The long run relationship between private consumption and wealth: Common and idiosyncratic effects

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Abstract. We investigate the long run relationship between private consumption, disposable income and wealth approximated by equity and house price indices for a panel of 15 industrialized countries. Consumption, income and wealth are cointegrated in their common components. The impact of house prices exceeds the effect arising from equity wealth. The long run vector is broadly in line with the life cycle permanent income hypothesis, if house prices are allowed to enter the relationship. At the idiosyncratic level, a long run equilibrium is detected between consumption and income, i.e. the wealth variable can be excluded. The income elasticity in the idiosyncratic relationship is significantly less than unity. Hence, the presence of wealth effects in consumption equations arises from the international integration of asset markets and points to the relevance of risk sharing activities of agents. Without sufficient opportunities, an increase in national saving rates would be expected, leading to a lower path of private consumption expenditures.

Keywords: Permanent income hypothesis, panel cointegration, wealth effects

JEL: C23, E21, E32, G15

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1 Introduction

Recent developments in international stock markets and large declines in house prices have brought wealth effects on consumption expenditures of private households back on the agenda. The stock market boom in the late 1990s and the vast acceleration in house prices before the financial crisis may have led to a rise in private consumption and subsequent growth in many countries. In the subsequent slowdown of economic activity, however, this process has been partially reversed. For an appropriate assessment of the business cycle, the reaction of private consumption to these changes is extremely important.

The examination of the wealth effect on private consumption has been stimulated by the seminal contributions of Lettau and Ludvigson (2001, 2004) for the US. In their VECM analysis, consumption (c), asset wealth (a) and income (y) cointegrate. Deviations from the common long run trend, captured by the *cay* residual, indicate changes in asset prices. Asset prices seem to carry the burden of adjustment after a shock, implying that wealth has huge 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 error correction term ratio can forecast changes in consumption or income. Thus, *cay* predicts business rather than stock market cycles.

The cointegration result has been confirmed in recent studies, like 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). In addition, 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 late 1990s

is crucial to explain the decrease of the savings rate in most euro area countries. An impact has been also detected 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 could have a different impact on private consumption spending that might be blurred at the aggregate level. For example, 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 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. Therefore, wealth effects may cancel out in the aggregate (Bajari, Benkard and Krainer, 2003). For every household that sells a house there is a household that buys it. The increase in consumption might be offset by a decrease in consumption by the buyer. These ambiguities do not play a role for stock market wealth.

According to Case, Quigley and Shiller (2005) an insignificant response of consumption to housing wealth might indicate multicollinearities of the wealth components in a time series setup. Therefore, the cross section has to be taken into account. In fact, the authors detected larger effects for housing wealth in panels of US states. In other studies, stock market wealth shows the larger impact (Dvornak and Kohler, 2003 and De Veir-

man and Dunstan, 2008). Ludwig and Slok (2004) and Carroll, Otsuka and Slacalek (2006) have emphasized that the long-run responsiveness of consumption to permanent changes in wealth depends on the institutional framework. In particular, the wealth effect 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 industrial countries including income, equity and housing wealth as explanatory variables for private consumption in the long run. An impact of stock market and housing wealth is reported, where the latter 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 variables is retained.

However, no paper has examined the role of international spillovers in determining the consumption pattern. In fact, a higher integration of financial markets offers the opportunity that agents can hedge consumption risk across countries by holding an internationally diversified portfolio of assets. While stock markets are highly integrated, housing markets may be largely driven by the domestic development. Nonetheless, national prices might be positively correlated because of a similar course of monetary policy. Hence, the first contribution of this paper is to look at the international dimension of aggregate consumption behaviour.

The second contribution refers to econometric methods. To explore the relation between consumption, disposable income and wealth more efficiently, many authors such as Ludwig and Slok (2004), Case, Quigley and Shiller (2005) and Carroll, Otsuka and Slacalek (2006) have chosen a paneleconometric environment, with countries or regions as the cross section. However, these studies have assumed that the cross section units are largely independent. This does not hold in the presence of common shocks, as exempli-

fied by the financial crisis. Therefore, this paper takes the recent developed panele-conometric techniques to control for cross section dependencies into account. The third contribution is related to the sample period, as the financial crisis is covered.

In principle, a long run equilibrium between consumption, income and wealth may occur because of the existence of international or national trends, or both. To explore these issues, each variable is decomposed 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 only 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 development. The more relevant the common relationship is, the less the ability to manage the national macroeconomic evolution.

This paper demonstrates that private consumption, income and wealth (measured, *inter alia*, by housing or equity wealth) are cointegrated in their common components. The long run vector is broadly in line with the life cycle permanent income hypothesis, if house prices are allowed to enter the relationship. It might be interpreted as a consumption equation, as the null hypothesis of weak exogeneity is rejected, at least in model variants based on house prices. In addition, wealth measures seem to react to deviations from the long run. The latter evidence, which reinforces the Lettau and Ludvigsson (2001, 2004) results is not general and depends on the concrete model specification. At the idiosyncratic level, cointegration can be established between consumption and income, i.e. the wealth variable is excluded. The income elasticity in the idiosyncratic relationship is significantly less than unity. Hence, the presence of wealth effects in

consumption equations arises from the international integration of asset markets and points to the relevance of risk sharing activities of agents. Without such opportunities, an increase in national saving rates would be expected, leading to a lower path of private consumption expenditures.

The rest of the paper is organized as follows. 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 the budget constraint. Thus, 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. However, if there is only limited access to credit, shocks in actual income might lead to corresponding shocks in private consumption. The introduction of wealth can weaken the relationship between actual income and actual 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. The deregulation in motgage markets made it 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 subject to high uncertainty, households may prefer a buffer stock of wealth to mitigate negative income shocks, see Carroll (1997). An increase in wealth raises the value of the buffer stock, and reduces 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.

The long-run relationship between consumption, income and wealth is derived from the intertemporal budget constraint, see Campbell and Mankiw (1990). 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 value. C_t denotes consumption and r_{t+1} is the stationary real interest rate. By dividing the equation through W_t, taking logs and a first order Taylor expansion around the consumption-wealth ratio, the relationship

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

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

can be obtained. The term $\rho=(W\text{-}C)/W$ 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 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 $\lambda=A/W$ is the average share of non labour wealth in total wealth. As H is not observable, a linear approximation is used. In particular, H is interpreted as the present or permanent value of labour income, Y. The consumption wealth relationship can be rewritten 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 integration 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 empirical work. In the presence of cross section dependencies, the tests suffer from to large size distortions, see Banerjee, Marcellino and Osbat (2004, 2005). The situation gets even 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. A similar approach is also relevant for cointegration. Banerjee and Carrion-i-Silvestre (2006) have presented residual based panel tests for unique long run relationship with weakly exogenous regressors. However, cross section spillovers are restricted to the error term.

If the dependencies between the cross sections are persistent, a cointegration finding might be interpreted in different ways. A long run equilibrium may exist between the cross sections and between the time series for single units in the panel. To account for this possibility, Gengenbach, Palm and Urbain (2006) have proposed a sequential testing strategy. They discuss the case where nonstationarities are 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 section 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.

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

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

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

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

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 (1995) 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 quite unlikely.

The panel cointegration tests do not provide an estimate of the long run relationship. 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 stages of development. Cross country differences reported in empirical studies might be traced back to measurement problems of wealth in various countries (Layard, Sterne and Young, 2005).

The cointegration relationship is estimated separately for the common and idiosyncratic components. For the common components, the reduced rank ML estimator suggested by

Johansen (1995) is appropriate. However, as this approach can produce extremely distorted and unreliable estimates if the sample size is not large, Brüggemann and Lütkepohl (2005) have recommended a two-step generalized least squares estimator, which appears to be more robust. This simple two step estimator is used as a cross-check to the ML results. For the idiosyncratic components, efficient estimation techniques like fully modified (FMOLS) and dynamic OLS (DOLS) are applied. As they control for potential endogeneity of the regressors and serial correlation, asymptotically unbiased estimates of the long run between the idiosyncratic components can be obtained. Cross section heterogeneity is limited to fixed effects, time trends and short run dynamics. The panel FMOLS estimator is obtained as the average of the country specific parameters (Pedroni, 1999). A panel DOLS estimator is obtained using the Mark, Ogaki and Sul (2005) procedure. After regressing out individual dynamic and deterministic elements, the residuals are stacked and a pooled regression is run. Note that the panel estimators do not need to control for cross section dependencies, as this part of the analysis focuses on the idiosyncratic components.

4 Data and results

The analysis is based on data for 15 industrial countries (Austria, Belgium, Canada, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Portugal, Spain, Sweden, the UK and the US) for the period from 1991Q1 to 2010Q2. The country sample includes the G7 and the EU old member states. 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. In addition to labour income, disposable income includes income received from wealth,

like interest payments, profits and dividends. Consistent labour income measures are not available in an international setting, as effective wages are not reported for some countries. 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, if real estate values are involved, see Lustig and Van Nieuwerburgh (2005). Some studies like Ludwig and Slok (2004) have used data on stock market capitalisation. However, stocks can be also owned by foreigners. Furthermore, not all the equities are listed, and housing wealth is neglected at all. Hence, the stock market capitalisation might not reflect the actual wealth of private households. The ECB (2009) has recommended to use price data instead of the stock of wealth, and this is the approach that is used in the subsequent analysis. Price series are readily available across countries, and are reported at the desired frequencies. Equity 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.

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. The sources of possible nonstationarities are relevant in this analysis, since the main aim is to distinguish cointegration between the common components and cointegration between the idiosyncratic components of the series.

The variables are decomposed into common and idiosyncratic factors by principal component analysis. As the components could be nonstationary, the decomposition is based on the differenced data, as suggested by Bai and Ng (2004). Once the factors have been

estimated, they are re-cumulated to match the stochastic properties of the original series. The idiosyncratic components 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 can be efficiently investigated by first generation panel unit root tests. Here, the test suggested by Im, Pesaran and Shin (2003) is applied. The test statistic is a standardized average of individual ADF statistics. Under the null hypothesis, the series under consideration is integrated for all panel members.

-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 involved. The first factor represents 75 percent of the overall variance of equity prices, thereby indicating high international integration of stock markets. For consumption, income, and house prices, however, less than 30 percent are captured, and in fact, some information criteria favour a higher number of factors for these variables. Nonetheless, the evidence is very robust to this choice[2]. While the common factors are nonstationary, the unit root hypothesis is rejected for the idiosyncratic components of

[2] For example, all important results can be replicated, if the common factors are obtained as a linear combination of the first two or three principal components for each variable, where weighting factors reflect the contribution to the overall variance.

the wealth series. Therefore, a long run relationship between consumption and wealth will be due to international developments. In contrast, a cointegrating finding between idiosyncratic components will not include wealth measures, as their idiosyncratic parts appear to be stationary.

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 share or house prices, the long run vector is unique. While the null hypothesis of no cointegration can be rejected at the conventional 0.05 level, if wealth is proxied by house prices, the evidence is somewhat weaker for the equity price variant. Two cointegration vectors seem to exist, if both wealth variables are included. This also implies no cointegration between the common components of the wealth measures. Regarding the idiosyncratic components, panel and group statistics suggested by Pedroni (1999, 2004) provide strong evidence in favour of cointegration between consumption and income, see Table 3.

-Tables 2 and 3 about here-

The cointegrating vector for the common components is exhibited in Table 4. If wealth is exclusively proxied by equity prices, the long run parameters do not seem to be reasonable, as the income elasticity exceeds unity. However, if house prices are considered, the cointegration vector is in line with the life cycle permanent income hypothesis, more or less. For example, if both wealth measures enter the analysis and the long run is estimated by the simple two step method (Brüggemann and Lütkepohl, 2005), the elastic-

ities with respect to income, equity and house prices are 0.77, 0.05 and 0.19, respectively. Thus, the impact of house prices exceeds the effect arising from equity wealth, implying that house prices fluctuations might cause a larger response in consumption. If both wealth measures are included, their effects have to be added, approximately. This coincides with the evidence on the cointegration rank. The wealth components appear to behave almost independently in the long run.

-Tables 4 and 5 about here-

Wealth elasticities are higher than the values reported by the IMF (2008) for the G7 countries or those of Labhard, Sterne and Young (2005). This is due to the fact that the results are restricted to the common components. The relative low income elasticity could also reflect measurement errors, since the analysis is done with a broader income concept than labour income. This can also explain the low income elasticity in the cointegration relationship for the evolution of the idiosyncratic components (Table 5). According to fully modified and dynamic OLS methods, this parameter is roughly equal to 0.5. Hence, the consumption response to an acceleration in income is less than proportional. If financial markets become less integrated, an increase in national saving rates should be expected, implying decelerating consumption expenditures. Because private consumption is the largest demand component in GDP, the long run growth prospects would be considerably lower.

Finally, it is explored whether the cointegrating relationship can be interpreted as a long run equation for private consumption. This would be the case if consumption is not

weakly exogenous with respect to the cointegrating relationship (Table 6). In fact, the null hypothesis can be rejected at very low significance levels, provided that housing is included as a measure of wealth. In contrast, consumption appears to be weakly exogenous, if wealth is solely approximated by equity prices. While income is weakly exogenous, wealth measures seem to react to deviations from the long run, but the latter result depends on the concrete model specification.

-Table 6 about here-

5 Conclusions

We investigate the long run relationship between private consumption, disposable income and wealth approximated by equity and house price indices for a panel of 15 industrialized countries. Consumption, income and wealth are cointegrated in their common components. The impact of house prices exceeds the effect arising from equity wealth. The long run vector is broadly in line with the life cycle permanent income hypothesis, if house prices are allowed to enter the relationship. At the idiosyncratic level, a long run equilibrium is detected between consumption and income, i.e. the wealth variable can be excluded. The income elasticity in the idiosyncratic relationship is significantly less than unity. Hence, the presence of wealth effects in consumption equations arises from the international integration of asset markets and points to the relevance of risk sharing activities of agents. Without sufficient opportunities, an increase in national saving rates would be expected, leading to a lower path of private consumption expenditures.

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Table 1: Unit root analysis

	Common component	Idiosyncratic component
Private consumption	-0.853	0.822
Disposable income	-3.229	-1.341
Share prices	-1.627	-1.729*
Housing prices	-2.887	-1.665*

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 a constant and a 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	25.79	29.83*	52.89*
r≤1	11.61	13.69	28.14
r≤2	1.80	4.70	16.15
r≤3			6.69

Note: Johansen (1995) trace statistics for the null hypothesis of no cointegration between the common components of private consumption, disposable income, and alternative wealth measures. 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

	Panel statistics	Group statistics
Variance ratio	2.403*	
Rho statistic	-3.755*	-2.806*
PP statistic	-2.940*	-2.508*
ADF statistic	-2.493*	-1.863*

Note: Pedroni (1999, 2004) tests for the null hypothesis of no cointegration between the idiosyncratic parts of private consumption and disposable income. The 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 at the 0.05 level of significance.

Table 4: Cointegration vector for common components

ML reduced rank estimator

	Equity prices	House prices	Equity and house prices
Disposable income	1.371 (0.154)	0.904 (0.307)	0.816 (0.120)
House prices		0.118 (0.092)	0.185 (0.035)
Share prices	0.129 (0.022)		0.053 (0.008)

Simple two step estimator

	Equity prices	House prices	Equity and house prices
Disposable income	1.516 (0.106)	0.986 (0.287)	0.772 (0.124)
House prices		0.130 (0.085)	0.189 (0.036)
Share prices	0.067 (0.015)		0.050 (0.008)

Note: Standard errors in parentheses. Cointegration parameters normalized by dividing through the coefficient of the common component of private consumption.

Table 5: Cointegration vector for idiosyncratic components

	Panel FMOLS	Panel DOLS
Disposable income	0.511 (0.028)	0.471 (0.028)

Note: Standard errors in parentheses. Cointegration parameters normalized by dividing through the coefficient of the idiosyncratic component of private consumption.

Table 6: Tests for weak exogeneity

	Equity prices	House prices	Equity and house prices
Consumptiom	1.690 (0.194)	7.489 (0.006)	7.986 (0.005)
Disposable income	0.063 (0.801)	0.027 (0.869)	0.025 (0.874)
House prices		1.452 (0.228)	8.828 (0.003)
Share prices	3.211 (0.073)		0.531 (0.466)

Note: Test statistics are distributed as chi-squared with 1 degree of freedom, *p*-values in parentheses. Under the null hypothesis the variable in the left column is weakly exogenous with respect to the cointegrating relationship. The long run is based on the common components.