The Long-run Determinants of Consumption in the Euro Area: Is there a Role for Uncertainty?

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Abstract

This paper examines the long-run (cointegrating) relationship between real consumption, real disposable income, real net financial wealth, real housing wealth, and uncertainties in future income (income uncertainty) and the rate of return on accumulated financial wealth (capital uncertainty) for a panel of 12 euro area countries. Using proxies for the unobservable housing wealth and income and capital uncertainty, we show that such a relationship does exist, but it is not homogenous for the euro area as a whole. Real disposable income and real net financial wealth are the main determinants of real consumption for the PIIGS (Portugal, Italy, Ireland, Greece, and Spain) and non-PIIGS (Austria, Belgium, Finland, France, Germany, Netherlands, and Slovenia) euro areas. Income and capital uncertainties are negatively associated with real consumption, but only in the PIIGS euro area.

Keywords: *consumption, disposable income, wealth, uncertainty, euro area*

JEL Classification: C51, E20, E21, E24

Introduction

The Great Recession and the sovereign debt crisis in the euro area elevated the uncertainty[1] in the euro area financial markets and were also associated with a drop in consumption among euro area households.[2] Thus, it is an important

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[1] As noted by the European Central Bank (ECB, 2016), it is difficult in empirical studies to separate between risk and (Knightian) uncertainty (see Knight, 1921). The first relates to situations in which economic agents can make decisions informed by the probability distribution of the possible future states of the matter, whereas in the latter situation no such probability distribution exists.

[2] Eurostat data show that the seasonally adjusted real final consumption of the euro area households and the non-profit institutions serving households steadily rose during 1999Q1 – 2007Q4, and then double-dipped in the third quarter of 2009 and the first quarter of 2013.

academic and, due to the size of the consumption in the gross domestic output (GDP), economic policy-relevant aim to empirically assess the association between uncertainty and consumption in the euro area.

The foundational contributions of Campbell and Mankiw (1989) and Lettau and Ludvigson (2001; 2004) to the theory of consumption have posited a stable long-run relationship among consumption, disposable income, and asset wealth,[3] yet the literature is silent regarding the role of uncertainty. The role of uncertainty as an important determinant of consumption has been stressed by theoretical contributions of, among others, Leland (1968), Sandmo (1970), Miao (2004), Eeckhoudt and Schleisinger (2008), Gunning (2010) and Vergara (2017). This strand of literature has in general identified two types of uncertainty that affect current consumption: uncertainty about future income (income uncertainty) and uncertainty about the rate of return on accumulated financial wealth (capital uncertainty). Under certain conditions, an increase in income uncertainty increases precautionary saving and decreases consumption (Leland, 1968; Sandmo, 1970), whereas the effect of an increase in capital uncertainty on consumption is uncertain. It can either increase or reduce consumption. Several studies have empirically investigated the association between consumption and income uncertainty (among others, e.g., Guiso, Jappelli and Terlizzese, 1992; Lyhagen, 2001; Costa et al., 2016), but the association between consumption and capital uncertainty has received much less attention (exceptions include Choudhry (2003) and Ibrahim and Law (2013)) and, to our best knowledge, is still undocumented for the euro area as a whole.[4]

This paper aims to fill this gap. Building on the theory and the existing empirical studies on the consumption function, we assess for the euro area as a whole the long-run (cointegrating) relationship between consumption and its determinants,

[3] It has become common in the empirical literature (see, e.g., Choudhry, 2003; Hamburg, Hoffmann and Keller, 2008; Slacalek, 2009; Costa et al., 2016) that builds on this theoretical premise to model consumption in a cointegrating framework.

[4] We note that recently an increasing body of empirical literature has shown that uncertainty contributed largely to the downturn in macroeconomic output around the world during the financial crisis and subsequent sluggish recovery (see, e.g., Bloom, 2009; Alexopoulos and Cohen, 2009; Stock and Watson, 2012; Holló, Kremer and Lo Duca, 2012; Bloom et al., 2012; Basu and Bundick, 2012; Gudmundsson and Natvik, 2012; Bonciani and van Roye, 2016; Moore, 2016; ECB, 2016). A few studies, without trying to estimate the consumption function, also analyze a dynamic, short-run association between uncertainty and consumption growth, typically in a vector autregression model. The majority of these studies concentrates on the U.S. (see, e.g., Alexopoulos and Cohen, 2009; Basu and Bundick, 2012) or other individual countries (e.g., Gudmundsson and Natvik, 2012). To our best knowledge, only Bonciani and van Roye (2016), applying aggregate euro area time series data, investigate a short-run association between uncertainty shocks, proxied by financial market variables in the euro area, and macroeconomic variables (including output, consumption, investment, capital, and inflation rate). They find that regardless of price rigidity, consumption growth responds negatively to an uncertainty shock.

including disposable income, net financial wealth, real house prices, unemployment rate (a proxy for income uncertainty), and a composite indicator of volatility in the financial markets of the euro area (a proxy for capital uncertainty).[5] Unlike the existing studies (e.g., Choudhry, 2003; Ibrahim and Law, 2013) that proxy for capital uncertainty by volatility in the stock market indexes,[6] we propose to use a more comprehensive proxy: the composite indicator of systemic stress (CISS), developed by Holló, Kremer and Lo Duca (2012), which is an aggregate euro area composite indicator of volatility in five segments of euro area financial markets (the money, equity, bond and foreign exchange markets, and financial intermediaries' risk profile). We utilize panel autoregressive distributed lag (ARDL) model, which enables us to explicitly account for heterogeneity of euro area countries. Unlike the existing consumption function studies (e.g., Skudelny, 2009; Sousa, 2009; Costa el al., 2016) for the euro area, we also verify whether the euro area can be treated as a homogenous group of countries.

Literature Review

The theoretical foundations of our empirical model are based on the works of Campbell and Mankiw (1989) and Lettau and Ludvigson (2001; 2004) that trace their foundations back to the life-cycle hypothesis of Modigliani and Brumberg (1954) and Ando and Modigliani (1963). In these models, the consumer faces an inter-temporal budget constraint, linking his or her consumption and aggregate wealth. Campell and Mankiw (1989) show that under certain assumptions (e.g., all the wealth being tradable and the stationarity of changes in the logarithms of consumption and net return on aggregate wealth) the difference in the logarithm of aggregate consumption and the logarithm of aggregate wealth should be stationary; that is, between them there should be a stable long-run (cointegrating) relationship. Lettau and Ludvigson (2001; 2004) extend the validity of the long--run relationship to a trivariate case by decomposing the aggregate wealth into human and asset (non-human) wealth. They argue for the validity of the long-run relationship even when the unobservable human wealth is substituted by the observable labor (disposable) income. Sinai and Souleles (2005) and Case, Quigley and Shiller (2005), among others, argue that asset wealth in the eyes of consumers is not homogenous. Consumers may respond differently to a change in housing wealth than to a change in financial wealth due to, among other things, bequest

[5] The latter three variables are used as proxies for the euro area panel unobservable variables of housing wealth and income and capital uncertainties.

[6] The same proxies are commonly used in the literature (see, e.g., Bloom, 2009; Basu and Bundick, 2012; Hirata et al., 2013).to analyze the short-run effect of uncertainty on macroeconomic variables.

motive, uncertainty about the actual value of the wealth, and the mental segregation of the group of asset wealth (ibidem). Following this reasoning, the aggregate wealth in Lettau and Ludvigson's (2001; 2004) model can be decomposed into human wealth and financial and nonfinancial (housing) wealth (see, e.g., De Veirman and Dunstan, 2008). The cointegrating relationship between consumption and its determinants put forward by theory[7] has been confirmed in several papers examining the wealth effect on consumption (e.g., Lettau and Ludvigson, 2004; Hamburg, Hoffmann and Keller, 2008; Muellbauer, 2008; De Bonis and Silvestrini, 2012; Costa et al., 2016).

The theoretical models of Campell and Mankiw (1989) and Lettau and Ludvigson (2001; 2004) assume a quadratic utility function. Allowing for a positive third derivative function of the utility function allowed Leland (1968) and Sandmo (1970) to show that income uncertainty is negatively (positively) related to consumption (saving)[8] and sparked empirical research on precautionary saving (e.g., Guiso, Jappelli, and Terlizzese, 1992; Lyhagen, 2001; Costa et al., 2016), confirming the theoretical prediction about the sign of association between consumption and income uncertainty. Lyhagen (2001), drawing on the studies of Caballero (1990), Weil (1993), and Guiso, Jappelli, and Terlizzese, (1992), shows that theoretically[9] a stable long-run relationship among consumption, wealth, and income uncertainty can be expected. Also, this theoretical prediction has received empirical support in the studies of, for example, Lyhagen (2001) and Costa et al. (2016).

Leland (1968), Sandmo (1970), Miao (2004), Eeckhoudt and Schleisinger (2008), Gunning (2010) and Vergara (2017) also investigated the role of capital uncertainty in consumption decisions. They note that an increase in capital uncertainty results in two conflicting effects: Because there is more to lose when more wealth is accumulation, the utility-maximizing consumer is likely to reduce saving and increase consumption (the substitution effect) in response to an increase in uncertainty about the rate of return on accumulated financial wealth; at the same time, to insure against a lower level of future consumption, the consumer is likely to increase saving and reduce consumption (the income effect). Which of the effects prevails is a matter for empirical research to investigate. The empirical evidence of the effect on the association between consumption and capital uncertainty is scarce, but includes Choudhry (2003) and Ibrahim and Law (2013). Choudhry (2003) uses stock market volatility as a proxy for capital

[7] Under certain assumptions, the cointegrating relationship among consumption, income, and wealth can also be explained within the theoretical framework of the permanent income hypothesis (see, e.g., Altissimo et al., 2005 and De Bonis and Silvestrini, 2012).

[8] See Kimball (1990) and Carroll and Kimball (1996) relating uncertainty to risk aversion.

[9] See also, e.g., Skinner (1988).

uncertainty and for the U.S. finds a negative association between consumption and capital uncertainty. Ibrahim and Law (2013), applying the same approach as Choudhry (2003), draw an identical conclusion for Malaysia.

Although the empirical literature on the consumption function is vast, the literature that focuses on the euro area as a whole is thin. The empirical studies for the euro area as a whole include Coenen and Straub (2005), Sousa (2009), Skudelny (2009), Jawadi and Sousa (2012) and Costa et al. (2016). Coenen and Straub (2005) apply aggregate euro area data for 1980 – 1999 and in the framework of the new-Keynesian dynamic stochastic general equilibrium modeling (DSGE) concentrate on analysis of the effects of shocks in government expenditure on consumption. They find that the prevalence of Ricardian households in the euro area causes a negative wealth effect (i.e., government spending shocks result in households' increased saving), thus suppressing consumption. Sousa (2009) uses quarterly aggregate euro area time series data for 1980Q1 – 2007Q4 to assess the long- and short-run elasticity of consumption to changes in disposable income and financial and housing wealth. His results show that all three variables are statistically significant, with expected positive slope coefficients. The elasticity of consumption to a disposable income increase is the highest, followed by financial wealth and housing wealth. The results are fairly robust to different measures of financial and housing wealth (gross/net wealth). Skudelny (2009) applies time series analysis on the aggregate euro area time series data for 1980 – 2006 and a panel data analysis covering eight countries for 1995 – 2006 to analyze the financial and housing wealth effects on consumption. She finds that changes in all wealth components are significantly positively related to changes in consumption in both the time series and panel data models. The marginal propensity to consume out of (net) financial wealth is higher than the marginal propensity to consume out of housing wealth. Jawadi and Sousa (2012) compare the wealth effects on consumption for the U.S., euro area, and U.K. The data sample for the euro area covers 1980:Q4 – 2008:Q1. The authors show that labor income, financial wealth, and housing wealth are cointegrated with consumption for all three economic areas. Applying cointegration analysis, the authors show that for the euro area the long-run elasticities of consumption with respect to labor income and financial wealth are significantly positive at 0.71 and 0.11, respectively. The elasticity of consumption with respect to housing wealth is positive (0.02) but statistically insignificant. The results for the U.K. are similar, with the exception of the housing wealth effect being statistically significant and larger than for the euro area, while for the U.S. the elasticity of consumption with respect to labor income exceeds 1 and the elasticity of consumption with respect to housing wealth is negative. The authors also apply quantile regression to show that during economic booms the sensitivity of consumption to changes in wealth and income are larger than during economic slowdowns. The study of Costa et al. (2016) is the one to which our study relates most closely. The authors estimate the consumption function for an unbalanced panel data sample of 11 euro area countries for 2000:Q1 – 2013:Q4. Applying the ARDL approach to co-integration analysis, they investigate long-run consumption elasticity with respect to disposable income, household indebtedness (measured by the loan-to-income ratio), gross public debt, real housing prices (used as a proxy for housing wealth), share prices (a proxy for financial wealth), deposit rate, and unemployment rate (proxying for uncertainty in income). The results show that the long-run elasticity of consumption with respect to disposable income is the largest (0.58), followed by real housing prices (0.08), the loan-to-income ratio (0.05), and real share prices (0.02). Other explanatory variables are insignificant at the 5% level and the elasticity of consumption pertaining to the unemployment rate is estimated to be zero.

Methodology

To assess the long-run (cointegrating) relationship between real consumption and explanatory variables, including real disposable income, real net financial wealth, real housing wealth, income uncertainty, and capital uncertainty, the following ARDL model is proposed (see Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999; Blackburne and Frank, 2007):[10]

$$c_{it} = \sum_{j=1}^{p} \eta_{ij} c_{i,t-j} + \sum_{j=0}^{q} \vartheta_{ij}' \boldsymbol{x}_{i,t-j} + \mu_i + \varepsilon_{it} \tag{1}$$

where i $(i = 1,\ldots,N)$ denotes a panel group (country), and t $(t = 1,\ldots,T)$ denotes time, c_{it} is the log of consumption, $\boldsymbol{x}_{it}$ is a $(kx1)$ vector of explanatory variables including the natural logarithms of the real disposable income, real net financial wealth, and real house price index.[11]

[10] An ARDL approach to consumption function modeling is used also by, e.g., Pesaran, Shin and Smith (1999), Blackburne and Frank (2007), Ciarlone (2012) and Costa et al. (2016).

[11] Data for housing wealth are not available for individual euro area countries. Therefore, we proxy for real housing wealth of euro area members by using the real house price index. This is a quite common approach in the empirical analysis of wealth effects on consumption (see, e.g., Aron and Muellbauer, 2008; Campbell and Cocco, 2007; Ciarlone, 2012; Costa et al., 2016). Applying this proxy entails assuming that the number of dwellings does not change (i.e., housing stock is fixed) and that the existing dwellings do not improve in quality. As Case, Quigley and Shiller (2005) note, the benefit of using this proxy is that consumption and housing wealth are not significantly related “merely because housing consumption is a component of aggregate consumption.”

Two additional variables are included in vector $\boldsymbol{x}_{it}$ – the unemployment rate and the CISS indicator that proxy for income and capital uncertainties, respectively k is the number of different explanatory variables, μ_i is the group (country) fixed effect, η_{ij} is a scalar of slope coefficients for the lagged dependent variable, ϑ_{ij}' is a ($kx1$) vector of slope coefficients for the lagged explanatory variables; p and q indicate the number of lags, and ε_{it} is the idiosyncratic error term.

The ARDL model (1) can be rewritten in the error-correction form (Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999).[12] We consider two possible specifications depending on the assumption of heterogeneity of the short- and long-run parameters: the mean group (MG) model (2), assuming heterogeneous short- and long-run parameters (see Pesaran and Smith, 1995 and Blackburne and Frank, 2007), and the pooled mean group (PMG) model (3), assuming heterogeneous short-run, but homogenous long-run parameters (see Pesaran, Shin and Smith, 1999 and Blackburne and Frank, 2007):

$$\Delta c_{it} = \gamma_i (c_{i,t-1} + \boldsymbol{\delta}_i' \boldsymbol{x}_{it}) + \sum_{j=1}^{p-1} \eta_{ij}^* \Delta c_{i,t-j} + \sum_{j=0}^{q-1} \vartheta_{ij}^* \Delta \boldsymbol{x}_{i,t-j} + \mu_i + \varepsilon_{it} \tag{2}$$

$$\Delta c_{it} = \gamma (c_{i,t-1} + \boldsymbol{\delta}' \boldsymbol{x}_{it}) + \sum_{j=1}^{p-1} \eta_{ij}^* \Delta c_{i,t-j} + \sum_{j=0}^{q-1} \vartheta_{ij}^* \Delta \boldsymbol{x}_{i,t-j} + \mu_i + \varepsilon_{it} \tag{3}$$

where Δ is the difference operator, $\gamma_i = -(1 - \sum_{j=1}^{p} \eta_{ij})$ is the error correction speed of adjustment, $\boldsymbol{\delta}_i = \sum_{j=o}^{q} \vartheta_{ij} / (1 - \sum_{k} \eta_{ik})$ is the vector of long-run parameter estimates (long-run elasticities), while $\eta_{ij}^* = -\sum_{m=j+1}^{p} \eta_{im}$ and $\vartheta_{ij}^* = -\sum_{m=j+1}^{q} \vartheta_{im}$ are the short--run parameter estimates. Pesaran, Shin and Smith (1999) note that γ_i in equation (2) (γ in equation (3)) must be significantly different from zero in the case of an existing long-run (cointegrating) relationship between the dependent and explanatory variables. The long-run parameters, in which our interest lies, are calculated from the short-run parameters.

[12] Introducing lags of the explanatory and dependent variables is a way to deal with serial correlation. Another advantage of this reparameterization of (1) is that the short-run deviations from the long-run equilibrium relationship are modeled together with the long-run relationship (see Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999; Birkel, 2014).

Therefore, the lag specification (p and q) can affect the long-run parameters (elasticities) estimation. It is common in the literature to set $p = q = 1$; in our case, this yields an ARDL (1,1,1,1,1,1) model,[13] which is the specification we apply in the paper. An appealing feature of the ARDL approach to cointegration analysis is that the variables in the model are allowed to be stationary or integrated (Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999). Regressors can be endogenous or exogenous (Chudik et al., 2013).

Pesaran and Smith (1995) show that in the dynamic panel models an assumption of homogeneity of short- and long-run parameters can yield inconsistent estimates, whereas the MG estimator, in which for each panel group member a regression model is estimated and then a simple arithmetic average of the parameters is calculated, yields consistent estimates. The PMG estimator is a between case, in which long-run parameters are constrained to be the same across panel groups, while short-run parameters are allowed to differ (Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999; Blackburne and Frank, 2007). The PMG estimates are consistent and more efficient than the MG estimates (which are always consistent), given that the assumptions on the slope constraints are true (see Blackburne and Frank, 2007). To select between the MG and PMG model, the Hausman test can be used (Pesaran, Shin and Smith, 1999; Blackburne and Frank, 2007). Estimates of (2) and (3) are obtained by applying the Stata routine xtpmg of Blackburne and Frank (2007).

Data and the Empirical Results

Empirical models (2) and (3) are estimated on quarterly data for 12 euro area countries, including Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Slovenia, and Spain.[14] The panel is unbalanced: it starts with 1999:Q1, at the earliest, and ends for all countries with 2016:Q4. A detailed description of the data used is in Table 1.

[13] The same lag specification was used in the study most related to ours (Costa et al., 2016). We also tried to estimate ARDL models (2) and (3) with other lag specifications. Following Pesaran, Shin and Smith (1999), we first estimated for individual countries in the panel the ARDL models setting 1 as the maximum lag specification for p and q, using the ardl Stata code of Kripfganz (see Kripfganz and Schneider, 2016). Aikake information criteria were then used to obtain the optimal lag specification for each explanatory variable. The results showed great diversity regarding optimal lag structure between countries, which was expected given a relatively large amount of explanatory variables and heterogeneity between countries; no unique specification was found that would be optimal for at least a third of the countries in the sample. This is another reason we prefer ARDL (1,1,1,1,1,1) to alternative lag specifications.

[14] Other euro area countries are not included due to unavailable data, particularly for disposable income.

Table 1

Description of Variables Used

Variable	Description of the primary data	Source
Natural logarithm of real consumption (c_{it})	Household and non-profit institution serving households (NPISH) final consumption expenditure, chain-linked (real) quarterly data, seasonally and calendar adjusted.	Eurostat
Natural logarithm of real disposable income (y_{it})	Gross adjusted disposable income of households and NPISH; real values are obtained by deflating nominal values with the HICP (harmonized index of consumer prices), and quarterly data are used; finally, seasonally adjusted data are obtained by using the X-13ARIMA-SEATS method within the JDemetra+ software developed by Eurostat (see Grudkowska, 2015).	Eurostat
Natural logarithm of real net wealth (nfw_{it})	Real net financial assets (real net financial wealth) of households and NPISH defined as a difference between total financial assets and total liabilities of the sector; quarterly data are used; finally, seasonally adjusted data are obtained by using the X-13ARIMA-SEATS method within the JDemetra+ software developed by Eurostat (see Grudkowska, 2015).	Eurostat
Natural logarithm of real house prices index (hp_{it})	Real house prices index, quarterly data.	Bank for International Settlements *Source*: National sources, BIS Residential Property Price database <http://www.bis.org/statistics/pp.htm>.
Unemployment rate (u_{it})	Unemployment rate, seasonally adjusted, quarterly data.	Eurostat
Composite indicator of systemic stress ($ciss_{it}$)	Composite indicator of systemic stress in the euro area as a whole (see Holló, Kremer and Lo Duca, 2012 for detailed description), quarterly data calculated as an arithmetic average of monthly data.	European Central bank

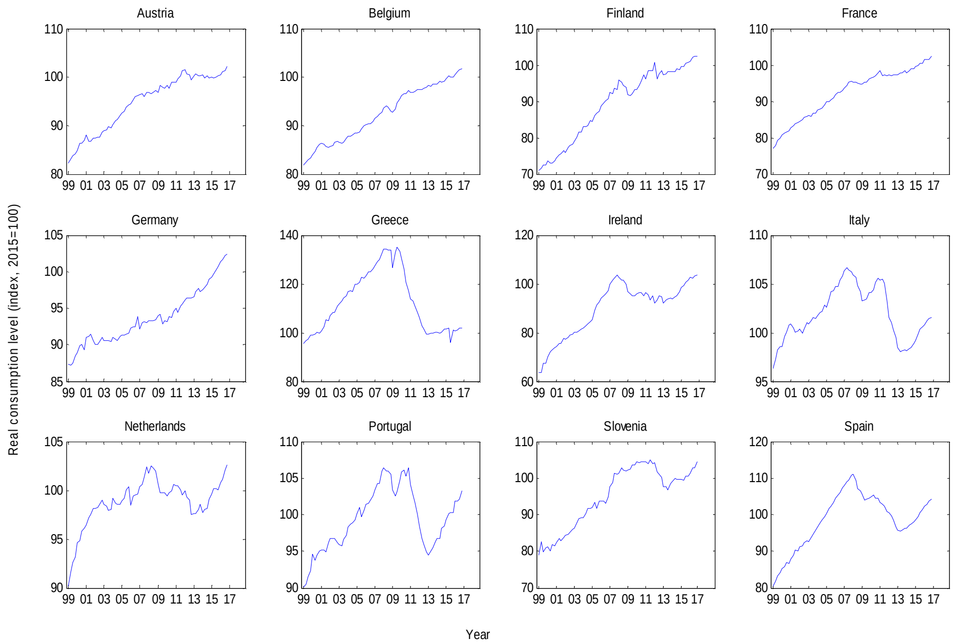
The dynamics of real consumption in individual euro area countries is presented in Figure 1. Real consumption level is presented as an index, with a reference year 2015. The figure conveys that the global financial and economic crisis and the euro area crisis were associated with a drop in real consumption in several euro area countries, the most significantly in the PIIGS countries. In seven countries (Austria, Belgium, Finland, France, Germany, Ireland, Netherlands) real consumption reached the highest level at the end of observation period, whereas in five, including 4 PIIGS countries, real consumption at the end of observation period was lower than before the global financial crisis.

The condition that the variables entering models (2) and (3) are either I(1) or (0) was checked by the IPS panel unit root test of Im et al. (2003) and the Fisher-type augmented Dickey-Fuller (ADF) test proposed by Choi (2001). The results, presented in Table 2, indicate that variables c_{it} , y_{it} , nfw_{it} , hp_{it} , and u_{it}

are integrated of an order of 1 (I(1)). It is less clear whether $ciss_{it}$ is I(1) or I(0). Regardless, the results show that the ARDL approach is valid (see Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999).

Figure 1

Real Consumption in Individual Euro Area Countries in the Period 1999:Q1 – 2016:Q4



Notes: Real (chain-linked) consumption level (expressed as index, 2015 = 100) of households and non-profit institution serving households (NPISH). The x-axis denotes years: 99 = 1999, 01 = 2001, etc.

Source: Own drawings based on Eurostat data.

In the ARDL approach to long-run analysis, it is not necessary to apply a formal panel cointegration test to ascertain empirically a long-run relationship between the dependent and explanatory variables (see Pesaran and Smith, 1995; Pesaran, Shin and Smith, 1999). Nonetheless, we apply the second-generation Westerlund's cointegration test to verify for cointegration between the logarithm of real consumption and a set of explanatory variables. We sequentially test for a cointegration relationship starting with the parsimonious cointegration test model [1] in which the logarithm of real consumption (c_{it}) is determined solely by the logarithm of disposable income (y_{it}), the logarithm of net financial wealth (nfw_{it}), and the logarithm of the real house prices index (hp_{it}). Next, the unemployment rate (u_{it}) is added to model [1], yielding cointegration test model [2]. Last, the composite indicator of systemic stress ($ciss_{it}$) is added to model [2], yielding model cointegration test model [3].

T a b l e 2

Results of the IPS Unit Root Test

Variable	IPS test		Fisher-type ADF test					
			Lag 1		Lag 2		Lag 3	
	Constant	Constant and trend	Constant	Constant and trend	Constant	Constant and trend	Constant	Constant and trend
c_{it}	0.8836 (0.8116)	2.0376 (0.9792)	1.6197 (0.9474)	2.5675 (0.9949)	1.9596 (0.9750)	2.6771 (0.9963)	0.9028 (0.8167)	1.2693 (0.8978)
y_{it}	1.0228 (0.8468)	0.1800 (0.5714)	0.6607 (0.7456)	0.0952 (0.5379)	1.9958 (0.9770)	1.7501 (0.9599)	1.8680 (0.9691)	1.6732 (0.9529)
nfw_{it}	–0.8298 (0.2033)	–0.3007 (0.3818)	0.3271 (0.6282)	–0.0255 (0.4898)	–0.5673 (0.2853)	–2.3523 (0.0093)	0.7287 (0.7669)	–0.1349 (0.4463)
hp_{it}	1.7076 (0.9561)	0.5281 (0.7013)	3.1533 (0.9992)	1.2727 (0.8984)	3.0001 (0.9987)	1.6165 (0.9470)	1.7052 (0.9559)	0.9668 (0.8332)
u_{it}	–0.4685 (0.3197)	1.9851 (0.9764)	1.1858 (0.8821)	3.9178 (1.0000)	0.4071 (0.6580)	2.7296 (0.9968)	–0.3839 (0.3505)	1.3196 (0.9065)
$ciss_{it}$	–4.0456 (0.0000)	–1.0352 (0.1503)	–4.3005 (0.0000)	–1.1216 (0.1310)	–3.0625 (0.0011)	0.2587 (0.6021)	–3.6071 (0.0002)	–0.3162 (0.3759)

Notes: The IPS test is based on the ADF regression for each panel group. The null hypothesis of the IPS and the Fisher-type ADF test is that for all the panel groups the process is a unit root. The rejection of the null implies that the series is stationary. To control for serial correlation in errors, lags are allowed (maximum of 3) in the IPS test and are determined by the Akaike information criteria. The W_{t-bar} test statistic is reported (averaged t-statistics of the ADF regression for each panel group) along with the corresponding p-value. In the Fisher-type unit root test, the ADF test is also performed for each panel group and then the p-values are combined in calculation of the overall test statistics (see Choi (2001) and Stata xtunitroot documentation). The p-values for the inverse normal Z statistics are reported (see Choi (2001) for details). To control for cross-section dependence, the variables in all tests were demeaned (except the variable $ciss_{it}$ which is common to all panel groups) before the tests were performed. Stata in-built routines were used for calculations.

Source: Author's computations.

T a b l e 3

Results of the Westerlund's (2007) Cointegration Test

Specification: variables in the cointegration relationship	Test statistics and the robust significance level			
	P_τ	P_α	G_τ	G_α
[1]: c_{it} , y_{it} , nfw_{it} , hp_{it}	–7.930 (0.000)	–11.243 (0.000)	–2.404 (0.000)	–12.794 (0.000)
[2]: c_{it} , y_{it} , nfw_{it} , hp_{it} , u_{it}	–9.322 (0.015)	–12.475 (0.015)	–2.628 (0.060)	–14.376 (0.010)
[3]*: c_{it} , y_{it} , nfw_{it} hp_{it} , u_{it} , $ciss_{it}$	–11.045 (0.005)	–16.474 (0.000)	–3.115 (0.020)	–18.842 (0.000)
[3]**: c_{it} , y_{it} , nfw_{it} , hp_{it} , u_{it} , $ciss_{it}$	–8.108 (0.095)	–12.400 (0.025)	–2.456 (0.140)	–14.487 (0.015)

Notes: For the specifics of the test, see Westerlund (2007). The optimal lag length in the error-correction specification of the test is determined automatically by the Akaike information criteria (whereby the maximum lag length is set to 1 due to time dimension limitations of the panel sample). We report the test statistics and the significance levels (in brackets) for rejection of the null hypothesis of no cointegration for the panel as a whole (P_τ and P_α) and the mean-group statistics (G_τ and G_α). The rejection of the first leads to the conclusion of a cointegration relationship between the dependent and explanatory variables for the panel as a whole, the second to a cointegration relationship for at least one member of the panel group (country in the panel). τ denotes standard errors and α the Newey-West standard errors in the error-correction specification. Bootstrapping (200 replications as Westerlund (2007) suggested) was performed to obtain p-values robust to cross-sectional dependence. Westelund (2007) argues that the P_α and G_α statistics should be more powerful than the P_τ and G_τ statistics when the time dimension exceeds substantially the cross-section dimension of the panel. The xtwest Stata routine of Persyn and Westerlund (2008) was used for computations. * The lag was set to 0 because the length of the series for Greece was too short; ** Greece was dropped from the sample and then the cointegration test was performed for 11 countries only.

Source: Author's computations.

The results (presented in Table 3) indicate a stable long-run (cointegrating) relationship between c_{it} and explanatory variables y_{it} , nfw_{it} , hp_{it} , and u_{it} (models [1] and [2]) at any conventional significance level. When $ciss_{it}$ is added to the cointegration relationship (model [3]), the null of no cointegration for the panel as a whole can be rejected at the 2.5% significance level. Assuming that the time dimension sufficiently exceeds the cross-section dimension of the panel sample (see notes to Table 3 and Westerlund (2007)), the contested cointegration between the natural logarithm of real consumption and all explanatory variables of models (2) and (3) can be confirmed.

A long-run relationship between consumption and all the explanatory variables is also supported by the significant negative error-correction speed of adjustment term of estimated models (2) and (3), γ_i / N and γ, respectively, as presented in Table 4. The obtained values indicate that the restoration of the long-run equilibrium after a shock is relatively fast for the MG model (about 47% in one quarter) and somewhat longer for the PMG model (about 24%). Along with the error-correction speed of adjustment term, the table lists the estimates of the long-run slope coefficients (long-run elasticities). It is apparent that the MG and PMG models yield different estimates. The Hausman test significantly rejects the null hypothesis of no significant differences in the long-run parameter estimates between the MG and PMG models. The MG model, the consistent model under the alternative hypothesis, is thus statistically more favorable.[15]

The results of the MG model show that only the long-run elasticities of consumption with respect to real adjusted disposable income (y_{it}) and real net financial wealth (nfw_{it}) are significant, while for the real house prices long-run elasticity is insignificant. A 1% increase in real adjusted disposable income (real net financial wealth) is associated with a 0.49% (0.12%) increase in real consumption. The estimated long-run parameter for house price index (hp_{it}) has the expected sign and is smaller in absolute terms than the estimated parameter for real net financial wealth, but statistically insignificant. The result is consistent with the extant literature (e.g., Sinai and Souleles, 2005; Case, Quigley and Shiller, 2005; Skudelny, 2009; Jawadi and Sousa, 2012; Costa et al., 2016) arguing that consumers respond differently to increases in housing wealth than to increases in financial wealth. The finding implies that asset wealth is not homogenous from the perspective of its effect on consumption decisions. We found that euro area consumers seem to consume more in response to an increase in their net financial wealth while this cannot be ascertained for an increase in their

[15] The Hausman test thus indicates that the euro area is a heterogenous group of countries; treating the group of countries as homogenous and estimating the long-run elasticities with pooled models may lead to inconsistent estimates.

housing wealth, which supports the findings of, for example, Skudelny (2009) and Jawadi and Sousa (2012), but not the findings of Costa et al. (2016), who reported the elasticity of real consumption with respect to house prices to be significant and greater than with respect to stock market prices. The slope coefficient for the unemployment rate (hp_{it}) and the composite indicator of financial stress ($ciss_{it}$) are statistically insignificant, implying that income uncertainty and capital uncertainty are not important determinants of real consumption in the euro area in the long run.

Table 4

The Results of the ARDL Model of Consumption: Euro Area as a Whole

Estimates of the long-run slope coefficients	**MG model (model 2)**	**PMG model (model 3)**
Error-correction term ($\frac{\gamma_i}{N}; \gamma$)	–.4735346*** (.0784032)	–.2422273*** (.0373452)
y_{it}	.4931356*** (.1492612)	.6303632*** (.052808)
nfw_{it}	.1235101*** (.0187619)	.0667333*** (.0151659)
hp_{it}	.0654401 (.0571952)	.017796 (.0211057)
u_{it}	–.002595 (.0017193)	–.002834*** (.0007023)
$ciss_{it}$	.0043858 (.0188634)	–.0325956 (.0118364)
Log-likelihood	1955.148	1887.407
Hausman test	17.60 (0.0035)	
No. of observations	490	

Notes: Models (2), i.e., MG, and (3), i.e., PMG, estimates for ARDL (1,1,1,1,1,1) are presented. The xtpmg Stata routine of Blackburne and Frank (2007) was used for computations. Only long-run slope coefficients are presented. Short-run estimates are obtainable from the author. Standard errors of the estimates are given in parentheses. ***/**/* – a 1%/5%/10% significance level. The (chi-square) test statistics of the Hausman test are presented; the significance level for rejection of the null hypothesis of no significant difference in the parameter estimated between MG and PMG models is noted in parentheses.

Source: Author's computations.

While the above results indicate that the sampled euro area countries cannot be treated as a homogeenous group, it is worth investigating whether more homogenous groups can be identified. It has become common in the recent empirical literature (e.g., IMF, 2012; Shambaugh, 2012; Ciccarelli, Maddaloni and Peydró, 2013; 2015) to segregate the euro area into two groups – a group of countries that was strongly negatively affected and a group of countries that was more moderately negatively affected by the global financial and the euro area crisis. In continuation, we follow this literature and divide the sampled euro area countries into the PIIGS (Portugal, Italy, Ireland, Greece, and Spain) and non--PIIGS (Austria, Belgium, Finland, France, Germany, and the Netherlands) euro areas. The results of models (2) and (3) estimated for the two groups of countries are presented in Table 5.

T a b l e 5

The Results of the ARDL Model of Consumption: PIIGS and Non-PIIGS Countries of the Euro Area

Estimates of the long-run slope coefficients	**PIIGS countries**		**Non-PIIGS countries**	
	MG (model 2)	PMG (model 3)	MG (model 2)	PMG (model 3)
Error-correction term ($\frac{\gamma_i}{N};\gamma$)	–.5978693*** (.1625404)	–.2422273*** (.0373452)	–.3847241*** (.0598647)	–.1676884*** (.0556117)
y_{it}	.414552** (.1819981)	.5786572*** (.0667253)	.5492667** (.2301871)	.7885385*** (.0746031)
nfw_{it}	.113997*** (.0202673)	.0902282*** (.0268099)	.1303052*** (.0298704)	.1657986*** (.0265687)
hp_{it}	.1265073 (.0844359)	.0478121 (.0339599)	.0218206 (.0782634)	.0007093 (.0180107)
u_{it}	–.0028404* (.0015419)	–.002232** (.0009928)	–.0024196 (.0028542)	.0056206 (.0015522)
$ciss_{it}$	.0217692 (.0454988)	–.035623* (.019792)	–.0080309 (.0084181)	–.0093023 (.0125485)
Log-likelihood	634.0975	605.3304	1321.05	1293.192
Hausman test	0.37 (0.9960)		7.80 (0.1677)	
No. of observations	161		329	

Notes: Models (2), i.e., MG, and (3), i.e., PMG, estimates for ARDL (1,1,1,1,1,1) are presented. The xtpmg Stata routine of Blackburne and Frank was used for computations. Only long-run slope coefficients are presented. Short-run estimates can be obtained from the author. Standard errors of the estimates are given in parentheses. ***/**/* – a 1%/5%/10% significance level. The (chi-square) test statistics of the Hausman test are presented; the significance level for rejection of the null hypothesis of no significant difference in the slope coefficients between MG and PMG estimates is noted in parentheses.

Source: Author's computations.

The Hausman test now shows that the estimates from the PMG model are not only more efficient than those from the MG model, but are also consistent, implying two relatively homogenous groups in the euro area. The long-run elasticity of real consumption with respect to real adjusted disposable income is higher in the non-PIIGS than in the PIIGS euro area. A similar can be noted for the long-run elasticity of real consumption with respect to real net financial wealth. The long--run elasticity of real consumption with respect to real house prices is insignificant in both parts of the euro area. An important distinction between the PIIGS and non-PIIGS countries can be detected regarding the long-run elasticity of consumption with respect to the unemployment rate and the composite indicator of financial stress. In the PIIGS countries, an increase in the unemployment rate (increase in uncertainty about future income) is significantly negatively related to real consumption: A one percentage point increase in the unemployment rate is associated with a 0.002% reduction in real consumption.[16] The association between the variables for the non-PIIGS euro area is positive but insignificant. Uncertainty in the return on accumulated financial wealth is an important long-run determinant of consumption only in the PIIGS euro area. A one-point increase in

[16] Unemployment enters models (2) and (3) in percentages.

the composite indicator of systemic stress is associated with a 0.036% drop in real consumption. A negative elasticity implies that the income effect of an increase in capital uncertainty dominates the substitution effect. A similar finding was obtained by Choudhry (2003) and Ibrahim and Law (2013) for a narrower measure of capital uncertainty and a sample of other countries.

The results have policy-relevant implications as they show that an increase in uncertainty affects consumer decisions not only in the short run (temporarily), as demonstrated by Alexopoulos and Cohen (2009), Basu and Bundick (2012), Gudmundsson and Natvik (2012) and Bonciani and van Roye (2016), but also in the long run. Our results show an increase in income and capital uncertainty is negatively associated with consumption in the long run in the PIIGS euro area countries where the stress of the global and the euro area crises was more protracted than in other part of euro area. The heterogeneity of uncertainty transmission in the PIIGS and non-PIIGS euro area indicates that common policy measures (undertaken for instance by the European central bank) may not be effective in alleviating uncertainty across the euro area. National policy measures (e.g. macroprudential policy, fiscal policy and structural reforms) that reassure markets may be necessary, especially when surges in uncertainty are not temporary.

Conclusion

The results of this paper demonstrate that for the investigated time period of 1999:Q1 – 2016:Q4 a long-run (cointegrating) relationship between real consumption, real disposable income, real net financial wealth, real house prices, and unemployment rate and the composite indicator of systemic stress in the euro area for a panel of 12 euro area countries can be identified. The euro area cannot be treated as a homogenous group: Long-run elasticities of consumption with respect to real disposable income and real net financial wealth are larger in the non-PIIGS euro area. Real house prices are not significant determinants of real consumption in the long run. Our results show that the unemployment rate and the composite indicator of systemic stress, used as proxies for income and capital uncertainty, are significantly negatively associated with real consumption in the long run, but only in the PIIGS euro area.

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