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THE EFFECTS OF FINANCIAL AND REAL WEALTH ON CONSUMPTION: NEW EVIDENCE FROM OECD COUNTRIES

by Riccardo De Bonis* and Andrea Silvestrini*

Abstract

In this paper we present new estimates of the effect of households' financial and real wealth on consumption. The analysis makes reference to eleven OECD countries and takes into account quarterly data from 1997 to 2008. Unlike most of the previous literature on European countries, we measure financial wealth using quarterly harmonized data on households' financial assets and liabilities, which have been gleaned from the flow of funds. For comparison, we also employ national share price indices as a proxy for financial wealth. We rely on 1) standard static panel and 2) single-country level autoregressive distributed lag estimations. Furthermore, we implement a recent econometric approach that allows for more flexible assumptions in the non-stationary panel framework under consideration. Our results show that both net financial wealth and real wealth have a positive effect on consumption. Overall, the influence of net financial assets is stronger than that of real assets.

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Keywords: consumption, household financial and real wealth, wealth effects, panel cointegration.

Contents

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* Bank of Italy, Economics, Research and International Relations.

1 Introduction[1]

The wealth effect has traditionally received considerable attention in the macroeconomic literature. The wide fluctuations in financial wealth and house prices in most industrialised countries over the last decade sparked new interest on this topic. In the US, the performance of the stock market was exceptional between 1995 and 2000, the years of the Internet bubble. The increase in household financial wealth contributed to the drop of the saving rate and to the support of consumer spending. The burst of the stock market bubble started at the beginning of 2000 and led to a decline in share prices until the first months of 2003. Then household consumption was mainly sustained by the increase in the value of houses, which began in the late Nineties but intensified after 2000.

On the contrary, in 2007 the financial turbulence interconnected with the start of falling house prices. The decline in house prices was bigger than the drop in 1932, at the worst point of the Great Depression. House prices downturns took place in other industrialised countries that had experienced a rapid rise in previous years. With regard to financial wealth, the crisis implied that in many countries the stock market capitalization was at the beginning of 2009 similar to the 1995 levels, when the great phase of surging prices began.

The goal of this paper is to provide new evidence on the link between financial wealth, real wealth and household consumption in a sample of OECD countries between 1997 and 2008. The novelty of the analysis is to rely on quarterly harmonized statistics on household financial assets, which have been gathered from the European flow of funds. Thus, we get a more accurate measure of household wealth compared to previous contributions in the literature. For sensitivity analysis, we also employ national stock

[1]We would like to thank an anonymous referee for helpful comments that helped improving the paper over an earlier version. For offering insights into this work, special thanks go to Piero Catte, Marco Magnani, Giovanni Mastrobuoni, Andrea Mercatanti and Franco Peracchi. This paper also benefited from comments made by participants at the 57th International Statistical Institute Conference, Durban, South Africa (16-22 August 2009) and at the 50th Riunione Scientifica Annuale della Società Italiana degli Economisti, Rome, (22-24 October 2009). The paper is the responsibility of its authors and the opinions expressed here do not necessarily reflect those of the Bank of Italy or the Eurosystem. Final version forthcoming in: Applied Financial Economics.

market indices. We implement a recent econometric approach that allows for more flexible assumptions in the non-stationary panel framework under consideration. Our results show that both net financial wealth - defined as the difference between financial assets and liabilities - and real wealth have a positive effect on consumption. Overall, the influence of net financial assets is stronger than that of real assets.

After this introduction, this paper is divided into six sections. Section 2 reviews the literature. Section 3 describes the data properties. Section 4 studies the integration and cointegration features of the variables involved. Section 5 presents the econometric specification and estimation strategy. Section 6 reports and discusses the empirical results. Lastly, Section 7 gives a summary and draws conclusions.

2 The literature on wealth effects and consumption

The empirical literature on the wealth effects is extensive and dates back to the early Sixties. It is difficult to summarise this literature because it refers to different periods, it employs different estimation techniques and uses different sources for real and financial wealth. In this Section we survey the most recent and important contributions, focusing on papers that used aggregate time series data.[2]

Most of the papers looked at the US experience. Poterba (2000) found that the marginal propensity to consume out of the stock market is between 0.01 and 0.05. Using the American flow of funds, Morris and Palumbo (2001) estimated a consumption response in the range of 3 to 6 cents-to-the dollar, according to different model specifications. Lettau and Ludvingson (2004) distinguished between cycle and trend changes in asset values. Consistent with previous studies, they found that in the US the marginal impact on consumption of a dollar increased in wealth is about 5 cents.

While in the past literature it was uncommon to distinguish between real and financial wealth, recently this subject received a careful attention. Carrol, Otsuka and Slacalek (2011) estimated that in the US the housing wealth effect is around 9 cents against 4 cents of the stock market wealth effect. Also according to Case, Quigley, and

[2]See Paiella (2007) and ECB (2009) for surveys which include the household-level evidence on wealth effects. See Guiso, Paiella and Visco (2005) for an analysis of the wealth effects using micro data on Italian households.

Shiller (2005) the housing wealth effect is larger than the stock market effect.

Focusing on other national cases, Blake (2004) studied the impact on consumption and retirement behaviour of various components of wealth in the UK. He found that wealth has a direct net positive effect on consumption, with a marginal propensity of about 0.01. Tang (2006) claimed that in Australia a permanent dollar increase in housing wealth leads to a six percent rise in consumption, three times the effect of financial wealth. Bassanetti and Zollino (2010) reached a different result analysing the Italian case: the size of the marginal propensity to consume out of housing wealth is about 1.5-2 cents, against values of 4-6 cents for the propensity to consume out of each euro increase in financial wealth. Similarly, studying Germany, Hamburg, Hoffmann and Keller (2008) assessed that a one-euro increase in asset wealth causes an increase in consumption by 4-5 cents.

Turning to papers that analysed panel of countries, Dreger and Reimers (2006) found a total wealth elasticity of consumption in a range of 3-5 percent for a group of EU countries. Byrne and Davis (2003) analysed the effect of disaggregated financial wealth on consumption functions for G7 countries, while Salotti (2010) focused on the link between wealth and household savings. Very recently, a set of papers tried to measure the effect of both financial and real assets on consumption. Dreger and Reimers (2009) found a consumption elasticity with respect to house prices of 2.5 percent and a consumption elasticity with respect to financial assets of 3 percent. For a panel of 16 major industrial countries, Slacalek (2009) estimated a long-run marginal propensity to consume out of total wealth averaged across countries of around 5 cents. There is some evidence that the housing wealth effect is smaller than the financial wealth effect, but the opposite seems to be true in the US and the UK, thus supporting the results of Carrol, Otsuka and Slacalek. The main explanation of these differences among countries is that in the UK and the US financial innovation is more sophisticated, letting people more easily to take cash out of their homes. For example, mortgage equity withdrawal is more common in the US and the UK than in the euro area. Also Catte *et al.* (2004) underlined the importance of the structural characteristics of housing and mortgage markets. Studying ten OECD countries, these authors claimed that the marginal propensity to consume out of housing wealth is on average stronger than that out of financial wealth in Australia, Canada, the Netherlands, the UK and the US. These countries have large and efficient mortgage markets, with a specific reference to the opportunities for housing equity withdrawal.

Labhard, Sterne and Young (2005, LSY thereafter) criticised the literature on the cross-country differences of the marginal propensity to consume out of wealth. According to LSY, these differences are misleading because they are not rooted in explainable structural differences among countries and are, on the contrary, attributable to data deficiencies. Using dynamic panel techniques, LSY found that the hypothesis of the long-run marginal propensity to consume out of financial wealth being the same across countries could not be rejected and estimated for this variable a value a little greater than 6 percent.

In summary, three issues are common to most of the contributions. First, a consensus emerged on the need to distinguish between real and financial assets when estimating the wealth effect on consumption: in the past literature aggregate household wealth was often used as determinant of consumption; in other cases the authors only employed financial assets. Second, previous empirical studies found that housing wealth became more important than financial assets in influencing consumption, but this effect is often different if we distinguish between Anglo-Saxon countries, where financial innovation and mortgage equity withdrawal were impetuous, and other European nations, where more traditional banking arrangements prevailed. Third, in the euro area financial assets often do not influence household consumption because financial deepening is smaller than in the United States or the United Kingdom and bank deposits remain important in household portfolio. These issues will be at the core of the empirical analysis presented in the following sections.

3 Descriptive data analysis

Our analysis is based on data assembled for a period ranging from 1997Q4 to 2008Q1, for eleven OECD countries: Austria, Belgium, Finland, France, Germany, Italy, Netherlands, Portugal, Spain, the United Kingdom and the United States. For these countries we have collected quarterly data on household consumption expenditure, income, financial wealth, household debt, real wealth and stock market indices. Below we provide a brief description of the data set.

Quarterly data on household private consumption are taken from the OECD's Quarterly National Accounts database. Quarterly data on personal disposable income are the same employed by Dreger and Reimers (2009).[3]

Aggregate financial wealth is obtained from the quarterly flow of funds data, which are constructed within a unified statistical framework (similarly, Case, Quigley, and Shiller, 2011, gather an estimate of financial wealth for the US from the Federal Reserve flow of funds). Specifically, financial assets include the four main instruments of household financial saving: deposits, securities other than shares, quoted shares and mutual funds units, and insurance technical reserves. We do not include unquoted shares in our measure of household financial wealth since the estimate of this item is not completely harmonized across countries. Net financial wealth is defined as the difference between total financial assets and financial liabilities (debt). To ensure robustness of our estimates, we also employ country-specific equity indices, which are a proxy for the stock market performance and hence for the valuation of household financial wealth. Stock market indices for Europe and the US are provided by Morgan Stanley Capital International (MSCI Indices) and cover a large part of the market capitalization. All these data are available, at quarterly frequency, through Datastream.

Household real wealth (or household real assets) refers to dwellings; because of data limitations, for most of the countries the other components of real wealth are not available. While some countries provide the quarterly value of household dwellings, in most of the cases only annual data are available: hence, when necessary, quarterly series are obtained by temporal disaggregation. The source of the real wealth data is mainly the OECD's Households Assets database. Yet, for a number of countries, household real wealth estimates are made available from the National Central Banks.[4] We stress that the measurement of household real wealth is not harmonized. Therefore, the estimates of the marginal propensity to consume out of real wealth should be interpreted with some caution.

[3]These data are taken from the World Market Monitor provided by Global Insight.

[4]We are grateful to Michael Andreasch (Austrian Nationalbank), Matti Okko (Bank of Finland), Francesco Zollino (Bank of Italy), and Pedro Abad Fernandez-Canaveral (Bank of Spain) for providing us with estimates of real wealth data. For Portugal, we refer to Cardoso, Farinha and Lameira (2008).

Private consumption, financial and real wealth data (expressed as a ratio to quarterly household disposable income) are depicted in Figure 1, which shows the evolution of the three time series separately, for each country in the panel. For some countries there appears to be a clear positive relation between wealth (real and financial) and consumption: this is the case, for instance, of Belgium, Italy, United States and United Kingdom.

[SEE FIGURE 1]

Furthermore, some useful information concerning the wealth to income ratios can be inferred from the plots. Let us focus, for instance, on Italy: the real wealth to income ratio roughly ranges from 12 to 22, whether the net financial wealth to income ratio is comprised between 8 and 12. Three important remarks are at order. First, as already explained, the stock of net financial wealth is defined as the sum of all financial assets minus liabilities, excluding unquoted shares. Second, by real wealth we mean the value of the stock of dwellings. Third, net financial wealth and real wealth are annual stock values recorded each quarter. Yet, consumption and income are flow variables, available quarterly in our database. This means that, to be meaningful and fully comparable with other empirical studies on the wealth effects, the previous values must be divided by four. As a consequence, in Italy, the real wealth to income ratio ranges from 3 to 5.5, whether the net financial wealth to income ratio from 2 to 3.

[SEE FIGURE 2]

Figure 2 plots the evolution of the stock market indices. In all graphs (except Austria), the high points of the dot-com bubble can be easily recognized in early 2000. At the same time, it can be observed a dramatic drop of the stock market indices at the end of 2000. Another drop is evident at the end of the sample (i.e., last quarter of 2007 and first quarter of 2008), clearly dating the acute period of the recent financial crisis. As a consequence, there might have been structure changes in the marginal propensity to consume in the long-run relationship between consumption and wealth: in other words, it is very likely that the marginal propensity to consume changed between an higher wealth regime up to March 2000, reflecting the high point of the dot-com bubble, and a lower wealth regime in the following years. This conjecture can also be tested statistically in the framework of panel cointegration, as we shall show in Section 4.

4 Testing for unit roots and cointegration

To identify long-run equilibrium relationships between consumption and wealth, cointegration tests may only be performed on panels that are known to be non–stationary. In this Section we implement panel unit root tests to check whether the variables under study contain zero frequency unit roots in their data generating process. It is widely acknowledged in the literature that panel unit root tests have higher power than unit root tests based on individual time series. We shall carry out a complete battery of unit root tests, with different model specifications. Such a check is important because it is well known that unit root analysis is sensitive to the choice of the model specification.

[SEE TABLE 1]

Table 1 reports the outcomes of several panel unit root tests for our variables of interest (consumption-income ratio, net financial wealth-income ratio, real wealth-income ratio, and the logarithm of stock market index): namely, the tests proposed by Levin, Lin and Chu (2002); Im, Pesaran and Shin (2003); Maddala and Wu (1999); Hadri and Larsson (2005). Levin, Lin and Chu (2002), Im, Pesaran and Shin (2003) and Maddala and Wu (1999) are commonly used unit root tests which basically combine, in different ways, individual unit root tests applied on each time series in the panel. These are all tests for the null of a unit root in the panel. Hadri and Larsson (2005) is instead a KPSS-type test. Therefore, the null hypothesis is that all time series in the panel are stationary against the alternative of a unit root. All the technical details are skipped and the interested reader is referred to the references above for additional explanations and a rigorous treatment.

The LLC, IPS, MW test results imply not rejection of the presence of a unit root for the series in the panel, assuming a constant and a time trend in the test regression. This holds true for consumption over income, net financial wealth over income, real wealth over income and the stock market index, at 5% significance level. HL test strongly rejects the null hypothesis of stationarity for the four variables, as well. Results are clear-cut and point to non-stationarity, which has to be taken into account at the modelling stage. These findings enable us to proceed with cointegration analysis.

To identify stable long-run equilibrium relationships among consumption-income ratio and wealth-income ratios or consumption-income ratio and the stock market index, we turn to the issue of panel cointegration. When the time dimension is relatively large with respect to traditional empirical studies based on panel data, a panel cointegration approach is useful since it allows for a more flexible modelling of heterogeneity within the panel (comparing to simple fixed or random-effects models). Moreover, panel cointegration can improve upon small samples limitations of conventional non-stationary methods (Pedroni, 2000).

Like in standard time series, in the panel setting there are different ways to test the null hypothesis of no cointegration. One possibility is to use residual based tests as suggested by Kao (1999), that extends the original Engle-Granger framework to account for panel data. In a nutshell, this approach requires first to estimate by pooled OLS to obtain the residuals, then to implement a pooled Dickey-Fuller regression. This test is based on the idea of deciding whether or not the error process of the estimated regression equation is stationary. If homogeneity and strict exogeneity assumptions hold, this residual based panel test for the null of no cointegration has the same asymptotic distribution as standard panel unit root tests.

[SEE TABLE 2]

Table 2 displays results of the residual panel cointegration test of Kao (1999). Two types of cointegration tests in panel data are presented: the Dickey-Fuller (DF) and the augmented Dickey-Fuller (ADF); in both cases the null hypothesis is no cointegration. The top part of the table refers to cointegration between consumption and financial and real wealth. The test statistics derived by Kao, on the basis of asymptotic results, imply in all cases rejection of the null of no cointegration. The bottom part of the table refers to cointegration between consumption and the stock market index (log). Similarly, test results point to strong rejection of the null hypothesis of no cointegration.

There is another issue concerning cointegration, especially in panels with a relatively large time dimension: if a structural break occurs in the cointegrating relation, this may lead to deceptive inference due a misspecified long-run relationship and misleading results in the cointegration tests. We know that in our sample, ranging from 1997Q4 to 2008Q1, there have been several financial distresses, namely, the 1997 Asian crisis, the stock market crash of 2000-2003 and the 2007 subprime crisis. In order to examine this conjecture, we use the panel cointegration test with structural changes developed by Westerlund (2006), who extends the panel LM cointegration test proposed by McCoskey and Kao (1998) to the case of multiple structural breaks in both the level and trend of a cointegrated panel regression. Using sequential limit arguments, this author shows that the test has a limiting Gaussian distribution which is free of nuisance parameters under the null hypothesis of cointegration. This limiting distribution is invariant with respect to the number and locations of break-dates and it is not necessary to compute different critical values for all possible patterns of break points. These latter are determined endogenously from the data. Additional details can be found in Westerlund (2006).[5] To identify the break-dates, we do not consider observations too close to the beginning or end of the sample (setting to 15 the trimming parameter and, consequently, not considering as possible break-date candidates all the observations lying in the 0–15 and 85–100 percent interior of the sample period). The maximum number of estimated break points is 3.

[SEE TABLE 3]

Table 3 reports the estimated break-points together with the fully modified OLS (FMOLS) based test statistics. For all countries, a cointegrated regression with at least one shift in the level is estimated. The top part of Table 3 refers to cointegration between consumption and financial and real wealth, while the bottom part to cointegration between consumption and the stock market index. The test provides evidence that there is a cointegrating relationship between consumption and real and financial wealth; when applied to consumption and the stock market index, the test brings up the same conclu-

[5] We are grateful to Joakim Westerlund for providing us with the GAUSS code to implement the panel LM test for the null of cointegration with multiple breaks.

sion. In general, there is a preponderance of breaks estimated between 2000 and 2003, when the stock market crash occurred. This is largely expected and is consistent with the evolution of the series (see Figures 1 and 2). Some break-dates are also identified in 2005, maybe reflecting the upward trend of the stock market that has emerged after 2003. At least one break-date is estimated for every country (except Austria, Netherlands and UK).

In summary, empirical evidence suggests that the variables in the panel are non-stationary. Furthermore, conventional tests and multiple breaks tests point to the existence of cointegration (with some regime shifts). These findings are relevant for a proper econometric specification and estimation strategy, as we shall describe in the following section.

5 The econometric methodology

Most macroeconomic theories of wealth effects are formulated according to a reduced-form consumption equation of the type

$$c_{i,t} = \alpha y_{i,t} + \beta^{FW} fw_{i,t} + \beta^{RW} rw_{i,t} + \varepsilon_{i,t}, \quad i = 1, 2, \ldots, N, \quad t = 1, 2, \ldots, T, \tag{1}$$

where $c_{i,t}$ represents private consumption, $fw_{i,t}$ financial wealth, $rw_{i,t}$ real wealth, $y_{i,t}$ disposable income, all expressed in levels. Equation (1) has been often suggested in the empirical literature and can be derived in the framework of the Permanent Income Hypothesis (PIH) developed by Friedman (1957), by making proper assumptions about the expected evolution of consumption. Equation (1) is also compatible with a steady-state form of the Life Cycle Hypothesis (LCH) exposited by Modigliani (1975).

A ratio specification has been often suggested as a possible transformation[6] of (1), meaning

$$\frac{c_{i,t}}{y_{i,t}} = \tilde{\alpha} + \tilde{\beta}^{FW} \frac{fw_{i,t}}{y_{i,t}} + \tilde{\beta}^{RW} \frac{rw_{i,t}}{y_{i,t}} + \varepsilon_{i,t}.$$

Or, with a different notation

$$C_{i,t} = \tilde{\alpha} + \tilde{\beta}^{FW} FW_{i,t} + \tilde{\beta}^{RW} RW_{i,t} + \varepsilon_{i,t}, \tag{2}$$

[6]Other transformations are possible. For instance, a log-linear specification was often suggested, if the interest centers on estimating elasticities of consumption with respect to wealth and income.

where $C_{i,t}$, $Wf_{i,t}$ and $Wr_{i,t}$ are all expressed as a ratio to income. Note that, in (2), $\tilde{\alpha}$, $\tilde{\beta}^{FW}$, $\tilde{\beta}^{RW}$ may be interpreted as marginal propensities to consume out of income, financial wealth and real wealth, respectively. In this paper we shall focus mainly on specification (2), which has been tested in recent empirical research.

An important issue is to choose the dynamic structure of the relationship between consumption and wealth. For empirical purposes, influential literature estimates autoregressive distributed lag (ARDL) models of consumption on income, introducing lag mechanisms to model the response of consumption to changes in income. Assume, for instance, that

$$c_{i,t} = \alpha y_{i,t} + \beta w_{i,t},$$

where $w_{i,t}$ is end of period private total wealth in levels and α, β are coefficients to be estimated. Assuming that dividends, interest and capital gains are compounded in income, the law of motion of the stock of wealth can be expressed as

$$w_{i,t} = w_{i,t-1} + y_{i,t-1} - c_{i,t-1}.$$

Simply re-arranging the last two equations, we get

$$c_{i,t} = \alpha y_{i,t} + (\beta - \alpha) y_{i,t-1} + (1 - \beta) c_{i,t-1},$$

which is an ARDL model of consumption on income.

In the paper, the specification in (2) will be given an ARDL structure. The model in (2), indeed, can be easily generalized introducing deterministic terms, an autoregressive lag polynomial for the dependent variable and complicated distributed lag schemes for the explanatory variables

$$C_{i,t} = \alpha_0 + \alpha_1 t + \sum_{j=1}^{p} \lambda_{i,j} C_{i,t-j} + \sum_{j=0}^{q} \boldsymbol{\beta}'_{i,j} \boldsymbol{W}_{i,t-j} + \varepsilon_{i,t}, \tag{3}$$

where $\boldsymbol{\beta}_{i,j} = (\tilde{\beta}^{FW}_{i,j},\ \tilde{\beta}^{RW}_{i,j})'$ and $\boldsymbol{W}_{i,t} = (FW_{i,t},\ RW_{i,t})'$, by definition. If the variables in (3) are integrated of order one, as recognized by Pesaran and Shin (1997), the traditional ARDL approach is no longer applicable (especially when working with large T, large N panels). Since, in Section 4, variables have been found to be difference stationary and cointegrated, this issue is particularly relevant for the case under study. Therefore, the stationary ARDL model in (3) has to be somehow re-parameterized to take care of the possible long-run relations among the variables.

To this aim, Pesaran, Shin and Smith (1999) show that (3) can be conveniently re-expressed as

$$\Delta C_{i,t} = \alpha_0 + \alpha_1 t + \phi_i \left(C_{i,t-1} - \boldsymbol{\theta}_i^{\prime} \boldsymbol{W}_{i,t} \right) + \sum_{j=1}^{p-1} \lambda_{i,j}^{*} \Delta C_{i,t-j} + \sum_{j=0}^{q-1} \boldsymbol{\beta}_{i,j}^{*\prime} \Delta \boldsymbol{W}_{i,t-j} + \varepsilon_{i,t}, \quad (4)$$

where $\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{i,j})$, $\boldsymbol{\theta}_i = \sum_{j=0}^{q} \boldsymbol{\beta}_{i,j} / (1 - \sum_{j=1}^{p} \lambda_{i,j})$, $\lambda_{i,j}^{*} = -\sum_{m=j+1}^{p} \lambda_{i,m}, j = 1, \ldots, p-1$ and $\boldsymbol{\beta}_{i,j}^{*} = -\sum_{m=j+1}^{q} \boldsymbol{\beta}_{i,m}, j = 1, \ldots, q-1$. Equation (4) represents the error correction re-parameterization of an ARDL model in the framework of dynamic single-equation regressions. Note that $\lambda_{i,j}^{*}$ and $\boldsymbol{\beta}_{i,j}^{*\prime}$ $(i = 1, 2, \ldots, N)$ are the parameters associated with the short-run differenced terms, ϕ_i are the coefficients that account for the speed of adjustment towards the long-run equilibrium, while $\boldsymbol{\theta}_i^{\prime}$ are the elements of the cointegrating vector.

In order to estimate the marginal propensity to consume from financial and real wealth, we apply the pooled mean group estimator suggested by Pesaran, Shin and Smith (1999). The estimator has been proposed in the large T, large N panels framework, whenever non-stationarity becomes an issue that can not be neglected. Broadly speaking, the pooled mean group estimator can be viewed as an intermediate procedure between pooling and averaging group estimates. Specifically, the estimator allows the intercepts, short-run coefficients and variances to differ across countries, while the long-run parameters are constrained to be identical across groups. This latter assumption is termed by Pesaran, Shin and Smith (1999) "long-run homogeneity" and requires to impose in (4)

$$\boldsymbol{\theta}_i = \boldsymbol{\theta} \quad (\forall i). \quad (5)$$

In Section 6 we shall present standard estimates of the marginal propensity to consume based on the pooled mean group estimator, relying on equation (4) with the ratio specification as in (2).

6 Results

We begin with some preliminary evidence based on static pooled estimation: in Table 4 we report the static fixed-effects estimates based on the ratio specification in (2). The top part of the table presents estimates of the marginal propensity to consume out of real and net financial wealth gathered from the flow of funds.

[SEE TABLE 4]

The point estimate of the marginal propensity to consume from net financial wealth, i.e., FW, is 0.0071, while the marginal propensity to consume from real wealth, i.e., RW, is 0.0008. These estimates are calculated on the basis of quarterly data, hence they have to be multiplied by four to get annualised values. This means 2.84 cents and 0.32 cents per euro of additional financial and real wealth, respectively. We include a time trend, which is statistically significant at 5 percent level.

In Section 2, we have surveyed papers that examined the effects of financial and real wealth on consumption: most of the times, financial wealth was approximated using stock market price data. Hence, as a sensitivity test, we also use this definition of financial wealth. The bottom part of Table 4 refers to marginal propensity to consume estimation when the stock market index (in log) is employed: the point estimate of the marginal propensity to consume is 0.0089 (3.56 cents per euro, in annualised terms). In general, the estimates are statistically significant working with conventional standard errors, although they are not significant if robust standard errors are used.

Yet, traditional static panel techniques do not allow to distinguish between the short-run and long-run dynamics. Furthermore, they are based on strong homogeneity assumptions among countries. For instance, fixed-effects models impose a single slope coefficient in the pooled estimation. The assumptions underlying static panel techniques appear to be too stringent in the case under study. Specifically, when the time dimension increases, potential country heterogeneity may be modelled in a richer way than using simple fixed (or random) effects models. This is done in the sequel by applying the pooled mean group estimator.

Hereafter, we consider the pooled mean group estimator, within an autoregressive distributed lag framework. The ARDL model in (3), re-parameterized introducing an error correction mechanism as in (4), is the starting point to perform pooled mean group estimation. For each country, the lag order of the ARDL model is chosen by applying the Schwarz Bayesian information criterion. Overall, the most commonly chosen representation is an ARDL(1,0,1), which in our framework reads as

$$C_{i,t} = \alpha_0 + \alpha_1 t + \lambda_{i,1} C_{i,t-1} + \tilde{\beta}_{i,0}^{FW} FW_{i,t} + \tilde{\beta}_{i,0}^{RW} RW_{i,t} + \tilde{\beta}_{i,1}^{RW} RW_{i,t-1} + \varepsilon_{i,t}, \quad i = 1, 2, \ldots, N,$$

that is, consumption over income is lagged once, real wealth over income is lagged once and only contemporaneous terms of net financial wealth over income are included.

Once the lag orders of the ARDL models have been selected for each country, the pooled mean group (PMG) estimator is used to obtain estimates of the marginal propensity to consume out of financial and real wealth.[7] Estimation of the long-run coefficients and of the group-specific error-correction coefficients is conducted by concentrated maximum likelihood, assuming Gaussianity of the innovations of the model in (4) with $\boldsymbol{\theta}_i = \boldsymbol{\theta} \quad (\forall i)$.

A number of different model specifications are used. The first estimates we present are relative to equation (4), using as a single regressor the net financial wealth measure gathered from the flow of funds. Table 5 reports the PMG estimation results. Formally, the homogeneity assumption in (5) holds: the variances and the short-run parameters are unrestricted, while the long-run coefficients are constrained to be identical across countries. We observe that the propensity to consume from net financial wealth has a statistically significant and positive coefficient, as expected: with quarterly data, the point estimate is 0.010, meaning 4 cents per euro of additional net financial wealth, on annual basis. The speed of adjustment coefficients are all negative and significant (except Portugal), supporting the evidence of cointegration among variables.

[SEE TABLE 5]

As a sensitivity test, we estimate the same model employing the log of the stock market index as a proxy for financial wealth. Results are given in Table 6. As it can be seen, the propensity to consume from financial wealth is positive and significant, 0.010

[7]Estimation is carried out by using a properly modified version of the GAUSS program provided by Professor M.H. Pesaran at the webpage http://www.econ.cam.ac.uk/faculty/pesaran/.

(although with a larger standard error, 0.03, than in Table 5). Hence, our results hold no matter whether we use the flow of funds definition of net financial wealth or the stock market index approximation. The speed of adjustment coefficients are all negative; they are statistically significant except in the case of Finland and Portugal. The adjusted R-squared statistics are, for the majority of the countries, lower than in Table 5.

[SEE TABLE 6]

As a further check, in Table 7 we propose a specification in which we use total wealth as a single regressor, which is defined as the sum of net financial wealth and real wealth. The long-run coefficient in the PMG estimation is found to be significant and equal to 0.009, hence 3.6 cents per euro of additional total wealth on an annual basis. The unconstrained speed of adjustment parameters, i.e., ϕ_i $(i = 1, 2, \ldots, N)$, are all significantly negative (except Portugal, which is positive but not statistically significant at conventional levels of confidence).

[SEE TABLE 7]

To further examine the relationship between household wealth and consumption, we come back to the distinction between the effect of financial and real wealth. Table 8 presents the results of PMG estimation, when both the long-run coefficients (propensity to consume from financial and real wealth) are constrained to be identical across countries. The variances and the short-run parameters are unconstrained, as usual. Both the marginal propensities are significant and positive, as expected. With quarterly data, the marginal propensity to consume out of real wealth is equal to 0.001, and the marginal propensity to consume out of net financial wealth is 0.009. On an annual basis, this corresponds to 0.4 cents per euro and 3.6 cents per euro, respectively. Thus, the marginal propensity to consume out of net financial wealth is found to be considerably larger the marginal propensity to consume out of real wealth. Concerning the speed of adjustment parameters, they are negative and significant in all countries except Portugal, France and Germany (although being positive, in these countries the estimates are not significantly different from zero).

[SEE TABLE 8]

Finally, in Table 9 we report single-country level estimation results. Also in this case, financial wealth measures are taken from the flow of funds. Yet, while up to now we have imposed a restriction that all the long-run coefficients are identical for all countries, in Table 9 we relax this assumption, using a completely unrestricted ARDL specification to examine the relationship between consumption and wealth. In this way, we can assess whether heterogeneity is present across individual countries.

[SEE TABLE 9]

Many studies have documented that the empirical results for UK and US are different from those obtained for other countries, indicating an heterogeneity of financial systems. Estimating single-country level ARDL equations, results in Table 9 point to similar conclusions: the empirical evidence of wealth effects on consumption is significant for the US and the UK, while it is weaker for other countries. Thus, we find a split between countries where mortgage equity withdrawal exists (typically the US and the UK) and other systems where this financial innovation is scarce or absent (euro area countries). In the majority of the countries the marginal propensity to consume coefficients take the correct sign, although quite often the significance of the parameter estimates is not achieved.[8] Therefore, group-specific OLS estimates using the ARDL approach may not provide us with precise estimates of the long-run marginal propensities to consume out of net financial and real wealth, yielding insignificant coefficients and counterintuitive signs.

7 Discussion and closing remarks

In this paper we have examined the effect of household financial wealth and real wealth on consumption, working with a panel of eleven OECD countries and quarterly data running from 1997 to 2008. According to panel unit root test results, all the series

[8]This is the case of Italy, for instance, for which the effects of real and financial wealth on consumption are not statistically significant. By contrast, as discussed in Section 2, Bassanetti and Zollino (2010) found significant wealth effects, laying in the range of 1.5-2 and 4-6 cents, respectively. This discrepancy may be due to the different time span considered, sample size, model specification and exogenous variables taken into account.

investigated are difference stationary and cointegrated. Therefore, we have adopted an estimation approach to make inference about the long-run relationships among consumption, financial and real wealth. Our findings can be summarized as follows.

First, in general, dynamic panel data regressions show that financial assets and real wealth positively influence household consumption. The estimate of the propensity to consume from financial wealth is larger than the propensity to consume from real wealth. We test the robustness of our estimates in a variety of ways. Two measures of household financial wealth are introduced: the first stems from flow of funds data, while the second is linked to share price indices. Furthermore, a number of estimation techniques are used: fixed-effects, pooled mean group estimation and single-country level ARDL estimation. Fixed-effects estimation allows intercept heterogeneity across countries, while assuming a single slope coefficient in the pooled estimation. Pooled mean group is an estimation procedure that constrains the long-run coefficients to be the identical across countries, while the short-run coefficients and error variances are left unrestricted. The approach based on single-country level ARDL allows long-run and short-run coefficients to differ across countries.

Overall, constraining the marginal propensity coefficients to be identical across countries (PMG estimation), we get estimates of the marginal propensity to consume out of net financial wealth in the range of 3.6-4 cents per euro, while the marginal propensity to consume out of real wealth is close to 0.5 cents per euro of additional wealth. In most of the cases the speed of adjustment coefficients are negative and significantly different from zero; furthermore, the point estimates are in general high, supporting the evidence of cointegration among variables. Estimation results are statistically significant and theoretically consistent. Clearly enough, restricting the marginal propensities to consume, the total number of unknown parameters is reduced; and this provides more degrees of freedom and hence more efficiency in estimation.

Second, looking at individual countries results (single-country level ARDL approach), the unrestricted coefficients of financial and real assets are most of the times positive, although not statistically significant for all countries. The empirical evidence supports the idea of a distinction between Anglo-Saxon financial systems and more traditional bank-oriented financial structures. Yet, it should be noted that the relatively small time series dimension of the data might have affected the estimation results, as it is somehow reflected by large standard errors and wide confidence intervals around the parameter estimates. This is an important caveat to our empirical findings on the heterogeneity across countries.

There is a great deal of future research stemming from this paper. We plan to extend our study to other factors often selected as co-determinants of aggregate consumption, such as demographic structure, distributional measures, interest rates, unemployment.

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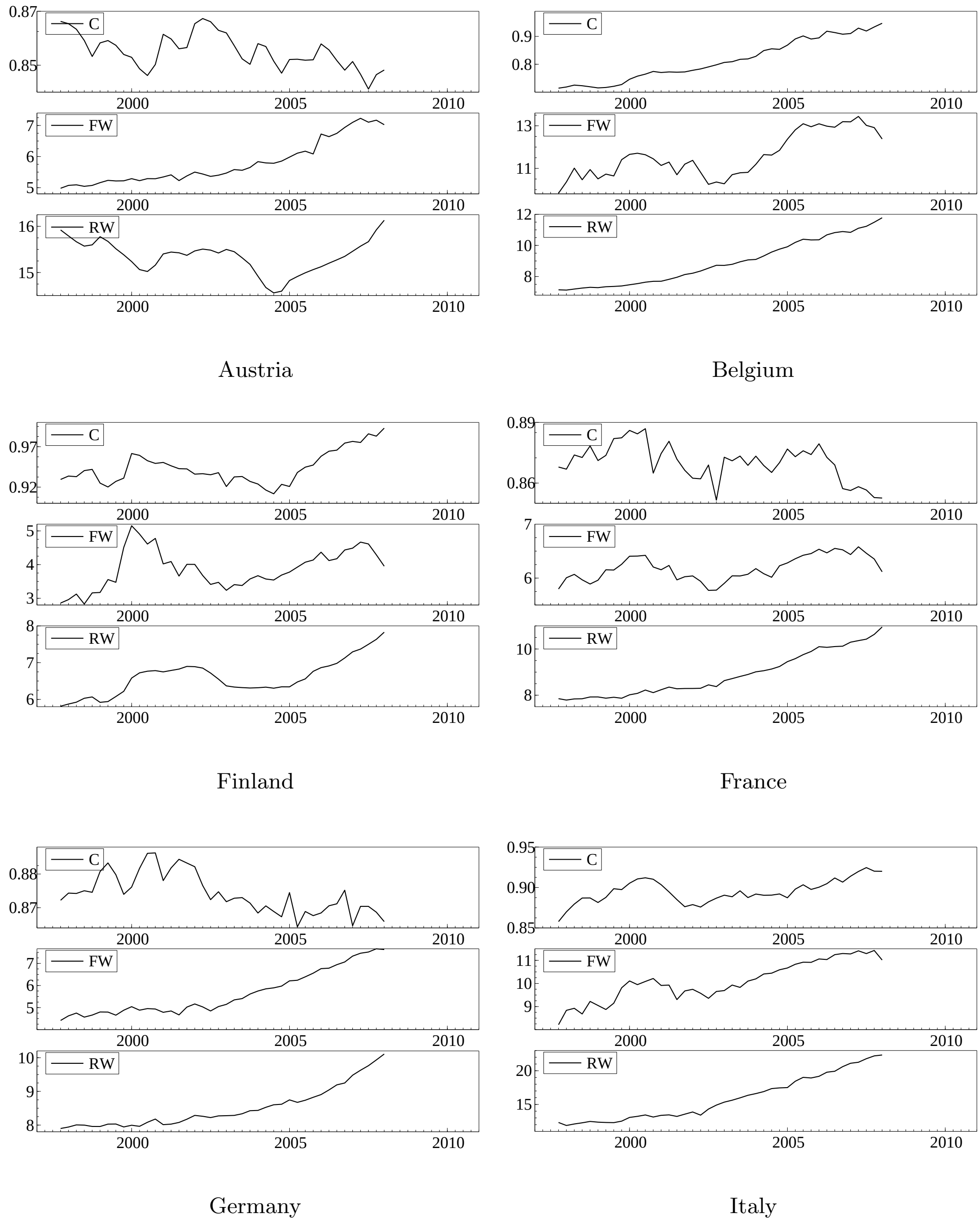
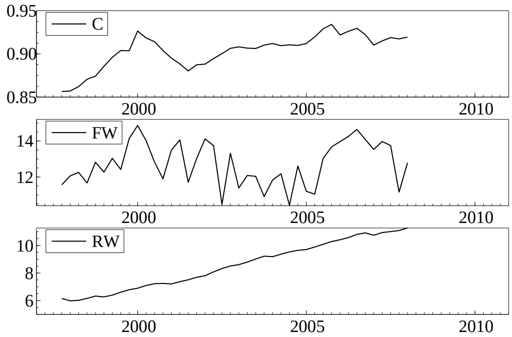
Figure 1: Consumption, financial and real wealth

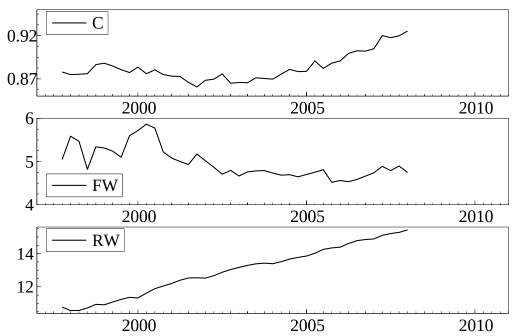


Quarterly household consumption expenditure (C), household net financial wealth (FW), household real wealth (RW). Note: all the variables have been expressed as a ratio to quarterly household disposable income. The sample goes from 1997Q4 until 2008Q1.

Figure 1: (continued) Consumption, financial and real wealth



Netherlands



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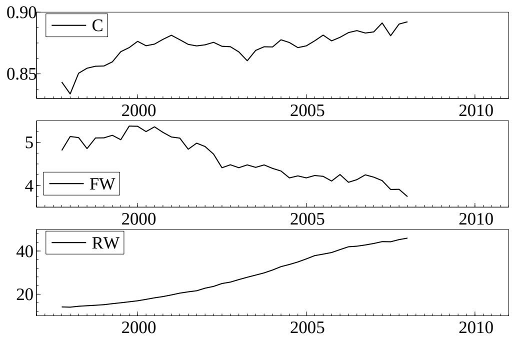
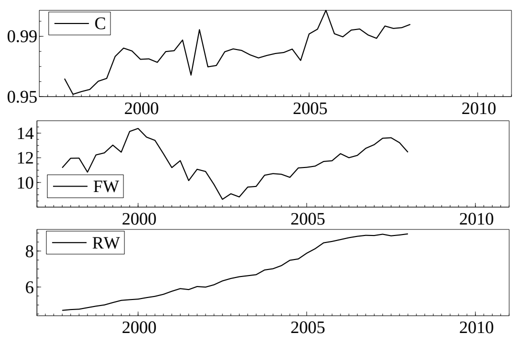
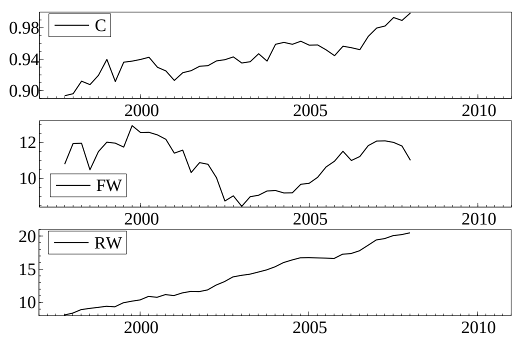


Spain



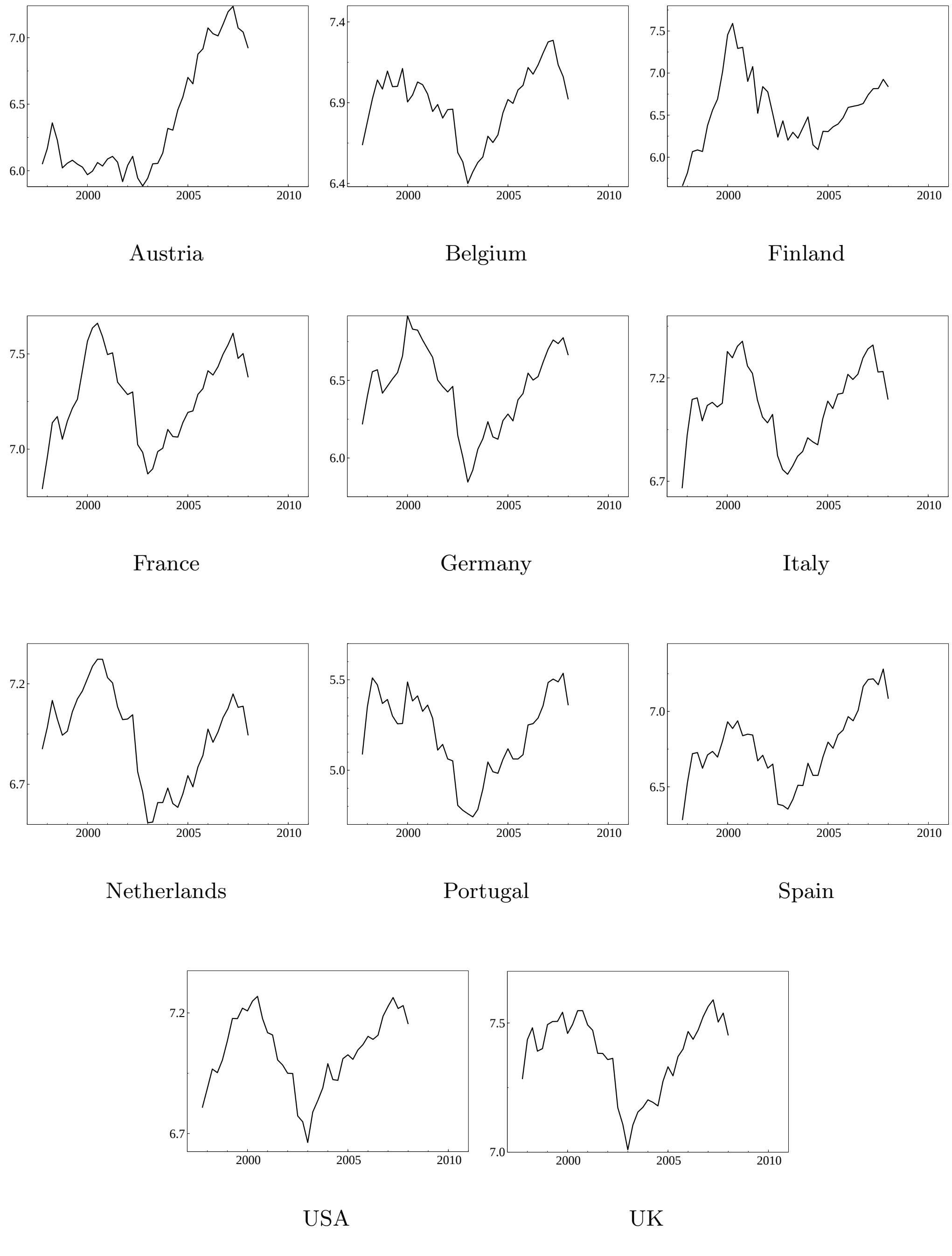
USA



UK

Quarterly household consumption expenditure (C), household net financial wealth (FW), household real wealth (RW). Note: all the variables have been expressed as a ratio to quarterly household disposable income. The sample goes from 1997Q4 until 2008Q1.

Figure 2: Stock market indices



Quarterly MSCI stock market indices (log scale). The sample goes from 1997Q4 until 2008Q1.

Table 1: Panel Unit Root Tests

Unit Root Test	Test Statistics	p-value
	Consumption over income	
Levin-Lin-Chu (LLC)	2.2397	0.9874
Im-Pesaran-Shin (IPS)	1.5739	0.9422
Maddala-Wu (MW)	15.1064	0.8576
Hadri-Larsson (HL)	12.4628	0.0000
	Net financial wealth over income	
Levin-Lin-Chu (LLC)	2.0063	0.9776
Im-Pesaran-Shin (IPS)	2.1695	0.9850
Maddala-Wu (MW)	14.3651	0.8880
Hadri-Larsson (HL)	10.7929	0.0000
	Real wealth (dwellings) over income	
Levin-Lin-Chu (LLC)	5.5897	1.0000
Im-Pesaran-Shin (IPS)	8.8461	1.0000
Maddala-Wu (MW)	1.7091	1.0000
Hadri-Larsson (HL)	18.8950	0.0000
	Stock market index (logs)	
Levin-Lin-Chu (LLC)	-0.1426	0.4433
Im-Pesaran-Shin (IPS)	-0.3209	0.3742
Maddala-Wu (MW)	18.1725	0.6957
Hadri-Larsson (HL)	2.5479	0.0054

Tests LLC, IPS are left-sided, while MW and HL are right-sided tests. All p-values are reported such that: H_0 is rejected if p-value < 0.05. The sample goes from 1997Q4 until 2008Q1.

Table 2: Panel Cointegration Test Results and Cointegration Estimates

The Panel Cointegration Test(Homogeneous): Kao (1999)

DF_ρ Test	-38.7577	Prob: 0.0000
DF_t Test	-18.4880	Prob: 0.0000
DF^*_ρ Test	-12.6275	Prob: 0.0000
DF^*_t Test	-13.1376	Prob: 0.0000

The ADF Panel Cointegration Test(Homogeneous): Kao (1999)

lags	ADF test statistic	prob:
1	-6.3498	0.0000

Group-mean Panel Dynamic OLS estimates

	coefficient	test statistic	p-value
FW	0.0108	6.4475	0.0000
RW	0.0108	14.0023	0.0000

The Panel Cointegration Test(Homogeneous): Kao (1999)

DF_ρ Test	-40.1462	Prob: 0.0000
DF_t Test	-18.7690	Prob: 0.0000
DF^*_ρ Test	-8.0492	Prob: 0.0000
DF^*_t Test	-10.3928	Prob: 0.0000

The ADF Panel Cointegration Test(Homogeneous): Kao (1999)

lags	ADF test statistic	prob:
1	-3.4730	0.0003

Group-mean Panel Dynamic OLS estimates

	coefficient	test statistic	p-value
ST.MKT. (logs)	0.0239	3.2825	0.0010

Outcome of the Kao (1999) test. FW is the coefficient of net financial wealth, RW is the coefficient of real wealth. ST.MKT. is the coefficient of the stock market index.

Table 3: Panel Cointegration Test Results with Multiple Structural Breaks

Regressors: FW & RW	
Country	Location of the estimated breaks
Austria	
Belgium	2000Q1; 2002Q3; 2005Q1;
Finland	2002Q4;
France	2005Q3;
Germany	2000Q1;
Italy	2000Q1; 2002Q3;
Netherlands	
Portugal	2000Q2; 2002Q4; 2005Q2;
Spain	2000Q1; 2002Q3;
USA	2001Q4;
UK	
Panel LM statistic	1.969

Regressor: ST.MKT. (logs)	
Country	Location of the estimated breaks
Austria	
Belgium	2000Q1; 2002Q3; 2005Q1;
Finland	2005Q3;
France	2000Q3; 2005Q3;
Germany	2003Q4;
Italy	2002Q2;
Netherlands	2002Q2;
Portugal	2002Q2; 2005Q3;
Spain	2000Q1;
USA	2000Q3; 2004Q4;
UK	2002Q2;
Panel LM statistic	1.277

The sample goes from 1997Q4 until 2008Q1.

Table 4: Preliminary evidence based on static fixed effects

	Static Fixed Effects Estimates				
	Coef.	St. Er.	t-ratio	Robust St. Er.	t-ratio
FW	0.0071	0.0015	4.8128	0.0062	1.1473
RW	0.0008	0.0004	2.1869	0.0008	1.0218
trend	0.0008	0.0001	6.5560	0.0005	1.6901
	Summary statistics and regression diagnostics				
	LL	SIGMA	AIC	SC	
	1075.89	0.023	1061.89	1033.11	
	Coef.	St. Er.	t-ratio	Robust St. Er.	t-ratio
ST.MKT. (logs)	0.0089	0.0042	2.1084	0.0109	0.8170
trend	0.0011	0.0001	11.3331	0.0005	2.0523
	Summary statistics and regression diagnostics				
	LL	SIGMA	AIC	SC	
	1066.17	0.023	1053.17	1026.44	

LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression, AIC for Akaike information criterion, while SC for Schwarz criterion. The sample goes from 1997Q4 until 2008Q1.

Table 5: Pooled Mean Group (PMG) estimates: FW constrained

Country	Parameter Estimates		Diagnostic Statistics		
	ϕ	FW	SIGMA	RBARSQ	LL
Austria	-0.273	0.010	0.004	0.08	165.80
	(0.098)	(0.002)			
Belgium	-0.546	0.010	0.007	0.30	143.62
	(0.116)	(0.002)			
Finland	-0.156	0.010	0.009	0.01	133.55
	(0.088)	(0.002)			
France	-0.538	0.010	0.006	0.26	147.81
	(0.169)	(0.002)			
Germany	-0.258	0.010	0.004	0.16	168.61
	(0.097)	(0.002)			
Italy	-0.227	0.010	0.006	0.06	152.84
	(0.093)	(0.002)			
Netherlands	-0.164	0.010	0.007	0.09	145.18
	(0.072)	(0.002)			
Portugal	-0.011	0.010	0.006	0.06	153.77
	(0.079)	(0.002)			
Spain	-0.490	0.010	0.004	0.37	167.48
	(0.101)	(0.002)			
US	-0.605	0.010	0.007	0.37	141.87
	(0.178)	(0.002)			
UK	-0.240	0.010	0.009	0.15	132.30
	(0.110)	(0.002)			

Estimation is conducted by pooled maximum likelihood, i.e., by concentrating the pooled maximum likelihood, under the assumption that the innovations are normally distributed (the Newton-Raphson optimization algorithm is employed). Figures in brackets are the standard errors of the coefficients. LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression. The sample goes from 1997Q4 until 2008Q1.

Table 6: Pooled Mean Group (PMG) estimates: Stock market index constrained

Country	Parameter Estimates		Diagnostic Statistics		
	ϕ	ST.MKT. (logs)	SIGMA	RBARSQ	LL
Austria	-0.219	0.010	0.004	0.04	168.87
	(0.098)	(0.003)			
Belgium	-0.280	0.010	0.008	0.11	142.10
	(0.103)	(0.003)			
Finland	-0.122	0.010	0.009	0.01	136.54
	(0.087)	(0.003)			
France	-0.459	0.010	0.007	0.15	148.49
	(0.141)	(0.003)			
Germany	-0.470	0.010	0.004	0.24	174.79
	(0.118)	(0.003)			
Italy	-0.217	0.010	0.006	0.08	155.25
	(0.084)	(0.003)			
Netherlands	-0.167	0.010	0.007	0.06	148.00
	(0.083)	(0.003)			
Portugal	-0.088	0.010	0.006	0.04	156.75
	(0.082)	(0.003)			
Spain	-0.340	0.010	0.005	0.17	163.46
	(0.097)	(0.003)			
US	-0.797	0.010	0.008	0.36	144.25
	(0.150)	(0.003)			
UK	-0.444	0.010	0.009	0.15	135.51
	(0.132)	(0.003)			

Estimation is conducted by pooled maximum likelihood, i.e., by concentrating the pooled maximum likelihood, under the assumption that the innovations are normally distributed (the Newton-Raphson optimization algorithm is employed). Figures in brackets are the standard errors of the coefficients. LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression. The sample goes from 1997Q4 until 2008Q1.

Table 7: Pooled Mean Group (PMG) estimates: Total wealth=FW+RW constrained

Country	Parameter estimates		Diagnostic Statistics		
	ϕ	RW+FW	SIGMA	RBARSQ	LL
Austria	-0.225	0.009	0.004	0.07	165.40
	(0.086)	(0.001)			
Belgium	-0.640	0.009	0.007	0.36	145.28
	(0.123)	(0.001)			
Finland	-0.205	0.009	0.009	0.03	134.04
	(0.100)	(0.001)			
France	-0.472	0.009	0.006	0.26	147.64
	(0.151)	(0.001)			
Germany	-0.199	0.009	0.004	0.12	167.88
	(0.085)	(0.001)			
Italy	-0.348	0.009	0.005	0.26	157.46
	(0.082)	(0.001)			
Netherlands	-0.197	0.009	0.007	0.12	145.84
	(0.075)	(0.001)			
Portugal	0.004	0.009	0.006	0.06	153.76
	(0.076)	(0.001)			
Spain	-0.153	0.009	0.004	0.31	165.49
	(0.039)	(0.001)			
US	-0.626	0.009	0.007	0.38	142.25
	(0.176)	(0.001)			
UK	-0.387	0.009	0.009	0.27	135.25
	(0.116)	(0.001)			

Estimation is conducted by pooled maximum likelihood, i.e., by concentrating the pooled maximum likelihood, under the assumption that the innovations are normally distributed (the Newton-Raphson optimization algorithm is employed). Figures in brackets are the standard errors of the coefficients. LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression. The sample goes from 1997Q4 until 2008Q1.

Table 8: Pooled Mean Group (PMG) estimates: FW and RW constrained (separately)

Country	Parameter estimates			Diagnostic Statistics		
	ϕ	FW	RW	SIGMA	RBARSQ	LL
Austria	-0.029	0.009	0.001	0.004	0.07	166.66
	(0.066)	(0.003)	(0.0002)			
Belgium	-0.002	0.009	0.001	0.007	0.41	148.18
	(0.019)	(0.003)	(0.0002)			
Finland	-0.097	0.009	0.001	0.007	0.49	147.81
	(0.060)	(0.003)	(0.0002)			
France	0.020	0.009	0.001	0.005	0.52	157.39
	(0.107)	(0.003)	(0.0002)			
Germany	0.009	0.009	0.001	0.004	0.12	168.94
	(0.059)	(0.003)	(0.0002)			
Italy	-0.168	0.009	0.001	0.006	0.08	154.45
	(0.094)	(0.003)	(0.0002)			
Netherlands	-0.160	0.009	0.001	0.007	0.11	146.89
	(0.060)	(0.003)	(0.0002)			
Portugal	0.091	0.009	0.001	0.006	-0.04	153.07
	(0.063)	(0.003)	(0.0002)			
Spain	-0.401	0.009	0.001	0.004	0.34	167.64
	(0.084)	(0.003)	(0.0002)			
US	-0.184	0.009	0.001	0.006	0.55	150.08
	(0.082)	(0.003)	(0.0002)			
UK	-0.092	0.009	0.001	0.008	0.39	140.01
	(0.056)	(0.003)	(0.0002)			

Estimation is conducted by pooled maximum likelihood, i.e., by concentrating the pooled maximum likelihood, under the assumption that the innovations are normally distributed (the Newton-Raphson optimization algorithm is employed). Figures in brackets are the standard errors of the coefficients. LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression. The sample goes from 1997Q4 until 2008Q1.

Table 9: Group-specific ARDL estimates (OLS): RW and FW unconstrained

Country	Parameter estimates			Diagnostic Statistics		
	ϕ	FW	RW	SIGMA	RBARSQ	LL
Austria	-0.393	-0.007	0.003	0.004	0.27	171.59
	(0.137)	(0.003)	(0.005)			
Belgium	-0.250	0.008	0.047	0.006	0.49	151.22
	(0.112)	(0.007)	(0.006)			
Finland	-0.281	-0.001	0.036	0.006	0.55	150.47
	(0.106)	(0.010)	(0.010)			
France	-0.453	0.034	-0.016	0.004	0.69	166.04
	(0.147)	(0.009)	(0.004)			
Germany	-0.399	-0.004	-0.008	0.003	0.32	174.26
	(0.160)	(0.006)	(0.013)			
Italy	-0.147	-0.007	0.003	0.006	0.11	155.14
	(0.106)	(0.027)	(0.006)			
Netherlands	-0.286	0.003	0.005	0.006	0.20	149.06
	(0.094)	(0.003)	(0.002)			
Portugal	0.007	0.038	-0.193	0.006	0.05	154.77
	(0.088)	(0.554)	(2.543)			
Spain	-0.464	0.028	0.002	0.004	0.43	170.36
	(0.090)	(0.008)	(0.000)			
US	-0.433	0.015	0.023	0.006	0.63	153.91
	(0.152)	(0.008)	(0.006)			
UK	-0.391	0.005	0.006	0.007	0.47	142.90
	(0.155)	(0.003)	(0.001)			

Group-specific estimates of the long-run coefficients based on ARDL specifications. Estimation is conducted by ordinary least squares. Figures in brackets are the standard errors of the coefficients. LL stands for log-likelihood of the model, RBARSQ for Adjusted R-squared, SIGMA for S.E. of regression. The sample goes from 1997Q4 until 2008Q1.

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