Erasmus University Rotterdam
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# The Saving Rate of Euro Area Households and the Interest Elasticity of Saving during the Great Recession

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## Abstract

Macroeconomic models and empirical evidence establish a positive relationship between real returns and the rate of saving. Nevertheless, in 2009, despite plummeting real interest rates to saving deposits, the saving rate of Euro Area households experienced a large increase. It is the aim of this paper to investigate whether this unexpected upsurge, could have been caused by a temporarily negative interest elasticity of saving, triggered in turn by a decrease in households' elasticity of intertemporal substitution. To do so, we make use of a two period model of life-cycle consumption. Contrary to what we hypothesised, our results suggest that households' interest elasticity of saving increased in the vicinity of 2009, in spite of a decrease in their elasticity of intertemporal substitution.

**Keywords:** Saving rate; real rate of return; interest elasticity of saving; structural breaks; intertemporal elasticity of substitution

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## 1. Introduction

### *1.1 Background*

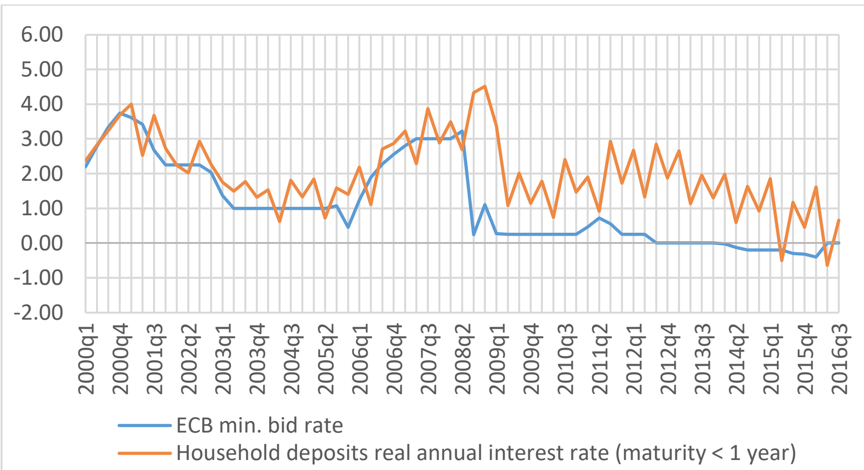
After a ten-year period of sustained economic growth, the burst of the financial crisis in early 2008 led the Euro Area to experience a severe economic slowdown. Only two quarters later, the economy of the Euro Area entered into what later would become a long-lasting recession (Figure 1).

*Figure 1 – Euro Area year-over-year (YOY) quarterly GDP growth rate*


Source: Eurostat

As a response to the rapid economic contraction, in late 2008, the European Central Bank (ECB) started an uninterrupted reduction of the minimum bid rate. In less than one year, the rate at which banks can borrow from the ECB went down by 92 percent. That reduction had one main goal, lowering market real returns in order to stimulate consumption and discourage saving (Figure 2).

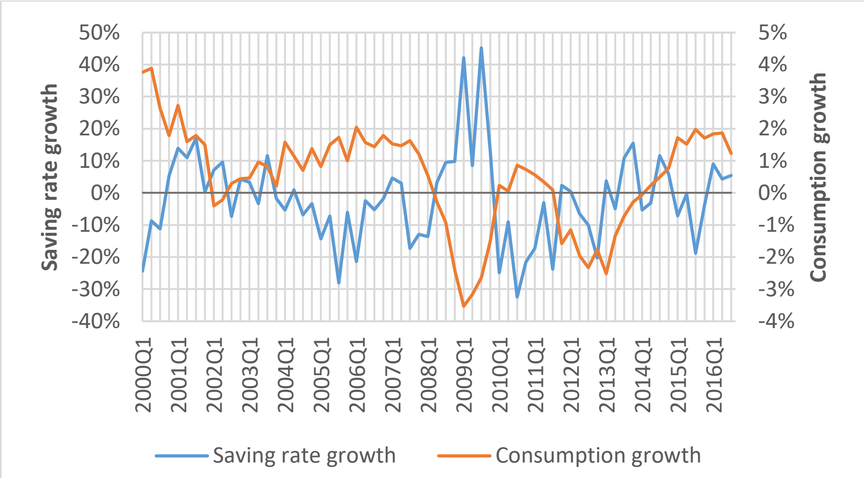
*Figure 2 – Euro Area interest rates*




Source: European Central Bank

As expected, the expansionary measures implemented by the ECB translated into a considerable decrease of the real interest rate on household saving deposits (Figure 2). However, households' immediate reaction to this decrease was highly unexpected. The saving rate of Euro Area households increased by 42 percent (year-over-year) in the first quarter of 2009 and by 45 percent in the third quarter of that same year. As a result, in 2009 the annual household saving rate returned to the high values registered before 2005, above 8 percent. Right after this sudden upsurge, in 2010 the ratio of net saving to net disposable income declined again by 19 percent, cancelling out the large increase of the previous year (Figure 3).

*Figure 3 – Euro Area YOY quarterly saving rate growth and consumption growth*




Source: Eurostat

### *1.2 Motivation*

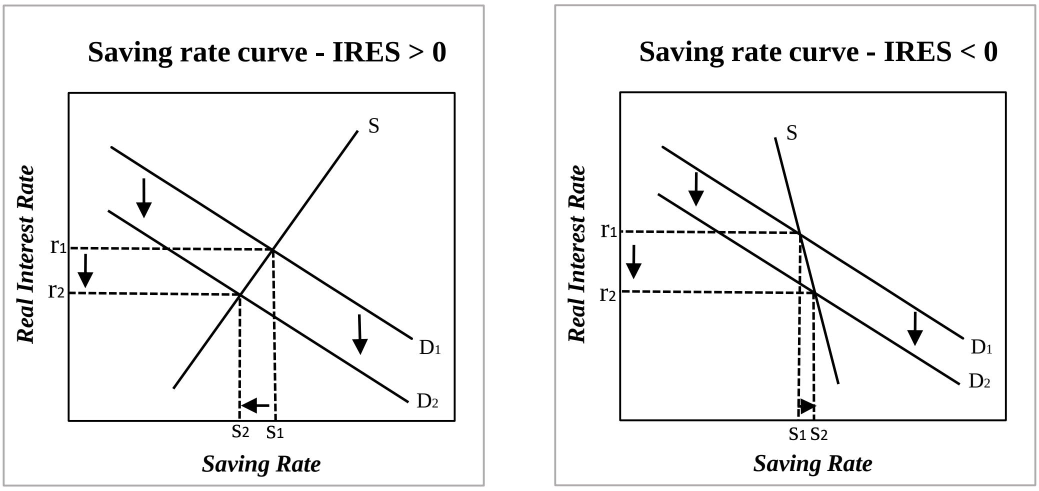
Mody et al. (2012) thoroughly investigated the determinants of 2009's temporary saving rate upsurge, finding the increase of labour income uncertainty to be one of its major causes. However, if this was indeed the case, we cannot avoid wondering why, despite the persistence of labour income uncertainty, the saving rate of Euro Area households decreased in 2010. As a consequence, we believe that the sharp fluctuation of Euro Area household saving rate between 2009 and 2010 demands further investigation.

In 2016, in an opinion article published in Project Syndicate, Daniel Gros, while providing no empirical evidence, suggested that the difficulties of the Euro Area to return to the consumption growth rates experienced before the outbreak of the financial crisis, could be explained by a backward bend in the lower end of the curve of saving supply. This would imply that, when the real rate of return tended to zero, individuals' interest rate elasticity of saving (IRES) would become negative, which would have led household saving to increase as a response to the expansionary monetary policy implemented by the European Central Bank.

We find the hypothesis proposed by Daniel Gros somewhat implausible given that, since early 2010, the saving rate of Euro Area households has a decreasing trend despite the present low real returns. Nevertheless, we do believe it worthwhile to investigate whether a temporarily negative IRES, caused by a decrease in the elasticity of intertemporal substitution (EIS), could explain part of 2009 increase in household saving rate.

As we can see in the first graph of Figure 4, with a positive IRES, that is, an upward sloping supply curve of saving, the decrease of the real rate of return from $r_1$ to $r_2$ would have caused the demand curve to move downwards, leading the equilibrium point to decrease from $s_1$ to $s_2$. However, as we can see on the second graph of this same figure, the reduction of the IRES, caused by a decline of households' EIS, could have led to the rotation of the supply curve of saving and, thus, a downward sloping supply curve of saving. If this had been the case, the reduction of the real rate of return could have led to an increase of the saving rate from $s_1$ to $s_2$.

*Figure 4: Market of loanable funds*




Our hypothesis, would be compatible with the positive relationship between the real rate of return and household saving rate found by Mody et al. (2009) for the period between 1980 and 2010. Since Mody's et al. (2009) empirical specification did not allow for time variation in the IRES, 2009’s temporary decrease in this parameter, if present, would have been overlooked.

*1.3 Preview*

It is our aim to shed some more light on this matter. To do so, we make use of a deterministic two-period model describing individuals’ consumption and saving decision under the life-cycle hypothesis. Based on this framework, we first regress the household saving rate on the real return and a series of complementary arguments intended to control for third factors, which are expected to affect the relationship between the main variables of interest.

In order to examine the presence of structural breaks in the relationship between the saving rate and the real rate of return in the vicinity of 2009, we employ the Quandt-Andrews test together with the Chow test for structural breaks. Then, following the break dates suggested by these tests, we add an interaction term between the real rate of return and a period dummy as a regressor. This interaction term intends to capture the sign and magnitude of the variation in the IRES during the inter-break period.

Finally, with the aim of analysing whether the hypothetical drop in the IRES could have been caused by a sudden temporary decrease in individuals’ EIS, we regress the rate of consumption growth on the same interaction term employed in the saving rate function.

## 2. Saving behaviour under the life-cycle hypothesis

*2.1 Theoretical analysis*

*2.1.1 Deriving the saving rate function to estimate the IRES*

The following section thoroughly analyses the effect of interest rate changes on households’ saving decisions under life-cycle hypothesis. In addition, it examines how changes in the EIS can affect the relationship between the saving rate and the real return to households’ wealth.

For the last half a century, the life-cycle hypothesis proposed by Modigliani and Brumberg (1954) has been the standard theoretical framework for analysing people’s saving decisions. According to this model, individuals form estimates of their life-time resources, that is, the actualized value of present and future earnings, and assign, bearing in mind their future needs and independently of their current income, a portion of that estimate to current consumption. The assumed presence of unconstrained capital markets enables individuals to smooth consumption at a significantly stable rate over the life-cycle: borrowing during youth, accumulating wealth along their work lives and dissaving after retirement. Per-period saving is determined by the difference between current income and permanent income.

Following the life-cycle hypothesis, the effect of an interest rate increase on aggregate saving can be disaggregated into three effects of differing sign and power:

(I) Firstly, the substitution effect lowers the price of future consumption relative to that in the present, driving individuals to increase current saving in order to enjoy a higher level of consumption in the future (Deaton, 1992).

(II) Secondly, counter to the substitution effect, the income effect decreases the present value of future consumption. Alternatively stated, it lowers the amount of saving required to attain a desired level of consumption in the future. This, makes individuals relatively wealthier and encourages them to increase current consumption at the expense of saving (Wilcox, 1993).

(III) Finally, complementing the substitution effect, the wealth effect lowers the present value of future labour and financial income, thus increasing saving at the expense of consumption. The revaluation of individuals' future capital-related earnings requires one to differentiate between fixed-rate and floating-rate assets. Concerning the former, when interest rates increase, the stream of future income remains constant. Consequently, as future cash flows are more heavily discounted, the value of fixed-rate assets decreases and so does individuals' wealth. In contrast, the future cash flows of variable-rate assets do increase, leaving their value unaffected by the rise of the discount rate or, in other words, no wealth effect (Elmendorf 1996).

To more clearly depict the relationship between households' consumption and saving decisions and the real rate of return, we make use of a two-period framework in which individuals, subject to an intertemporal budged constraint, determine per period consumption in order to maximize life-time utility. Consumer preferences are described by Constant Relative Risk Aversion (CRRA) utility function given by:

$$u(c) = \frac{c^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}}$$

where $\sigma$ is individuals' EIS. This parameter, of major importance in our later discussion, captures the responsiveness of the rate of growth in consumption to innovations in the real rate of return. Larger values of the EIS imply a higher willingness of households to substitute consumption across periods to benefit from interest rate changes. It can be formally defined as:[1]

$$EIS = \frac{\partial ln \, {}^{c_2}/_{c_1}}{\partial \ln(1+r)} = \sigma \tag{1}$$

Once defined the EIS, we proceed to derive the saving rate function.[2] Using CRRA utility, the two period consumer's problem is:

$$\max. \quad \frac{c_1^{\,1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}} + \beta \frac{c_2^{\,1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}} \quad \text{subject to} \quad c_1 + \frac{c_2}{(1+r)} = a_0 + y_1 + \frac{y_2}{(1+r)}$$

where $c_t$ and $y_t$ respectively denote the amount of consumption and income in period t (t = 1, 2), $a_0$ stands for consumer's initial endowment of financial assets, r is the real rate of return to wealth and β is agent's subjective discount factor, which captures how much they care about future consumption.

Setting-up and solving the Lagrangian, the first order condition for the consumer's problem can be summarized as:

$$\frac{c_1^{\,-\frac{1}{\sigma}}}{1+r} = \beta \, c_2^{\,-\frac{1}{\sigma}} \tag{2}$$

Solving for consumption in period 2, the Euler equation implies:

$$c_2 = [\beta \, (1+r) \, ]^{\sigma} c_1 \tag{3}$$

Plugging Equation 3 inside the intertemporal budget constraint and solving for $c_1$, we can obtain the consumption function in period one. Furthermore, making use of the fact that actual saving equals current income minus current consumption ($s_t = y_t - c_t$), we can obtain savings in t=1:

$$s_1 = y_1 - \left[a_0 + y_1 + \frac{y_2}{(1+r)}\right] * \frac{1}{[1 + \beta^{\sigma} \, (1+r)^{\sigma-1} \, ]}$$

Finally, dividing both sides of the saving function by $y_1$, we can find the rate of saving:

$$\frac{s_1}{y_1} = 1 - \left(1 + \frac{a_0}{y_1} + \frac{y_2/y_1}{1+r}\right) * \frac{1}{1 + \beta^{\sigma} * (1+r)^{(\sigma-1)}}$$

[1] The full derivation of the EIS is presented in Appendix A.3.
[2] The full derivation of the saving rate function is presented in Appendix A.1.

Now, before taking the first order Taylor approximation of this equation, we can simplify the notation by substituting $\frac{a_0}{y_1}$ by A, $\frac{y_2}{y_1}$ by Y and $(1+r)$ by R:

$$\frac{s_1}{y_1} = 1 - \left(1 + A + \frac{Y}{R}\right) * \frac{1}{1 + \beta^{\sigma} * R^{(\sigma-1)}}$$

In addition, taking the natural logarithm of both sides and using the fact that, since the rate of saving is considerably close to zero, $ln(1 - \frac{s_1}{y_1})$ is approximately equal to $(-\frac{s_1}{y_1})$, this expression can be further simplified to:

$$\frac{s_1}{y_1} = ln\left(1 + \beta^{\sigma} R^{(\sigma-1)}\right) - ln\left(1 + A + \frac{Y}{R}\right) \qquad (4)$$

Finally, we take a first order Taylor approximation of Equation 4 around $A = \bar{A}$, $Y = \bar{Y}$ and $R = \bar{R}$. This, will allow us to obtain households' saving rate as a linear function of relative wealth, relative future income and the real rate of return, each of them multiplied by their partial effect on household saving rate. This equation is given by:

$$\frac{s_1}{y_1} \approx c + \beta_1(R - \bar{R}) + \beta_3(Y - \bar{Y}) + \beta_4(A - \bar{A})$$

where:

$$c = f(\bar{A}, \bar{Y}, \bar{R}) = ln\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right) + \beta^{\sigma} \bar{R}^{(\sigma-1)}$$

$$\beta_1 = \frac{\partial\, s_1/y_1}{\partial R} = \underbrace{(\sigma - 1)\frac{\beta^{\sigma} \bar{R}^{\sigma-2}}{1 + \beta^{\sigma} \bar{R}^{(\sigma-1)}}}_{\text{(I) subst. and (II) income effect}} + \underbrace{\frac{1}{\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right)} * \frac{\bar{Y}}{\bar{R}^2}}_{\text{(III) wealth effect}} \qquad (5)$$

$$\beta_3 = \frac{\partial\, s_1/y_1}{\partial Y} = -\frac{1}{\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right)} * \frac{1}{\bar{R}}$$

$$\beta_4 = \frac{\partial\, s_1/y_1}{\partial A} = -\frac{1}{\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right)} * 1$$

Moving $\bar{R}$, $\bar{Y}$ and $\bar{A}$ inside the constant term (α) we can obtain what is going to be our empirical specification, later used to estimate the interest elasticity of saving:

$$\frac{s_1}{y_1} \approx \alpha + \beta_1 R + \beta_3 Y + \beta_4 A \tag{6}$$

Equation 5 summarizes the three effects that define the interest elasticity of saving: whereas the (I) substitution and (II) income effects, which are collapsed into a single term[3], are related to the first term on the right hand side of this equation, the (III) wealth effect is captured by the second term.

Carefully observing Equation 5, we can see that the EIS plays a major role in determining the response of households' saving rate to changes in the real return. More concretely, this parameter governs the sign and magnitude of the net effect resulting from the aggregation of the opposing substitution and income effects. If the EIS is larger than one, the positive substitution effect outweighs the negative income effect. In contrast, if the EIS is less than one, the income effect outweighs the substitution effect. Finally, if the EIS is equal to one, these two effects neutralize each other, which results in a positive IRES solely determined by the wealth effect.

In order to show the theoretical feasibility of a negative interest elasticity of saving as a result of a decrease in the elasticity of intertemporal substitution, in Table 1 we present some preliminary calculations using Equation 5 and quarterly time-series data for the Euro Area. The real return to individuals' wealth ($R_t$) is approximated using the real return to deposits held by households and NPISH with an agreed maturity up to one year, relative future income ($Y_t$) is the ratio of net disposable income in t+1 to that in t and relative financial wealth ($A_t$) is defined as the ratio of net financial wealth to net disposable income. With regards to the consumers' preferences, we assume a constant subjective discount factor (β) of 0.98, an elasticity of intertemporal substitution (σ) of 1 before and after 2009, and a 50 percent decrease of the EIS during 2009. Plugging this values inside Equation 5, the estimates obtained for the interest elasticity of saving are as follows:

*Table 1 – Hand calculations using Equation 5*

| Column # | *1* | *2* | *3* | *4* | *5* | *6* | *7* | *8* |
|---|---|---|---|---|---|---|---|---|
| Quarter | *σ* | *β* | *R* | *A* | *Y* | *SE + IE* | *WE* | *IRES* |
| *2008Q1* | 1.00 | 0.98 | 1.03 | 8.72 | 1.09 | 0.00 | 0.09 | 0.094 |
| *2008Q2* | 1.00 | 0.98 | 1.03 | 7.83 | 0.91 | 0.00 | 0.07 | 0.089 |
| *2008Q3* | 1.00 | 0.98 | 1.04 | 8.42 | 1.04 | 0.00 | 0.08 | 0.092 |

[3] Substitution and income effect could not be disentangled.

| Column # | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 |
|---|---|---|---|---|---|---|---|---|
| Quarter | $\sigma$ | $\beta$ | *R* | *A* | *Y* | *SE + IE* | *WE* | *IRES* |
| *2008Q4* | 1.00 | 0.98 | 1.05 | 7.74 | 0.95 | 0.00 | 0.04 | 0.090 |
| *2009Q1* | 0.50 | 0.98 | 1.03 | 7.92 | 1.08 | -0.25 | 0.11 | -0.137 |
| *2009Q2* | 0.50 | 0.98 | 1.01 | 7.51 | 0.92 | -0.75 | 0.52 | -0.150 |
| *2009Q3* | 0.50 | 0.98 | 1.02 | 8.50 | 1.05 | -0.63 | 0.41 | -0.147 |
| *2009Q4* | 0.50 | 0.98 | 1.01 | 8.07 | 0.94 | -1.32 | 1.17 | -0.153 |
| *2010Q1* | 1.00 | 0.98 | 1.02 | 8.68 | 1.08 | 0.00 | 0.34 | 0.097 |
| *2010Q2* | 1.00 | 0.98 | 1.01 | 7.82 | 0.93 | 0.00 | 0.19 | 0.094 |
| *2010Q3* | 1.00 | 0.98 | 1.02 | 8.62 | 1.05 | 0.00 | 0.07 | 0.094 |
| *2010Q4* | 1.00 | 0.98 | 1.01 | 8.28 | 0.95 | 0.00 | 0.06 | 0.090 |

Notes: (1) SE stands for substitution effect; (2) IE stands for income effect; (3) WE stands for wealth effect; (4) the IRES is the result of aggregating SE, IE and WE.

As we can see, with σ = 1 both before and after 2009, the aggregation of the substitution and income effect is equal to zero (Column 6). As a result, the IRES (Column 8) is solely determined by the wealth effect (Column 7) and thus positive. In contrast, with σ = 0.5 during 2009, the positive strength of the substitution effect diminishes. Consequently, the net effect resulting from the aggregation of the substitution and income effect (Column 6) becomes negative, which in turn results in a lower and negative interest elasticity of saving (Column 8). This potentially negative IRES, could help explaining why in 2009, despite a 49 percent decrease in the real rate of return, the saving rate of Euro Area households rose by 20 percent.

### *2.1.2 Deriving the consumption growth function to estimate the EIS*

In order to find the expression to estimate the EIS we have to go back to Equation 3:

$$c_2 = \left[\beta\,(1+r)\,\right]^{1/\theta} c_1$$

Moving $c_1$ to the left hand side of the equation and taking natural logarithms of both sides we can find consumption growth. Then, taking the first order Taylor approximation of this expression around r=0 we can find consumption growth as a function of the real return and the partial effect of this latter variable on the former, that is, the EIS[4]:

$$\ln\left(\frac{c_2}{c_1}\right) = f(0) + \frac{\partial\, ln\, c_1/c_2}{\partial r}\,(r-0)$$

[4] The full derivation can be found in Appendix A.2.

$$\ln\left(\frac{c_2}{c_1}\right) = c + \sigma r \quad (7)$$

*2.1.3 Disentangling the inverse relationship between EIS and RRA*

Before concluding this section, we would like to make a remark on an important implication derived from the use of CRRA utility. This specific utility function imposes an inverse relationship between the elasticity of intertemporal substitution and individuals' degree of relative risk aversion (RRA), as can be shown by solving the expression for the coefficient of relative risk aversion proposed by Arrow and Pratt:

$$RRA = -\frac{u''(c)}{u'(c)} * c = = -\frac{-\theta c^{-\theta-1}}{c^{-\theta}} * c = \theta c^0 = \theta$$

As a result of this inverse relationship, when the σ tends to infinity, individuals exhibit relative risk neutrality (θ → 0); when the EIS tends to zero, individuals present an infinitely large RRA (θ → ∞); and when the EIS is smaller than zero, individuals show to appreciate relative risk (θ < 0). This implication, however, conflicts with some of the empirical findings on these two parameters.

Even though the EIS is usually found between 0.5 and 1, evidence presented by Hall (1988) suggests that it could lay between 0 and 0.1. While this latter estimate, despite being extremely low, may be feasible from a consumer behaviour point of view, as explained by Hall (1989), "the corresponding conclusion that the RRA is close to infinity is incompatible with the observed willingness of consumers to take on risk."

To avoid the shortcomings of a reciprocal relationship between risk attitudes and intertemporal substitution, Epstein and Zin (1989) develop a recursive utility function that allows to break the link between these two parameters. In a stochastic two-period setting, utility in t=1 can be expressed as:

$$U_1 = \left[(1-\gamma)c_1^{\rho} + \gamma(E_1[c_2^{\alpha}]^{\rho/\alpha})\right]^{1/\rho}$$

where RRA is captured by α, the EIS is captured by 1/(1-ρ) and the rate of time preference is captured by 1/($\gamma$-1). Assuming no uncertainty, this expression boils down to:

$$U_t = \left[(1-\gamma)c_1^{\rho} + \gamma c_2^{\rho}\right]^{1/\rho}$$

Interestingly, some parallelism can be established between this expression and CRRA utility. Raising both sides to 1/ρ and dividing them by (1- $\gamma$), this utility function becomes:

$$\frac{U_t^{\rho}}{1-\gamma} = c_1^{\rho} + \frac{\gamma}{(1-\gamma)} c_2^{\rho}$$

and, since ρ = 1- θ and β = $\gamma/(\gamma$-1), making use of these equalities and dividing both sides by (1-θ) the above equation can be expressed as:

$$\frac{U_t^{\rho}}{(1-\gamma)(1-\theta)} = \frac{c_1{}^{1-\theta}}{1-\theta} + \beta \frac{c_2{}^{1-\theta}}{1-\theta}$$

being the right hand side of this function equivalent to two-period CRRA utility. This framework, allows us to analyse the EIS without incurring in imprecisions due the inverse relationship between this parameter and relative risk aversion established by CRRA utility.

Having now defined both the IRES and the EIS, we first review the empirical findings on these two parameters and then we investigate what could make them to vary over time.

*2.2 Empirical findings on the IRES under the life-cycle hypothesis*

Until the mid-twentieth century, the belief that aggregate saving was unresponsive to variations in the rate of return was widely accepted among researchers. This notion, was fuelled by an extraordinary stability of the US saving-income ratio during the postwar period, which was characterized by large interest rate fluctuations. In opposition to this current of thought, along the subsequent decades, several authors would present empirical evidence illustrating the importance of the rate of return as a determinant of individuals' saving behaviour.

One of the first attempts to assess the IRES is made by Wright (1967), who estimates the interest elasticity of consumption by introducing an interest rate variable into a consumption function, similar to that presented by Friedman (1957) and Ando and Modigliani (1963). Then, this author employs the parameters found to compute the IRES of life-cycle consumers, obtaining estimates on the range 0.19-0.24. Through a quite more extensive analysis that involves different estimation techniques and alternative measures of the real after-tax return to wealth, Boskin (1976) finds evidence of a positive relationship between the rate of return and saving, with interest elasticity estimates clustering around 0.4. In addition, this author notes that the low estimates obtained by earlier authors were quite certainly due to the downward bias caused by the use of nominal before-tax interest rates instead of real after-tax ones.

Unlike the authors presented above, Summers (1982) recognizes the importance of identifying differences across individuals when investigating the IRES, as these can lead to differing saving propensities. Assuming an initial rate of return equal to 0.02, a rate of time preference equal to 0.02 and an EIS equal to one third, the author finds an IRES of 1.3, considerably larger than that found by earlier empirical investigation. In addition, Summers reports some analytical calculations, from which three main observations can be drawn: firstly, as suggested earlier in this section, lower values of the EIS lead to a lower IRES, because the positive influence of the substitution effect is diminished; secondly, the IRES is somewhat decreasing on individuals' age; and thirdly, there is a negative relationship between the interest sensitivity of aggregate saving and the interest rate level, effect that is enlarged when high initial interest rates are combined with higher values of subjective time preference.

Contrasting with Summers' (1982) high estimates, Elmendorf (1996) finds an IRES ranging between -0.33 and 0.40, being the average estimate equal to 0.12[5]. In line with the results obtained by Summers (1982), differences across these suggest a positive relationship between the IRES and both the EIS and the rate of time preference. In addition, the IRES shows to be decreasing along individuals' work lives, being this trend moderately reverted after retirement. According to Elmendorf (1996), this happens because the revaluation of younger individuals' human wealth takes place before consumption; as a result, when the rate of return appreciates, their budget constraint is relaxed, encouraging them to increase consumption at the expense of saving. However, as individuals age, the financial wealth component of savers' total wealth goes up. If this wealth is held in form of physical capital, payments will most likely take place after consumption. Thus, when the rate of return goes up, individuals' wealth will decrease as physical capital depreciates, creating an incentive for them to save more.

The empirical evidence presented in this section suggests a generally positive IRES. Nevertheless, the results obtained by Elmendorf (1996) show that in the presence of extremely low values of the EIS, expansionary monetary policy in the form of lower interest rates could as a matter of fact lead to higher saving rates. In the coming sections

[5] These estimates are obtained with the same parameters employed by Summers (1982).

we present a compilation of the EIS estimates obtained by the empirical literature and investigate the feasibility and potential causes of such low EIS values.

*Table 2 - Interest elasticity of saving estimates (IRES) - Equation 5*

| *Author(s)* | $\delta(s_1/y_1) / \delta R$ |
|---|---|
| Wright (1967) | 0.2 |
| Boskin (1976) | 0.4 |
| Summers (1982) | 1.3 |
| Elmendorf (1996) | 0.1 |
| *Other authors* | |
| Blinder (1975) | 0 |
| Gylfason (1981) | 0.3 |
| Friend and Hasbrouck (1983) | 0 |
| Makin (1986) | 0.4 |
| Bovenberg and Evans (1990) | 0.5 |

Notes: (1) Wright (1967), Summers (1982) and Elmendorf (1996) present a range of estimates obtained using different parameter assumptions and model specifications. The estimates presented in this table reflect the average values obtained by these three authors.

## *2.3 Empirical findings on the EIS*

### *2.3.1 Challenges and estimates*

As we have seen, the interest elasticity of saving is to a large extent determined by the elasticity of intertemporal substitution. Numerous authors have attempted to estimate the actual value of this latter parameter. However, the results obtained are highly sensitive to the methodology applied. As a result, there is wide range of EIS estimates available (see Table 3).

The differences across estimates arise from three main sources:

(1) Probably the largest of these three is the choice of the measure used to approximate the return to households' total wealth. Due to the unobservable nature of this variable, it is common practice to employ a stock market index, assuming it to be representative of individuals' wealth portfolio. This approach is followed by Epstein and Zin (1991), who find EIS estimates between 0.2 and 0.9.

Other authors employ the returns of different financial assets present in households' asset portfolio, as it is the case of Summers (1982). This author makes use of the returns to treasury bills, long term government bonds, saving deposits and corporate stocks. His results suggest important differences across returns, as the median estimate obtained using the return to long term bonds (0.96) is six times larger than that obtained using stock returns (0.14). Interestingly, the results obtained using the return to saving deposits, with a median of 0.49, are the closest to the overall median estimate (0.50).

In opposition to the extended practice of using financial asset returns, Lustig et al. (2013) maintain that these are unlikely to illustrate the true wealth portfolio of the representative consumer, as the share of human wealth is found above 90 percent of individuals' total wealth. To solve this issue, Thimme and Völkert (2015) approximate the rate of return by means of a variable that "captures deviations from the common trend in consumption, asset holdings, and labour income". Deviations from this trend are assumed "to produce movements in the consumption-aggregate wealth ratio and thus predict future asset returns". Employing this measure, their estimates suggest an EIS well above unity.

(2) Another source of differences across estimates is the choice of the measure used to approximate consumer expenditure. Weber (1970) uses US aggregate data on the consumption of nondurables to study the relationship between consumption growth and the real rate of return. His findings suggest an EIS below one-half, with estimates ranging between 0.13 and 0.41. In contrast, in a later paper published in 1975, using US data on both durable and non-durable goods, Weber finds considerably higher estimates, comprised between 0.56-0.75. These results suggest that, when the intertemporal substitution between durables and non-durables is allowed, the elasticity of intertemporal substitution increases.

*Table 3 - Elasticity of intertemporal substitution estimates (EIS) – Equation 1*

| *Author(s)* | $\sigma$ |
|---|---|
| Weber (1970) | 0.1 – 0.4 |
| Weber (1975) | 0.6 – 0.8 |
| Summers (1982) | 0.1 – 2.7 |
| Hall (1988) | -0.4 – 0.4 |
| Epstein and Zin (1991) | 0.2 – 0.9 |
| Thimme and Völkert (2015) | 1.5 – 2.0 |

(3) As important as the measures that are used to approximate the rate of return and consumption, is the selection of the instrumental variables. When estimating the elasticity of intertemporal substitution, an instrumental variable approach is a widespread choice among researchers, as it avoids potential endogeneity issues.

The estimates by Summers (1982) that we have analysed earlier in this subsection, were obtained using two-period lagged values of consumption growth as an instrument.

However, contrary to the use of this specific instrument, Hall (1988) maintains that lagged consumption growth is unlikely to be exogenous due to data time-aggregation. According to this author, the endogeneity of consumption growth would be the main source of the high estimates obtained by earlier researchers. With different instruments, Hall (1988) finds a significantly lower EIS ranging between -0.4 and 0.35. In his own words, “the elasticity is unlikely to be much above 0.1, and may well be zero”.

To conclude this section, in spite of the considerable differences across estimates that result from their high sensitivity to methodology differences, the average EIS appears to be between 0.5 and 1. This, suggests that the income effect might outweigh the substitution effect. In turn, this implies that the sign of the interest elasticity of saving will be determined by the difference between the positive net effect resulting from the aggregation of the substitution and income effects, and the negative wealth effect.

### *2.3.2 Time variation of the EIS and thus of the IRES*

When estimating the interest rate elasticity of saving, the elasticity of intertemporal substitution is usually assumed to be constant. This, as long as other parameters are also kept fixed, implies a constant IRES. However, contrary to this widespread assumption, the results obtained by various authors suggest that the EIS may not only be subject time-variation, but also follow a pro-cyclical path. That is, increasing when the level of consumption expands and decreasing when it contracts.

The study that most explicitly shows this pattern is that presented by Attanasio and Browning (1995). These two authors, by means of a flexible parameterization of the consumer preferences, find that the absolute value of the EIS is an increasing function of the level of consumption.

In this same line, Atkeson and Ogaki’s (1996) findings suggest that wealthier households find it easier to substitute consumption across periods. Using Indian panel data, they find that the wealthiest five percent of households have an average EIS of 0.8, whereas that of the poorest five percent would be up to 0.3 points lower. In the same line, using US and Indian time-series data, Atkeson and Ogaki (1996) find the average EIS of US households to be 0.13 points larger than that of their Indian counterparts. Similarly, Lawrance (1991) finds that the EIS of above-median labour income households (0.8) would be 1.6 times as large as that of below-median labour income households (0.5).

Finally, Zeldes (1989) divides the households of his sample into two different subgroups based on their financial wealth to income ratio – those likely to be constrained and those not –. His results show shows liquidity unconstrained households (1.4) to have a significantly higher EIS than those subject to liquidity constraints (-1.5).

In light of the evidence pointing towards the existence of a consumption-varying elasticity of intertemporal substitution, what could justify it? According to Atkeson and Ogaki (1996), there exist at least two intuitive reasons why the EIS might be smaller for the poor than it is for the rich. "First, if there are positive subsistence consumption requirements, then poor consumers have a smaller portion of their budget left over after satisfying subsistence requirements to save or consume at their discretion. Second, the consumption in excess of subsistence of necessary goods (such as food) may be less substitutable across time than is the consumption of luxury goods. Since the poor spend a higher fraction of their total expenditure on subsistence and necessary goods than do the rich, their EIS of total consumption expenditure may be smaller than the EIS of the rich. Thus the EIS may rise with the level of wealth."

Based on the findings presented in this subsection, a constant elasticity of intertemporal substitution appears to be rather implausible. In addition, the positive correlation observed between this parameter and consumption levels, suggests that the EIS may follow a pro-cyclical pattern, which would, in turn, enable a lower IRES during consumption downturns. As a result, was the drop in household consumption strong enough, the IRES could become negative, which means that decreasing real rates of return would lead to higher household saving rates.

### *2.4 Alternative saving motives*

This section introduces a series of relevant deviations from the life-cycle hypothesis observed by the empirical literature and examines how these could affect the aggregate IRES and thus the validity of our hypothesis.

#### *2.4.1 Introducing the alternative saving motives*

The life-cycle hypothesis presented by Modigliani and Brumberg (1954) relies on a series of simplifying assumptions that keep it highly tractable:

(1) One of the main implications derived from this framework is that expected increases in transitory income should have no effect on current consumption, which leaves no room for rule-of-thumb consumers.

(2) In addition, individuals' wealth is expected to be equal to zero by the end of the life-cycle or, what is the same, they are assumed to have no bequest saving motives.

(3) Furthermore, concerning capital markets, the life-cycle hypothesis assumes that individuals are free to borrow and lend at a fixed rate of interest. This, implies that they are capable of financing current consumption against their future labour income.

(4) Finally, the life-cycle hypothesis assumes economic agents to have certainty about the size and distribution of their lifetime income, the present and future rate of interest, their future consumption needs and the extent of their lifetime.

*2.4.1 Evidence on the alternative saving motives*

In spite of their convenience, the assumptions presented in the above subsection have been widely criticized for their strength and lack of empirical foundation, as presented in the following paragraphs.

(1) Campbell and Mankiw (1989) find that between 23 and 50 percent of changes in consumption follow from changes in current earnings. Along the same line, Wilcox (1989) shows that "fully anticipated increases in social security benefits cause large increases in consumption expenditure at the time when the increases are paid". This evidence, suggests that whereas a share of consumers are indeed forward looking as predicted by the life-cycle hypothesis, the rest may have short-term planning horizons.

(2) On the other end of the spectrum of consumer behaviour we find evidence indicating that some people are even more far-sighted than the life-cycle hypothesis suggests. Bernheim (1991) shows that "many people would stop short of converting all their assets into annuities, even in the presence of perfect insurance markets". Kotlikoff and Summers (1981) maintain that intergenerational bequests may be one of the major determinants of capital formation in the US.

(3) Concerning capital markets, Zeldes (1989) finds the Euler equation of liquidity constrained households to be violated and that of unconstrained households to hold, which supports the view that capital market imperfections should be taken into account when modelling individuals' consumption and saving behaviour. According to Hall and Mishkin (1982) this description fits 20 percent of the population.

(4) Finally, some authors emphasize the importance of uncertainty for aggregate capital formation. Skinner (1988) and Carroll et al. (1992) find labour income uncertainty to increase the slope of individuals' consumption path, observing higher saving along individuals' early years and lower saving at later stages of their lifetime.

### 2.4.2 *Effect of these alternative saving motives on the IRES*

In view of the evidence presented in the above subsection, what are the implications derived from the presence of these deviations for the interest elasticity of saving?

(1+3) Elmendorf (1996) examines the elasticity of saving of current-income consumers – group comprising both individuals with short-term planning horizons and individuals that, in spite of having long term planning horizons, find themselves unable to smooth consumption over time due to the presence of liquidity constraints – by analysing how interest rate changes affect their cash flows.

Within this group of economic agents, the author distinguishes between individuals with and without financial wealth. In the simplest scenario, current-income consumers hold no financial wealth and set consumption equal to income, which makes their cash-flows unaffected by interest rate changes. However, if current-income consumers do hold a small amount of financial wealth, changes in the rate of interest are likely to affect their consumption by changing both their income and wealth.

Elmendorf (1996) holds that even though households are observed to receive slightly less interest than they pay out, their receipts respond much more quickly than their payments to interest rate changes. Thus, the combined effect is that an increase in the interest rate generally raises households' interest receipts more than their interest payments. As the cash flow increases, consumption rises and the saving-to-income ratio falls. This, establishes that current-income consumers have a negative elasticity of saving. However, it is important to keep in mind that, since the share of aggregate wealth held by current-income consumers is rather small, their saving decisions are unlikely to have a large impact on the aggregate elasticity of saving.[6]

(2) Summers (1982) introduces intergenerational transfers by making use of interdependent utility functions, in which the utility of parents depends on either the utility or the consumption of their descendants. The author, suggests that adopting this formulation has radical implications for the long-run elasticity of savings with respect to the rate of return. Bequest leavers are found to save whenever the after-tax interest rate exceeds their rate of time preference and dissave whenever the former falls below the latter. This, would imply a long term IRES tending to infinity and thus risk neutrality.

[6] Supporting this idea, Summers (1982) extends the life-cycle hypothesis by allowing for the presence of liquidity constrained consumers, however, his results are not significantly different from those reported by the standard model.

What about the short term? Chamley (1981) holds that convergence to a new steady state is rather fast. His estimates suggest that a permanent one percent increase in the rate of return would increase savings by up to 40 percent. As bequest leavers are shown to account for a large share of aggregate saving, it seems clear that their presence should increase the IRES quite substantially.

(4) Finally, Engen (1992) finds the IRES to be considerably lower in the presence of uncertainty. Even though the extent of this reduction is affected by the assumptions concerning the level of uncertainty and the consumers' rate of time preference, the introduction of uncertainty is found to lower the IRES by almost 70 percent on average.

As we have seen in this section, the effects of the different deviations presented show to take both positive and negative signs. Whereas the presence of bequest leavers seems to increase the aggregate IRES, the presence of current-income consumers and uncertainty show to lower this parameter. Due to the difficulty to determine their net effect on the aggregate IRES, for simplicity, we assume the IRES of life-time consumers to be representative of the aggregate parameter.

## 3. Data

In order to establish the relationship between the variables of interest, this study employs time-series data on the Euro Area and four of its member states: Germany, Finland, Italy and Spain. The sample includes 55 quarterly observations comprised between 2003Q1 and 2016Q3, containing both data on households and non-profit institutions serving households (NPISH).

Due to the absence of a centralized data base reporting net macroeconomic data with quarterly periodicity for this specific institutional sector of the Euro Area and its member states, we resort to a variety of different data sources.

For Euro Area, data on net saving, net disposable income, final consumption expenditure and the return has been retrieved from the Statistical Data Warehouse of the European Central Bank (ECB), whereas that on financial wealth and the harmonized consumer price index (HCPI) has been obtained from the Statistical Office of the European Communities (Eurostat).

For individual member states, data on net saving and net disposable income has been extracted from their respective statistical national offices. Data covering financial wealth and the HCPI has been obtained from the Eurostat, except for Italy's financial wealth, which is retrieved from the Statistical Office of the Organization for Economic

Cooperation and Development (OECD). Finally, data on the member states' rate of return, like for the Euro Area aggregate, has been extracted from the ECB's data base.

Each of the measures employed has been converted to real per capita terms using the relevant HCPI and the population size. In addition, some of them have been modified to fit the theoretical model presented in section 2.

The definitions of the variables used in the empirical section of this paper are as follows: the saving rate ($s_t/y_t$), used to estimate the IRES, is the ratio of net saving to net disposable income, whereas consumption growth ($ln(c_t - c_{t-1})$), used to estimate the EIS, is defined as the rate of growth of final consumption expenditure. With regards to the main independent variable, the real return to individuals' wealth ($R_t$) has been approximated using the real return to deposits held by households and NPISH with an agreed maturity up to one year. Finally, regarding the control variables, relative future income ($Y_t$) is the ratio of net disposable income in t+1 to that in t; relative financial wealth ($A_t$) is defined as the ratio of net financial wealth to net disposable income; and ex-post inflation ($i_t$) is the rate of growth of the quarterly HCPI. Table 4 presents the descriptive statistics for each of the variables used in the empirical model:

*Table 4: Descriptive statistics*

| | *Mean* | *SD* | *Max.* | *Min.* | *ADF test.* | *MADF test* |
|---|---|---|---|---|---|---|
| *Euro Area* | | | | | | |
| $s_t/y_t$ | 0.068 | 0.031 | 0.131 | 0.025 | 0.011* | - |
| $ln(c_t - c_{t-1})$ | 0.000 | 0.017 | 0.030 | -0.041 | 0.011 | - |
| $R_t$ | 0.018 | 0.010 | 0.045 | -0.006 | 0.123 | 0.081* |
| $Y_t$ | 1.002 | 0.063 | 1.086 | 0.913 | 0.177 | - |
| $A_t$ | 8.769 | 0.707 | 10.563 | 7.509 | 0.215* | - |
| $i_t$ | 0.004 | 0.006 | 0.016 | -0.01 | 0.092 | - |
| *Finland* | | | | | | |
| $s_t/y_t$ | 0.008 | 0.037 | 0.086 | -0.072 | 0.426* | - |
| $ln(c_t - c_{t-1})$ | 0.005 | 0.052 | 0.079 | -0.095 | 0.012 | - |
| $R_t$ | 0.016 | 0.011 | 0.044 | -0.001 | 0.475 | 0.060* |
| $Y_t$ | 1.009 | 0.103 | 1.200 | 0.876 | 0.000 | - |
| $A_t$ | 4.665 | 0.618 | 5.818 | 3.304 | 0.006 | - |
| $i_t$ | 0.004 | 0.005 | 0.018 | -0.005 | 0.111 | - |
| *Germany* | | | | | | |
| $s_t/y_t$ | 0.101 | 0.021 | 0.144 | 0.076 | 0.165 | - |
| $ln(c_t - c_{t-1})$ | 0.002 | 0.036 | 0.047 | -0.069 | 0.003 | - |
| $R_t$ | 0.013 | 0.011 | 0.046 | -0.003 | 0.416 | 0.012* |
| $Y_t$ | 1.001 | 0.011 | 1.031 | 0.977 | 0.001 | - |
| $A_t$ | 7.442 | 0.734 | 8.774 | 5.766 | 0.045* | - |
| $i_t$ | 0.004 | 0.004 | 0.009 | -0.007 | 0.000 | - |

| | *Mean* | *SD* | *Max.* | *Min.* | *ADF test.* | *MADF test* |
|---|---|---|---|---|---|---|
| *Italy* | | | | | | |
| $s_t / y_t$ | 0.056 | 0.058 | 0.169 | -0.062 | 0.355* | - |
| $ln(c_t - c_{t-1})$ | -0.001 | 0.021 | 0.035 | -0.059 | 0.027 | - |
| $R_t$ | 0.015 | 0.018 | 0.049 | -0.015 | 0.027 | 0.057 |
| $Y_t$ | 1.002 | 0.093 | 1.183 | 0.864 | 0.056 | - |
| $A_t$ | 11.644 | 0.895 | 13.562 | 9.594 | 0.079 | - |
| $i_t$ | 0.004 | 0.017 | 0.037 | -0.027 | 0.314 | - |
| *Spain* | | | | | | |
| $s_t / y_t$ | 0.030 | 0.07 | 0.148 | -0.105 | 0.173 | - |
| $ln(c_t - c_{t-1})$ | 0.000 | 0.029 | 0.053 | -0.058 | 0.127 | - |
| $R_t$ | 0.017 | 0.016 | 0.048 | -0.02 | 0.561 | 0.010* |
| $Y_t$ | 1.010 | 0.152 | 1.252 | 0.825 | 0.185 | - |
| $A_t$ | 6.014 | 1.100 | 8.470 | 3.709 | 0.190 | - |
| $i_t$ | 0.005 | 0.013 | 0.024 | -0.023 | 0.116 | - |

Notes: (1) SD stands for standard deviation; (2) ADF denotes the Augmented Dickey-Fuller test; (3) MADF stands for modified Augmented Dickey-Fuller test; (4) * denotes that a trend has been included in the test equation.

As we can observe in this table, the Augmented Dickey-Fuller test generally rejects or is considerably close to rejecting at 0.1 significance level the null hypothesis that the time-series data used to conduct our empirical analysis has a unit root. There are, however, a few important and worrisome exceptions, as this same test fails to reject the presence of a unit root in the level of the real rate of return for Finland, Germany and Spain, with probability values ranging between 0.41 and 0.56; and also in the level of the saving rate of Finland and Italy, with probability values of 0.38 and 0.36 respectively.

The presence or absence of a unit root in the real rate of return has been widely discussed by econometric literature. Especially interesting for the matter that concerns us are the findings by Garcia and Perron (1996) and Clemente et al. (1998), who maintain that conventional unit root tests like the Augmented Dickey-Fuller test may be biased towards non-rejection of the null hypothesis when the real rate of return shows to experience a structural break.

Observing the development of the real rate of return in the Euro Area as a whole, Finland, Germany, Italy and Spain (Appendix B, Figure 8) we can see what appears to be a structural break taking place between late 2008 and early 2009. For this reason, it may be more appropriate to use a modified unit root test that does allow for levels and trends to differ across subsamples.

The results obtained by means of the modified test depict an entirely different picture. These, suggest that real rate of return follows a stationary process in the Euro

Area and each of the member states.[7] As a result, we use as a working assumption that first difference transformation of the data may not be needed.

With regards to the saving rate of Finland and Italy, even though the Augmented Dickey-Fuller test suggests the contrary, we believe that if the sample comprised a larger time span, the mean and variance of this variable would actually show to remain largely constant over time, as suggested by the life-cycle hypothesis.

## 4. Methodology

This section presents the methodology used to examine whether 2009 large increase in the saving rate of Euro Area households, could be partly explained by a sudden temporary drop in individuals' IRES, caused in turn by a temporary decrease in their EIS. More concretely, we first expose the reasons behind the choice of the Quandt-Andrews test as the preferred method to examine the presence of structural breaks in the IRES, and explain how this test, despite being a single breakpoint test, can be used to find multiple structural breaks. Secondly, we describe the model that is used to determine the sign of the potential temporary change in the IRES. And finally, we describe the model used to analyse whether the change in the IRES, could have been caused by a change in the EIS.

### *4.1 Analysing the presence of structural breaks in the IRES*

Visual analysis of the evolution of the saving rate and the real rate of return throughout the sample period, provides a considerably clear idea of the location of the potential break quarters – first break in late 2008 and second break in early 2010 –. As a result, we could be tempted to employ the test for known structural breaks presented by Chow (1960). However, as suggested by Hansen (2001), arbitrarily picking the break quarters may lead us to miss the true break dates. For this reason, we find it more appropriate to use a test for unknown structural breaks.

The two foremost methods for unknown breakpoint testing are the Quandt-Andrews and the Bai-Perron. Based on the fact that we want to test for two potential structural breaks, the method that would suit us best is the Bai-Perron test, as it allows for multiple unknown breakpoints. However, to the best of our knowledge, this test is not

[7] Whereas for the Euro Area, Finland, Germany and Italy only the intercept of the real rate of return has been allowed to vary across subsamples, for Spain both the intercept and the trend have been allowed to differ before and after the break.

compatible with instrumental variables estimation. This, leaves us with the Quandt-Andrews test as the only remaining alternative.

The Quandt-Andrews test performs a Chow test at every date comprised between $t_{min}$ and $t_{max}$[8], presenting two different F-statistics for each of the Chow tests conducted: a Wald F-statistic and a Likelihood F-statistic. As we have said earlier in this section, the Quandt-Andrews is a single breakpoint test – it tests the break date with the largest F-statistic against the null hypothesis of no structural breaks –. Nevertheless, since larger F-values imply a higher likelihood that the null hypothesis of no structural breaks is rejected, the collection of F-statistics obtained from this test can be used to identify more than one break. Once the potential break dates are identified with more precision, we can conduct a Chow test[9] on each of them to determine their level of significance.

### *4.2 Determining the sign of the change in the IRES*

Once the break dates have been found, in order to determine the sign of the change in the IRES during the inter-break period, we regress the saving rate on the real return and a series of complementary arguments. These complementary arguments are intended to control for third factors that could affect the relationship between the main variables of interest. The estimated saving rate regression, constructed following Equation 6, is given as follows:

$$\frac{s_t}{y_t} = \alpha_0 + \beta R_t - \alpha_1 Y_t - \alpha_2 A_t - \alpha_3 i_t + \mu Q_t + \gamma D_t + u_t \tag{8}$$

$s_t/y_t$ is the rate of saving, $R_t$ is the real return, $Y_t$ is relative future income, $A_t$ is relative financial wealth and $i_t$ stands for inflation. In addition, $Q_t$ is a (3x1) vector of quarterly dummies and $D_t$ is a (4x1) vector including first, second, third and fourth lags of the dependent variable.

Concerning the main parameter of interest, the coefficient measuring the interest elasticity of saving in Equation 8 is:

$$\beta = (\beta_1 + \beta_2 \delta_t)$$

[8] $t_{max}$ and $t_{min}$ are the lower and upper bounds of the analysed interval and are obtained using 15 percent trimming, as recommended by Andrews (1993).

[9] The Chow test compares the RSS of the pooled (restricted) regression with that of the unrestricted alternative, that is, the RSS resulting from regressing separately on each of the subsamples. The larger the difference between the sums, the larger the probability to find a structural break.

where $\delta_t$ is a period dummy that takes value 1 if t is comprised within the break dates suggested by the Quandt-Andrews test, and 0 otherwise. The interaction term between this country-specific period dummy and the rate of return ($R_t$), allows the IRES to fluctuate across subperiods. Thus, if $\delta(t)=0$, the IRES ($\beta$) is just captured by $\beta_1$. Per contrast, if $\delta(t)=1$, the IRES is captured by both $\beta_1$ and $\beta_2$. This latter parameter reflects the variation in the IRES during the inter-break period.

Based on Equation 6 and the literature analysed, we would expect the IRES before and after the inter-break period ($\beta_1$) to be positive, whereas the change of the IRES during the inter-break period ($\beta_2$) is expected to be negative. In addition, $\beta_1$ is expected to be smaller in absolute terms than $\beta_2$, which would enable a negative IRES between the first and the second break. With regards to the control variables, $\alpha_1$ and $\alpha_2$, which capture the partial effect of relative future income and relative financial wealth on the saving rate respectively, are both expected negative. Similarly, the effect of inflation on the saving rate ($\alpha_3$) is also expected to be negative.

Finally, since we are studying elasticities, the optimal specification would involve logarithmic transformations. Nevertheless, the output of such specification showed a very low F-statistic and a largely negative $R^2$. As a result, we decided to use a non-logarithmic specification. To the best of our knowledge, using a non-logarithmic specification should not affect the validity of our results, as we are interested in the sign of the change in the IRES, rather than in the absolute magnitude of this parameter.

### *4.3 Analysing the cause of the change in the IRES*

To analyse whether the hypothetical drop in the IRES could have been caused by a sudden temporary decrease in individuals' EIS, following Equation 7, we regress the rate of consumption growth ($ln(c_t$ - $c_{t-1}$)) on the real rate of return ($R_t$) and a (3x1) vector of quarterly dummies:

$$ln(c_t - c_{t-1}) = \gamma_0 + \sigma R_t + \mu Q_t + u_t \tag{9}$$

The elasticity of intertemporal substitution is captured by σ, which, like the interest elasticity of saving (β), can be decomposed as follows:

$$\sigma = (\sigma_1 + \sigma_2 \delta_t)$$

As we can see, this parameter is allowed to differ across sub-periods with the introduction of an interaction term between a period dummy (δ) and the real rate of return ($R_t$). Thus, if $\delta(t)=0$, the EIS ($\sigma$) is just captured by $\sigma_1$. Per contrast, if $\delta(t)=1$, the EIS is captured by both $\sigma_1$ and $\sigma_2$. This latter parameter reflects the variation in the EIS during the inter-break period.

In addition to the base specification (9), in order to increase its explanatory power, we also regress an alternative model that includes the control variables used in the saving rate function:

$$ln(c_t - c_{t-1}) = \gamma_0 + \sigma R_t + \gamma_1 Y_t - \gamma_2 A_t - \gamma_3 i_t + \mu Q_t + u_t \tag{10}$$

where, as in Equation 8, $Y_t$ is relative future income, $A_t$ is relative financial wealth and $i_t$ stands for inflation.

With regards to the sign of the parameters, the effect of relative future income on the saving rate ($\gamma_1$) is expected positive. In contrast, both the effect of relative financial wealth ($\gamma_2$) and inflation ($\gamma_3$) are expected negative. Finally, no changes are expected on $\sigma$ with respect to the results obtained using the base specification (9).

### *4.4 Estimation method*

Equations (8), (9) and (10) are estimated using an instrumental variable approach, i.e., Two-Stage Least Squares (2SLS), which aims to correct for the potential endogeneity of the rate of return. The most likely causes of this endogeneity are omitted variable bias and simultaneous causality. Following Summers (1982) and Hall (1988), we use as instruments two-quarter lagged values of the real rate of return and the interaction term between the rate of return and the period dummy.

With the aim of correcting for heteroskedasticity and autocorrelation, we make use of heteroskedasticity and autocorrelation consistent standard errors. In addition to this, since the autocorrelation in the saving rate regression appears to be quite severe, we include the first, second, third and fourth lags of the dependent variable on the right hand side of the regression. Finally, since the data employed is seasonally unadjusted, we add first, second and third quarter dummies to the regression to correct for seasonality.

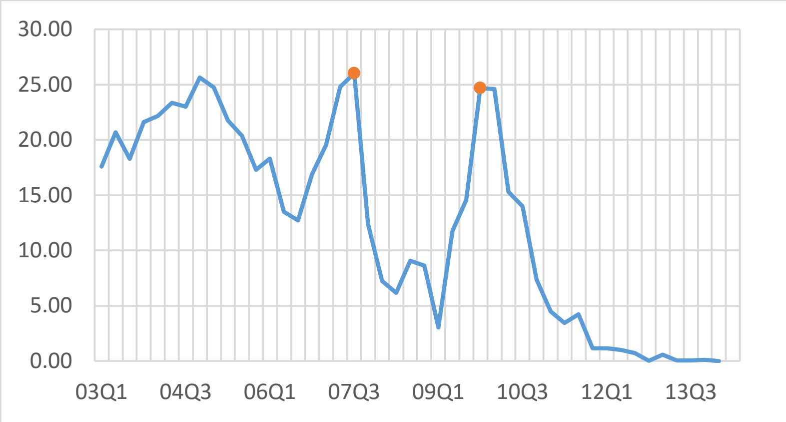
## 5. Results

This section is divided in three different subsections. The first subsection presents the results obtained on the presence of structural breaks in the IRES. Subsequently, the second subsection discusses the estimates found on the IRES. And finally, the third subsection discusses the estimates found on the EIS.

*5.1 Breaks on the IRES*

Figure 5 presents the time plot of the Wald F-statistic reported by the Quandt-Andrews test, conducted on the $\beta_1$ of Equation 8 before the introduction of the interaction term. If we examine this figure, we can see that at the Euro Area level, as hypothesized at the beginning of this paper, there appear to be two structural breaks in the relationship between the saving rate and the real return to wealth in the vicinity of 2009: the first break would have taken place in 2007Q3, three quarters before the Euro Area entered into recession, whereas the second break would have done so in 2009Q4, two quarters after the first recession was overcome. Both break dates are shown to be statistically significant at the 1 percent level by the Chow test.

*Figure 5 – Euro Area IRES Wald F-statistic*

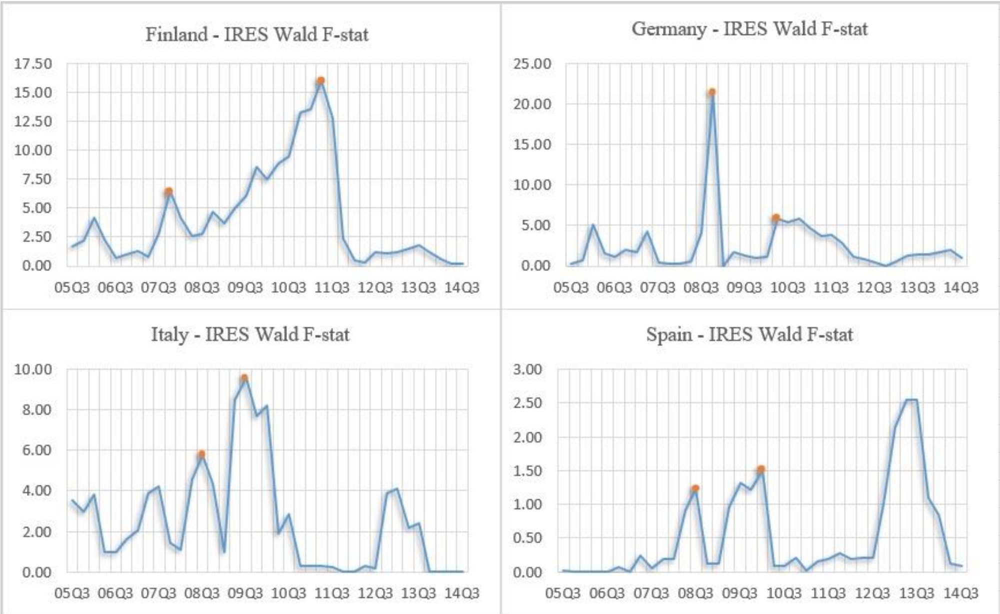



Source: Author's calculations

At the member state level, there appears to be substantial heterogeneity concerning the break dates and the duration of the inter-break period (Figure 6). The evolution of the Wald F-statistic presents considerable similarities for Germany, Italy and Spain. These three countries experienced a slight increase of the instability in the relationship between the variables of interest in mid-2007, when the first break at the Euro Area level happened to take place, however, their largest Wald F-statistics are found in late 2008 and late 2009/ early 2010. The exact peak quarters are the following: 2008Q4 and 2010Q2 for Germany, 2008Q3 and 2009Q3 for Italy and 2008Q3 and 2010Q1 for Spain. During the period comprised between the first and the second peak the Wald F-statistic decreased significantly, which makes the potential break dates fairly easy to detect.

With regards to Finland, identifying the first break is not so straight forward, as the Wald F-statistic follows an almost uninterrupted increasing trend between 2007Q2 and 2011Q2. With a certain component of arbitrariness, we assume the first break to take place in 2007Q4 – since it is the only quarter within this period that was preceded (followed) by two consecutive quarters of increasing (decreasing) Wald F-statistic – and the second break in 2011Q2. The Chow test shows the country level breaks mentioned above to be significant at 5 percent level, except for Germany's second break, which is not found significant. Finally we must say that the presence of structural breaks in Germany and Italy, unlike in the case of Spain and Finland, is quite surprising since their respective household saving rates did not experience important changes during the inter-break period.

*Figure 6 – Country IRES Wald F-statistics*




Source: Author's calculations

## *5.2 Estimates of the IRES*

Table 5 contains the estimates for the slope of the saving curve obtained using the model presented in Equation 8. This table includes results both before and after the introduction of the interaction term between the real return and the country-specific period dummy. As a little reminder, the period dummy has been created using the breaks suggested by the Quandt-Andrews test, regardless of their significance.

As we hypothesized ($\beta_1 > 0$), before the introduction of the interaction term, the regression results show the real rate of return to affect positively households' saving rate at the Euro Area level, with an estimate of 0.52. This, suggests that a one percentage point increase in the former variable would lead to an increase in the latter of half a percentage point. The relationship between the real rate of return and households' saving rate appears to differ considerably across countries, with generally positive estimates ranging between -0.32 and 1.56. Quite interestingly, the sensitivity of households' saving rate to changes in real rate of interest shows to be larger in Spain and Finland, countries where household financial wealth per capita is lower.

Regarding the time variation of the interest elasticity of saving, in opposition to what we hypothesized ($\beta_2 < 0$, $\beta_1 < |\beta_2|$ and $\beta < 0$), at the Euro Area level the introduction of the interaction term does not result in a negative IRES during the inter-break period (2007Q3-2009Q4). As we hypothesized ($\beta_1 > 0$), the base estimate for the IRES is positive (0.38). However, contrary to what we expected ($\beta_2 < 0$), the coefficient for the interaction term (0.15) suggests an increase of the IRES during the inter-break period. These estimates, point towards a positive IRES during the inter-break period (0.53) in the Euro Area.

At the country level, the results are considerably heterogeneous. The estimates obtained for Finland and Spain depict a similar situation to the one just observed for the Euro Area, with both a positive base term and a positive interaction term. In the case of Germany, even though the interaction term suggests a decrease of the IRES during the inter-break period (-0.02), due to the larger magnitude of the positive base term (0.09), the IRES would have stayed positive during the period between breaks. Unlike the other countries analysed, Italy presents both a negative base term (-0.31) and a negative interaction term (-0.09), which suggests a considerably negative IRES during the inter-break period (-0.40).

To conclude, the positive sign of the IRES during the inter-break period in the Euro Area as a whole, Finland and Spain would rule out a temporarily negative IRES as the cause of the large saving rate increase that they experienced in the vicinity of 2009. In addition, we should express our surprise for the negative IRES during the inter-break period in Italy because, despite the decreasing real returns, this country did not experience an increase of the saving rate in 2009.

Finally, it is important to comment that these results should be taken with caution for four main reasons. Firstly, the interaction term is not significantly different from zero in any of the regressions estimated. Secondly, the interaction term presents considerably large standard errors, almost as large as the parameter coefficients, which could lead to important imprecisions. Thirdly, the Wald test used to test the joint restriction $\beta_1 + \beta_2\delta_t = 0$, only fails to reject the null hypothesis for the Euro Area as a whole, while rejecting it for each of the individual member states analysed. Fourth, despite the inclusion of a 4x1 vector including first, second, third and fourth lags of the dependent variable, the of model used to obtain the results for Italy suffers from severe autocorrelation.

*Table 5: Slope of the saving rate curve (IRES) - Dependent variable: saving rate - Estimation method: 2SLS*

| *Regressor/Area* | *Euro Area* | | *Finland* | | *Germany* | | *Italy* | | *Spain* | |
|---|---|---|---|---|---|---|---|---|---|---|
| $c$ | 0.106 | 0.111 | 0.481*** | 0.446*** | -0.001 | -0.001 | 0.324*** | 0.325*** | -0.093 | 0.076 |
| | (0.09) | (0.08) | (0.11) | (0.09) | (0.05) | (0.05) | (0.10) | (0.09) | (0.22) | (0.19) |
| $r_t$ | 0.529*** | 0.378* | 0.681** | 0.151 | 0.081 | 0.090 | -0.323 | -0.310 | 1.565 | 0.172 |
| | (0.19) | (0.21) | (0.29) | (0.45) | (0.10) | (0.06) | (0.59) | (0.70) | (2.51) | (1.59) |
| $r_t$*$dummy_t$ | - | 0.155 | - | 0.937 | - | -0.021 | - | -0.085 | - | 2.354 |
| | - | (0.16) | - | (0.96) | - | (0.17) | - | (0.83) | - | (2.27) |
| $inc_{t+1}/inc_t$ | -0.160 | -0.168* | -0.492*** | -0.484*** | 0.026 | 0.026 | -0.302*** | -0.304*** | -0.014 | -0.185 |
| | (0.11) | (0.10) | (0.13) | (0.10) | (0.05) | (0.05) | (0.11) | (0.09) | (0.11) | (0.26) |
| $wea_t/inc_t$ | 0.002 | 0.003 | -0.012** | -0.005 | 1.28E-04 | 1.52E-04 | 0.002 | 0.001 | 0.004 | 0.012 |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00) | (0.01) | (0.02) | (0.03) |
| $inf_t$ | -0.423 | -0.358 | -0.903*** | -1.376* | 0.186 | 0.162 | -0.080 | -0.068 | -0.103 | 1.154 |
| | (0.41) | (0.33) | (0.56) | (0.80) | (0.18) | (0.25) | (0.53) | (0.66) | (1.18) | (2.15) |
| $savr_{t-1}$ | 0.083 | 0.184 | 0.285*** | 0.220 | 0.534*** | 0.518** | 0.314** | 0.293* | 0.646* | 0.700 |
| | (0.17) | (0.13) | (0.15) | (0.16) | (0.19) | (0.20) | (0.13) | (0.17) | (0.39) | (0.47) |
| $savr_{t-2}$ | 0.528*** | 0.480*** | 0.257*** | 0.250** | 0.028 | 0.039 | 0.114 | 0.125 | 0.244 | -0.052 |
| | (0.17) | (0.13) | (0.08) | (0.10) | (0.17) | (0.21) | (0.11) | (0.12) | (0.21) | (0.36) |
| $savr_{t-3}$ | 0.018 | -0.018 | 0.265 | 0.349** | -0.074 | -0.074 | 0.200 | 0.217 | -0.335* | -0.393* |
| | (0.15) | (0.14) | (0.09) | (0.15) | (0.13) | (0.15) | (0.16) | (0.20) | (0.19) | (0.21) |
| $savr_{t-4}$ | 0.389*** | 0.425*** | 0.052 | 0.103 | 0.223*** | 0.232** | 0.341** | 0.336** | 0.598* | 1.011 |
| | (0.13) | (0.11) | (0.12) | (0.14) | (0.08) | (0.11) | (0.14) | (0.15) | (0.31) | (0.73) |
| *JPR* | - | 0.036 | - | 0.146 | - | 0.686 | - | 0.771 | - | 0.461 |
| *R-squared* | 0.960 | 0.960 | 0.870 | 0.840 | 0.980 | 0.980 | 0.960 | 0.960 | 0.890 | 0.880 |
| *Prob. model* | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *SSR* | 0.003 | 0.003 | 0.009 | 0.011 | 0.000 | 0.000 | 0.007 | 0.007 | 0.028 | 0.032 |
| *LM test (prob.)* | 0.461 | 0.184 | 0.287 | 0.196 | 0.434 | 0.326 | 0.007 | 0.002 | 0.787 | 0.554 |

Notes: (1) standard errors are presented in parenthesis; (2) * indicates parameter significance at 10 percent level, ** at 5 percent level and *** at 1 percent level; (3) the results for the regression including the interaction term have been obtained using two-period lagged values of return and the interaction term as instruments; (4) quarterly dummy variables have been used to tackle seasonality; (5) JPR, which stands for joint probability of the return, tests whether the joint restriction $r_t + r_t*dummy_t = 0$; (6) the estimates for the quarterly dummies have not been reported so as to be able to fit the table in a single page; (7) LM test refers to the Breusch-Godfrey test for serial correlation, whose null hypothesis is that there is no serial correlation of any order up to 20; (8) it has not been possible to solve or diminish the severe autocorrelation suffered by the standard errors of the estimates obtained for Italy, the model used was the best of all the tested specifications.

### *5.3 Estimates of the EIS*

In the previous subsection we have observed rather contradictory evidence regarding the time-variation of the interest rate elasticity of saving: whereas the Wald F-statistic – obtained by means of the Quandt-Andrews test – and the Chow test indicated the presence of structural breaks both at the Euro Area and country level around 2009, the interaction term between the real rate of return and the period dummy suggested the opposite, being not significantly different from zero in any of the regressions estimated. What does evidence say regarding the elasticity of intertemporal substitution?

This section first presents the EIS estimates obtained using the real rate of return as the only explanatory variable for consumption growth (Equation 9), as established by the theoretical framework derived in Section 2, and then compares these results with the results obtained by means of a more comprehensive expression that includes relative financial wealth, relative future income and inflation as complementary regressors (Equation 10).

#### *5.3.1 Base model estimates*

Table 6 contains the estimates for the elasticity of intertemporal substitution obtained using the model presented in Equation 9. At the Euro Area level, before the introduction of the interaction term, the real rate of return presents a negative and significant effect on the rate of consumption growth (-0.22). However, once the interaction term is included in the regression, the coefficient for the rate of return ($\sigma_1$) is close to zero (-0.01).[10] This parameter suggests that, before and after the inter-break period, the EIS of the Euro Area would have been rather small.

Interestingly, the negative effect of the base term seems to be absorbed by the interaction term between the period dummy and the rate of return ($\sigma_2$ = -0.24). This, indicates that during the inter-break period (2007Q3-2009Q4), the EIS decreased substantially. As a result of this decrease, the EIS during the inter-break period would have become considerably negative[11]. It is important to remark that the interaction term, which presents a rather small standard error, is significant at the 1 percent level. In addition, the Wald test used to test the joint restriction $\sigma_1 + \sigma_2\delta_t = 0$, rejects the null hypothesis with a p-value of 0.00.

[10] As a reminder, the period dummy contained in the interaction term is based on the breaks observed in the interest elasticity of saving.

[11] The net effect of the real return on the rate of consumption growth during the inter-break period is the sum of $\sigma_1$ and $\sigma_2$.

*Table 6: Elasticity of intertemporal substitution (EIS) - Dependent variable: consumption growth - Estimation method: 2SLS*

| *Regressor/Area* | *Euro Area* | | *Finland* | | *Germany* | | *Italy* | | *Spain* | |
|---|---|---|---|---|---|---|---|---|---|---|
| *c* | 0.005* | 0.002 | 0.051*** | 0.048*** | 0.018*** | 0.018*** | -0.015 | -0.015 | 0.052*** | 0.050*** |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| $r_t$ | -0.224** | -0.016 | -0.230 | 0.322 | -0.168** | -0.159* | 0.017 | -0.022 | -0.485 | -0.337 |
| | (0.11) | (0.11) | (0.25) | (0.29) | (0.08) | (0.09) | (0.53) | (0.48) | (0.31) | (0.33) |
| $r_t$*$dummy_t$ | - | -0.239*** | - | -0.708*** | - | -0.022 | - | 0.148 | - | -0.185 |
| | - | (0.08) | - | (0.20) | - | (0.22) | - | (0.23) | - | (0.36) |
| *dq1* | 0.002 | 0.001 | -0.117*** | -0.118*** | -0.070*** | -0.070*** | 0.016** | 0.016** | -0.012 | -0.014 |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| *dq2* | -0.007*** | -0.006*** | 0.006 | 0.006 | 0.022*** | 0.022*** | 0.024* | 0.024* | -0.102*** | -0.098*** |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.02) |
| *dq3* | 0.008*** | 0.007*** | -0.061*** | -0.065*** | -0.005 | -0.005 | 0.018*** | 0.018*** | -0.058*** | -0.058*** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| *JPR* | - | 0.005 | - | 0.156 | - | 0.234 | - | 0.832 | - | 0.148 |
| *R-squared* | 0.402 | 0.503 | 0.943 | 0.944 | 0.954 | 0.954 | 0.859 | 0.855 | 0.876 | 0.872 |
| *Prob. model* | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *SSR* | 0.002 | 0.001 | 0.008 | 0.008 | 0.003 | 0.003 | 0.003 | 0.003 | 0.006 | 0.006 |
| *LM test (prob.)* | 0.763 | 0.544 | 0.645 | 0.777 | 0.253 | 0.236 | 0.335 | 0.250 | 0.191 | 0.210 |

Notes: (1) standard errors are presented in parenthesis; (2) * indicates parameter significance at 10 percent level, ** at 5 percent level and *** at 1 percent level; (3) the results for the regression including the interaction term have been obtained using two-period lagged values of return and the interaction term as instruments; (4) in addition, the results for Italy and Spain have been obtained using 4 and 2 lags of the dependent variable as regressors to correct for autocorrelation; (5) JPR, which stands for joint probability of the return, tests whether the joint restriction $r_t + r_t*dummy_t = 0$; (6) the estimates of quarterly dummies have not been reported so as to be able to fit the table in a single page; (7) LM test refers to the Breusch-Godfrey test for serial correlation, whose null hypothesis is that there is no serial correlation of any order up to 20.

At the member state level, Germany, Spain and especially Finland present a similar situation to the one depicted by the Euro Area. The results obtained for these three countries suggest negative EIS before the inclusion of the interaction term, with estimates ranging between -0.17 and -0.34. Once the interaction term is added to the regression, the base effect of the rate of return on consumption growth increases, while the interaction term ($\sigma_2$), with a negative coefficient – Finland (-0.70), Germany (-0.02) and Spain (-0.18) –, appears to absorb part of the negativity of the base term. This, indicates that the EIS decreased during their respective inter-break period.

Unlike the other countries analysed, Italy presents a positive EIS before the introduction of the interaction term. Once this latter variable is added to the regression, it suggests an increase of the EIS during the inter-break period (0.12) which, again, contrasts with the negative change of the EIS observed for Finland, Germany and Spain.

Regarding the country-level results, we must express our surprise for the generally negative sign of the EIS, which implies a negative relative risk aversion and, thus, risk loving consumers. In addition, we must warn the reader that, as those obtained for the IRES, the country-level estimates obtained on the EIS should be taken with caution for three main reasons: (1) the base term is only significant for Germany, (2) the interaction term is only significantly different from zero for Finland and (3) the Wald test fails to reject the null hypothesis ($\sigma_1 + \sigma_2\delta_t = 0$) for each of the member states analysed.

*5.3.2 Comprehensive model estimates*

Table 7 contains the estimates for the elasticity of intertemporal substitution obtained using the model presented in Equation 10. After including relative financial wealth, relative future income and inflation as control variables for possible omitted variable bias, we can see that the estimates obtained for the Euro Area as a whole, Finland and Germany are substantially similar to those presented by the theoretical model. Even though some differences can be seen on the magnitudes of the estimates, the pattern described by their respective EIS throughout the sample period is highly similar to that presented by the previous model: their EIS appears to be considerably smaller during the inter-break period, as suggested by the negative sign of the interaction term between the period dummy and the real rate of return.

On the contrary, the EIS estimates for Italy and Spain present important changes across models. Whereas the theoretical model showed Italian willingness to substitute consumption across periods to increase during the inter-break period and that of Spain to decrease, the estimates obtained with the more comprehensive model illustrate exactly the opposite: $\sigma_2$ negative for Italy and positive for Spain. Important changes can also be

*Table 7: Elasticity of intertemporal substitution (EIS) - Dependent variable: consumption growth - Estimation method: 2SLS*

| *Regressor/Area* | *Euro Area* | | *Finland* | | *Germany* | | *Italy* | | *Spain* | |
|---|---|---|---|---|---|---|---|---|---|---|
| *c* | -0.029 | -0.033 | -0.020 | 0.019 | 0.021 | 0.017 | -0.244*** | -0.230*** | -0.012 | -0.009 |
| | (0.03) | (0.03) | (0.03) | (0.08) | (0.04) | (0.03) | (0.08) | (0.09) | (0.03) | (0.03) |
| $r_t$ | -0.114 | -0.005 | -0.132 | 0.277 | -0.245* | -0.200 | 0.322 | 0.337 | -0.090 | -0.308 |
| | (0.11) | (0.12) | (0.21) | (0.69) | (0.15) | (0.13) | (0.30) | (0.28) | (0.35) | (0.38) |
| $r_t$*$dummy_t$ | - | -0.124 | - | -0.582 | - | -0.146 | | -0.398 | - | 0.292 |
| | - | (0.09) | - | (0.98) | - | (0.35) | | (0.39) | - | (0.45) |
| $ln(inc_{t+1}/inc_t)$ | -0.012 | -0.007 | 0.054 | 0.068 | -0.226* | -0.217** | -0.199** | -0.197*** | -0.017 | -0.033 |
| | (0.10) | (0.08) | (0.08) | (0.09) | (0.12) | (0.12) | (0.07) | (0.06) | (0.06) | (0.06) |
| $ln(wea_t/inc_t)$ | 0.015 | 0.016 | 0.050** | 0.026 | -0.001 | 0.000 | 0.082** | 0.076** | 0.029*** | 0.032** |
| | (0.01) | (0.01) | (0.02) | (0.05) | (0.02) | (0.02) | (0.03) | (0.03) | (0.01) | (0.01) |
| $ln(inf_t)$ | 0.000*** | 0.000*** | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | -0.001 | 0.000 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| *dq1* | -0.002 | -0.003 | -0.116*** | -0.122*** | -0.072*** | -0.071*** | 0.045** | 0.046*** | -0.018 | -0.019 |
| | (0.01) | (0.01) | (0.02) | (0.02) | (0.00) | (0.00) | (0.02) | (0.02) | (0.02) | (0.02) |
| *dq2* | -0.008*** | -0.007** | -0.015 | -0.019* | 0.022*** | 0.022*** | 0.028*** | 0.027*** | -0.088*** | -0.095*** |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.02) |
| *dq3* | 0.006 | 0.005 | -0.053** | -0.065** | -0.002 | -0.002 | 0.069*** | 0.069*** | -0.058*** | -0.064*** |
| | (0.01) | (0.01) | (0.02) | (0.03) | (0.00) | (0.00) | (0.02) | (0.02) | (0.02) | (0.02) |
| *consg(-1)* | -0.022 | -0.037 | -0.397** | -0.422*** | - | - | - | - | -0.103 | -0.116 |
| | (0.17) | (0.15) | (0.16) | (0.15) | - | - | - | - | (0.15) | (0.14) |
| *consg(-2)* | 0.346*** | 0.279** | -0.077 | -0.154 | - | - | - | - | 0.312** | 0.381** |
| | (0.13) | (0.14) | (0.13) | (0.15) | - | - | - | - | (0.15) | (0.20) |
| *JPR* | - | 0.230 | - | 0.463 | - | 0.328 | - | 0.916 | - | 0.970 |
| *R-squared* | 0.573 | 0.623 | 0.960 | 0.955 | 0.957 | 0.957 | 0.771 | 0.810 | 0.886 | 0.885 |
| *Prob. model* | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *SSR* | 0.001 | 0.001 | 0.006 | 0.006 | 0.003 | 0.003 | 0.005 | 0.004 | 0.005 | 0.005 |
| *LM test (prob.)* | 0.591 | 0.693 | 0.371 | 0.616 | 0.283 | 0.168 | 0.065 | 0.103 | 0.200 | 0.170 |

Notes: (1) standard errors are presented in parenthesis; (2) * indicates parameter significance at 10 percent level, ** at 5 percent level and *** at 1 percent level; (3) the results for the regression including the interaction term have been obtained using two-period lagged values of return and the interaction term as instruments; (4) the estimates for Germany and Italy have been obtained without using 2 lags of the dependent variable as regressors because, unlike for the other countries, this specification did not suffer from autocorrelation; (5) JPR, which stands for joint probability of the return, tests whether the joint restriction $r_t + r_t$*$dummy_t = 0$; (7) LM test (prob.) refers to the Breusch-Godfrey test for serial correlation, whose null hypothesis is that there is no serial correlation of any order up to 20.

observed when it comes to the significance of the estimates. The joint restriction that $\sigma_1 + \sigma_2\delta_t = 0$ loses significance. As a result, the null hypothesis fails to be rejected for the Euro Area and each of the member states analysed.

The estimates presented in subsections 5.3.1 and 5.3.2 indicate that, as hypothesized, the willingness of Euro Area households to shift consumption across periods as a response to real interest rate changes would have decreased in the vicinity of 2009, with the only exception of Italy (in the base model) or Spain (in the extended model). However, this decrease is only significant for the Euro Area as a whole and Finland.

## 6. Conclusion

To conclude, we relate the results presented in the previous section to the twofold hypothesis raised at the beginning of this paper. As a reminder, the first part of our hypothesis suggested that, 2009's increase in the saving rate of Euro Area households, could have been caused by a decrease in their interest elasticity of saving (IRES). If, as a result of this decrease, the IRES had become negative, households would have increased relative saving as a response to the decreasing real returns. With regards to the second part of our hypothesis, it proposed that the decrease in households' IRES, could have been caused by a drop in their elasticity of intertemporal substitution (EIS).

Starting with the second part of our hypothesis, as predicted, the results suggest that the EIS of Euro Area households decreased in the vicinity of 2009. This same situation is observed at the country level, with the only exception of Spain or Italy, depending on the empirical specification used to estimate the EIS.

However, relating now to the first part of our hypothesis, did the decrease of the EIS lead to a negative IRES?

(1) At the Euro Area level, our estimates indicate that the considerable decrease of the EIS did not result in a negative IRES. In fact, not only is the interaction term of the IRES smaller than its base term in absolute magnitude but, in addition to that, the interaction term, which captures the change in the IRES during the inter-break period, shows a positive sign. This, rules out a negative IRES as the cause of the large saving rate increase experienced by the Euro Area in 2009.

(2) At the country level, our results are highly heterogeneous. The connection between the EIS and the IRES in Finland resembles that of the Euro Area. The estimates show a positive and larger IRES during the inter-break period, despite the concurrent decrease in households' EIS. In Germany, whereas the decline in the EIS during the inter-

break period coincided with a decrease in the IRES, our results show that the decrease of the IRES was not strong enough for this parameter to become negative.

Concerning Italy and Spain, based on the full model, which shows a somehow larger explanatory power and a generally lower sum of squared residuals, the results indicate that, in Italy, the decline of the EIS coincided with a lower and negative IRES. With regards to Spain, the changes in the EIS and the IRES during the inter-break period relate to each other as predicted by the theoretical model. However, the positive sign of the change in the EIS and the IRES during the period between 2008Q3 and 2009Q4 contradicts our hypothesis.

*Table 8: Summary table*

| | *EA* | *FI* | *DE* | *IT* | *ES* |
|---|---|---|---|---|---|
| *Period* | 07q3-09q3 | 07q4-11q1 | 08q4-10q1 | 08q3-09q2 | 08q3-09q4 |
| Δ *EIS inter-break period* | - | - | - | - | + |
| Δ *IRES inter-break period* | + | + | - | - | + |
| *Sign IRES inter-break period* | + | + | + | - | + |

Notes: (1) Δ EIS inter-break period is based on the full model (Equation 10); (2) the sign of the IRES is determined by the sum of the base term ($\beta_1$) and the interaction term ($\beta_2$) of Equation 8.

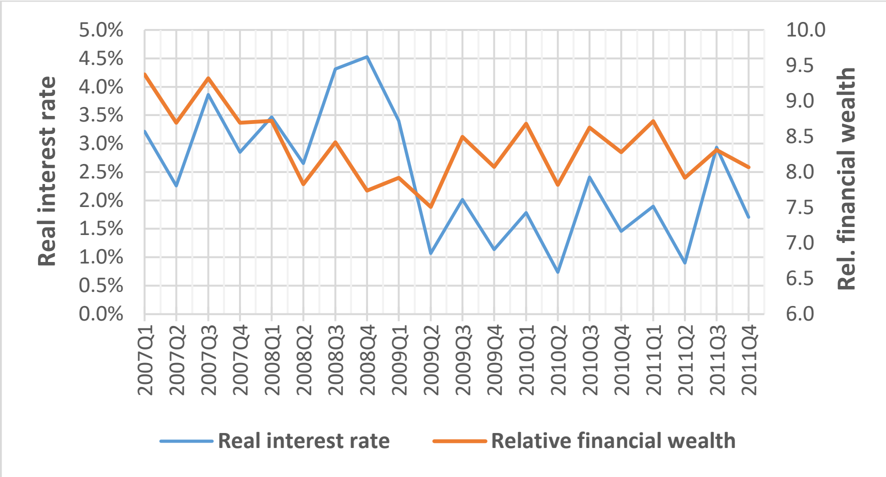
If, as our results suggest, the EIS of Euro Area households decreased between 2007Q3 and 2009Q3, why did their IRES concurrently increase? To answer this question, we have to go back to the theoretical expression for the IRES:

$$IRES = \frac{\partial\, s_1/y_1}{\partial R} = \underbrace{(\sigma - 1)\frac{\beta^{\sigma}\bar{R}^{\sigma-2}}{1+\beta^{\sigma}\bar{R}^{(\sigma-1)}}}_{\text{(I) subst. and (II) income effect}} + \underbrace{\frac{\bar{Y}}{\bar{R}^2\left(1+\bar{A}+\frac{\bar{Y}}{\bar{R}}\right)}}_{\text{(III) wealth effect}} \tag{5}$$

Equation 5 shows the importance of the magnitude of relative wealth ($\bar{A}$), relative future income ($\bar{Y}$) and the real rate of interest ($\bar{R}$) in determining the strength of the wealth effect and, thus, of the IRES. If we analyse the evolution of these three variables, we can observe that, in 2009, whereas relative future income remained fairly constant, relative financial wealth and the rate return decreased by 2.2 percent and 57.6 percent respectively (Figure 7).

Since both relative financial wealth and the real rate return are present in the denominator of a positive fraction, it becomes apparent that their decline, could have enlarged the strength of the wealth effect. This, would have led to an increase in the IRES during the inter-break period, despite the drop in the EIS. In other words, the increased positive strength of the wealth effect, could have been larger than the increased negative power of the income effect, leading to an increase of individuals' IRES.

*Figure 7: Euro Area real interest rate and relative financial wealth*




Source: ECB, Eurostat and OECD

Finally, we refer back to the hypothesis raised by Daniel Gros. This author, suggested that the difficulties of the Euro Area to return to the consumption growth rates experienced before the financial crisis, could be explained by a backward bend in the lower end of the curve of saving supply. However, as we have seen in the above paragraph, our theoretical model shows the IRES to be increasing in the real rate of interest. Consequently, from a theoretical point of view, the possibility of a backward bend in the presence of life-cycle consumers appears to be quite remote. In addition, since 2010, both the real rate of interest and the saving rate of Euro Area households have shown a decreasing trend. In light of this evidence, a sustained backward bend in the IRES seems rather improbable.

## Appendix A

### *A.1 Derivation of the saving rate function*

The two period consumer's problem is:

$$\text{max.} \quad u(c_1) + \beta u(c_2) \quad \text{subject to} \quad c_1 + \frac{c_2}{(1+r)} = a_0 + y_1 + \frac{y_2}{(1+r)}$$

where $c_t$ and $y_t$ respectively denote the amount of consumption and income in period t, $a_0$ stands for consumers' initial endowment of financial assets, r is the real rate of return to wealth and β is agents' subjective discount factor, which captures how much they care about future consumption. The consumer's problem can be solved setting up a Lagrangian function, which allows us to determine the optimal consumption plan:

$$\mathrm{L} = u(c_1) + \beta u(c_2) + \lambda\left[\left(a_0 + \mathrm{y}_1 - \mathrm{c}_1\right)(1 + \mathrm{r}) + \mathrm{y}_2 - \mathrm{c}_2)\right]$$

The partial derivatives of the Lagrangian function with respect to $c_1$ and $c_2$ are given by:

$$\frac{\partial \mathrm{L}}{\partial \mathrm{c}_1} = u^{'}(c_1) - \lambda(1 + \mathrm{r}) = 0 \;\; ; \; \frac{\partial \mathrm{L}}{\partial \mathrm{c}_2} = u^{'}(c_2) - \lambda = 0$$

Solving for λ, the first order conditions can be summarized by:

$$\frac{u'(c_1)}{1 + r} = \beta\, u'(c_2)$$

Now, using constant relative risk aversion utility

$$u(c) = \frac{\mathrm{c}^{1-\frac{1}{\sigma}}}{1 - \frac{1}{\sigma}} \;\; with \;\; u'(c) = c^{-\frac{1}{\sigma}}$$

and solving for $c_2$, the Euler equation can be expressed as:

$$c_2 = [\beta\, (1 + r)\, ]^{\sigma} c_1$$

Now, solving the budget constraint for $c_1$ and plugging the above Euler equation inside the budget constraint, we can obtain:

$$c_1 = a_0 + y_1 + \frac{y_2}{(1 + r)} - \frac{[\beta\, (1 + r)\, ]^{\sigma} c_1}{(1 + r)}$$

and, solving for $c_1$ we can find consumption in period one as a function of $a_0$, $y_1$, $y_2$, and r:

$$c_1 = \left[a_0 + y_1 + \frac{y_2}{(1 + r)}\right] * \frac{1}{\left[1 + \frac{[\beta\, (1 + r)\, ]^{\sigma}}{(1 + r)}\right]}$$

Then, making use of the fact that actual saving equals current income minus current consumption ($s_1 = y_1 - c_1$), we can obtain saving in t=1:

$$s_1 = y_1 - \left[a_0 + y_1 + \frac{y_2}{(1 + r)}\right] * \frac{1}{[1 + \beta^{\sigma}\, (1 + r)^{\sigma - 1}\, ]}$$

and, dividing both sides of the saving function by y1 we can find the rate of saving:

$$\frac{s_1}{y_1} = 1 - \left[1 + \frac{a_0}{y_1} + \frac{y_2/y_1}{(1+r)}\right] * \frac{1}{[1 + \beta^{\sigma}(1+r)^{\sigma-1}]}$$

Now, taking the natural logarithm of both sides and using the fact that, since the rate of saving is considerably close to zero, $ln(1 - \frac{s_1}{y_1})$ is approximately equal to $(-\frac{s_1}{y_1})$, this expression can be simplified to:

$$\ln(1 - \frac{s_1}{y_1}) = ln\left(1 + \frac{a_0}{y_1} + \frac{y_2/y_1}{1+r}\right) + ln\left(\frac{1}{[1 + \beta^{\sigma}(1+r)^{\sigma-1}]}\right)$$

$$\ln\left(1 - \frac{s_1}{y_1}\right) = ln\left(1 + \frac{a_0}{y_1} + \frac{y_2/y_1}{1+r}\right) + ln(1) - ln(1 + \beta^{\sigma}(1+r)^{\sigma-1})$$

$$\frac{s_1}{y_1} = -ln\left(1 + \frac{a_0}{y_1} + \frac{y_2/y_1}{1+r}\right) + ln\left(1 + \beta^{\sigma}(1+r)^{(\sigma-1)}\right)$$

Before taking the first order Taylor approximation we simplify the notation of this equation by substituting $\frac{a_0}{y_1}$ by A, $\frac{y_2}{y_1}$ by Y and $(1+r)$ by R:

$$\frac{s_1}{y_1} = -ln\left(1 + A + \frac{Y}{R}\right) + ln\left(1 + \beta^{\sigma}R^{(\sigma-1)}\right)$$

Finally, we take the first order Taylor approximation of this expression around $A = \bar{A}$, $Y = \bar{Y}$ and $R = \bar{R}$ to obtain households' saving rate as a linear function of relative wealth, relative future income and the real rate of return, each of them multiplied by their partial effect on household saving rate:

$$\frac{s_1}{y_1} \approx c + \beta_1(R - \bar{R}) + \beta_3(Y - \bar{Y}) + \beta_4(A - \bar{A})$$

where:

$$c = f(\bar{A}, \bar{Y}, \bar{R}) = ln\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right) + \beta^{\sigma}\bar{R}^{(\sigma-1)}$$

$$\beta_1 = \frac{\partial\, s_1/y_1}{\partial R} = \frac{1}{\left(1 + \bar{A} + \frac{\bar{Y}}{\bar{R}}\right)} * \frac{\bar{Y}}{\bar{R}^2} + (\sigma - 1)\frac{\beta^{\sigma}\bar{R}^{\sigma-2}}{1 + \beta^{\sigma}\bar{R}^{(\sigma-1)}}$$

$$\beta_3 = \frac{\partial\, s_1/y_1}{\partial Y} = -\frac{1}{\left(1+\bar{A}+\frac{\bar{Y}}{\bar{R}}\right)} * \frac{1}{\bar{R}}$$

$$\beta_4 = \frac{\partial\, s_1/y_1}{\partial A} = -\frac{1}{\left(1+\bar{A}+\frac{\bar{Y}}{\bar{R}}\right)} * 1$$

and, moving $\bar{R}$, $\bar{Y}$ and $\bar{A}$ inside the constant term (α) we can obtain what is going to be our empirical specification to estimate the interest elasticity of saving:

$$\frac{s_1}{y_1} \approx \alpha + \beta_1 R + \beta_3 Y + \beta_4 A$$

where:

$$\alpha = c + \frac{1}{\left(1+\bar{A}+\frac{\bar{Y}}{\bar{R}}\right)}\left(\bar{A}+\frac{\bar{Y}}{\bar{R}}+\frac{\bar{R}\bar{Y}}{\bar{R}^2}\right) - (\sigma-1)\frac{\beta^\sigma \bar{R}^{\sigma-2}}{1+\beta^\sigma \bar{R}^{(\sigma-1)}} * \bar{R}$$

*A.2 Derivation of the function of consumption growth*

$$c_2 = [\beta\,(1+r)\,]^\sigma c_1$$

Moving $c_1$ to the right hand side of the Euler equation and taking natural logarithms of both side, consumption growth is given by:

$$\ln\left(\frac{c_2}{c_1}\right) = ln(\,[\beta\,(1+r)\,]^\sigma)$$

which, using the properties of logarithms, can be rewritten as:

$$\ln\left(\frac{c_2}{c_1}\right) = \sigma\,[ln\beta + \ln(1+r)]$$

Finally, taking the first order Taylor approximation of this expression around r= 0 we can find consumption growth as a function of the real return and the partial effect of this latter variable on the former:

$$\ln\left(\frac{c_2}{c_1}\right) = f(0) + \frac{\partial\, ln\, c_1/c_2}{\partial r}\,(r-0)$$

$$\ln\left(\frac{c_2}{c_1}\right) = \sigma[ln\beta + \ln(1+0)] + \sigma\frac{1}{1+0}(r-0)$$

$$\ln\left(\frac{c_2}{c_1}\right) = c + \sigma r$$

*A.3 Derivation of the elasticity of intertemporal substitution*

$$\ln\left(\frac{c_2}{c_1}\right) = \sigma\ ln\beta + \sigma \ln(1+r)$$

Taking the first difference of both sides of the consumption growth function derived in Appendix A.2 we obtain:

$$\text{dln}\left(\frac{c_2}{c_1}\right) = d[\sigma ln\beta + \sigma \ln(1+r)]$$

which, can be transformed to:

$$\text{dln}\left(\frac{c_2}{c_1}\right) = \sigma dln\beta + \sigma\ \text{dln}(1+r)$$

Now, since β is a constant, dln(β) is equal to zero, which leaves us with:

$$\text{dln}\left(\frac{c_2}{c_1}\right) = \sigma\ \text{dln}(1+r)$$

Finally, bringing dln(1+r) to the left hand side of the equation we obtain the EIS:

$$EIS = \frac{\partial \ln\left(\frac{c_2}{c_1}\right)}{\partial \ln(1+r)} = \sigma$$

## Appendix B

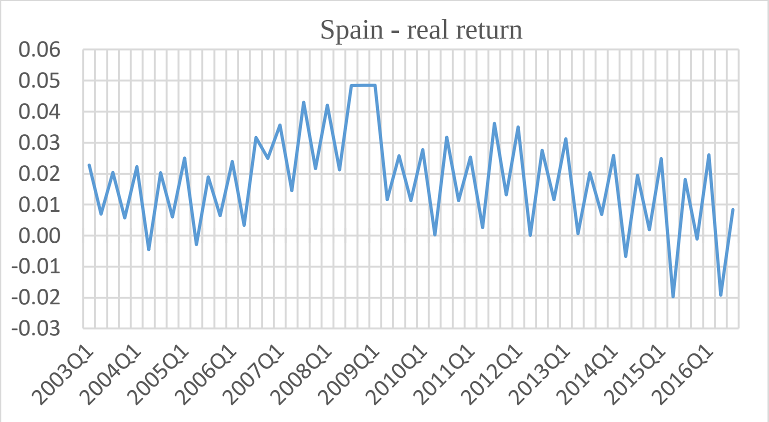
*Figure 8: Evolution of the real rate of return by member state*

As shown by this figure, the real return to household deposits with a maturity of up to 1 year experienced an unprecedented decrease between late 2008 and early 2009.


Source: European Central Bank


Source: European Central Bank


Source: European Central Bank




Source: European Central Bank

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