The role of asset markets for private consumption. Evidence from paneleconometric models

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Abstract. We explore the long and short run relationship between private consumption, disposable income and housing and financial wealth approximated by price indices for a panel of industrialized countries. Consumption, income and wealth are cointegrated in their common, but not in their idiosyncratic components. This stresses the relevance of international spillovers to explain aggregate consumption behaviour. The cointegrating vector is robust and in line with the life cycle permanent income hypothesis. The income elasticity does not differ from unity, and wealth elasticities are within a range of 2 to 5 percent. According to the error correction mechanism, consumption could not be interpreted as a weakly exogenous series.

Keywords: Permanent income hypothesis, panel cointegration, wealth effects

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1 Introduction

Recent developments in the stock market and changes in the valuation of house prices have brought wealth effects on consumption expenditures of private households back onto the agenda. The stock market boom in the late 1990s and the vast acceleration in house prices over the past years may have led to a rise in private consumption and subsequent growth in many countries. In the current slowdown of economic activity, however, this process is partially reversed. For example, housing starts and permits declined by 50 percent since 2006 in the US, and the Dow Jones industrial index lost more than 30 percent over that period. The deterioration in financial markets has already affected the real economic performance. For an appropriate assessment of the business cycle, the reaction of private consumption to these changes is extremely important. A huge response of consumption due to the presence of international spillovers in asset markets can actually deepen the recession in the world economy.

The wealth effect on consumption is often modelled by the marginal propensity to consume out of wealth. Much work has been initiated by the seminal papers of Lettau and Ludvigson (2001, 2004) for the US. In their VECM analysis, consumption, asset wealth and income cointegrate. Deviations from the common long run trend, captured by the *cay* residual, signal changes in asset prices. Hence, asset prices seem to carry the burden of adjustment after a shock, implying that wealth has large transitory components that are uncorrelated with consumer spending. However, the size of short run movements in the wealth variable is controversial. According to De Veirman and Dunstan (2008) and Hamburg, Hoffmann and Keller (2008) the consumption wealth ratio is able to forecast changes in consumption or income, respectively. Thus, *cay* predicts business rather than stock market cycles.

The cointegration result has been basically confirmed in recent studies, such as Davis and Palumbo (2001), Palumbo, Rudd and Whelan (2002), Bertaut (2003), Fernandez-Curegedo, Price and Blake (2003), Tan and Voss (2003) and Labhard, Sterne and Young (2005). More or less, wealth effects in the US and the UK exceed those in continental Europe. However, Dreger and Reimers (2006) have argued that the rise in stock markets in the 1990s is crucial to explain the decrease of the savings rate in most euro area countries. An impact is also reported for Japan, but since household wealth has changed little on balance in Japan in recent years, it has been less important to explain the consumption pattern.

Several authors have also distinguished between wealth components. Stock market and housing wealth can have a different impact on private consumption spending that is blurred by looking at the aggregate. In particular, housing represents both an asset and a consumption item. If house prices increase, wealth may rise, but also the cost of housing services (Poterba, 2000). Increases in the value of owner-occupied housing do not foster the ability of a household to consume more of other goods and services unless that household is willing to realise the increased value, for example, by moving into a less expensive flat. Many households are not willing to do this, including those who intend to leave their homes as bequests. For homeowners planning to increase their consumption of housing services (by moving into a more expensive home) or renters waiting to enter the housing market, the net effect is negative. Housing wealth effects may cancel in the aggregate, see Bajari, Benkard and Krainer (2003). For every household that sells a house there is a household that buys it. The increase in consumption (through the seller) might be offset by a decrease in consumption by the buyer. These ambiguities do not play a role for financial wealth.

According to Case, Quigley and Shiller (2005) an insignificant response of consumption to housing wealth might indicate multicollinearities of the two wealth components in a time series setup. Therefore, the cross section has to be taken into account. In fact, Case, Quigley and Shiller (2005) detected a larger effect for housing wealth in panels of US states and developed countries. In other studies, stock market wealth shows a larger impact, see Dvornak and Kohler (2003) for Australian states and De Veirman and Dunstan (2008) for New Zealand. Ludwig and Slok (2004) and Carroll, Otsuka and Slacalek (2006) have emphasized that the long-run responsiveness of consumption to permanent changes in wealth is higher for countries with a market-based than for countries with a bank-based financial system. The IMF (2002) has estimated an error correction model for OECD countries comprising income, equity and housing wealth as explanatory variables for consumption in the long run. An impact of stock market and housing wealth is detected, where the latter effect dominates in the US and the UK. In an update the IMF (2008) extended the short run specification by including inflation, while the impact of wealth effects is retained.

The main contribution of this paper is to establish the cointegration property in more precise terms. This is achieved by recent developed paneleconometric techniques designed to control for cross section dependencies. In particular, a long run equilibrium between consumption, income and wealth may occur due to the existence of international or national trends, or both. To explore these issues, each variable is separated into common and idiosyncratic components. Cointegration between the common components refers to the presence of international spillovers that dominate the relationship. In contrast, cointegration between idiosyncratic components may arise due to developments relevant on the national level. This distinction has huge implications for policymakers. If the common components cointegrate, international business cycles are expected to have a huge impact on the national economic evolution. In fact, this paper demonstrates that consumption, income and wealth (measured, *inter alia*, by housing or financial wealth) are cointegrated in their common components. In contrast, the evidence in favour of cointegration is rather limited for the idiosyncratic components. Furthermore, the long run vector is in line with the life cycle permanent income hypothesis. The income elasticity is not different from unity, and the marginal propensity to consume out of wealth is in a range of 3 to 5 per cent. The *cay* residual is crucial to explain subsequent consumption growth. Hence, consumption cannot be considered as a weakly exogenous variable.

The rest of the paper is organized in several subsections. The next section (Section 2) reviews the main transmission channels running from wealth to private consumption, and derives the empirical model. Section 3 discusses the panel cointegration techniques applied in the analysis. Section 4 describes the data and holds the results. The last section (Section 5) concludes.

2 Impact of wealth on private consumption

The life cycle permanent income hypothesis provides the theoretical framework to relate consumption to income and wealth. According to this hypothesis, private consumption responds to permanent income, the latter defined as the present value of expected lifetime resources, see Ando and Modigliano (1963). These resources include physical wealth, like housing and financial wealth, and human wealth, i.e. current labour income plus the present discounted value of the expected future labour income stream. An increase in wealth will raise consumption, because of its impact on expected lifetime income. If the resources become more valuable, the household can shift its consumption plans upward without violating its budget constraint. Hence, an increase in consumption is predicted in each period over the remaining lifetime. In the long run, the cumulated response of consumption is equal to the rise in permanent income.

Additional channels come into play if households are faced by liquidity constraints, see Muellbauer (2008) and De Veirman and Dunstan (2008). According to the permanent income hypothesis, households can borrow or lend to smooth consumption over the business cycle. But, if there is only limited access to credit, shocks in actual income might lead to corresponding shocks in private consumption. When a household experiences an increase in current or expected wealth, the value of the collateral it can offer to banks is higher. This means that banks are less reluctant to increase their loans. Therefore, the household can borrow more in order to finance extra consumption. Due to deregulation in the mortgage market in recent years, it became easier and cheaper for consumers to borrow against housing collateral to finance consumption. Cheaper access to home equity means that, for a given increase in asset prices, more borrowing is devoted to private consumption.

Furthermore, if future income and asset values are uncertain, households may prefer a buffer stock of wealth to mitigate negative income shocks, see Carroll (1997), among others. An increase in wealth raises the value of the buffer stock, and thereby reduce the need for precautionary saving. In that case, financial market liberalisation will weaken the relationship between consumption and wealth, as it will lower the fraction of liquidity constrained consumers.

A long-run relationship between consumption, income and wealth can be derived solely from the intertemporal budget constraint, see Campbell and Mankiw (1989). The starting point is the decomposition of total household wealth, W into asset wealth A and human wealth H and a wealth accumulation equation

(1) $$W_t = A_t + H_t$$

(2) $$W_{t+1} = (1 + r_{t+1})(W_t - C_t)$$

where the stock of wealth refers to the beginning of period t. Moreover, C_t denotes consumption and r_{t+1} is the net return on W_t. By dividing through W_t, taking logs and a first order Taylor expansion around the consumption-wealth ratio, the equation

$$w_{t+1} - w_t \approx r_{t+1} + (1 - 1/\rho)(c_t - w_t) \quad \rightarrow$$

$$c_t - w_t \approx \rho(r_{t+1} - \Delta c_{t+1}) + \rho(c_{t+1} - w_{t+1})$$

can be obtained, where $\rho=(W-C)/W$ is the steady state ratio of investment to wealth, lower case letters c and w denote the logs of the respective variables and constants are omitted. Solving the last equation forward yields

$$c_t - w_t = E_t \sum_{i=1}^{\infty} \rho^i (r_{t+i} - \Delta c_{t+i}) .$$

The transversality (no bubble) condition is fulfilled ($\rho<1$). The logarithm of total wealth is approximated by a weighted average of the logarithm of its two components, i.e. asset and human wealth

$$w_t \approx \lambda a_t + (1 - \lambda) h_t$$

where λ=A/W is the average share of non labour wealth in total wealth. Despite that H is not observable, a linear approximation is employed by interpreting H as the present or permanent value of labour income, Y. Then, the consumption wealth relationship can be re-written as

$$(3) \qquad cay_t = c_t - \lambda a_t - (1-\lambda) y_t \approx E_t \sum_{i=1}^{\infty} \rho^i (r_{t+i} - \Delta c_{t+i}) + (1-\lambda) z_t$$

where z is a white noise error term from the income approximation, see Lettau and Ludvigson (2001). Since all variables on the right hand side of the equation are stationary, the *cay* residual should be stationary too. Therefore, the intertemporal budget constraint implies a cointegrating relationship between consumption, asset wealth and labour income, where the cointegration parameters of asset wealth and income add to unity. Because λ is not time varying, *cay* denotes the consumption-wealth ratio. Fluctuations in this measure reflect expected future changes in consumption, asset wealth, and labour income.

3 Panel cointegration

The cointegration properties of the variables involved determine the appropriate specification of the consumption function. If the series cointegrate, the relationship between consumption, income and wealth should be interpreted as a long run equilibrium, as deviations are mean reverting. However, it has been widely acknowledged that standard unit root and cointegration tests can have low power against stationary alternatives, see for example Campbell and Perron (1991). Panel tests make progress in this respect. Since the time series dimension is extended by the cross section, inference relies on a broader information set. Therefore, gains in power are expected, and more reliable evidence can be obtained.

However, first generation panel unit root and cointegration tests are often based on the assumption of independent panel members. Because of common shocks, this condition is hardly fulfilled in applied work. In the presence of cross section dependencies, the tests are subject to large size distortions, see Banerjee, Marcellino and Osbat (2004, 2005). The situation gets worse if the number of cross sections is increased. To overcome these deficits, panel unit root tests have been developed that control for the dependencies via a common factor structure, see Breitung and Pesaran (2005). A similar approach is also relevant for cointegration. Banerjee and Carrion-i-Silvestre (2006) have presented residual based panel tests for a single cointegrating relationship with weakly exogenous regressors. In their model

(4) $$Y_{it} = \alpha_i + \beta_i X_{it} + u_{it}$$

(5) $$u_{it} = \lambda_i ' F_t + E_{it}$$

the index i denotes the cross section and t the time dimension. The error u can follow a common factor structure. In particular, F and E are the common and idiosyncratic components, respectively, that can be either integrated of order 1, I(1) or mean reverting, I(0). Cointegration implies the stationarity of both the common and idiosyncratic parts of the error term.

If the dependencies between the cross sections are persistent, the cointegration finding might be interpreted in different ways. A long run equilibrium can exist between the cross sections and between the time series for single units in the panel. Gengenbach, Palm and Urbain (2006) have proposed a sequential testing strategy. They discuss the case where non-stationarity is solely driven by a reduced number of common stochastic trends, and the case where both common and idiosyncratic stochastic trends are present in the data.

The starting point is a decomposition of each variable into common factors and idiosyncratic components, as suggested by Bai and Ng (2004). If the common factors are I(1), but the idiosyncratic components are I(0), the nonstationarity in the panel could be entirely driven by a reduced number of international stochastic trends. This would be the case of cross member cointegration. Cointegration between the series can occur only if the common factors of the variables cointegrate. If both the common factors and idiosyncratic components are I(1), cointegration is examined separately for the common and the idiosyncratic components. Suppose that the series *Y* and *X* have a single I(1) common factor, i.e.

(6) $$Y_{it} = \lambda_{1i} F_t^Y + E_{it}^Y$$

(7) $$X_{it} = \lambda_{2i} F_t^X + E_{it}^X$$

where *F* denote the common factors and *E* the idiosyncratic elements of the respective variables. A panel cointegrating relationship between *Y* and *X*

(8) $$Y_{it} - \beta_i X_{it} = \lambda_{1i} \left(F_t^Y - \lambda_{2i} \frac{\beta_i}{\lambda_{1i}} F_t^X \right) + E_{it}^Y - \beta_i E_{it}^X$$

requires that the null of no cointegration is rejected for both the common and the idiosyncratic components. Cointegration between common factors can be examined by the usual time series tests such as the Johansen reduced rank approach. As the idiosyncratic components are independent by construction, their analysis is done by standard panel tests such as those of Pedroni (1999, 2004). It should be noted, however, that the existence of cointegrating relationships that annihilate both the common and idiosyncratic trends is very unlikely, see equation (8).

Note that the panel cointegration tests do not provide an estimate of the long run relationship. Mark, Ogaki and Sul (2005) have proposed a dynamic seemingly unrelated regression estimator for the cointegration vector in case of correlated errors. It is based on a two step procedure. In the first step, the endogenous variable in each equation is regressed on the leads and lags of the first difference of the regressors from all equations to control for possible endogeneities. The leads and lags from other equations are included to take cross section dependencies into account, see also Saikkonen (1991). In the second step, the SUR method is applied using the residuals from the first step regression. As a drawback, this procedure is only feasible when the cross section dimension is rather small.

The cointegration vector should be identical for all panel members, more or less, as fundamental economic principles are involved. In fact, there is only little theoretical rationale for a wide dispersion of the cointegration parameters if the countries are at a similar stage of their development. Cross country differences reported in many empirical studies might be traced back to measurement problems of wealth in various countries (Layard, Sterne and Young, 2005). Therefore, after testing for cointegration, the pooled mean group estimator suggested by Pesaran, Shin and Smith (1998) is applied to reveal a common cointegration vector. It restricts the long run coefficients to be identical across the countries, but allows short run coefficients and error variances to vary across groups.

4 Data and results

The analysis is based on data for 14 industrial countries for the period from 1991Q1 to 2007Q4. The data are taken from the World Market Monitor provided by Global Insight. Consumption refers to total consumption expenditures of private households. Income is proxied by personal disposable income. Apart from labour income, this measure includes income received from wealth, like interest payments, profits and dividends. Consistent labour income measures are hardly available in an international setting. In some countries, effective wages are not reported. In other countries, they refer not to the entire economy, but only to the industrial sector. Therefore, the analysis is done with the broader income concept.

The appropriate definition of wealth is more critical. Measurement errors in a cross section of countries are likely, especially, when real estate values are involved, see Lustig and Van Nieuwerburgh (2005). Some studies like Ludwig and Slok (2004) have used stock market capitalisation data. However, stocks can be also owned by foreigners. Furthermore, not all the equities are actually listed, and housing wealth is neglected at all. Hence, the stock market capitalisation might not reflect the actual financial wealth of private households. The ECB (2009) has recommended to use price data instead of the stock of wealth, and this is the approach we followed in the subsequent analysis. Price series are readily available across countries, and are reported at the desired frequencies. Share prices refer to the national stock market index, and house prices are price indexes for new houses. All series are deflated by the CPI and measured in logs. Inflation is the annualized logarithmic difference of the CPI.

The first step is to examine the unit root properties of the variables involved. Although the vast majority of studies has already detected stochastic trends in consumption, income, and asset prices, this is not a trivial task. In particular, the sources of possible nonstationarities are relevant in the analysis presented here, since the main aim is to distinguish cointegration between the common components and cointegration between the idiosyncratic components of the series. Thus, the presence of random walks has to be explored for the two components in a separate way.

In particular, the variables are decomposed into common and idiosyncratic elements by principal component analysis. Since the principal components could be nonstationary, the decomposition is based on differenced data, as suggested by Bai and Ng (2004). Once the factors have been estimated, they are re-cumulated to match the integration properties of the original series. The idiosyncratic elements arise from a projection of the variables on their common components. Inference on the unit root properties is obtained by standard time series tests for the common factors. As the defactored series are independent by construction, stochastic trends in the idiosyncratic components are efficiently explored by first generation panel unit root tests.

-*Table* 1 *about here*-

The number of common factors in the principal component analysis is estimated using the BIC3 criterion, see Bai and Ng (2002). Since the cross section and time series dimensions of the panel are approximately of the same magnitude, this criterion tends to be superior over the alternatives. The results in Table 1 refer to the single factor model for all variables. For consumption and income, some other criteria would favour a higher number of factors. However, the evidence is very robust to this choice. While the common factors appear to be nonstationary, the unit root is rejected for the idiosyncratic components of the series, apart from consumption. Hence, random walks in the data are mostly driven by international developments. As a consequence, cointegration may hold between the common components. But a long run relationship is rather unlikely for the idiosyncratic components.

-*Tables* 2 *and* 3 *about here*-

The outcome of the cointegration tests is broadly consistent with the unit root finding, see Table 2. According to the Johansen (1995) trace statistic, there is strong evidence for a long run relationship between the common factors of consumption, income and wealth. In models with either share or house prices, the long run vector is unique. Two cointegration vectors exist, if both wealth variables are considered. This implies no cointegration between the common parts of the wealth ingredients. Furthermore, some of the Pedroni (1999, 2004) tests point to cointegration even between the idiosyncratic components, see Table 3. This is particularly true for the PP, but not for the ADF type statistics. Note that a long run relationship cannot be not expected from the unit root analysis (Table 1). A cointegration result between the idiosyncratic components should be therefore interpreted with care.

The cointegrating parameters are revealed by the panel mean group estimator, see Table 4 for the regression results. The long run vector is in line with the life cycle permanent income hypothesis. The income elasticity is not different from unity, and the wealth elasticities are about 0.02 and 0.05. The pooled estimates are lower the individual values reported by the IMF (2008) for the G7 countries or values reported by Labhard, Sterne and Young (2005).[2] The impact from the stock market is slightly larger than its counterpart from the housing market. If both wealth measures are considered, their effects have to be added, at least approximately. This coincides with the evidence on the cointegration rank (see Table 2). The wealth components behave almost independently in the long run.

-Tables 4 *about here-*

Whether or not the long run can be interpreted in terms of an equation explaining private consumption can be inferred from the error correction model. If consumption adjusts to restore the equilibrium in response to a shock, it could not be weakly exogenous with respect to the cointegrating relationship, i.e. the error correction term should affect subsequent consumption behaviour in a significant way. Results obtained with different wealth specifications are exhibited in Table 5.

-Table 5 *about here-*

Besides the error correction term the equations include the contemporaneous change in income, stock and housing wealth and inflation. Inflation serves as a measure of consumption uncertainty, see IMF (2008). IV methods are used to control for possible en-

[2] If the common components are cointegrated, but the idiosyncratic components are not, the long run vector could be alternatively estimated solely on the grounds of the common components. Similar results can be obtained in this case. However, the income elasticity slightly exceeds unity.

dogeneity. In a first step, the contemporaneous regressors are explained by lagged consumption, income, wealth and inflation on a country-by-country basis. Then, the fitted values from this regression are used as instruments for the contemporaneous variables in the second step.

If the feedback coefficient is restricted to be homogeneous across countries, the *t*-statistic reveals a highly significant impact of the long run relationship on subsequent consumption growth. Households adjust their expenditures to restore the long run equilibrium. Note that the parameters are substantially lower in absolute value than the estimates of -0.5 recently reported by the IMF (2008). The use of quarterly instead of annual data might explain some part of the difference.

The homogeneity restriction can be suspended. Country individual feedback parameters are reported in the lower part of the Table. For most countries the feedback parameter turns out to be negative, as one can expect from a stable model. Some exceptions should be noted. The long run vector does not play any role to explain aggregate consumption behaviour in Austria and Canada. For Belgium, Germany, and Portugal, the error correction term appears to significant only at higher (0.10) levels, at least in some of the specifications.

5 Conclusions

This paper explores the long and short run relationship between private consumption, disposable income and housing and financial wealth for a panel of industrialized countries. Consumption, income and wealth are cointegrated in their common, but probably not in their idiosyncratic components. The cointegrating vector turns out to be rather robust across different model specifications and in line with the life cycle permanent income hypothesis. The income elasticity does not differ from unity, and the wealth elasticities are small but positive. According to the error correction mechanism, consumption could not be interpreted as a weakly exogenous series. This implies that a correction of asset prices would reduce to some extent consumption expenditures. Our findings are also related to the literature on international business cycles. The cointegration finding for the common components stresses the relevance of international spillovers to explain aggregate consumption behaviour.

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Table 1: Unit root analysis

	Common component	Idiosyncratic component
Private consumption	-2.939	-0.293
Disposable income	-2.681	-1.881*
Share prices	-1.621	-1.704*
Housing prices	-2.849	-2.219*

Note: The optimal lag length in the regressions is determined by the general-to-simple approach suggested by Campbell and Perron (1991). Unit roots are examined via the ADF regression (with linear time trend) in case of the common component, and via the IPS test for the idiosyncratic component, see Im, Pesaran and Shin (2003). An asterisk denotes the rejection of the unit root hypothesis at least at the 0.05 level.

Table 2: Cointegration of common components

Rank null hypothesis	Equity prices	House prices	Equity and house prices
r≤0	30.55*	40.48*	63.49*
r≤1	12.95	15.27	34.45*
r≤2	1.96	4.69	14.02
r≤3			2.99

Note: Johansen (1995) trace statistics. Lag length of VAR determined by Schwarz criterion and equal to 2 for the VAR in levels. To correct for finite sample bias, the trace statistic is multiplied by the scale factor $(T\text{-}pk)/T$, where T is the number of the observations, p the number of the variables and k the lag order of the underlying VAR model in levels, see Reimers (1992). Critical values are taken from MacKinnon, Haug and Michelis (1999), and are also valid for the finite sample correction. A * indicates the rejection of the null hypothesis of no cointegration at least on the 0.05 level of significance.

Table 3: Cointegration of idiosyncratic components

	Equity price		House price		Equity and house prices	
	Panel	Group	Panel	Group	Panel	Group
Variance ratio	1.623		1.957*		1.258	
Rho statistic	-1.688*	-2.367*	-1.503	-2.000*	-1.005	-1.595
PP statistic	-1.830*	-2.466*	-1.639	-2.256*	-1.578	-2.409*
ADF statistic	-0.229	-0.363	-1.146	-0.343	-0.042	-0.308

Note: The Pedroni (1999, 2004) test statistics are asymptotically distributed as standard normal. The variance ratio test is right-sided, while the other tests are left-sided. Maximum truncation lags are set to 4 and determined using data dependent criteria. A * indicates the rejection of the null hypothesis of no cointegration at least on the 0.05 level of significance.

Table 4: Cointegration vector, pooled mean group estimator

	Model 1	Model 2	Model 3
Constant	-0.824 (0.032)	-0.615 (0.034)	-0.444 (0.037)
Income	1.081 (0.005)	1.016 (0.013)	1.000 (0.006)
House prices	0.019 (0.005)		0.024 (0.003)
Share prices		0.029 (0.001)	0.030 (0.001)

Note: Standard errors in parentheses.

Table 5: Error correction specification

	Model 1	Model 2	Model 3
EC	-0.036 (0.006)	-0.049 (0.006)	-0.051 (0.007)
Country individual feedback mechanism			
AT	-0.030 (0.040)	0.003 (0.024)	-0.012 (0.026)
BL	-0.026 (0.020)	-0.036 (0.022)	-0.046 (0.025)
CA	-0.001 (0.015)	0.002 (0.017)	0.004 (0.016)
FI	-0.092 (0.016)	-0.112 (0.024)	-0.108 (0.025)
FR	-0.078 (0.030)	-0.131 (0.036)	-0.093 (0.031)
GE	-0.129 (0.079)	-0.094 (0.091)	-0.139 (0.093)
IR	-0.099 (0.040)	-0.178 (0.058)	0.027 (0.08)
IT	-0.038 (0.013)	-0.053 (0.016)	-0.051 (0.016)
JP	-0.039 (0.19)	-0.034 (0.018)	-0.030 (0.017)
NL	-0.072 (0.30)	-0.072 (0.027)	-0.109 (0.032)
PO	-0.062 (0.048)	-0.119 (0.062)	-0.114 (0.060)
SP	-0.057 (0.025)	-0.069 (0.024)	-0.080 (0.024)
UK	-0.082 (0.031)	-0.065 (0.020)	-0.090 (0.026)
US	-0.123 (0.034)	-0.154 (0.029)	-0.164 (0.030)

Note: Sample period 1991.1-2007.4, IV estimation with fixed effects. Dependent variable is the change in private consumption. Specification of error correction term (EC) according to Table 4, EC lagged one period. AT=Austria, BL=Belgium, CA=Canada, FI=Finland, FR=France, GE=Germany, IR=Ireland, IT=Italy, JP=Japan, NL=Netherlands, PO=Portugal, SP=Spain, UK= United Kingdom, US=United States. Standard errors in parentheses.