food, non-durable and total household spending as well as on the current/last job and the number of years of contributions towards the public retirement pension scheme. We evaluate the change in consumption caused by retirement by exploiting the exogenous variability in pension eligibility to correct for the endogenous nature of the retirement decision. Our identification strategy rests upon a rather standard assumption made in the treatment evaluation literature (see Imbens and Lemieux, 2007) and in the context we deal with in this paper amounts to assuming that consumption would be the same around the threshold for pension eligibility if no individual retired.

The remainder of this paper is organized as follows. Section 2 provides an overview of the literature and the motivation for this paper. Section 3 presents how we deal with the endogeneity problem arising from self-selection of eligible individuals into retirement. Section 4 deals with data-related issues, in particular with the definition of the pension eligibility. In Section 5 we show that pension eligibility is a variable that is measured with error, and we discuss the implications of this on our estimates. Section 6 presents our results and some robustness checks. The economic implications of our findings are presented in Section 7, while Section 8 concludes.

2 Literature Review

The life-cycle model of Modigliani and Brumberg (1954) predicts that individuals save to smooth consumption over time: in its simplest version, they save during their working lives to keep their consumption level constant once they retire. Hamermesh (1984) was the first paper to argue that consumers apparently do not save enough to achieve this aim. If households enter retirement with inadequate savings, they must cut their consumption level, contrary to the life-cycle model predictions.

As we have already pointed out in the previous section, the recent literature has focused on estimating how consumption levels change around retirement. The existence of a consumption fall around retirement is documented for the UK (Banks, Blundell and Tanner, 1998, BBT in what follows), for the US (Bernheim, Skinner and Weinberg, 2001, BSW), and for Italy (Miniaci, Monfardini and Weber, 2003, MMW) and is known as the retirement consumption puzzle (or retirement savings puzzle).

BBT use British cohort data and show that the standard Euler equation, in which consumption growth is a function of intertemporal prices and changes in demographics, over-predicts the level of consumption by as much as 1.5% on an annual basis for ages between 60 and 67. The cumulated consumption shortfall over this age band, where most people retire, is around 10%. BBT argue that only a fraction of this drop can be attributed to the increased leisure time that accompanies retirement. Later work by Smith (2006) uses information on food for UK households who retired over the sample period, and stresses the importance of distinguishing between voluntary and involuntary retirement: a significant drop for food consumption is observed only for those who retire early because of poor health or job loss.

BSW use PSID data to estimate Euler equations for food consumption. The retirement status is instrumented by taking age-specific predicted probabilities conditional on demographics. The sample is split in groups: low wealth-to-income households drop their consumption most. BSW estimate a median drop of 14%, but higher drops for low wealth, low income replacement households. BSW conclude that "31% of the sample reduce their consumption by at least 35 percentage points". The evidence they provide is consistent with the notion that consumers do indeed enter retirement with inadequate savings. A number of papers have further investigated the issue on US data. For example Haider and Stephens (2007) estimate a smaller consumption drop for those who retire at the expected time. Fisher *et al.* (2005), who use CEX data, deflate expenditure by the squared root of household size and estimate a smaller drop (around 2.5%) for total expenditure than for food consumption (around 5.7%).

Recent papers by Aguiar and Hurst (2005 and 2007) and Hurd and Rohwedder (2006) stress that the drop in expenditure at retirement does not necessarily imply a drop in utility. For instance, work-related expenditure (transport to and from work, canteen meals and business clothing) is no longer needed - whether they account for a large enough part of pre-retirement consumption is an open issue. Also, home production of services (laundry, gardening, house-cleaning, cooking) may become advantageous, and the extra leisure time may allow consumers to shop more efficiently. This last channel has recently been stressed by Aguiar and Hurst (2005 and 2007) in their careful analysis of food consumption around retirement, whilst the increase in home production of services by recent retirees has been documented by Hurd and Rohwedder (2006) who exploit time-use data. The literature has investigated as further reasons for this drop unexpectedly low pensions or liquidity problems as well as time-inconsistent behavior (Angeletos *et al.*, 2001) .

The only available evidence for Italy is provided in MMW and is based on diary-level data covering the 1985-96 period. MMW estimate a fall in non-durable consumption at retirement of 5.4% and argue that this is a lower bound if there is heterogeneity in work-related expenses and individuals with higher expenses retire earlier. However, the data MMW use does not contain any retrospective information on work-histories, and MMW cannot identify what part of the consumption fall at retirement is as planned and what other part is instead due to the realization of bad health outcomes or other shocks to the consumer's environment that affect the retirement decision.

The Italian case is of particular interest because liquidity problems are less likely to play a role: Italian employees receive a large lump-sum payment upon retirement. If cash considerations matter, we would expect a surge in consumption at retirement rather than a drop. To be more precise, private sector employees contribute a non-negligible fraction of gross annual earnings (6.91%) to a severance pay fund, that earns 1.5% return plus 75% of inflation for each year, and is paid as a lump-sum when the job is terminated. Public sector employees contribute the same, but their final payment is based on their last pay and tenure (it can be worth up to forty times their gross monthly salary). Both types of employees can however borrow against their severance fund payment an amount that cannot exceed 70% of its current value, but only for exceptional health expenses or for the purchase of their own home or their children's. There is no such fund for self-employed workers.

3 Identification

This section presents the basic features of regression discontinuity analysis following the discussion in Hahn, Todd and van der Klaauw (2001), to which the interested reader is referred for further details. The relationship with the literature on programme evaluation is straightforwardly established by letting the retirement decision be the ”treatment” and the retired individuals be the ”treated”.

Following the notation of the potential outcome approach to causal inference, let (Y_1, Y_0) be the two potential outcomes one would experience by retiring and not retiring, respectively. In the context of this paper, Y_1 and Y_0 represent household expenditure (”consumption”) corresponding to the head being retired and not being retired, respectively. The causal effect of retirement on consumption is then defined as the difference between these outcomes, $\beta = Y_1 - Y_0$, which is not observable at the household level since being retired reveals Y_1 but conceals Y_0. Accordingly, though not observable, β represents the change in consumption corresponding to a change in the retirement status of the household head, which is our quantity of interest.

Let R be the binary variable denoting the retirement status, with $R = 1$ for retired heads and $R = 0$ otherwise. A discontinuity design (Thistlethwaite and Campbell, 1960) arises when R depends on an *observable* variable S and there exists a *known* point in the support of S where the probability of being treated changes discontinuously. Formally, if $\bar{s}$ is the discontinuity point, then a regression discontinuity is defined if

$$Pr\{R = 1|\bar{s}^+\} \neq Pr\{R = 1|\bar{s}^-\}. \tag{1}$$

Here and in the following $\bar{s}^+$ and $\bar{s}^-$ refer to those individuals *marginally* above and below $\bar{s}$, respectively. Throughout this paper, S is assumed to be continuous on the real line.

In the context of this paper, the expression in (1) implies that the probability of retirement varies discontinuously with an *observable* variable S. Throughout our analysis, S will denote the distance from the point in time the household head becomes eligible for retirement. To fix ideas, let the *eligibility status* be established according to the deterministic rule $\mathbb{1}(S \geq 0)$. That is, individuals are eligible for retirement if and only if they exhibit a value of the variable S not below the threshold $\bar{s} = 0$. Of course, such a variable can take on negative values (if individuals are not yet eligible for retirement) as well as positive values (if individuals, regardless of their retirement status, already are).

The potential of using eligibility rules to overcome the selection problems arising in the study of causal effects has been already pointed out by several papers in the literature (see, amongst others, Battistin and Rettore, 2008). In the context of this paper, pension eligibility does not necessarily imply that individuals are actually retired; on the other hand, individuals not eligible cannot be retired, thus inducing a discontinuity in the probability of retirement around the threshold for eligibility. As we will show in what follows, such a discontinuity can help solve the endogeneity problem arising from the analysis of the retirement status.

Following Trochim (1984), the distinction between *sharp* and *fuzzy* designs depends on the size of the discontinuity. The former design occurs when the probability of participating conditional on S steps from zero to one as S crosses the threshold $\bar{s}$. That is, the treatment status deterministically depends on whether individuals' values of S are above $\bar{s}$. For example, a *sharp* design would correspond to the hypothetical situation in which retirement is mandatory conditional on eligibility: is such an extreme case, we would have $R = \mathbb{1}(S \geq 0)$.

A *fuzzy* design occurs when the size of the discontinuity at $\bar{s}$ is smaller than one. According to this definition the retirement decision neatly fits a fuzzy design conditional on S. As a result of the eligibility rule and of self-selection, the probability of retirement for those scoring a value of S below the threshold $\bar{s}$ is zero by definition, since they are not yet eligible. The probability of retirement for those scoring above $\bar{s}$ is smaller than one because retirement is not mandatory. This implies that the probability of retirement is discontinuous at the threshold for eligibility and the size of the discontinuity is less than one.

It can be shown that the average causal effect of retirement on consumption can be recovered for those retirees around $\bar{s}$ (see Battistin and Rettore, 2008, and the Appendix for further details). This effect is obtained as the ratio between the difference of mean outcomes for individuals marginally above and below $\bar{s}$ and the proportion of retired at $\bar{s}^+$, namely the following relationship holds

$$E\{\beta|R = 1, \bar{s}^+\} = \frac{E\{Y|\bar{s}^+\} - E\{Y|\bar{s}^-\}}{E\{R|\bar{s}^+\}}. \tag{2}$$

The condition for this result to hold is the following, which will be maintained throughout the remainder of this paper.

Condition 1. The mean value of Y_0 conditional on S is a continuous function of S at $\bar{s}$.

Condition 1 requires that in the counterfactual world in which no one retire, no discontinuity would take place at the threshold for selection. Intuitively, in order to give a causal interpretation to discontinuities of Y around the threshold for eligibility $\bar{s} = 0$ it has to be the case that in the absence of retirement no discontinuity would be observed in the outcome Y around $\bar{s} = 0$. This amounts to assuming a smooth consumption profile at $\bar{s} = 0$ in a counterfactual world of no retirement.

4 Data

In our analysis, information on consumption and pension status is obtained from the Bank of Italy Survey on Household Income and Wealth (SHIW) for the period 1993-2004. In what follows we describe how the variables of interest are defined from raw data. As it will be clear from the remainder of this section, the panel component of the survey is not exploited in our empirical exercise, as identifications of the discontinuity in consumption does not require panel data. However we will make use of the longitudinal dimension of the survey to retrieve some of the key variables and/or to improve data quality.

4.1 Information on consumption

Information on expenditure is derived from recall questions on consumption which underwent several changes over this period. In the 1995 wave (collected in 1996) respondents were first asked "What was the monthly average spending of your household in 1995 on all consumer goods?" - they were instructed to exclude mortgage payments, rent, major house renovation as well as purchases of listed consumer durables (cars, furniture, appliances, jewels etc.). They were then asked "What instead is the monthly average figure for just food consumption? Consider spending on food products in supermarkets and the like and the spending on meals eaten regularly outside the home". Finally, they were asked questions on purchases and sales of consumer durables over the whole year. The same set of questions was asked in the 1998 wave.

In the 2000, 2002 and 2004 waves the questions on durable purchases and sales were asked first, followed by the non-durable consumption and food questions, but the wording and contents was otherwise identical. The 1993 wave, instead, asked for a different definition of food consumption: respondents were not instructed to include meals regularly consumed out of the home, but to consider expenditure at grocery stores and subtract home and personal cleaning products.[1]

4.2 The definition of retirement status

Retirement status is defined on the basis of two questions. In each survey wave, respondents were asked whether each household member was employed for the most part of the year. If the answer was negative, they were then asked whether the household member was first-time work seeker, unemployed, home-maker, job pensioner, non-job pensioner, student, conscript or other. Non-job pensions were defined as disability, survivor and social pensions. In this paper we consider a person as retired if

[1] Battistin, Miniaci and Weber (2003) compare consumption data across SHIW and diary-based SFB for one particular year, 1995.

Table 1: Composition of the SHIW sample

	1993	1995	1998	2000	2002	2004
Males						
Worker	66.8%	61.0%	60.1%	58.5%	54.9%	54.8%
Retired	28.4%	32.1%	31.8%	35.3%	39.0%	40.0 %
Females						
Worker	42.9%	43.4%	44.3%	42.9%	41.9%	43.8%
Retired	21.5%	26.3%	22.3%	24.4%	29.1%	29.7%

she/he is classified as job-pensioner. This distinction contributes to the result that the sum of the percentages of workers and pensioners is in some years much below one hundred (see Table 1).

4.3 The definition of pension eligibility

The aim of this section is to summarize how we derive the variable S that measures the time to (or from) pension eligibility . As we have shown Section 3, pension eligibility is a crucial variable in our analysis. This is measured both on the basis of *age* and on the basis of *seniority* (accrued contributions years). The SHIW sample can be used to compute eligibility: it is a large cross-sectional sample and covers several years (1987 to 2004), though it has been run every two years. SHIW also has a small panel component (rotating panel) that can be used to study actual transitions from work to retirement.

Table 1 provides a brief description of the SHIW data available for workers and retired individuals - the residual category includes other conditions such as disabled or homemaker. In this table we have considered only household heads and their spouses/partners. It should be noted that the percentage of individuals who are active or have been active in the past is far much larger for males than for females. This result is largely dominated by the labour market behaviour of older cohorts: women were characterized by lower educational attainments and lower labour market participation. An indication of this is the fact that over time there is a non negligible increase in the percentage of females who are either working or retired.

In the remainder of the paper, we focus our analysis on males who are or have been active in the labor market, as they are the ones for whom job-pension eligibility can be defined.

4.3.1 Working life and pension claims

A crucial feature of many pension systems is the design of pension benefits: in most European countries this is the defined benefit (DB) variety and it is related to some average of lifetime earnings. An extreme version of this is a final salary computation method: this was basically adopted in Italy until the 1992 pension reform as well as during the transitional phase, which Italy is still going

Table 2: Retirement eligibility rules: age and years of contributions

	Private Sector		**Public Sector**		**Self-employed**	
	age and years	only years	age and years	only years	age and years	only years
1996	54 and 35	36	53 and 35	36	57 and 35	40
1997	54 and 35	36	53 and 35	36	57 and 35	40
1998	54 and 35	36	53 and 35	36	57 and 35	40
1999	55 and 35	37	53 and 35	37	57 and 35	40
2000	55 and 35	37	54 and 35	37	57 and 35	40
2001	56 and 35	37	55 and 35	37	58 and 35	40
2002	57 and 35	37	55 and 35	37	58 and 35	40
2003	57 and 35	37	56 and 35	37	58 and 35	40
2004	57 and 35	38	57 and 35	38	58 and 35	40
2005	57 and 35	38	57 and 35	38	58 and 35	40
2006	57 and 35	39	57 and 35	39	58 and 35	40
2007	57 and 35	39	57 and 35	39	58 and 35	40
2008	57 and 35	40	57 and 35	40	58 and 35	40

through. In particular, until 1992 the pension benefit was based on the average of the last five years earnings, during the transitional phase these became the last 10 years earnings. The 1995 Pension reform changed radically this system as benefit should be computed according to a Notionally Defined Contribution (NDC) method. In the latter case, pension benefits are automatically linked to an average of lifetime earnings, adjusted by some actuarial coefficients.

In this paper we are particularly interested in eligibility conditions: until 1992-1993 they were quite simple as people could retire at age 60 (55 women) in the private sector, or any age if they had completed 35 years of contributions. The early retirement option was quite generous because it did not attract any actuarial penalty and a large fraction of workers retired before the reforms through this route. After the 1992 reform the normal retirement age has been set at 65 for men and 60 for women (to be reached gradually by the year 2001). Both age and seniority requirements for early retirement grew over time, starting essentially in 1995 (according to the sector of employment) as shown in Table 2 (rules prevailing after 1998 according to the Law 449/1997; these rules apply to white- collar employees, they differ only slightly for blue-collar employees). There was enough flexibility offered by these criteria which we can model explicitly in our data. One final problem to be mentioned is that, with some interruptions, the Government has imposed a delay-window on retirees after they became eligible, which, depending on the month of birth, could postpone retirement by 3 or even 6 months. While we cannot observe the effects of these windows directly, we do not think they would introduce too much noise in our data because we measure consumption and eligibility at yearly frequency.

4.3.2 The measurement of years of contributions

In some waves years of contribution have been explicitly recorded in SHIW for workers and for people currently retired; however this question was not asked in 1993 and missing values are occasionally found also in other waves. Because the age of the respondent does not provide enough information to measure eligibility, we have adopted a simple imputation method, by distinguishing subjects already retired at the interview time from those still at work. In fact, we make use of retrospective information on (*i*) self-reported age at entrance in the labour market and (*ii*) self-reported age of retirement (if retired). The imputation is carried out also taking into account whether the individual belongs to the panel component or not.

We can distinguish the following cases. First, if a worker or a retiree is interviewed *only once* in 1993, then we have no choice then to drop the observation. This halves the sample size of 1993. Second, if the individual is observed in 1993 and she belongs to the *panel* sub-sample of SHIW, an imputation of years of contribution is made on the basis of the 1995 recorded figure on contributions. Third, for a *retired* person who is observed only once between 1995 and 2004 and has a missing value for the contribution years an imputation is made on the basis of retrospective questions on the age at retirement and the age of entry into the labour market. If the information is missing for a *worker* instead (and she is observed only once between 1995 and 2004) the imputation is made on the basis of the difference between current age and age of entry into the labour market. Fourth, if a retiree or a worker belongs to the panel, missing values of contributions can be recovered from the previous or subsequent waves.[2]

4.3.3 Years of contributions and the retirement status

To conclude this section, in Table 3 we present the percentage of individuals self-reporting to be retired by the values of the eligibility variable S. For brevity we only report these percentages for values of S between -5 and $+5$. As already discussed above, since retirement can only be entered conditional on eligibility, it has to be the case that no retired individuals are observed for negative values of the eligibility variable. Observing retirees at negative values of S would be symptomatic of errors in the recording of the eligibility variable or, possibly, misreporting of the retirement status (or both).

Despite the sharp design implied by the eligibility rule, we find that a non-negligible proportion of individuals whose imputed value of the eligibility variable is negative report to be retired. Notably, something like 2.5 percent of males and 14 percent of females marginally ineligible for retirement self-

[2]This imputation could result in a noisy measure because the respondent can have a vague recollection (especially if far in the past) of the events. However the number of cases where an imputation had to be carried out is not large (apart from the year 1993). The distribution of years of contributions as well as that of S resulting from this strategy is reported in the Appendix.

Table 3: Percentage of retired individuals by eligibility status

eligibility	**males**	**females**
-5	0.045	0.013
-4	0.013	0.023
-3	0.015	0.024
-2	0.015	0.035
-1	0.025	0.140
0	0.333	0.568
1	0.626	0.704
2	0.613	0.800
3	0.652	0.627
4	0.740	0.781
5	0.666	0.605

report to be retired. These figures basically decrease as the time to the first eligibility year increases, though proportions are still non-negligible when $S = -5$.

Sadly enough, it is somehow immediate to conclude that the identification strategy outlined in Section 3 is no longer valid. More precisely, estimators of the causal effect of retirement on consumption based on discontinuities around $S = 0$ are in general biased for the parameter of interest. In Section 5, we will address this problem and propose a strategy which - under certain assumptions on measurement errors - identifies the causal parameter of interest.

4.4 Sample selection

In estimation we take all observations that are within a 10-year band from eligibility. However we drop those that are exactly at eligibility, $S = 0$, because the recall question on consumption could, for those who do retire, cover both pre and post-retirement periods. We investigated different selection criteria: even though the estimated key parameter is not much affected by the bandwidth choice, we find that with smaller bands (such as 5-years bands) estimates are affected by outliers in the age profile; with larger bands (15-year and more) composition effects start playing an important role, because of mortality.

In our empirical exercise, we take the head of the household to be the male. We select couples and single males, and do not use information on the wife's employment position to classify a household as retired or otherwise.

5 Accounting for measurement error

In what follows we will allow for measurement error in the variable S, but we will maintain the assumption that R in not mismeasured.[3] The motivation for allowing for measurement error in the time-to(from)-eligibility variable S builds on the evidence presented in Section 4.3, in that a non-negligible fraction of non-eligibles for pension indeed self-report to be pensioners. Under the maintained assumption that the pension status is not misreported, it follows that S is affected by measurement error. It is worth noting that the crucial assumption of correct classification of pensioners follows from our definition of a job-pensioner: the respondent has to both be retired and draw a job-related pension. Individuals who do not work and do not draw a job-pension are excluded from the analysis in the original sample selection.

In the Appendix we derive the conditions on the measurement error that allow to retrieve the causal parameter $E\{\beta|R = 1, \bar{s}^+\}$ from raw data. The main results can be summarized as follows. First, we show that the evidence provided in Section 4.3 is *not* consistent with the hypothesis of having classical measurement error in S. A more general model for measurement error is therefore needed. We do that by assuming that individuals whose observed value of the eligibility variable is τ are in fact a *mixture* of individuals whose true value, S^* say, is $S^* = \tau$ and individuals whose reported value $S = \tau$ is affected by measurement error. Formally, the observed value S_{obs} is

$$S_{obs} = S^* Z + S(1 - Z),$$

where Z is a binary variable equal to one for the exact reporters and equal to zero otherwise and S is the value contaminated by a measurement error. This is known as the *contaminated sampling model* discussed, amongst others, by Horowitz and Manski (1995).

Second, we show that even if in the presence of the measurement error the sample analogue of (2) is inconsistent for the parameter of interest, the latter is nonetheless identifiable provided that conditional on S^* the process generating measurement errors is orthogonal to the process of interest. In particular, if the latter condition is satisfied (see the Appendix for further details) it is immediate to see that the following ratio

$$\frac{E\{Y|S_{obs} = \bar{s}^+\} - E\{Y|S_{obs} = \bar{s}^-\}}{E\{R|S_{obs} = \bar{s}^+\} - E\{R|S_{obs} = \bar{s}^-\}}, \tag{3}$$

identifies the causal effect of retirement on consumption at $\bar{s}$. This result will be heavily used in the estimation section below, as it implies that consistent estimates of the causal effect of retirement on

[3]For a detailed analysis of the impact of measurement error in S for the identification of the causal effects of a binary treatment see Battistin and Chesher (2007). For the impact of misclassification (and/or misreporting) of R see, for example, Battistin and Sianesi (2006). Note, however, that in the context of this paper having measurement error in S implies that also the eligibility status is potentially misrecorded, thus inducing a more complicated structure of the error.

consumption can be recovered by a simple instrumental variable strategy where the eligibility status is used to solve for the endogeneity of the retirement status (see Imbens and Angrist, 1994, and Hahn, Todd and van der Klaauw, 2001).

6 Empirical analysis

6.1 Estimation issues

The causal effects of retirement on consumption can be estimated along the lines of what discussed in the last section. Though the effect of interest can be non-parametrically identified (see the Appendix for further details), the analysis presented in what follows builds upon a fully parametric approach, which - we think - represents a reasonable approximation to the data. As the sample analogue of (3) coincides with an instrumental variable estimator – locally at $\bar{s}$ – where the endogenous variable R is instrumented by the eligibility status $\mathbb{1}(S_{obs} \geq \bar{s})$ (see Imbens and Angrist, 1994), the regressions presented in the next section will all take the following form

$$Y_{s,t} = \beta_{0t} + \beta_1 R_{s,t} + \beta_2 S_{obs,t} + \beta_3 S^2_{obs,t} + \varepsilon_{s,t},$$

where the first-stage regression is run on $\mathbb{1}(S_{obs} \geq \bar{s})$. All variables are indexed by t and s to emphasize that we take sample averages by calendar year, as well as by years to/from eligibility. We allow for year-specific intercepts to take into account the changes in the wording of consumption questions that were discussed in Section 4.

6.2 Results

The estimation procedure that we take can be described as follows. First, for each survey year, we compute averages of household non-durable expenditure and proportions of retired male heads by values of S_{obs} (between -10 and 10). Second, we regress averages of household non-durable expenditure on proportions of retired heads and a quadratic polynomial in S_{obs}, restricting the sample to "husband and wife" and "single males" household only. In all specifications we add time dummies as regressors. Implicit in this specification is the assumption that the leisure of the spouse is separable from that of the male head. Retirement is instrumented with the eligibility status, yielding good results for the first stage.[4]

The first stage regression is defined from the regression of the retirement status on pension eligibility, a second order polynomial in S_{obs} and time dummies. Such regression presents a R^2 of 0.92, and

[4]Clustering is explicitly taken into account in the computation of the standard errors to allow for correlation among mean expenditures at the same value of S in different years.

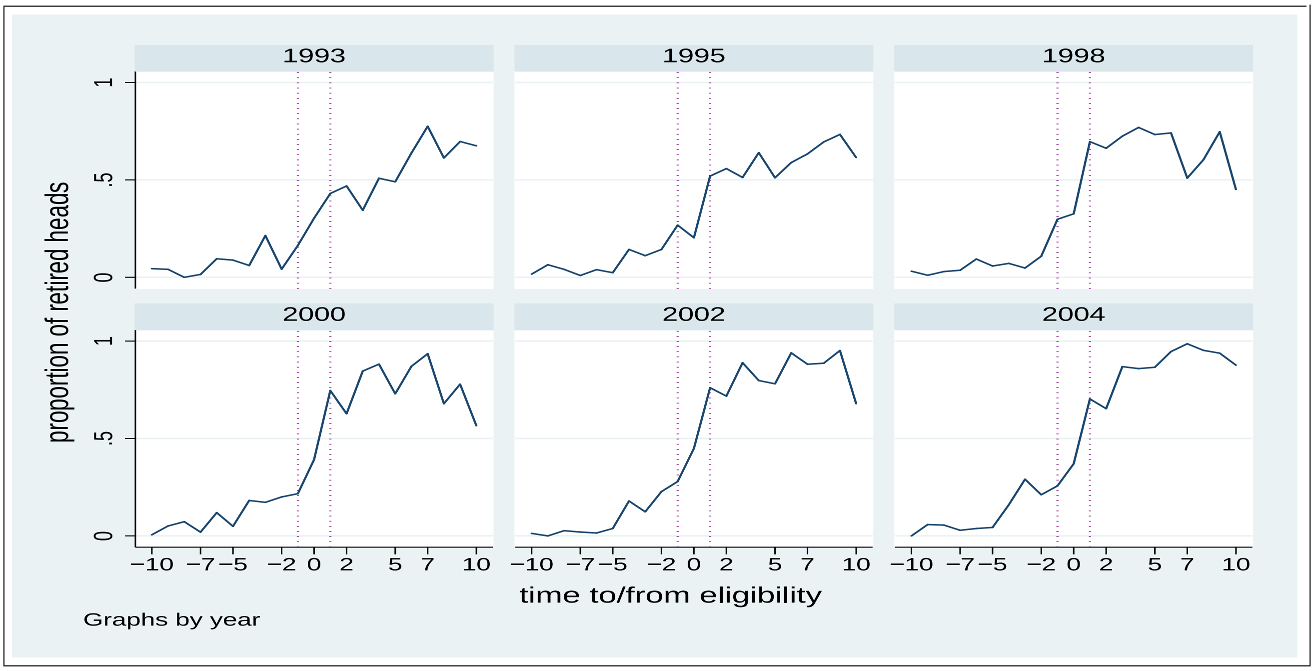

Figure 1: First stage - by year

the coefficient on eligibility is estimated at 0.435 with a standard error of 0.038. The relation between retirement status and eligibility is relatively stable over the years (see Figure 1).

Two sets of results are presented: for non-durable consumption and for food. Results for non-durable consumption are reported in the top panel of Table 4, suggesting a consumption drop of around 9.8 percent, that is significantly different from zero at the 10% level. Results for food are reported in the bottom panel, suggesting a larger consumption drop (14.1 percent), that is significantly different from zero at the 5% level.

The lack of precision in the estimation of the drop for non-durable consumption is likely to reflect the noise-to-signal ratio to be expected in this type of recall question (see Browning, Crossley and Weber, 2003, for an appraisal). Recall questions on food consumption are less heavily affected by memory problems, and therefore more informative (Battistin, Miniaci and Weber, 2003). But if we take point estimates at face value, our finding that food expenditure falls more at retirement than non-durable expenditure may appear at odds with expectations. However, for all years but one food consumption is defined to include meals regularly consumed out of the home, and this is a typical example of work-related expenses that are likely to vanish when an individual retires. Also, Aguiar and Hurst (2005) make a convincing case that food expenditure drops at retirement do not translate into food consumption drops, because of better shopping opportunities associated with more leisure time.

Table 4: Estimation results

Non-durable expenditure				
	coefficient	**std. err.**	**t**	**p-value**
Male retired	-0.0983	0.0567	-1.74	0.085
S_{obs} male	-0.0055	0.0027	-2.05	0.043
S_{obs}^2 male	-0.0003	0.0002	-1.91	0.059
Food expenditure				
	coefficient	**std. err.**	**t**	**p-value**
Male retired	-0.1409	0.05442	-2.59	0.011
S_{obs} male	-0.0028	0.0026	-1.07	0.287
S_{obs}^2 male	-0.00008	0.00014	-0.58	0.561

6.3 Robustness Checks

To gather evidence on the validity of Condition 1 (on which our identification strategy crucially relies), we implemented an over-identification test following Lee (2008). The idea behind this test is to exploit outcomes satisfying the following two conditions. First, it has to be known on a logical ground that they are not affected by the eligibility status, thus there must be no causal effect of becoming eligible on these outcomes. Second, they have to be correlated with the unobservables which are likely to affect the level of consumption.

Consider for instance the case of education. It is well known that education is a good proxy for life-time access to economic resources and should therefore affect consumption. However, educational choices are determined much before eligibility to retirement, thus the causal effect of retirement on education must be zero. It therefore follows that if household heads marginally below and above the threshold $S = 0$ featured significant differences in their education level, we would have evidence against the validity of our identification strategy. Discontinuities in outcome variables determined prior to eligibility are consistent with the hypothesis that households at $S = 0^-$ are not fully comparable to households at $S = 0^+$, so that any discontinuity detected in the outcome of interest (consumption) may be signal of other sources of selection.

In what follows, we implement the general idea in Lee (2008) by applying the same estimation procedure described in Section 6.1 to detect discontinuities on a battery of outcomes determined prior to eligibility. They include dummies for educational qualifications, age of the head, years of pension contributions of the heads and size of the main residence. Results are reported in Table 5 and in all cases we consider they are not against our identifying restriction.

Table 5: Overideintification Tests

variable	coefficient	se	**variable**	coefficient	se
college	0.0219	0.0238	**years of contributions**	-1.9376	1.3585
high – school	-0.0380	0.0469	**age**	-0.5662	1.5568
middle – school	0.0300	0.0503	**age**^{2}	-8.9669	169.65
primary – school	0.0075	0.0607	**homesize**	0.1138	5.7213

7 Economic Interpretation

We have found that non-durable consumption drops at retirement by a relatively large amount: 9.8%. The question is whether this drop should be taken as evidence against life-time optimizing behaviour.

A number of papers have emphasized that a consumption drop at retirement is compatible with the life-cycle model if leisure affects the marginal utility of consumption. Specific ways in which this non-separability may come into play at retirement are that work-related expenditures are no longer needed (Banks, Blundell and Tanner, 1998, emphasize this channel) or some goods and services may start being produced at home (as argued in Aguiar and Hurst, 2005, and Hurd and Rohwedder, 2006, who report anticipated drops around 13%).[5]

A simple way to find out whether our estimated consumption drop is consistent with the life cycle model without uncertainty is to perform some back-of-the-envelope calculations. Let us take the simplest possible utility function, a power utility defined over a Cobb-Douglas composite good made of non-durable consumption, C, and (male) leisure, l

$$U_t = \frac{\left(C_t^{\alpha} l_t^{1-\alpha}\right)^{1-\gamma}}{1-\gamma},$$

where $\gamma > 0$ is the reciprocal of the elasticity of intertemporal substitution (EIS) and α measures the within period consumption share for periods when leisure is not at a corner (hence $0 < \alpha < 1$)

$$\alpha = \frac{p_t C_t}{p_t C_t + w_t l_t},$$

where p_t is the price of the consumption good at time t and w_t the nominal wage at time t.

Life-time optimization implies that households should keep the marginal utility of consumption, U_c, constant

$$U_c = \alpha C^{\alpha-1-\alpha\gamma} l^{(1-\alpha)(1-\gamma)}.$$

This implies that consumption must react to changes in leisure according to the following relation

$$\frac{dC}{dl}\Big|_{U_c=\overline{U}} = \frac{C(1-\alpha)(1-\gamma)}{l(1-\alpha+\alpha\gamma)}.$$

[5] Blau (2008) stresses that consumption drops at retirement can also be reconciled with life-time optimization if there is uncertainty over layoffs, job offers, health and mortality, and retirement is a discrete event that is freely chosen by the household. However, in Blau's model the causal effect of retirement on consumption is zero - our IV technique that takes into account the discrete nature of retirement estimates precisely this causal effect.

We see that consumption should drop as a result of an increase in leisure if $\gamma > 1$, it should increase otherwise. A great attraction of the Cobb-Douglas lies in its analytical tractability: the formula above can be used to derive an elasticity as a function of just two preference parameters, α and γ.

We can estimate α in our data as follows. We take a sample of prime-age workers and assume that l is defined as $(T - h)$, where T is the maximum number of hours a person could possibly work (16 a day, say) and h is hours of work. In the case where $h = 8$, for instance, $l = 8$, and $wl = wh =$ earnings. The average propensity to consume out of earnings is roughly 80%, so $\alpha = 0.44$. No estimate for the elasticity of intertemporal substitution exists in Italy (to our knowledge), but studies on micro data run in other countries suggest that the elasticity should lie in the $0.5 - 0.8$ range. The real business cycle literature as well as recent studies on aggregate consumption and asset returns (Bansal and Yaron, 2004) emphasize the elasticity should be in excess of unity. We shall therefore consider a range for γ between 0.5 and 1.5.

We have estimated the average consumption drop associated to retirement to be 9.8%, with a standard error of 5.7%. In our calculations of the α parameter, retirement doubles leisure. Thus the estimated elasticity is indeed -0.098. Such a drop is consistent with utility maximization if $\gamma = 1.19$ ($EIS = 0.84$). If we consider a one-standard error range around the point estimate, we find that a 0.155 drop is consistent with $\gamma = 1.315$ ($EIS = 0.76$) and a 0.041 drop is consistent with $\gamma = 1.076$ ($EIS = 0.93$). In this context, a zero drop obtains if $\gamma = 1$ (EIS =1), whereas consumption should *increase* at retirement for smaller values of γ.

This back-of-the-envelope exercise can be used as a benchmark, but cannot resolve whether the estimated drop in consumption induced by retirement is due to increased consumption of leisure or to some other reason. Some more specific investigations may help shed further light on what driving factors lie behind our estimated retirement consumption drop.

7.1 Poverty subsample

In an influential paper, BSW relate the size of the consumption drop to wealth prior to retirement. They find that the drop is largest amongst the relatively poor, and this strongly supports their conclusion that the retirement consumption drop is due to inadequate provision for old age by "more impatient" or less informed consumers. We do not expect lack of resources to be a problem in Italy for individuals who retired during our sample period, partly because the pension system provisions were generous and well understood for these cohorts, and partly because employees receive a large lump-sum payment upon retirement (see the literature review section).

However, we have also run a formal test similar to that proposed by BSW. BSW have panel data,

Table 6: Overideintification Tests

poverty subsample variables		
variable	coefficient	se
w_fit	2170.5	17999.7
w_poor	0.0033	0.0471

so can condition on pre-retirement wealth. Since the panel component of SHIW comprises too few individuals, in what follows we will proceed by exploiting only repeated cross-sections of data. We have taken a sub-sample of individuals not yet eligible for retirement and regressed their total wealth on variables that correlate with life-time wealth, but do not change abruptly as people become eligible. The specification includes a set of dummy variables for education and for couples, plus a second order polynomial in the size of the home and a set of zero-sum year dummies. On the basis of this equation (that explains 15.38% of the variance of the dependent variable) we can assign a predicted wealth value (labelled *w_fit*) for each sample observation, whether eligible or not.

We then select the lowest third of the *w_fit* distribution - and call this is the poverty sub-sample (labeled *w_poor*). We have checked that the variables we use to split the sample can be considered to be genuinely pre-determined with respect to eligibility to retirement variables by running the same over-identifying restrictions test that we presented for education, age and home size in Section 6.3. In this case, we have run IV regressions for both the continuous variable *w_fit* and the discrete zero-one indicator *w_poor* on retirement and found insignificant coefficients, as shown in Table 6.

If the BSW mechanism is at play, we expect households in the poverty sub-sample to drop their consumption most at retirement. Our findings go in the opposite direction. Despite a good fit of the first stage regressions ($R^2 = 0.935$; coefficient on eligibility of 0.407 with a s.e. of 0.0382), the effect of retirement on non-durable consumption is effectively zero, on food consumption is -0.027, with a standard error of 0.088.

Thus for the sample of relatively poor households, retirement is strongly associated to eligibility, but does not lead to reductions in consumption. This corroborates our claim, that liquidity problems at retirement are not an issue in Italy, but calls for an explanation. A possible reason for the negligible impact of retirement on consumption of the relatively poor is that work-related expenses are much less important for this group, that is mostly made of blue collar workers. Blue collar workers normally eat meals at factory canteens for free, use heavily subsidized public transport to go to work, and do not pay for their work clothes. White collar workers, on the contrary, tend to eat in bars and restaurants, drive to work and buy expensive suites and dresses to wear at the office. Thus retirement has much more of an impact on the latter than on the former group's consumption.

Table 7: Effect of retirement on demographics

variable	coefficient	se
famsize	-0.3041	0.1259
kids0-18	0.0372	.02491
kids 18+	-0.2539	0.0926
couple	-0.0222	0.0457

7.2 Cohabitation of parents and children

Italian youths live with their parents well into their twenties and thirties. This remarkable feature of Italian society is at least partly explained by economic considerations: Becker *et al.* (2004) have focused on the inability of young Italians to find secure jobs, Manacorda and Moretti (2006) have emphasized parents' taste for cohabitation, whilst Alessie, Brugiavini and Weber (2006) have stressed the role played by inadequate financial resources to move out and enter the housing market.

This wide-spread cohabitation, together with the rules of the Italian pension system, suggests that an additional likely causal effect of the household head being eligible for retirement may be a change in household composition. In fact, as we have seen, employees receive a severance pay upon retirement, that is related to the number of years on the job. The anecdotal evidence is that as fathers (and mothers) retire they use it to (help) buy a house for their sons and daughters who then leave the parental home (Guiso and Jappelli, 2002, document the role that inter-vivos transfers play in Italy on home-ownership). In this sense the decision of the household head to retire and the decision of the son/daughter to leave the household may be co-determined and affected by the eligibility status of the household head. Since the consumption pattern of the household depends on the size of the household itself it is clear that retirement induced by eligibility affects consumption both directly and via its effect on household size.[6]

In this case, we have run IV regressions for overall household size (*famsize*) and the number of cohabiting grown children (*kids 18+*) on retirement and found highly significant coefficients, see Table 7. The estimated causal effect of the eligibility status on the household size is as large as -0.30, statistically significant. By breaking down this causal effect by type of household membership it is clear that it is driven by the negative causal effect of the household head's retirement status on the number of children older than 18, that is the kind of effect we mentioned above. No effect is instead found on the number of younger children or the proportion of couples in the sample.

The immediate consequence of these results is that the estimated retirement consumption drop

[6] Manacorda and Moretti (2006) also stress the importance of parents' pension eligibility on the home-leaving decision by the children. They claim that pension reforms of the nineties forced parents to stay at work longer and hence raised their income. In their model parents have a taste for co-residence: the additional income due to postponed retirement would have been used to bribe the children into co-residence.

Table 8: Estimation Results

	non-durable consumption		*food*	
dependent variable	coefficient	s.e.	coefficient	s.e.
ln(C)-ln(famsize)	0.0165	0.0583	-0.0271	0.0526
ln(C)-.5 x ln(famsize)	-0.0411	0.0536	-0.0840	0.0512

does not translate into a similar change in marginal utility of consumption, if utility depends on family size. The simplest way to allow for the effects of family size on utility is to express consumption in *per capita* terms, either by dividing it directly by the number of household members, or, more sensibly, by some function (the squared root for instance) of the number of household members that takes into account economies of scale. In Table 8 we provide point estimates and standard errors of the coefficient of interest when the dependent variable is re-defined to take into account family size (we do not report the coefficients on the constant, S, S^2 and year dummies). The first row of numbers corresponds to the case where consumption is defined in *per-capita* terms, the second row to the case where it is deflated by the square root of family size (along the lines of many widely adopted equivalent scales, such as the OECD scale - see also Fisher *et al.*, 2005). We see that the retirement consumption drop disappears in the former case, it is halved in the latter case.

Given that there is a strong case for taking economies of scale into consideration, we shall consider the estimates presented in the bottom row as our preferred estimates. We can thus conclude that retirement induces a 4.1% drop in non-durable consumption (insignificantly different from zero), and a 8.4% drop in total food (borderline significant).

7.3 Work-related expenses

The simple, one-good case presented in our back-of-the-envelope calculations hides the fact that the impact of retirement on total (non-durable) consumption is ambiguous, because some goods may be leisure substitutes and some other leisure complements. MMW show examples of both, notably food out and transport as substitutes, food at home and heating fuel as complements. Aguiar and Hurst (2008) show that entertainment expenses also increase after retirement, in agreement with their nature of leisure complements.

We do not have detailed expenditure information in our main data set, SHIW, as we only know non-durable consumption and a food item that is the sum of food at home and meals regularly consumed out of the home. However, we gained access to different diary-level data on consumer spending for 2002. This large data set, collected by the Italian statistical office (ISTAT), contains records of current employment, household composition, size of the main residence and a few other household characteristics, but no information on years of contributions, or past employment histories.

Table 9: Consumption drop in SFB diary data

	unadjusted		p-score adjusted		partially adjusted	
Category	**difference**	s.e.	**difference**	s.e.	**difference**	s.e.
Non-durable	-510.35	22.70	-241.02	29.74	-305.91	24.45
Food at home	-89.70	6.74	-5.96	8.90	-25.57	7.20
Meals out	-44.37	2.95	-35.68	3.46	-39.38	3.26
Alcohol	-5.24	0.96	-2.75	1.13	-3.46	1.05
Tobacco	-13.35	0.95	-7.98	1.26	-9.56	.03
Clothing	-99.09	7.26	-58.05	9.28	-68.48	7.64
Personal services	-7.96	2.73	-5.29	3.01	-5.65	2.81
Transport	-153.67	7.56	-76.04	9.91	-96.27	8.09
Heating	-19.75	2.56	-8.73	3.19	–10.31	2.72
Phones	-15.95	0.96	-8.99	1.14	-10.20	1.03
Housing services	-7.82	4.34	-2.55	4.47	-3.01	4.39
Other	-53.43	4.02	-28.99	4.93	-34.01	4.32

Thus years to and from pension eligibility are not known, and our identification strategy cannot be applied.

What we can do is to compare two groups of households, those whose head's age lies between 50 and 54 and 65 to 69, respectively. Heads of household in the former group are mostly employed (81.8% are employed, 9.6% are retired, the others are either unemployed or out of the labour force), in the latter they are mostly retired (82.7% are retired, 8.0% are employed, all the others are out of the labour force).

In Table 9 we report the difference between average spending of the older group and average spending of the younger group, and its standard error. The first column lists the various commodities considered, the second and third columns present a straight comparison, whereas the fourth and fifth columns refer to a comparison that corrects for composition effects in terms of region of residence, number of equivalent adults and size of the main residence.[7]

The first row of the table depicts that non-durable consumption falls by 510 euros a month (−31.1%) between the early fifties and late sixties. However, once composition effects are taken into account, this drop is reduced to 241 euros a month (−15.6%). This is larger than our estimate of the consumption drop at retirement, suggesting that age and composition effects play a role that is not fully accounted for in our adjustment procedure. One possible interpretation of our estimates is that almost 60% of the overall consumption drop over this period of the life cycle is due to retirement, the remaining 40% reflects changes in preferences due to poorer health and other unobservable age-related characteristics.

[7] The adjustment was made via propensity score weighting so to make the distribution of these characteristics for the younger and the older groups be equal to the distribution of households whose head is aged between 59 and 64.

If we look at the adjusted drops, we see that the largest items are meals out, clothing and transport, that account together for 169.77 euros, that is for over two thirds of the overall drop. At least two of these items, clothing and transport, are typically considered work-related expenses, whereas the drop in meals out is consistent with both a reduction in work-related expenses (meals regularly consumed out of the home) and the home production hypothesis. In fact, there is a switch from meals out to food at home in relative terms: food at home stays constant once composition effects are taken into account (and increases its budget share considerably), whereas meals out are reduced by 35.68 euros - around 41% in relative terms! - and so does its budget share.

As we know from the previous section, eligibility-induced retirement causes a significant drop in the number of grown children living within the household. So far we have neglected this drop, on the assumption that children who are induced to leave home by retirement would leave home at some later stage anyhow. However, we can check what would happen if this assumption was incorrect, and all such children stayed home, by evaluating the p-score correction for an increased number of cohabiting children in the older age group (in a sense, in doing so we partially adjust the number of equivalent adults). Given the estimates presented in Table 7, we attribute to all households aged 65-69 a fourth of an adult more, and obtain the results shown in the last two columns of Table 9. These are reasonably close to the ones presented in the previous two columns.

The conclusion that we draw from this exercise is that our estimated retirement consumption drop could well be due entirely to a reduction of work-related expenses and a substitution away from market goods to home-production of food.

8 Conclusions

In this paper we have investigated the size of the consumption drop due to retirement in Italy. We have used micro data covering the 1993-2004 period on food, and non-durable household spending, and evaluated the causal effect of retirement on consumption by exploiting the exogenous variability in pension eligibility to correct for the endogenous nature of the retirement decision. We have taken a regression discontinuity design approach, and made the identifying assumption that consumption would be the same around the threshold for pension eligibility, if the individual would not retire. We have shown that a non-negligible fraction of individuals retire as soon as they become eligible, and estimated the part of the consumption drop that is associated with retirement induced by eligibility. Given that pension eligibility is a variable that is measured with error, we have also evaluated the impact of measurement error on our estimates.

Our key result is that non-durable consumption drops by 9.8% because of (male) retirement. We show that such fall is not driven by liquidity problems for the less well off in the population, and can be accounted for by drops in goods that are work-related expenses or leisure substitutes, such as clothing, meals out and transport. However, we also show that retirement induces a significant drop in the number of grown children living with their parents, and this can account for a large fraction of the retirement consumption drop - if expenditure is deflated by a standard family size measure that takes into account economies of scale in consumption, the estimated drop is more than halved, and its point estimate is not significantly different from zero.

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Appendix

This Appendix is organized as follows. First, we derive the identification result discussed in Section 3. Second, we present the distribution of S resulting from the imputation strategy discussed in Section 4.3. Finally, we derive the conditions on the distribution of the measurement error in S that allow to recover the drop in consumption around retirement within a standard instrumental variables approach.

Identification through discontinuities in the eligibility rule

As the decision to retire is entirely up to eligible individuals, it is crucial to discuss how the resulting endogeneity problem can be accounted for in the context of discontinuity designs. We will consider the case of a sharp discontinuity first, and discuss the fuzzy case further below in this section. Let

$$Y = Y_0 + R(s)\beta$$

be the observed outcome as it results from the retirement decision: it follows from the last expression that we have either $Y \equiv Y_1$ or $Y \equiv Y_0$ depending on the retirement status of individuals ($R = 1$ and $R = 0$, respectively). The dependence of the retirement status R on S is stressed by writing $R(s)$. The difference of mean outcomes for individuals marginally above and below the threshold $\bar{s}$

$$E\{Y|\bar{s}^{+}\} - E\{Y|\bar{s}^{-}\} \tag{4}$$

can be written as

$$E\{Y_0|\bar{s}^{+}\} - E\{Y_0|\bar{s}^{-}\} + E\{R(s)\beta|\bar{s}^{+}\} - E\{R(s)\beta|\bar{s}^{-}\}, \tag{5}$$

which simplifies to

$$E\{Y_0|\bar{s}^{+}\} - E\{Y_0|\bar{s}^{-}\} + E\{\beta|\bar{s}^{+}\}$$

since $R = \mathbb{1}(S \geq \bar{s})$. Condition 1 in Section 3 is then sufficient for the mean impact of the treatment at $\bar{s}^{+}$ to be identified with a sharp discontinuity. This condition requires that in the counterfactual world of no retirement, no discontinuity would take place at the threshold for selection. Intuitively, in order to give a causal interpretation to discontinuities of Y around the threshold for eligibility $\bar{s} = 0$ it has to be the case that in the absence of retirement no discontinuity would be observed in the outcome Y around $\bar{s} = 0$. This amounts to assuming a smooth consumption profile at $\bar{s} = 0$ in a counterfactual world of no retirement. If this condition holds, we can write

$$E\{\beta|\bar{s}^{+}\} \equiv E\{Y|\bar{s}^{+}\} - E\{Y|\bar{s}^{-}\},$$

so that the difference in expected consumption expenditures above and below the threshold for eligibility identifies the average causal effect of retirement on consumption for marginally eligible subjects.

When the treatment status is not the result of a sharp design, the discontinuity in the probability to retire around the threshold is less than one thus defining a fuzzy design. It follows that the mean impact at $\bar{s}$ cannot be identified by simply comparing the mean outcome for marginal retired to the mean outcome for marginal non-retired households. In general, additional conditions are required to recover meaningful causal parameters from (4), thus loosing much of the simplicity of the design. Hahn, Todd and van der Klaauw (2001) as well as many other authors in the literature point out that assumptions can be made to recover causal effects for a particular group of individuals around the threshold $\bar{s}$. Such assumptions qualify S as an instrumental variable for R around $\bar{s}$, so that a LATE (Imbens and Angrist, 1994) parameter can be estimated for the group of compliers. Heckman, Lalonde and Smith (1999) emphasize this point by saying that much of the simplicity of the design is lost moving from a sharp design to a fuzzy design.

In the context of this paper self-selection of households into retirement fits the partially fuzzy design described by Battistin and Rettore (2008). As a result of the eligibility rule and of self-selection, the probability of retirement for those heads scoring a value of S below the threshold $\bar{s}$ is zero by definition, since they are not eligible for retirement. The probability of retirement for those scoring above $\bar{s}$ is smaller than one because retirement is not mandatory. This implies that the probability of retirement is discontinuous at the threshold for eligibility and the size of the discontinuity is less than one (i.e. according to the terminology introduced in the previous section, a fuzzy design is defined). As pointed out by Battistin and Rettore (2008), despite the fuzziness of this design the existence of a sharp eligibility rule can help recover much of the simplicity of the design.

To recover the regularity conditions required for identification consider again the difference in (4). Since participation is precluded to marginally ineligibles ($R(\bar{s}^-) = 0$), the expression in (5) can be written as

$$E\{Y_0|\bar{s}^+\} - E\{Y_0|\bar{s}^-\} + E\{R(s)\beta|\bar{s}^+\}.$$

If Condition 1 holds, by using the law of iterated expectations and by noting that $E\{R(s)\beta|R = 0, \bar{s}^+\} = 0$ the previous expression yields

$$E\{R(s)\beta|\bar{s}^+\} = E\{\beta|R = 1, \bar{s}^+\}Pr\{R = 1|\bar{s}^+\},$$

so that the mean impact on retirees in a right-neighborhood of $\bar{s}$ is identified by

$$E\{\beta|R = 1, \bar{s}^+\} = \frac{E\{Y|\bar{s}^+\} - E\{Y|\bar{s}^-\}}{E\{R|\bar{s}^+\}}. \tag{6}$$

In other words, Condition 1 is sufficient for the effect of retirement on the retirees to be identified locally at the threshold for eligibility $\bar{s}$.

It turns out that, despite the *prima facie* fuzzy nature of this set-up, the LATE (Imbens and Angrist, 1994) at the discontinuity point is identified under the same condition used to estimate the average treatment effect in the sharp design. The result rests upon the fact that the probability of retirement on the left-hand side of $\bar{s}$ is zero by design, and this simplifies the expression in (5) without further assumptions on individuals' behavior. It also follows that (6) can be estimated following an instrumental variable procedure, where eligibility is used as an instrument for the actual status R conditional on S.

The measurement of years of contributions

Figure 2 shows the distribution of years of contributions, separately for men and women, resulting from the imputation procedure described in Section 4.3. For men there is a high percentage of cases with contribution spells above 30 years - with a relevant spike at 35 years. For women the distributions is much more dispersed and in many cases there are only 5 or 10 years of contributions completed.

It follows that the variable which measures the distance from the first eligibility year (S) is computed by first establishing on the basis of recorded age and years of contributions the eligibility year and then taking the difference with the survey year. The distribution of this variable in the range (-10, 10) for males and females is reported in Figure 3. Negative values of this variable imply that eligibility for retirement has not yet been attained. Positive values measure the time from the first year of eligibility. While for men the distribution of S is rather even over the range of negative and positive values, for women there is a prevalence of positive values of S, indicating that a large percentage of women have past their first eligibility year (and they are presumably retired).

Allowing for measurement error in S

Consider the regression of R on S_{obs}, where

$$S_{obs} = S^* Z + S(1 - Z).$$

To begin with consider the case in which S_{obs} is equal to $S^* + u$, with u a classical measurement error. Since $E\{R|S_{obs}\}$ is equal to

$$\int_{\bar{s}}^{+\infty} E\{R|S^*\} f_{S^*|S_{obs}}(S^*) dS^*$$

this measurement error model implies that no discontinuity results at $\bar{s}$ which is in sharp contrast with the apparent discontinuity we do observe at the threshold for eligibility (see Table 3 and Figure 1).

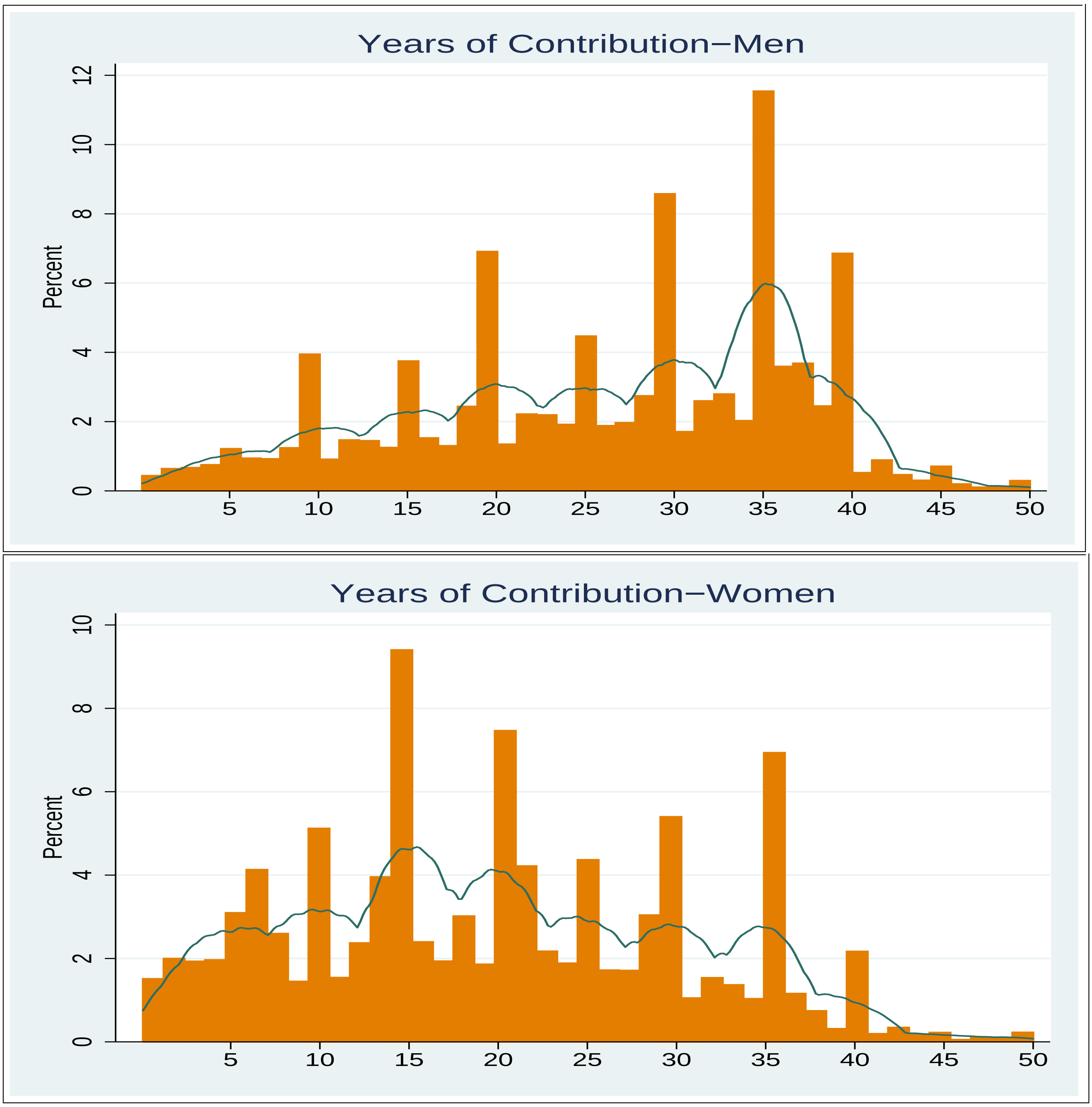

Figure 2: Contributive years - waves 1993 to 2004

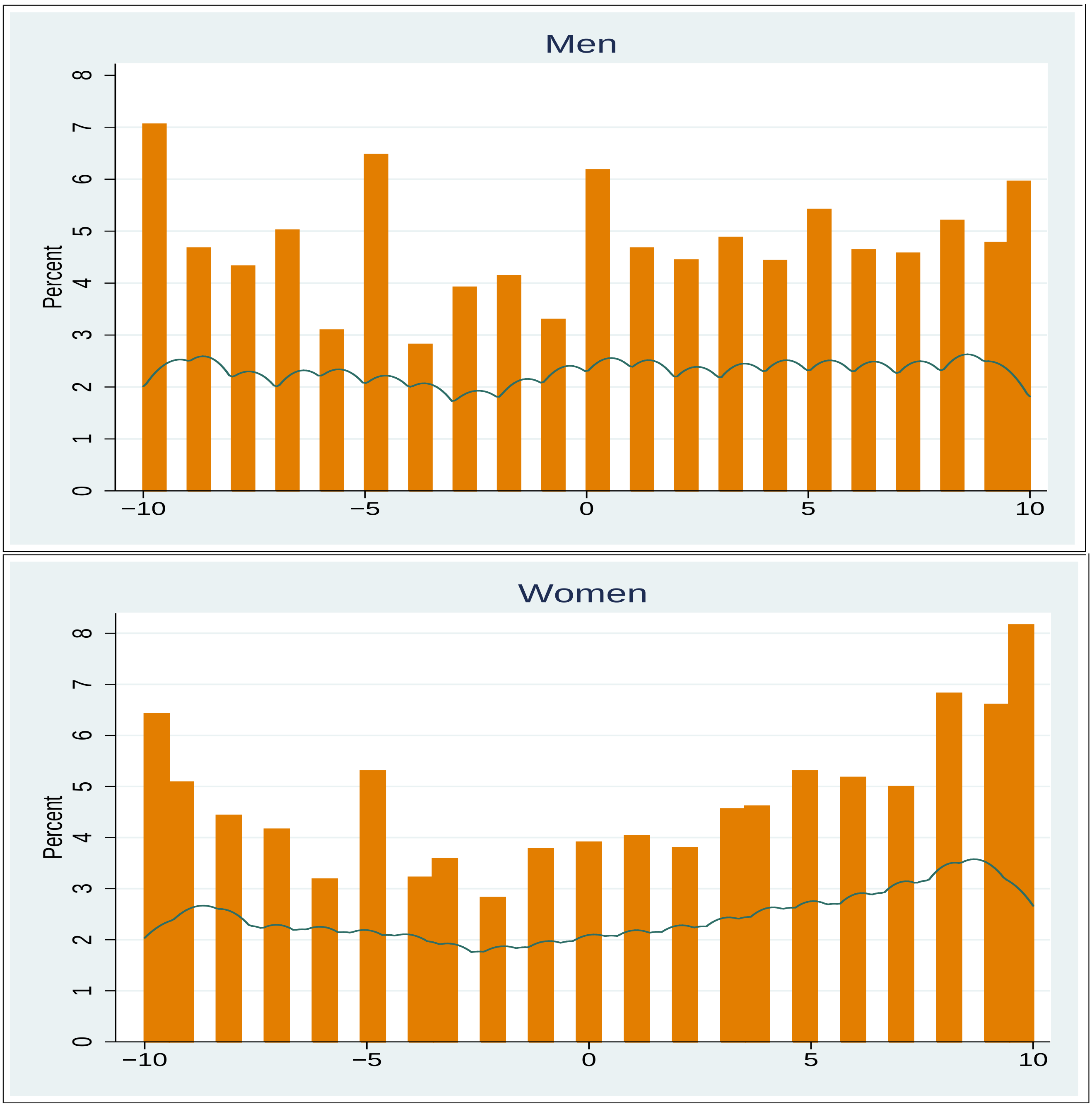

Figure 3: Distribution of S

From the definition of S_{obs} and by using the law of iterated expectations we have

$$\begin{aligned}
E\{R|S_{obs} = s_{obs}\} &= E\{R|S^* = s_{obs}, Z = 1\}E\{Z|S_{obs} = s_{obs}\} \\
&+ E\{R|S = s_{obs}, Z = 0\}(1 - E\{Z|S_{obs} = s_{obs}\}), \\
&= E\{R|S^* = s_{obs}, Z = 1\}E\{Z|S_{obs} = s_{obs}\} \\
&+ (1 - E\{Z|S_{obs} = s_{obs}\}) \\
&\int E\{R|S = s_{obs}, S^* = \tau, Z = 0\} f_{S^*|S,Z}(\tau|s_{obs}, 0)d\tau.
\end{aligned}$$

Under the assumption $(Z, S)\perp R|S^*$ which states the ignorability of measurement error for the generating process of R conditional on the true value of the eligibility variable S^* the last expression becomes

$$\begin{aligned}
E\{R|S_{obs} = s_{obs}\} &= E\{R|S^* = s_{obs}\}E\{Z|S_{obs} = s_{obs}\} \\
&+ (1 - E\{Z|S_{obs} = s_{obs}\}) \\
&\int_{\bar{s}}^{+\infty} E\{R|S^* = \tau\} f_{S^*|S,Z}(\tau|s_{obs}, 0)d\tau,
\end{aligned}$$

as $E\{R|S^* = \tau\} = 0$ when $\tau < \bar{s}$. Note that, in general, the measurement error in S does not need to be classical, though it has to be non-differential, i.e. it must contain no information on R (and on Y, see below) once the true value S^* has been controlled for (see Bound, Brown and Mathiowetz, 2001).

Under smoothness conditions of the distribution of (S, S^*) around $(\bar{s}, \bar{s})$, it follows from the last expression that

$$E\{R|S_{obs} = \bar{s}^+\} - E\{R|S_{obs} = \bar{s}^-\} = E\{R|S^* = \bar{s}^+\}E\{Z|S_{obs} = \bar{s}^-\}, \tag{7}$$

implying that the discontinuity in the retirement probability observed around the threshold for eligibility understates the true discontinuity by the factor $E\{Z|S_{obs} = \bar{s}^-\}$. It therefore follows that the estimated discontinuity is downward biased for the true discontinuity. The bias term can be estimated from the proportion of heads who self-report being retired though marginally ineligible according to S_{obs}, $P\{R = 1|S_{obs} = \bar{s}^-\}$.

By applying a similar argument to the regression function of Y on S_{obs} under the condition $(Z, S)\perp Y|S^*$ we obtain

$$E\{Y|S_{obs} = \bar{s}^+\} - E\{Y|S_{obs} = \bar{s}^-\} = (E\{Y|S^* = \bar{s}^+\} - E\{Y|S^* = \bar{s}^-\})E\{Z|S_{obs} = \bar{s}^-\}, \tag{8}$$

which implies that the discontinuity in consumption expenditures estimated around $S_{obs} = \bar{s}$ is still downward biased for the true discontinuity by the same factor $E\{Z|S_{obs} = \bar{s}^-\}$ responsible for the bias in the estimation of discontinuity in the retirement probability.

The parameter in (6) can then be written as

$$E\{\beta|R=1,S^*=\bar{s}^+\}=\frac{E\{Y|S^*=\bar{s}^+\}-E\{Y|S^*=\bar{s}^-\}}{E\{R|S^*=\bar{s}^+\}},$$

which depends on the joint distribution of (Y, R, S^*). Because of measurement error in S_{obs}, (7) and (8) imply that the estimator constructed by using the empirical analogues of the quantities in the last expression from raw data (Y, R, S) is *not* consistent for the parameter of interest.

However, under the condition $(Z, S)\perp(R, Y)|S^*$ on the measurement error it is immediate to see how the following ratio

$$\frac{E\{Y|S_{obs}=\bar{s}^+\}-E\{Y|S_{obs}=\bar{s}^-\}}{E\{R|S_{obs}=\bar{s}^+\}-E\{R|S_{obs}=\bar{s}^-\}}, \tag{9}$$

identifies the causal effect of retirement on consumption since the bias factor cancels out. As pointed out by Imbens and Angrist (1994), the latter expression can be interpreted as an instrumental variable estimand, where the eligibility status is used to correct for the endogeneity of R. The sample analogue of (9) can be obtained by taking the ratio of the discontinuity pictured in the top panel of Figure 4 to the discontinuity in the bottom panel of the same figure (see Section 6.2 in the paper for a description of how data points are derived).

Additional references

- Bound, John, Brown, Charlie and Nancy A. Mathiowetz, N. (2001), “Measurement error in survey data”, in J.J. Heckman and E. Leamer (eds.), *Handbook of Econometrics. Vol. 5*, Amsterdam: North-Holland, 3705-3843

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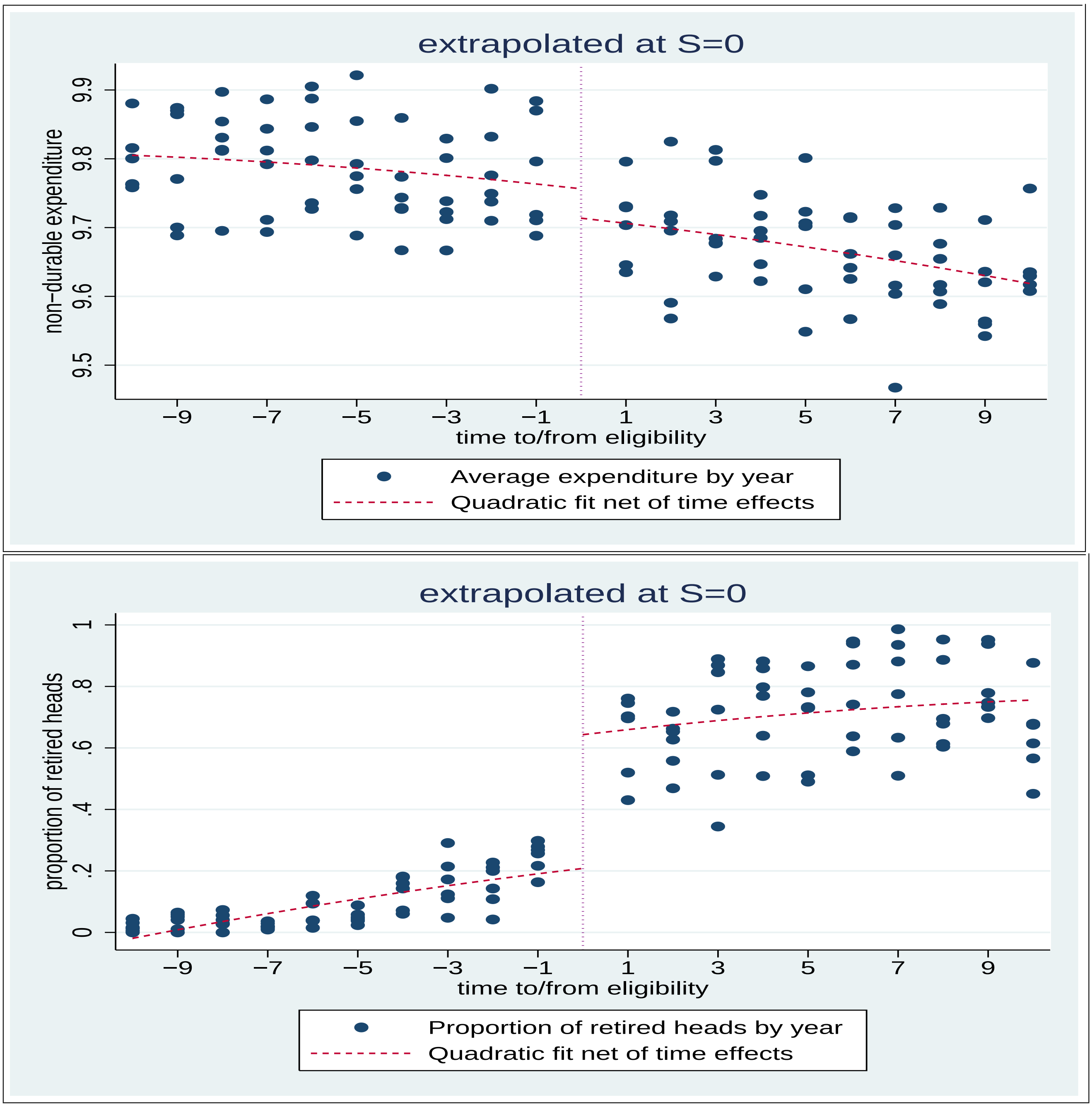

Figure 4: Estimation of the causal effect of eligibility on retirement and on non-durable consumption expenditures