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Research article

The impact of population ageing on consumption in Spain: Revisiting the retirement – consumption puzzle

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Abstract: As the proportion of the Spanish population entering retirement continues to grow, the sustainability of the pay-as-you-go social security system is facing increasing pressure. This study examined the implications of this demographic shift on consumption patterns in Spain from 2002 to 2017, focusing specifically on the retirement consumption puzzle using data from the Survey of Households Finances (SHF). Our findings, obtained from various estimation methods, suggest that the average decline in non-durable consumption at retirement is approximately 20%. This is in line with results from other developed economies and exceeds previous estimates for Spain. However, our analysis provided inconclusive results regarding spending on food and beverages. Finally, we observed no significant differences in the responses of working individuals and retirees to changes in interest rates, suggesting that stabilisation policies do not necessarily have to be adjusted due to population ageing.

Keywords: consumer spending; panel data; regression discontinuity; instrumental variables; intertemporal substitution for consumption

JEL Codes: C23, C26, D12, D15, E21, J26

1. Introduction

Since the beginning of this century, population ageing in developed economies has posed a significant challenge to the sustainability of the welfare state. This is not only due to growing pressures on pension and social security systems (the most well-known and studied aspect of this issue) but also because of broader implications for the design, management, and effectiveness of economic policy. As the share of the inactive population rises, policy responses and long-term planning are more complex.

One of the earliest areas in which the effects of population ageing were empirically observed was in the analysis of aggregate consumption. Researchers found that the spending behaviour of retirees was not consistent with the predictions of the economic theory. According to the permanent income hypothesis, rational individuals smooth consumption over their lifetimes based on total wealth and expected income, implying that retirement itself should not cause a significant change in consumption. However, many empirical studies showed a significant decline in spending at the point of retirement, challenging the implications of the model, even though they did not assess whether this decline persists thereafter.

In this context, the objective of this work is twofold. First, it aims to quantify the impact of the retirement of the household's main breadwinner on non-durable consumption expenditure. We use various methodological approaches applied to panel data from the Survey on Household Finances (*Encuesta Financiera de las Familias*, SHF) provided by the Bank of Spain, covering the period from 2002 to 2017. Specifically, we first estimate this impact using a regression discontinuity approach proposed by Battistin et al. (2009), applied to grouped individual data based on retirement eligibility. Second, we estimate the impact of retiring directly on the individual-level panel data. The latter approximation allows us to control for unobserved heterogeneity and assess the influence of other variables relevant to the retirement decision. As an alternative strategy, we compare actual household non-durable expenditure in the first year following the retirement of the household head with a counterfactual prediction generated from a consumption function estimated using data from individuals who remain active in the labour force.

Our findings reveal that Spanish households reduce their non-durable consumption by approximately 20% in the year the main earner retires, which is consistent with evidence from other developed economies and at odds with the predictions of standard life-cycle consumption theory. This finding is robust as it holds across various estimation methods. Furthermore, consistent with results from other countries, our evidence on changes in food and beverage expenditures following retirement is inconclusive. Second, we analyse the intertemporal behaviour of non-durable consumption in households with retired versus non-retired heads, finding no significant difference in the estimated elasticity of intertemporal substitution between the two groups. Based on this, we conclude that the ageing of the Spanish population is unlikely to diminish the effectiveness of monetary stabilisation policies.

Additionally, we provide informal evidence suggesting that the decline in non-durable consumption observed at retirement is temporary, or at least partially recovers over time, as indicated by the similar average expenditure levels of retirees and non-retirees reported in the survey. Furthermore, this recovery in consumption coincides with increases in both liquid and illiquid wealth

during retirement, similarly to what happens in other economies, which further challenges the explanatory power of the permanent income hypothesis in accounting for observed behaviour.

The remainder of the paper is organised as follows: Section 2 outlines the various estimation methods used to assess changes in Spanish household expenditure after retirement, as well as potential differences in intertemporal optimisation between retired and non-retired households. In Section 3, we describe the empirical implementation of our estimation approach and provide an overview of the data employed. Section 4 presents and discusses the empirical findings, comparing them with previous studies on the Spanish economy. Finally, Section 5 concludes the paper.

2. Literature review

The empirical observation that expenditures decline upon retirement, commonly referred to as the *retirement consumption (or saving) puzzle*, following the seminal contributions of Banks et al. (1998) and Bernheim et al. (2001), poses a significant challenge to the predictions of the permanent income hypothesis. Banks et al. (1998) documented a substantial reduction in British household consumption at retirement. In contrast, Bernheim et al. (2001) demonstrated that the saving and wealth trajectories of American retirees align more closely with rule-of-thumb behaviour, mental accounting, or hyperbolic discounting than with the standard life-cycle model.

Given that retirement represents one of the most predictable life-cycle events, particularly in countries with pay-as-you-go pension systems, the observed decline in consumption is difficult to reconcile with the core assumptions of the permanent income model. Moreover, surveys that investigate expected patterns of consumption and saving at retirement indicate that individuals anticipate a significant drop in spending despite the predictability of the event (see Hurd and Rohwedder, 2003; Ameriks et al., 2007; Haider and Stephens, 2007). Consequently, the puzzle has attracted substantial scholarly attention and has been examined across diverse countries and time periods. Early contributions include Hurd and Rohwedder (2003), Fisher et al. (2005), and Miniaci et al. (2010), with evidence specific to Spain provided by Labeaga and Osuna (2007), Christensen (2008), and Luengo-Prado and Sevilla (2013).

Initial explanations within the life-cycle framework attributed the decline in consumption to reductions in work-related expenses, such as fuel, transportation, and food (see Aguiar and Hurst, 2005, 2013; Hurst, 2008; Agarwal et al., 2015). Subsequent studies emphasised time allocation adjustments, arguing that retirees substitute market expenditures with home production, often at lower monetary cost, due to changes in the opportunity cost of time.[1] Blau (2008) evaluated these mechanisms alongside potential non-separabilities between consumption and leisure, concluding that a standard

[1] Schwerdt (2005), using German data, found that the increase in home production cannot be entirely attributed to a substitution effect. Overall, however, changes in food and beverage expenditures at retirement do not appear to be consistently significant for most households. As Rønnow et al. (2024) noted, some studies reported declines in food spending ranging from 3.9% to 29.3%, whereas others found no statistically significant effect, such as Hurd and Rohwedder (2013). Moreover, food expenditure data are often noisy, which complicates empirical analysis and frequently leads to inconclusive findings (see Haider and Stephens, 2007).

life-cycle model with retirement timing uncertainty cannot fully account for the magnitude of the observed decline.[2]

More recent research explores alternative explanations, including heterogeneity in longevity, medical expenditure risk, and bequest motives (see De Nardi et al., 2009, 2010, 2016, 2025). Some studies also highlight retirees' reluctance to decumulate housing wealth, for instance, through reverse mortgages, as a contributing factor (see van Ooijen et al., 2015; Suari-Andreu et al., 2019; French et al., 2023).

The explanation that has recently received the greatest attention attributes the retirement consumption puzzle to medical expenditures and the health status of retirees, following the influential contribution of Amy Finkelstein et al. (2013). This emphasis is likely motivated by increasing concerns regarding elderly healthcare costs in the United States. Finkelstein et al. (2013) developed a framework in which the marginal utility of consumption depends on health status. Using data on permanent income, health indicators, and proxies for utility such as subjective well-being, they documented that the marginal utility of consumption declines as health deteriorates.[3] Within this framework, changes in health can alter the optimal life-cycle profile of consumption and saving. However, this explanation is less applicable in countries where social security systems cover healthcare costs comprehensively, as is typical in many European welfare states, particularly the Nordic countries (see Christensen et al., 2022, for Denmark; Olafsson and Pagel, 2024, for Iceland; van Ooijen et al., 2015; Suari-Andreu, 2019, for the Netherlands). In the Spanish case as well, Labeaga and Sánchez-Robles (2026) found no evidence that the decline in consumption at retirement can be attributed either to changes in health expenditures or to reductions in work-related spending.

Most empirical research to date has concentrated on the impact of the transition from employment to retirement on individual consumption, often neglecting other important dimensions of the phenomenon. As population ageing intensifies, attention has shifted towards broader questions, including whether retirees' patterns of consumption, saving, and wealth accumulation differ systematically from those of the working population, and whether such differences persist beyond the initial drop in spending at retirement.

There are clear reasons to expect that retirees' consumption behaviour diverges from that of working individuals, given their distinct life-cycle positions and expenditure needs. Yet there is no a priori reason to assume that the initial decline in spending persists throughout retirement, especially in the context of rising life expectancy. Accordingly, partial recovery in consumption following the initial decline is plausible, as households adapt to a lower but stable income level. This perspective frames the puzzle as a "retirement illusion", consistent with behavioural economics and self-control theory, which suggests that individuals may initially overreact to income changes before gradually adjusting

[2] In contrast to Aguiar and Hurst (2005), who documented a decline in non-durable and food expenditures in the United States, Been and Goudswaard (2023) found no such decline in the Netherlands. This occurs despite evidence of substantial increases in home production and leisure following retirement, and despite similar estimates of the intertemporal elasticity of substitution of leisure in the two countries.

[3] A priori, the relationship between health and the marginal utility of consumption is ambiguous, since some consumption expenditures (such as travel) can complement good health, while others (like prepared meals or personal assistance) may serve as substitutes.

consumption patterns (see Sheffrin and Thaler, 1981; Thaler, 1986). An alternative interpretation regards the post-retirement decline in consumption as reflecting time-inconsistent preferences, explainable via hyperbolic discounting (see Laibson et al., 1998).

Recent advances in data availability, particularly administrative and survey sources capturing variables beyond expenditure and income, have enabled research to move beyond the immediate consumption effects of retirement. Evidence challenges the view of uniform wealth decumulation implied by the canonical permanent income model. Using HRS data, Love et al. (2008) found that the median annualised wealth of retired households increases. Van Ooijen et al. (2015) showed that elderly Dutch retain substantial assets even at advanced ages, while Christensen et al. (2022) reported rising wealth among a significant share of retired Danes. Olafsson and Pagel (2024) found that Icelandic households tend to increase savings or reduce indebtedness after retirement, corroborated by data from the CEX, SCF, and German household surveys. Extending this evidence across Europe, Horioka and Ventura (2024), using SHARE microdata covering 11 countries (including Spain), found that less than half of retired Europeans actively decumulate wealth. They term this phenomenon the *wealth decumulation puzzle* and attribute it primarily to the bequest motive, generous public pension systems, and reluctance to sell or leverage owner-occupied housing, highlighting the persistence of wealth among retirees.

3. Methodology

In most developed countries, retirement timing is, with few exceptions, a personal decision made by individuals within legal boundaries. This makes it particularly challenging to analyse the impact of retirement on consumption. Ideally, to assess this impact, one would compare a household's consumption in the same period under two alternative scenarios: when the household head is retired and when they are not. However, this direct comparison is not possible, as we can only observe one of these two states at any given point in time.[4]

Battistin et al. (2009) proposed the following approach to solve this problem. Let R be a binary variable indicating the retirement status of the household head, where $R = 1$ if retired, and $R = 0$ otherwise. Define *D, distance to eligibility,* as the time (in months or years) from the current moment to the point at which the household head becomes eligible for retirement under current legislation. This variable D can take both positive and negative values: $D \geq 0$ indicates that the individual is eligible to retire, while $D < 0$ implies they are not yet eligible. Let C_R and C_{NR} represent household non-durable consumption expenditures at time t, corresponding to the retirement and non-retirement statuses of the household head, respectively. The difference $\delta = C_R - C_{NR}$ captures the causal effect of retirement on household consumption. However, as noted by Battistin et al. (2009), δ is not directly observable because only one status can be observed for each individual at a given time.

According to Battistin and Rettore (2008) and Battistin et al. (2009), the effect of retirement on consumption can be recovered from the sample of individuals whose D value is around 0, using the following ratio:

[4] This challenge disappears in contexts where retirement is mandatory at a certain age, as it allows for clearer identification of its effects. See Li et al. (2016).

$$E\{\beta|R=1,D=0^{+}\}=\frac{E\{C|D=0^{+}\}-E\{C|D=0^{-}\}}{E\{C|D=0^{+}\}} \quad (1)$$

where the values 0^{+} and 0^{-} indicate, respectively, situations marginally above and below $D = 0$.[5]

Assuming that the measurement error in R is negligible, and attributing all measurement error to D, upon detecting a non-negligible fraction of retired individuals showing $D < 0$ in the data, Battistin et al. (2009) derived the conditions that allow the parameter δ to be estimated from the assumption that the distribution of the D observations is a mixture of its measured values with and without error, i.e., what is known as a *contaminated sample model*. In this regard, they showed that the basic condition for identification is that the statistical process that generates the measurement error in D is orthogonal to the process being analysed, which allows a consistent estimator of δ to be obtained by instrumenting the retirement status variable R through a variable that captures the eligibility status, thereby solving the endogeneity problem.

To this end, Battistin et al. (2009) implemented a procedure that consists of averaging non-durable expenditure and the share of retired household heads for each value of D and for each year for which data is available, which we refer to as $C_{d,t}$ and $R_{d,t}$, respectively. Thus, they regress $C_{d,t}$ on $R_{d,t}$, D, and D^2. In this regression, $R_{d,t}$ is instrumented with a prediction of the eligibility status. This prediction is based on a regression of $R_{d,t}$ on a dummy variable E that takes the value 1 if the individual is eligible for retirement and 0 otherwise, and, again, a second-order polynomial of D.

It is worth noting that, as Battistin et al. (2009) pointed out, the estimation procedure described does not, in principle, require panel data. For this reason, both Battistin et al. (2009) and, more recently, Marini (2024), applied the method directly to the averages of $C_{d,t}$ and $R_{d,t}$ for a limited range of values for D, both positive and negative.

However, both studies overlooked a key issue: the computed averages are subject to measurement error, which is likely to increase as the number of individuals within each cell decreases. This number varies considerably with the value of D, without a clear pattern, between the different periods available, making its impact on the results difficult to predict.[6] In order to avoid this measurement error issue, we apply the same procedure to the SHF panel data between 2002 and 2017.[7] Thus, following Battistin et al. (2009), we first perform an auxiliary regression as a basis for predicting the eligibility status, which is given by the following expression:

$$R_{i,t}=\beta_0+\alpha_i+\alpha_t+\beta_1E_{i,t}+\beta_2D_{i,t}+\beta_3D_{i,t}^2+\beta_4X_{i,t}+\varepsilon_{i,t} \quad (2)$$

where $R_{i,t}$ is a dummy variable equal to 1 if the individual is retired and 0 otherwise; α_i and α_t are, respectively, an individual and time effects. $D_{i,t}$ is the distance to eligibility, and $E_{i,t}$ is a dummy

[5] This result requires that function D is continuous. See Battistin and Rettore (2008) and Battistin et al. (2009).

[6] The minimum cell size in Battistin et al. (2009) was 33 individuals, and 38 in Marini (2024). The average number of individuals per cell in these studies, 90 and 109, respectively, is in line with the commonly accepted minimum cell size of around 100 per group in pseudo-panel analyses. This implies that roughly half of the cells in both studies fall below this recommended threshold.

[7] Unlike us, Li et al. (2016) estimated the same model using the entire available sample with OLS, without accounting for fixed and time effects.

variable that takes the value 1 if $D_{i,t}$ is non-negative and 0 otherwise. $X_{i,t}$ is a set of variables that may affect the retirement decision. In our empirical implementation, $X_{i,t}$ includes the following variables: the ages of the household head and their partner; a dummy variable indicating whether the individual has been predominantly self-employed during their working life; the present value of the pension fund and annual contributions to it by both the individual and their employer (when available); and two dummy variables indicating whether the household head or their partner is unemployed. The self-employment dummy is included because, in Spain, self-employed individuals typically extend their working lives longer, on average, than salaried employees. The number of years of social security contributions has been excluded from the set *Z*, as it constitutes a key component in the construction of *D*. Finally, $\varepsilon_{i,t}$ is a classical error term. It is important to note that we do not use averages, and our data consist of variables that retain individual-level variability. Furthermore, Equation (2) enables us to control for both individual fixed effects, which may be particularly relevant in the context of retirement decisions, and the effects of other demographic variables that could also play a significant role.

Based on the parameters obtained from the estimation of Equation (2), we can estimate the retirement status of each individual, $\hat{R}_{i,t}$, which, following Battistin et al. (2009), can be used to estimate individual non-durable consumption, according to the following expression:

$$C_{i,t} = \delta_0 + \gamma_i + \gamma_t + \delta_1 \hat{R}_{i,t} + \delta_2 D_{i,t} + \delta_3 D_{i,t}^2 + \delta_4 X_{i,t} + \eta_{i,t} \quad (3)$$

where γ_i is an individual effect, γ_t accounts for time effects, and $\eta_{i,t}$ is another error term with usual properties. In Equation (3), $\hat{\delta}_1$ can be interpreted as an estimate of the value of ($C_{i,R} - C_{i,NR}$) for each individual *i*, i.e., the magnitude of the change in their non-durable consumption resulting from the transition from active employment to retirement, since $\hat{R}_{i,t} = 0$ implies that the individual decides not to retire. Note that the relevant variables in determining the level of consumption, collected in $X_{i,t}$, do not necessarily have to be the same as in Equation (2), although both will be closely related.

Following Marini (2024), our analysis employs the full sample of household heads, including those with a female head. This approach is motivated by three considerations. First, in Spain, legal retirement ages are identical for men and women, rendering eligibility conditions gender-neutral. Second, the number of retired female breadwinners in the survey is too small to yield statistically reliable averages in a gender-specific analysis. Third, although the determinants of retirement decisions differ between men and women, potentially complicating joint analysis, the uniform eligibility rules allow us to assume that gender differences primarily affect retirement timing rather than the decision to retire per se.[8] Consistent with Battistin et al. (2009) and Marini (2024), the variable *D* is computed for all individuals in the dataset; however, estimation is restricted to those with *D* values

[8] The gender difference in pension access in Spain is primarily driven by greater discontinuities in women's professional careers, as well as the persistent gender pay gap, both of which result in lower average pensions for women compared to men. To address this disparity, Spain introduced a gender gap supplement in 2021, which is granted when it can be demonstrated that either partner's professional career has been adversely affected by childcare responsibilities.

between −12 and +12, excluding $D = 0$.[9] This procedure results in 144 groups (6 time periods × 24 values of D) used in the estimation.

It is important to highlight that the above methodology overlooks the potential impact of involuntary early retirement due to economic reasons. However, if the proportion of this type of retirees is significant, estimates using this approach might be biased, as these individuals would be misclassified in the survey, showing $D < 0$. Given the high incidence of involuntary early retirement in Spain and the inability to distinguish these cases in the SHF data, we propose an alternative estimation method that addresses this issue.[10]

The alternative approach to estimate this variation takes advantage of the panel structure of our data. Assuming accurate identification of individuals who transition from active employment to retirement during the sample period, our method compares two values of their non-durable consumption once retired. The first value is a prediction based on the estimated parameters of the consumption function using the full sample of active individuals, applying panel data techniques. The second is their observed consumption at the time they are first observed as retired.

According to the permanent income hypothesis under rational expectations, individuals fully anticipate both the timing and consequences of retirement. Therefore, no significant change in consumption should be observed at retirement, and the predicted and actual values should be the same. Using this alternative procedure, the estimate of δ for each individual who retires during the sample period is calculated as the difference between their observed non-durable consumption in the year of retirement $t = R$, $(C_{i,R})$, and the predicted consumption for that same year $(\hat{C}_{i,R})$. This prediction is based on the estimated parameters of the consumption function derived from the sample of all economically active individuals in the survey. Assuming rational expectations and that household income follows an autoregressive process, the consumption function used for prediction depends on household wealth, current income, and a broad set of demographic and labour-related variables.[11]

Our final exercise involves estimating a standard first-order condition for the intertemporal optimisation of non-durable consumption, using SHF data for three subsamples: non-retirees, retirees, and the total population. The goal is to assess whether intertemporal consumption behaviour differs across these groups. This analysis contrasts with our previous exercises, which focused on identifying changes in consumption levels associated with the transition into retirement. It is important to note that an individual's response to changes in the real interest rate, i.e., their intertemporal elasticity of

[9] The reason for this exclusion is to avoid situations in which individuals have covered consumption in advance immediately before and after retirement. Individuals with a D value outside this range do not appear to be seriously considering retirement. For comparison, Battistin et al. (2009) and Marini (2024) restricted their analysis to the range D = [−10+10], also excluding $D = 0$. In any case, our results remain robust to variations in the chosen range of D.

[10] Although no official statistics are available, a report by the Spanish Congress of Deputies on access to early retirement pensions, prepared by the Secretariat of State for Social Security and Pensions (2022), estimated that involuntary early retirements accounted for 39% (36%) of early retirees and 18% (16%) of all retirees in 2016 (2017). This type of retirement is particularly prevalent in Spain due to the high unemployment rates among older workers approaching the legal retirement age.

[11] A similar procedure for estimating the effect of the pandemic on Spanish non-durable consumption can be found in Cutanda and Sanchis (2025b).

substitution, may remain unchanged, regardless of whether there is a reduction in consumption at the time of retirement.

Throughout the study, we employ multiple estimation strategies. For comparative purposes, we estimate the model by Battistin et al. (2009) using a cross-sectional approach based on clusters defined by distance to retirement and time periods. In addition, we implement a panel data fixed-effects specification at the individual level, which allows us to control for unobserved heterogeneity and to evaluate the effects of variables beyond distance to retirement.

The consumption function and the first-order conditions of intertemporal optimisation are estimated using individual-level data via the generalised method of moments with instrumental variables (GMM-IV), following Hansen and Singleton (1982). Both the consumption function and the Euler equations are specified in a fixed-effects panel framework.[12] In both approaches, lagged values of the explanatory variables serve as instruments, mitigating potential endogeneity and measurement error concerns (see Griliches and Hausman, 1986).

4. Data and empirical issues

The empirical work in this study is conducted using data from the Spanish Survey of Household Finances (SHF) for the period 2002–2017,[13] which we treat either as repeated cross-sections or as a panel, depending on the analysis. The survey provides detailed information on non-durable household consumption, as well as expenditures on food and beverages consumed (both at home and away from home), which are reported directly by respondents. In addition, the SHF contains data on total and regular household income, labour supply of the members of the household, and a comprehensive number of assets and liabilities, enabling the construction of various measures of household wealth, both liquid and illiquid. The SHF maintains a consistent format for expenditure-related questions across the entire sample period. For the estimates of the first-order equation of intertemporal optimisation, we use the commercial banks' borrowing rate for terms between one and three years, sourced from the Statistical Bulletin of the Bank of Spain.

Retirement status is based on respondents' self-reported current employment situation, chosen from the following categories: employed, self-employed, unemployed, retired or early retired, permanently disabled, student, engaged in housework, or any other type of inactivity. The survey also allows for verification of retirement status through reported pension income received by any household member.

Regarding the eligibility for retirement of the household head, we follow the approach of Battistin et al. (2009) and consider both the respondent's age and the number of years of social security

[12] However, while we take first differences of the consumption function when generating predictions, we do not do so when estimating the first-order condition of intertemporal optimization. This follows the standard approach in the literature, since in this context, the endogenous variable is already the first difference of consumption.

[13] Although data from the 2020 and 2022 waves are available, we exclude them from the analysis. The 2020 wave corresponds to the year of the COVID-19 pandemic, during which consumption dropped sharply, potentially distorting the results (see Cutanda and Sanchis, 2025b). The 2022 wave, meanwhile, marks a shift in the survey frequency from triennial to biennial, reducing comparability with previous waves.

contributions; such information is directly available in the survey. These variables are evaluated in light of the eligibility criteria established by the Spanish pension system.

The Spanish pension system operates on a pay-as-you-go basis, funded through social security contributions from both employers and employees. Pension benefits are closely linked to both the amount and duration of contributions, although a non-contributory scheme exists to guarantee a minimum income for individuals without sufficient contribution histories.[14] However, increasing financial pressures have led to more reliance on state contributions to cover the growing deficit of the pension system. In this context, most measures introduced in the 2013 reform were dismantled by 2023, many even before their implementation. Notably, the so-called sustainability factor, which would have linked initial pension benefits to life expectancy, was abolished, while the principle of pension indexation to the Consumer Price Index (CPI) was reinstated. As of 2025, the legal retirement age is 66 years and 6 months, and individuals must have contributed for at least 15 years, including a minimum of two years within the 15 years preceding the pension claim.[15] Voluntary early retirement is permitted from the age of 63, provided the individual has accumulated at least 38 years and 3 months of contributions.[16] We incorporate all of these eligibility criteria to determine the retirement status of each household's main breadwinner, based on their age and reported years of contributions.

Both the legal retirement age and the required number of years of social security contributions have increased over the sample period. In the case of the retirement age, the change did not take effect until the final year of the sample. In contrast, the contribution requirement increased gradually, from around 30 years in 2002, in some cases, to 36 years and 3 months by 2017. In practice, the values obtained for the number of contribution years for each value of D were rounded down to the nearest whole number to create groups with a larger number of observations, thereby ensuring that the computed averages remained representative.[17]

Laitner et al. (2018) developed a model of post-retirement behaviour that incorporates health status uncertainty. Their model includes means-tested Medicaid assistance for nursing home care and successfully replicates empirical patterns through simulation techniques. More recently, Blundell et al. (2024) documented the prevalence of transitory income and health shocks in old age and analysed their impact on consumption. They found that negative health shocks reduce the marginal utility of consumption, while out-of-pocket medical expenses appear independent of transitory income shocks. They interpreted this evidence as suggesting that many elderly individuals are close to satiation in their consumption due to medical expenses. Furthermore, they showed that the satiation point varies with health, as out-of-pocket medical costs increase with worsening health shocks. However, it is important to note that while these findings may plausibly explain the U.S. context (given the characteristics of the U.S. healthcare system), they do not seem directly applicable to the Spanish case.

[14] Voluntary private pension plans, introduced in Spain in 1987, can supplement public pensions as a way to enhance retirement income and address anticipated future funding challenges in the public pension system.

[15] The legal retirement age will gradually increase to 67 by 2027.

[16] See Labeaga and Sánchez-Robles (2026) for a more detailed description of the Spanish pension system.

[17] The calculation of this variable is also problematic in the case of Battistin et al. (2009), who resort to imputation methods to estimate its values for years in which the survey does not provide the information directly.

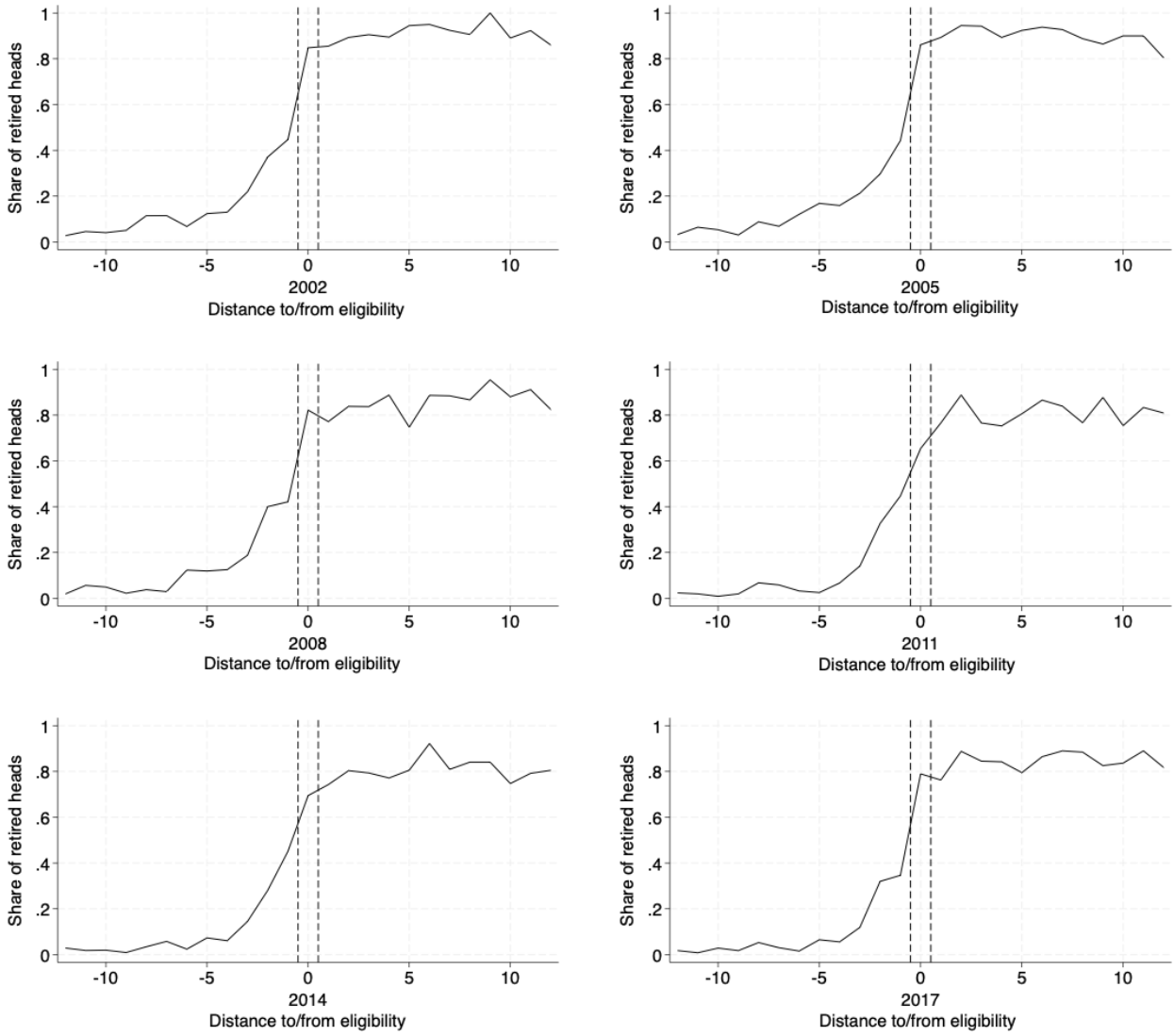

Figure 1. Share of retired household heads per year.

Figure 1 shows the share of retired household heads relative to their distance from eligibility, limited to ±12 years, for each year of observation. As can be seen, the relationship between retirement status and eligibility remains quite stable over the years. Further, there is a relatively small, but non-negligible, proportion of individuals who retire before reaching eligibility. However, as anticipated, this does not prevent us from identifying the causal effect of retirement on consumption when using eligibility status as an instrument for retirement. This is confirmed by the sharp discontinuity in the share of retired household heads at the eligibility threshold (distance from eligibility ≥0), as shown in Table 2 and Figure 1. This clear jump demonstrates that eligibility status strongly influences the decision to retire.

5. Results

In this section, we start by presenting some descriptive statistics that contain interesting retirement facts, and continue with the main estimation results. The description of the variables we use in our estimation exercises is reported in Table A.2 in the Appendix.

Table 1. Number and share of retired and non-retired individuals.

Year	Total	Non-retired (%)	Retired (%)	Retired in the sample period (%)
2002	4,714	3,211	1,503	-
		(0.68)	(0.32)	
2005	5,227	3,535	1,692	209
		(0.68)	(0.32)	(0.12)
2008	5,354	3,578	1,776	281
		(0.67)	(0.33)	(0.16)
2011	5,254	3,415	1,839	310
		(0.65)	(0.35)	(0.17)
2014	5,268	3,482	1,786	220
		(0.66)	(0.34)	(0.12)
2017	5,663	3,687	1,976	286
		(0.65)	(0.35)	(0.14)
Total	31,480	20,908	10,572	1,306
		(0.66)	(0.34)	(0.14)

As for the descriptive statistics, in Table 1, we report the number and proportion of retired and non-retired individuals over the sample period. As shown, retirees make up 30–35% of the total sample, with a slight upward trend over time. These figures are broadly consistent with those reported by Battistin et al. (2009) for the same period, although the proportion of retirees in the SHIW rises to nearly 40% in the final years (2002–2004). Table 1 also reports the number of new retirees each year, based on the SHF panel data. On average, new retirees represent approximately 15% of the total retired population in the survey.[18] This information is particularly relevant, as it allows us to isolate changes in consumption among retirees that occur after the retirement date but are not necessarily caused by retirement itself.

Table 2 reports the number of individuals eligible for retirement, along with the share of retired and non-retired individuals within this group. Note that the number of eligible individuals nearly doubles at $D = -2$ and $D = -1$ but remains relatively stable at around 800 for other values of D. Interestingly, the share of retirees increases steadily across the distribution of D, doubling at $D = 0$ and $D = 1$, and remaining above 80% for subsequent values. This pattern suggests that the majority of individuals retire as soon as they become legally eligible. However, a small subset, approximately 20–25% of those eligible, chooses to remain in the workforce beyond that point.

[18] Note that these figures do not take into account those individuals who begin participating in the survey in the year they retire.

Table 2. Eligible individuals and their share of retired/non-retired depending on D values.

	−4	−3	−2	−1	0	1	2	3	4
Non-retired	737	636	1,052	901	220	170	127	141	125
Retired	63	111	482	591	653	668	710	748	683
	0.08	0.15	0.31	0.40	0.75	0.80	0.85	0.84	0.85

Finally, Table 3 presents average expenditures on non-durable goods and on food and beverages (both at home and away), as well as total household income, liquid wealth, and illiquid wealth. In terms of expenditure, both non-durable consumption and spending on food and beverages exhibit a similar pattern. Interestingly, average expenditures in both categories are higher among retired individuals than among non-retired ones. This is somewhat surprising, given the well-documented evidence that peak consumption in the life cycle typically occurs well before retirement. However, when focusing specifically on individuals who retired between 2005 and 2017, whom we can track using the panel, their average expenditures in both categories are lower than those of both the retired and non-retired groups. This holds true whether we consider only the year of retirement (which is the most relevant for our analysis) or all subsequent years in which these individuals continue to participate in the survey (except in the case of food and beverage expenditure, where their spending is higher than that of the non-retired group).

Table 3. Non-durable and food and beverage expenditures, total income, and liquid and illiquid wealth for different samples.

			Retired in 2005–2017	
	Non-retired	Retired	First year retired	2005–2017
Non-durable expenditure	14,393.16	14,478.13	12,673.61	13,965.59
Food and beverages	6,628.916	7,028.624	6,587.096	6,975.611
Total income	57,131.94	48,016.04	48,081.12	49,152.16
Liquid wealth	95,493.32	183,895.1	82,697.9	117,476.7
Illiquid wealth	351,148.5	537,522.3	460,588.5	497,398

The observed decline in expenditure around the time of retirement may serve as informal evidence of the so-called *retirement consumption puzzle*. The subsequent recovery in spending does not necessarily contradict this interpretation, as it may reflect a rational response to changing economic circumstances.[19] In this regard, a comparison of total household income across groups appears to support this hypothesis. Specifically, the income of pensioners increases over the sample period, and remarkably, the average income observed in the year of retirement closely matches the average income of the broader retired sample.

Nevertheless, the comparison of average wealth across these groups provides more robust evidence against the predictions of the permanent income hypothesis. As expected, given typical life-cycle wealth profiles, both liquid and illiquid wealth are lower among non-retired individuals

[19] Nevertheless, the history of the Spanish economic cycle during the sample period makes this hypothesis implausible a priori.

than among retirees. For individuals observed in their first year of retirement, average liquid wealth is lower than that of non-retired individuals, while average illiquid wealth is higher. This pattern is, at first glance, consistent with the observed decline in expenditure upon retirement. However, the increase in both measures of wealth after retirement is inconsistent with the predictions of the canonical permanent income model. According to the model, there is no clear justification for such an increase (except possibly a desire to leave a legacy). Even so, it is difficult to rationalise postponing this goal until retirement, especially when it coincides with a decline in consumption during the initial years of retirement.

Now, we present the main results of our analysis. First, we estimate the eligibility for retirement using the SHF data, grouped by distance to retirement and year, as well as at the individual level. Second, we examine the impact of retirement on non-durable goods and food and beverage expenditures (both at home and away from home) using predicted eligibility based on the previous estimates. We then compare these findings with an alternative estimate based on predicted individual non-durable consumption, derived from parameters of an estimated individual consumption function. Third and last, we investigate whether the intertemporal behaviour of retired and non-retired individuals differs, by comparing the results from the estimation of the first-order condition for intertemporal optimisation across both groups of households.

Table 4 presents the auxiliary regressions used to determine retirement eligibility. Columns (2) and (3) show the results based on group averages of individual observations, classified according to the value of *D*, following the methodology of Battistin et al. (2009). In contrast, Column (1) reports the results from a fixed-effects model estimated on the same sample, exploiting the panel structure of the SHF. To facilitate a comparison with the results of Battistin et al. (2009), Column (3) presents the estimates based on the sample of male household heads, while Column (2) reports the results for the full sample, including both men and women. As shown, the estimates are nearly identical across the two samples, despite the male-only sample containing approximately 20% fewer observations.

The estimated coefficient for eligibility in the total sample is 0.534 (and statistically significant), with an R^2 of 97%. This estimate is slightly higher than the value reported by Battistin et al. (2009) for a male-only sample (0.435) but lower than the value obtained by Marini (2024) for a sample that does not distinguish by gender (0.775). This result suggests that, in Spain, becoming eligible for retirement increases the probability of actually retiring by 53.4%. Additionally, the estimated coefficients for the distance to retirement and its squared term are also statistically significant, though very small in magnitude. This indicates that, in the Spanish case, the distance to eligibility has a limited impact on the likelihood of retirement, a pattern consistent with Marini's (2024) findings for Italy.

As previously noted, the within-groups estimator controls for unobserved heterogeneity and allows the inclusion of variables beyond eligibility and distance to retirement. Using this method, the estimated coefficient for eligibility drops to 0.338, although it is still significant, reflecting the expected reduction when additional controls are included. Distance to retirement and its square remain negligible in effect. Notably, the coefficient for distance turns negative, as theoretically expected. Overall, both the individual-level and grouped data analyses suggest that distance to eligibility is not a particularly relevant factor in explaining retirement behaviour in Spain.

Regarding the other explanatory variables, the age of the main breadwinner is, as expected, highly significant and positively associated with the probability of retirement. In contrast, the partner's age is

not statistically significant, suggesting that each partner's retirement decision is largely independent and not influenced by the age, eligibility, or retirement status of the other.

Table 4. First-stage regression results.

	Panel data	Cross section	
	All	All	Only males
Eligibility	0.338***	0.534***	0.515***
	(0.016)	(0.024)	(0.024)
D	−0.005*	0.016***	0.022***
	(0.003)	(0.002)	(0.002)
D^2	−0.000***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)
Age head	0.030***	-	-
	(0.004)		
Age wife	−0.000	-	-
	(0.001)		
Self-employed head	−0.014	-	-
	(0.015)		
Current value of the pension plan	−0.000**	-	-
	(0.000)		
Annual contribution to the pension plan (individual)	−0.000***	-	-
	(0.000)		
Annual contribution to the pension plan (firm)	−0.000	-	-
	(0.000)		
Unemployed head	−0.232***	-	-
	(0.022)		
Unemployed wife	−0.014	-	-
	(0.022)		
Constant	−1.432***	0.311***	0.330***
	(0.219)	(0.020)	(0.019)
Year dummies	YES	YES	YES
Observations	12,473	144	144
R-squared	0.221	0.970	0.976
Number of households	7,581	7,581	5,859

Notes: 1. The dependent variable in the first column is the interest rate ($R_{i,t}$), and the share of pensioners in the total sample in the second and third columns. 2. Standard errors are in parentheses. 3. ***, **, and * mean statistically significant differences at 1%, 5%, and 10% levels, respectively.

Variables related to private pension plans also show high statistical significance, though with unexpected signs.[20] Except for the employer's annual contribution, which is not significant, the present value of the pension fund and the individual's annual contributions negatively affect the likelihood of

[20] The small size of the estimated parameters in this case can be explained by the different numerical dimensions of these variables and the endogenous variable.

retirement. The association between private pension funds and delayed retirement suggests that individuals may view these funds more as a savings instrument than as a primary means of financing retirement, unless, alternatively, this pattern reflects the necessity of extending their working lives to increase pension savings, when individuals make inadequate forecasts and insufficient provisions. The stringent tax treatment of cashing out the pension fund at retirement in Spain gives some support to this hypothesis. This withdrawal from the pension plan is classified as labour income rather than savings, subjecting them to a higher marginal tax rate than would apply in the latter case. Finally, employer contributions to private pension plans are not statistically significant, which likely reflects their limited importance and relatively small amounts.[21,22]

As for labour supply variables, *self-employed head*, a dummy indicating whether the main breadwinner has been predominantly self-employed, does not show a statistically significant effect, despite the general trend of longer working lives among the self-employed in Spain. The commonly accepted explanation is that they have contributed on a particularly low taxable base, which forces them to extend their working lives to increase the value of their initial pension. Finally, unemployment status variables show that being unemployed significantly reduces the probability of retirement for the main breadwinner, while the partner's unemployment status has no effect. Together with the non-significance of the partner's age, this reinforces the view that the retirement decision of the main breadwinner is largely autonomous and unaffected by the spouse's employment situation.

Table 5 reports the regression results from Equation (3) for non-durable expenditure and spending on food and beverages (both at home and away), using predicted retirement status based on the parameters estimated in Table 4. Columns (1) and (2) show the results using individual-level data; columns (3) and (4) present estimates based on grouped data for the total sample; and columns (5) and (6) report the grouped results for the male-only sample.

Regarding the grouped data estimates, our findings indicate that non-durable expenditure by Spanish households declines by 19.9% at retirement in the total sample and by 16% in the male-only sample. In contrast, the estimated reduction in food and beverage spending is not statistically significant for men, but shows a 12.6% decline in the total sample, significant at the 10% level.[23] As

[21] The sample includes over 9,100 observations of individual contributions to private pension schemes, compared to just over 2,200 individuals reporting employer contributions. Moreover, the average employer contribution is approximately 30% lower than the average individual contribution.

[22] The complexity of retirement decisions, particularly regarding timing, implies that policies aimed at influencing incentives may yield unintended and potentially counterproductive effects. Using Swedish administrative data, Kolsrud et al. (2024) demonstrated that efforts to extend working lives can reduce consumption. Skinner (2007) offers an insightful work of this complexity.

[23] Marini (2024) also reported a lower estimate of the decline in non-durable expenditure for the total sample (12.3%) compared to the sample of men (8.2%), and a higher estimate for women (19.9%). Our estimated decline in non-durable expenditure is twice that reported by Battistin et al. (2009), 9.83%, although their estimated reduction in food and beverage expenditure (14.09%) is broadly comparable to ours, both in the men's sample. Finally, a direct comparison of our results on food and beverage spending with theirs is not feasible, as they distinguish between food consumed at home and food consumed away from home.

in Battistin et al. (2009), both distance to eligibility and its square exhibit negative coefficients, although only the squared term is statistically significant.

Table 5. Second-stage regression results for non-durable and food and beverage expenditures.

	Non-durable consumption	Food and beverage consumption	Non-durable consumption	Food and beverage consumption
Pensioner	−0.174**	−0.097	-	-
	(0.067)	(0.071)		
Share of pensioners	-	-	−0.199**	−0.126*
			(0.078)	(0.072)
D	−0.001	−0.002	−0.006	−0.005
	(0.002)	(0.002)	(0.004)	(0.004)
D^2	−0.000***	−0.000**	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Age head	0.001	0.003	-	-
	(0.003)	(0.003)		
Male head	0.053***	0.060***	-	-
	(0.020)	(0.021)		
Self-employed	−0.006	−0.028	-	-
	(0.025)	(0.026)		
Employee	−0.012	−0.040	-	-
	(0.024)	(0.026)		
Years contributing PPF	0.003***	0.003***	-	-
	(0.001)	(0.001)		
Years contributing SS	−0.000	−0.000	-	-
	(0.001)	(0.001)		
Years as full-time worker	0.001	0.000	-	-
	(0.001)	(0.001)		
Unemployed head	−0.129***	−0.102***	-	-
	(0.029)	(0.030)		
Unemployed wife	−0.114***	−0.104***	-	-
	(0.023)	(0.024)		
Constant	5.315***	4.563***	5.303***	4.658***
	(0.142)	(0.149)	(0.053)	(0.048)
Time dummies	YES	YES	YES	YES
Observations	25,625	25,616	144	144
R-squared	0.026	0.044	0.572	0.629
Number of households	13,904	13,902		

Notes: 1. The dependent variable in this specification is total non-durable expenditure (first and third columns) and total food and beverages at home and away from home (second and fourth columns). 2. Self-employed: Dummy variable equal to 1 if the head of household was self-employed for the majority of their working life, and 0 otherwise. Employee: Dummy variable equal to 1 if the head of household was an employee for the majority of their working life, and 0 otherwise. 3. PPF: Private pension fund. SS: Social security. 4. Standard errors are in parentheses. 5. ***, **, and * mean statistically significant differences at the 1%, 5%, and 10% levels, respectively.

These results are within the range reported in the comparative literature, although they are closer to the higher end of the estimated effects of retirement on consumption. For example, Hurd and Rohwedder (2003) found that British and American households anticipate a 20% drop in consumption around retirement, with observed reductions ranging from 12% to 17%; Laitner and Silverman (2005) found that the drop is around 16% in the U.S. case. Similarly, Li et al. (2015) estimated a 19% decline in non-durable expenditure in China following the implementation of mandatory retirement.

Columns (1) and (2) of Table 5 present the fixed-effects model estimates based on individual-level data. The estimated decline in non-durable expenditure is broadly consistent with the results obtained using group-level averages, although slightly smaller in magnitude, at 17.4%. This difference can be attributed to the broader specification used in the individual-level analysis. Regarding food and beverage expenditure, the coefficient associated with retirement status is not statistically significant.[24]

Regarding distance and its squared term, the results are consistent with those obtained using group-level averages. In contrast, the age of the main breadwinner does not show a significant effect, likely because its influence is already captured by the distance variable. The dummy variable for the gender of the main breadwinner is statistically significant at the 1% level, indicating that households with a male head tend to have higher non-durable expenditure.

We also include dummy variables indicating whether the main breadwinner has been predominantly self-employed or employed throughout their working life. Neither variable is statistically significant. The number of years the main breadwinner has contributed to private pension plans has a positive and statistically significant effect on non-durable expenditure. In contrast, the number of years contributing to social security is not significant, likely due to the limited variation imposed by legal requirements for accessing public pensions. Similarly, the number of years worked full-time does not have a significant impact.

However, the unemployment status of both the head of the household and their partner is statistically significant at the 1% level, with both negatively affecting non-durable household expenditure, as expected, and with coefficients of similar magnitude. It is important to note that while the spouse's unemployment status does not influence the retirement decision, it does emerge as an important determinant, alongside the breadwinner's unemployment, of household spending on non-durable goods.

As explained in the previous section, we also assess the magnitude of the decline in non-durable expenditure at retirement by comparing the actual expenditure of each individual in the panel who retired during the sample period with the predicted expenditure based on a consumption function. This function depends on total wealth, regular income, and a comprehensive set of demographic and labour-related variables, and has been estimated using a panel of all active individuals, employing a GMM-IV approach. The estimation results for this consumption function are reported in Table A.1 of the Appendix.

Our estimate indicates that, in the first year of retirement, predicted non-durable expenditure exceeds actual expenditure by an average of 20.4% of pre-retirement spending. This finding is generally consistent with results from prior analyses. Ideally, the decline in spending should be

[24] Olaffson and Pagel (2025), using a fixed-effects model with Icelandic data, estimated that total spending decreases by 21.6%, and by 26.8% when controlling for income. For fuel and transport, the estimated decline exceeds 30%.

compared to the expenditure level individuals would have maintained had they not retired, which is unobservable. We expect this level of expenditure to be higher than the previous observed spending, provided that the three-yearly inflation rate has been positive. Specifically, our estimate measures the proportion of the prediction error for first-year retirees' non-durable expenditure (calculated using the specified consumption function) relative to the expenditure they would have incurred had they not retired. To construct this counterfactual, we project retirees' pre-retirement spending forward using the three-year average growth rate of non-durable expenditure among active individuals in the SHF (22.43%) over the 2002–2017 period.[25]

The related literature also reports mixed findings for Spain. For instance, Luengo-Prado and Sevilla (2012), analysing data from the Continuous Household Budget Survey (*Encuesta Continua de Presupuestos Familiares*, ECPF) between 1985 and 2004, identified a small decline in non-durable expenditure driven primarily by reduced spending on work-related items such as clothing, transport, and dining out, which they interpret as evidence that there is no puzzle in Spain, supporting the hypothesis of a substitution of market goods with home-produced goods following retirement. Labeaga and Osuna (2007), using the same dataset from 1985 to 1997, observed a decline of around 5% in subsidised categories like health and public transport, which they interpreted as evidence of the puzzle for the Spanish case. Finally, Christensen (2008) found no significant drop in spending except for medicines, another subsidised category. Overall, the evidence from these early studies links changes in spending at retirement primarily to work-related or health-related expenses. With respect to the former, it appears unlikely that such costs could account for a substantial decline in overall spending, particularly in food expenditure (see note 25).

In contrast, Labeaga and Sánchez-Robles (2026), using data from both the ECPF and the Permanent Consumption Survey (*Encuesta Permanente de Consumo*, EPC) across two earlier sub-periods and a more recent one (2016–2022), found a negative correlation between consumption growth and retirement status. They interpreted this as additional evidence of the retirement consumption puzzle in Spain. These mixed findings are related to differences in data sources, methodologies, and sample periods.[26]

Finally, our last exercise examines whether retirees and non-retirees differ in their intertemporal non-durable consumption behaviour. It is important to note that the response of non-durable expenditure to changes in interest rates may not necessarily be influenced by the reduction in spending that occurs upon retirement. Individuals might simply adjust their overall expenditure level to reflect changes in income while continuing to optimise consumption over time. In our analysis, which follows Labeaga and Sánchez-Robles (2026) approach, we estimate a first-order intertemporal optimisation equation for non-durable expenditure across three groups: the full sample, active main earners, and

[25] Although this figure may seem high, it is important to consider the context: between 2002 and 2017, the Spanish economy underwent a significant housing bubble, which burst in 2008. During the boom years, average three-year individual consumption growth rates exceeded 30%, followed by low growth rates afterwards. Notably, Spain also has one of the highest home ownership rates in the world, which further amplifies the impact of housing market fluctuations on consumption patterns.

[26] It should also be noted that most of these results are based on the ECPF and EPC surveys, which are conducted quarterly and follow individuals for a maximum of two years. This limited tracking period may not be sufficient to accurately capture and analyse consumption changes triggered by retirement.

retirees. The estimation of this equation allows us to estimate the intertemporal elasticity of substitution (IES), which measures the response of the growth rate of non-durable consumption to the real interest rate; it measures how changes in the real interest rate induce individuals to substitute present consumption for future consumption.

To address potential endogeneity issues, we employ a fixed effects model estimated using the generalised method of moments (GMM) with instrumental variables. The specification used includes a broad set of demographic and labour-related variables, plus yearly dummies and a time trend. Table 6 presents these results, where we report the estimates for the total sample (Column 1), the estimates for the sample of active individuals (Column 2), and the estimates for the retirees' sample (Column 3).

For the total sample, we get an estimate for the intertemporal elasticity of substitution of 0.839, statistically significant at the 1% level (see Column 1).[27] This specification controls for a broad set of demographic and labour variables, a time trend, and year fixed effects. Among the explanatory variables, the dummy variable that accounts for a retired main breadwinner is not statistically significant, suggesting that retirement status does not affect the growth rate of non-durable expenditure. This result is at odds with Labeaga and Sánchez-Robles (2026), although they used different data sets and sample periods.[28]

To further investigate this issue, in Column (2), we replicate the baseline specification for the active main breadwinner sample. The estimated IES slightly increases to 0.847, and the set of statistically significant variables remains consistent. The Hansen test *p*-value improves to 91%, likely reflecting greater sample homogeneity.

Finally, in Column (3), we report the results for retired household heads. Here, the estimated IES decreases marginally to 0.829, with a highly significant Hansen test *p*-value of 99.8%. Age and its square lose significance, while only the change in the number of adults remains significant. Among other variables, only the partner's retirement dummy is significant.

Table 6. Estimation of the first-order intertemporal optimization condition for consumption.

	All	Non-retirees	Retirees
IES	0.839***	0.847***	0.829***
	(0.019)	(0.022)	(0.038)
Number of members	0.015	0.036	−0.002
	(0.026)	(0.029)	(0.039)
Number of adults	0.106***	0.078***	0.144***
	(0.024)	(0.027)	(0.046)
Age head	0.022***	0.019***	0.014
	(0.005)	(0.005)	(0.011)
Age head squared	−0.000***	−0.000***	−0.000*
	(0.000)	(0.000)	(0.000)

Continued on next page

[27] We obtain a *p*-value of 64% of the Hansen test of overidentification restrictions, which indicates that our instrument set is valid.

[28] Our estimates are consistent with previous findings on the elasticity of intertemporal substitution in Spain for the same sample period (see Cutanda and Sanchis, 2025a).

	All	Non-retirees	Retirees
Education 1	0.028	0.093**	−0.034
	(0.025)	(0.036)	(0.034)
Education 2	0.049*	0.124***	−0.039
	(0.026)	(0.038)	(0.037)
Education 3	0.056*	0.132***	−0.034
	(0.029)	(0.041)	(0.045)
Education 4	0.075***	0.159***	−0.013
	(0.027)	(0.040)	(0.035)
Education 5	0.018	0.087*	−0.046
	(0.034)	(0.047)	(0.054)
Employee head	−0.031	−0.017	-
	(0.032)	(0.031)	
Unemployed head	0.086	0.133	-
	(0.104)	(0.098)	
Self-employed head	−0.094**	−0.098***	-
	(0.038)	(0.037)	
Pensioner head	0.028	-	-
	(0.027)		
Multiple job employee head	−0.399*	−0.432**	-
	(0.211)	(0.213)	
Part time job head	0.053	0.056	-
	(0.091)	(0.088)	
Employee wife	−0.076***	−0.094***	−0.114
	(0.028)	(0.031)	(0.073)
Unemployed wife	−0.039	−0.111	0.045
	(0.113)	(0.137)	(0.186)
Self-employed wife	0.032	0.064	−0.173*
	(0.043)	(0.048)	(0.101)
Pensioner wife	−0.060**	−0.051	−0.062**
	(0.026)	(0.042)	(0.031)
Part-time job wife	0.096	0.066	0.297
	(0.082)	(0.085)	(0.274)
Time trend	YES	YES	YES
Year dummies	YES	YES	YES
Observations	14,379	9,189	5,190
Hansen test of overidentifying restrictions			
Chi2(3)	1.686	0.512	0.125
p-value	(0.640)	(0.916)	(0.998)

Notes: 1. The dependent variable in this specification is the increase in the logarithm of total non-durable expenditure. 2. Standard errors are in parentheses. 3. The educational variables are as follows: Education 1 corresponds to illiterate individuals (reference category); Education 2 includes primary education and education for labour market insertion (requiring less than 300 hours of secondary school instruction); Education 3 corresponds to individuals with secondary education; Education 4 captures specialized courses and labour insertion programmes in Spain; Education 5 corresponds to college education; Education 6 corresponds to any

special education requiring to have previously completed secondary education. 4. The instrument variable set used is the lagged value of all explanatory variables and the lag of all the dummies for the labour status of the household head and their partner. This instrument set is expanded, in columns (1) and (3), with the lag of the household's debt, a dummy on future saving, and of the total household's net worth, and, in column (2), with the lags of the household's debt and of the available credit to the household. 5. ***, **, and * mean statistically significant differences at the 1%, 5%, and 10% levels, respectively.

The results presented in Table 6 show that there are distinct consumption patterns between households with active main breadwinners and those whose main breadwinners are retired, reflecting their respective stages in the life cycle. It is worth noting that the only statistically significant demographic variable for the retiree sample is the change in the number of adults, consistent with findings from other economies.[29] In households with active earners, a greater number of variables show statistical significance, and the model specification is much broader, with demographic and labour-related factors playing a more prominent role. In contrast, retired households exhibit fewer significant variables and a narrower specification. The estimated values of the IES across the different samples are of a similar order of magnitude to those recently obtained using U.S. data and lie within the range reported in previous studies based on SHF data (see Cutanda and Sanchis-Llopis, 2025a). Notably, the differences across samples are very small, suggesting that the responsiveness of Spanish consumption to changes in the real interest rate does not vary substantially with retirement status. Despite these small differences, the response to changes in the real interest rate is largely consistent across both groups, suggesting that intertemporal optimisation behaviour remains unaffected by the change in non-durable expenditure observed at retirement.[30] Given the anticipated growth in the retired population in the near future, this finding offers reassuring implications for the effectiveness and management of monetary policy. However, it should be noted that this does not diminish the serious challenges that fiscal policy will continue to face.

6. Conclusions

In this paper, we have examined the relevance of the so-called *consumption retirement puzzle* in Spain, using data from the SHF. Since the early 2000s, a growing body of empirical evidence has documented a significant decline in non-durable consumption following retirement in developed countries. This phenomenon poses a serious challenge to the permanent income/life cycle model, particularly given the predictable nature of retirement and its expected consequences. As population

[29] Empirical evidence reveals significant differences in the economic behaviour of retired couples versus singles. For instance, Poterba et al. (2011) identified a substantial decline, or at least no increase, in assets among households undergoing a family status transition. Similarly, van Ooijen et al. (2015) emphasized the importance of bequests and transfers following the death of the first spouse. Additionally, De Nardi et al. (2025) found that savings of retired singles tend to decrease with age, whereas those of retired couples generally increase. See also Addoum (2017) for further insights.

[30] We conducted a formal test to assess the equality of the estimated IES between households with an active head and those with a retired head. We cannot reject the null hypothesis of equality of the IES. These results are available from the authors upon request.

ageing accelerates across these countries, the puzzle has gained increasing importance in academic and policy discussions.

Previous empirical evidence regarding the *consumption retirement puzzle* in Spain has produced mixed evidence. In general, most of these early studies either downplayed the existence of the puzzle in Spain or concluded that its magnitude was significantly lower than in comparable countries. However, their findings were generally based on outdated sample periods and relied on datasets that were not well-suited for analysing intertemporal optimisation models, limiting the robustness of their conclusions.

In this study, we estimate the decline in Spanish non-durable consumption at retirement using various estimation methods for the period 2002–2017. Specifically, we apply the regression discontinuity approach proposed by Battistin et al. (2009) to grouped data (clustered by distance to eligibility and time period). Further, we apply a similar approach to individual-level data, which allows us to assess the relevance of variables beyond distance to eligibility in influencing the retirement decision. Our results suggest a decline in non-durable expenditure upon retirement in Spain between 17% and 20%. However, we find no statistically significant evidence of a reduction in spending on food and beverages. Regarding other determinants, retirement seems to be influenced more by age and the employment or unemployment status of the household's main earner than by proximity to eligibility. The existence and value of a private pension plan do not appear to accelerate retirement (likely due to the tax penalties associated with cashing out the pension plan at retirement under the Spanish tax law). Finally, we find no evidence that the age or retirement status of a spouse affects the individual's retirement decision.

When we estimate the decline in non-durable expenditure as the prediction error derived from applying the estimated parameters of a consumption function (based on the sample of active individuals), the result is slightly higher but remains within a similar range. This finding reinforces our earlier conclusions regarding the existence and magnitude of the *consumption retirement puzzle* in Spain.

The final exercise conducted involves analysing the intertemporal optimisation behaviour of active and retired individuals by comparing the estimated elasticities of intertemporal substitution (IES) for both groups. These estimates are derived from the standard first-order condition of intertemporal optimisation in non-durable consumption. The resulting IES values, although aligned with recent empirical literature, are not statistically different. We interpret this as positive news for the effectiveness of monetary policies in the context of population ageing, as it suggests that individuals' responsiveness to changes in interest rates does not change after retirement. Notably, the response to interest rate changes appears similar for both groups, despite their otherwise distinct patterns of intertemporal behaviour, as highlighted in our findings. In practical terms, beyond the effect of interest rate changes, only variations in household size are statistically significant for retirees. This result is consistent with previous empirical evidence from other comparable economies.

Therefore, we consider that our empirical methodology does not provide a definitive explanation for the decline in Spanish expenditure upon retirement. However, several factors suggest that conventional explanations are insufficient. The fall in expenditure appears largely temporary, making it unlikely that reductions in work-related expenses or the substitution of out-of-home consumption with home production can fully account for the observed decline. Indeed, the expenditure that could

plausibly be replaced by home-produced goods is relatively small. Similarly, medical expenses are an improbable driver, given the protection offered by the Spanish social security system, which is similar to that of some European countries. The limited relevance of these explanations in the Spanish context is confirmed by recent empirical evidence for Spain (see Labeaga and Sánchez-Robles, 2026).

A more plausible explanation lies in patterns of consumption, saving, and wealth accumulation that deviate from predictions of the canonical intertemporal choice model. In particular, the bequest motive and a pronounced reluctance to liquidate real estate assets, especially relevant in Spain due to its high home ownership rate, may play a central role. Psychological factors and bounded rationality may further contribute to specific phenomena, such as the marked initial drop in expenditure. Overall, these findings suggest that retirement consumption in Spain is shaped by a complex interplay of economic and behavioural factors, highlighting the need for further research.

Our findings reveal a significant decline in non-durable consumption at retirement, coupled with a broadly unchanged responsiveness to interest rate movements before and after retirement. Taken together, these results point to a more complex underlying phenomenon than is typically assumed. While much of the existing literature has concentrated on the immediate effect of retirement, our evidence suggests that this initial adjustment may be less relevant than commonly believed. In particular, the persistence of post-retirement consumption patterns following the initial drop appears to pose a more serious challenge to the life-cycle/permanent income model than the drop itself. Moreover, the fact that none of the explanations proposed in the literature can independently account for the evidence presented here highlights the multifaceted nature of the issue. This complexity, in turn, underscores the need for larger samples and longer observation periods to fully understand the dynamics of consumption around retirement.

Use of AI tools declaration

The authors declare they have not used Artificial Intelligence (AI) tools in the creation of this article.

Author contributions

All authors contributed equally to the design, conceptualisation, methodology, empirical estimation, discussion, and writing of the article.

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Conflict of interest

All authors declare no conflicts of interest in this paper.

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