The Impact of a Change in Real Estate Value on Private Consumption in Estonia

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Abstract

This paper examines the link between the value of residential real estate stock and private consumption using data from Estonia. Estonia has a high share of owner-occupancy due to the property reform that was begun after regaining independence in 1991, and has also seen large changes in real estate prices. A vector error correction model was constructed with private consumption, real estate stock value, GDP and household debt as the variables. The results suggest the presence of a long-run relationship, showing that changes in the value of residential real estate stock positively affect private consumption in the long run, i.e. the effect is permanent and has the expected positive sign.

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Keywords: house prices, real estate wealth effect, owner-occupancy, cointegration, vector error correction model

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1. Introduction

The value of real estate changes over time. Europe, North and Central America and parts of Asia witnessed a rapid increase in house prices before 2007 and subsequently large declines during the global financial crisis (GFC) of 2008–2010. An example of a very large drop in real estate prices is in the post-Soviet country Estonia, where house prices declined severely during the GFC: the fall compared to the boom price was 60–70%, which is much higher than it was in most other EU countries except Latvia and Lithuania. In 2004–2006 the Estonian real estate price index tripled and future prospects for the market were optimistic. In 2007, the number and value of real estate transactions started to decrease slowly as the global economy headed towards crisis. There was a slowdown in economic growth and a soft landing was predicted for Estonia, but the GFC of 2008–2010 was much more severe. The peak of the GFC for the real estate market was in July 2009. By the end of the period of the data sample in this paper, real estate prices had almost doubled.

Since housing influences the business cycle, the boom in house prices in many developed countries before the GFC and the following sharp price declines have attracted the attention of policymakers and researchers. It is important to study the topic because, if there is a significant impact of a change in real estate value on economic activity, then this would indicate that policymakers should keep a close eye on the dynamics of the housing market to identify and possibly prevent large booms and busts. The problem is that shocks to house prices have not been studied extensively in the case of post-Soviet countries, including the periods during and after the GFC.

According to theory, there are four channels through which asset prices can be considered to affect economic activity: 1) the wealth effect on consumption; 2) Tobin's Q effect on investment; 3) the balance sheet effect on private spending (via the credit channel); and 4) the confidence effect on private spending (Altissimo et al., 2005). Ludwig and Slok (2002) divide the first channel (the wealth effect on consumption) into five sub-channels: the realized wealth effect, the unrealized wealth effect, the budget constraint effect, the liquidity constraint effect and the substitution effect. The realized and unrealized wealth effects cause an increase in house prices and so increase private consumption. The budget constraint and substitution effects work in the opposite direction. The liquidity constraint effect depends on how well the financial market is functioning.

Total wealth is not homogenous, but consists of several components. The three main components of household wealth are tangible wealth (mainly housing), non-equity financial wealth and equity wealth. This paper focuses on housing wealth, specifically residential real estate wealth. Most of the literature has not divided wealth into these components for calculating elasticity or marginal propensity to consume (MPC). The current paper focuses on housing wealth, as in, for example, Case et al. (2005), Pachebo and Barata (2005) and Campbell and Cocco (2007). Housing wealth is an important component of household wealth. As shown below, Estonia has a remarkably high rate of owner-occupied residential real estate, while only approximately 1.5% of the population has an account in the stock market.[1]

Estonia is unique for its large share of owner-occupied residential real estate, largely because of the housing and property reform that was begun in 1991. During this reform,

[1] According to Talpsepp (2010), from 1 January 2004 to 30 June 2008, transactions from 20,758 different individual accounts were made on the Nasdaq OMX Tallinn.

much of the property that had been unlawfully expropriated during the first years of the Soviet occupation in Estonia (1940) and after World War II was returned to its former owners or their successors. In addition to this restitution, state-owned residential real property was privatized – government tenants had the chance to buy their homes using privatization vouchers (every permanent resident of Estonia was entitled to a privatization voucher according to their length of employment). This led to a very high share of owner-occupied homes[2]; nearly 90% of the housing stock was in private ownership in Estonia by 2012. According to Eurostat, more than half of the population in each EU member state lived in owner-occupied dwellings in 2012, ranging from 53.2% in Germany up to 96.6% in Romania. It appears that it is characteristic of the new member states to have high owner-occupancy rates. The owner-occupancy rate has been slightly declining in recent years in Estonia. Estonia was chosen as the subject of this research because of its uniqueness due to its historical background and also due to the existence of a good transaction register (described in the data section of this paper).

This paper seeks to ascertain the impact of a change in real estate value on aggregate private consumption. The aim of the paper is to investigate empirically how private consumption is affected by changes in real estate value. The research questions that this paper aims to answer are as follows: Will a change in the value of property lead to a change in consumer spending? How much will it change consumption?

Essentially, wealth is not exogenous with respect to consumption but, rather, jointly endogenously determined. Most existing literature, however, implicitly assumes that a large proportion of the fluctuations in housing wealth are exogenous and that the dynamics are not substantially affected by decisions on consumption and other macroeconomic variables (Slacalek, 2009). In this paper I use four variables (private consumption, real estate stock value, GDP and household debt) in a vector error correction model (VECM) approach and hence all the variables are first considered endogenous and are subsequently tested for weak exogeneity. Based on the impulse response functions, I estimate the long-term elasticity of private consumption with respect to a 1% change in real estate stock value. The results show that the estimated elasticity is higher than in most previous studies referred to in the literature review, but it is in agreement with the upper end of the range in the results of, for instance, Pachebo and Barata (2005).

This paper contributes to the literature in the following way. Using a good database of real estate transactions in a transition country which has experienced a large boom and bust, I examine how the developments in the housing market may have affected consumption behaviour. The topic has not been studied widely in post-Socialist transition countries and the results may be applicable to other CEE countries that have witnessed dramatic changes in the value of housing stock, serving as an example of the impact of a change in real estate value on consumers in post-Socialist transition countries. The topic of the present paper has been studied by Paabut and Kattai (2007) and Šonje et al. (2012), but their time series are shorter and the methods used are different in my paper. Impulse response functions make it easier to see the effect of different shocks on the adjustment path of the variables.

The remainder of the paper is organised as follows: in the next section, a brief literature review is presented. In section 3 the econometric methods used are described, in section 4

[2] In 1994, around 71% of all housing stock was owned by the state or local municipalities in Estonia, while at the beginning of 2002 the percentage was only 4.2%.

an overview of the data sources is given, and in section 5 the results are presented and discussed. The last section concludes.

2. Literature review

The effect of wealth on private consumption has traditionally been analyzed within the framework of the permanent income hypothesis or the life-cycle model (Ando and Modigliani, 1963; Friedman, 1957; Modigliani, 1971). Pachebo and Barata (2005) depart from the life-cycle theory of Modigliani, explaining consumption as a variable depending on wealth, beyond income. They argue that the traditional life-cycle model given in Ando and Modigliani (1963) implies that aggregate consumption is linearly related to labour income and wealth, but says nothing about the cointegration properties of the variables. Pachebo and Barata (2005) also argue that in the long run, trends in consumption are closely related to trends in income and wealth; in the short run, household consumption can deviate from this long-run equilibrium but will tend gradually to revert to equilibrium over time.

The life-cycle model of Gali (1990) implies that consumption, income and wealth variables may share a common trend, showing through the life-cycle factors that consumption changes should be predictable and smoother than in the infinite-horizon model. Modigliani (1971) advocates the significance of wealth effects on consumption. As indicated by Boone et al. (1998), subsequent evidence has presented some criticism of life-cycle theory. It is argued that this model takes no account of uncertainty in the future stream of revenue, as well as that the strength of any wealth effect should be linked to the distribution of wealth and the existence of liquidity constraints. The conventional analysis does not take into account the possibility that the variables are non-stationary or that there is reverse causality between, for instance, wealth and consumption (Pachebo & Barata, 2005). In the current paper, the order of the integration of the variables is tested and, since the variables are found to be I(1), cointegration analysis is performed.

Consumer intertemporal utility maximization under the lifetime resource constraint states that current consumption is proportional to total wealth (Altissimo et al., 2005):

$$C = MPC_W[A+H(Y)] \approx MPC_W A + MPC_Y Y \quad (1)$$

where C is consumption, A is real non-human wealth, H is real human wealth (the present value of expected labour income (net taxes)), MPC_W is marginal propensity to consume wealth and MPC_Y is marginal propensity to consume income. As can be seen in equation (1), within the framework of the life-cycle theory of permanent income, consumption is a function of human and non-human capital wealth (including labour income, transfer income, property income and financial wealth). Additionally, lags are put into the model. Equation (1) also shows that the weighted averages of non-human wealth and labour income can be presented as related to consumption in the long run because, by definition, labour income and human wealth are cointegrated (Altissimo et al., 2005).

Typically, the behaviour of utility-maximizing consumers is analyzed assuming rationality, perfect capital markets and the absence of distorting taxes or rigidities. In these ideal circumstances, where the composition of wealth can be changed without friction or cost, MPCs of different wealth components should be equal. In imperfect capital markets,

the properties associated with different wealth components may well affect their respective MPCs; the MPC of each wealth component depends positively on the liquidity of the asset. For example, if the only way for a homeowner to increase consumption, as a response to a rise in housing value, is first to sell the house, then it is clear that most homeowners will not react to the increase in the market value of housing (Altissimo et al., 2005). The possibility of taking out a new mortgage on the house affects MPC positively. A rise in house prices increases the value of the collateral available to households, stimulating consumption and housing investment by making it easier and less expensive to borrow against the value of a house (Aoki et al., 2002).

The negative effect comes from renters. When housing value rises, the value of potential future homes for renters to buy also rises, and so they must save more, and hence consume less, in order to buy a house someday. In addition, the rental payments for their current home may rise, which will also decrease the tenant's private consumption. It goes the opposite way in the case of a decrease in house prices. Muellhauer (1994), Altissimo et al. (2005) and Ludwig and Slok (2002) suggest that differences between assets based on liquidity and the distribution of homeownership could imply different aggregate MPCs – the higher the proportion of homeowners and the lower the proportion of households in the rental market, the larger the consumption response to a rise in house prices. This means that an increase in the homeownership rate could increase the probability that a positive wealth effect will outweigh negative income and substitution effects on consumption – the greater the share of homeowners, the greater the MPC of wealth. Therefore, Estonia should have quite a large MPC of real estate wealth, due to the large proportion of owner-occupancy.

Altissimo et al. (2005) argue that the MPC of non-equity financial wealth is likely to be the highest, compared to other types of wealth. Since equity prices are more volatile than house prices, households may find it more difficult to assess whether a change in their equity wealth is permanent or temporary. They are likely to be more cautious in adapting consumption plans to changes in equity wealth than to changes in housing wealth. In most countries, housing wealth is more evenly distributed than equity holdings, which are concentrated at the upper end of income distribution, meaning that, at the aggregate level, the wealth effect from housing may be expected to be more important than the effect from equity wealth. Case et al. (2005) argue that research to quantify the effect of changes in wealth on changes in consumption has largely used aggregate measures of wealth that emphasize the stock market, and does not attempt to measure housing wealth with any accuracy. House purchases are generally largely financed with a mortgage loan, but financial asset purchases are not. As a result, increases in house prices will give a greater investment return to households than an equal percentage change in, for example, equity prices (Tang, 2006). Since only a very small share of the population is active in the stock market in Estonia, it is not justified to use stock wealth to measure MPC of wealth and hence, in this paper, residential real estate market data is used.

Some theories imply that an increase in housing wealth due to higher real estate prices leads to increased consumption, while other theories suggest that this is not the case (Aben et al., 2012). Ludwig and Slok (2002) say that there is no a priori reason to expect a positive impact from changes in housing prices on consumption but, due to the deregulation of financial markets, it should be expected that the positive influence of housing prices on consumption has become stronger over time. Pachebo and Barata (2005) and Muellbauer and Lattimore (1995) state that an increase in house prices may or may not make households

better off because the positive effect for homeowners could be offset by the negative impact on renters. This means that when the value of residential real estate, for example, rises, the change in private consumption depends on how easily a property could be sold. If liquidity is high, private consumption rises more (since the homeowner can sell the house easily and hence cash out the increased value). Gan (2010) concludes that the main effect of housing wealth on consumption appears to stem from a reduction in precautionary saving.

It has been argued in the literature that the length of planning horizon is also important in calculating MPC of wealth. Following Poterba (2000), Altissimo et al. (2005) show that the shorter the horizon, the higher the MPC of wealth. On the other hand, a strong bequest motive affects the MPC of wealth in the same way as the lengthening of the planning horizon (Altissimo et al., 2005). Campbell and Cocco (2007), using data on individual households, also find evidence that the housing wealth effect is higher for older households.

There are alternative views regarding the effect of house prices on consumption – for example, Cristini and Sevilla (2014) discuss and compare the wealth effect hypothesis and the common factor hypothesis. They compare the results of two influential papers, Campbell and Cocco (2007) and Attanasio, Blow, Hamilton and Leicester (2009), which use the same data set but reach opposite conclusions as to the impact of changes in house prices on consumption. Browning et al. (2013) find, using Danish data, little evidence of a housing wealth effect on consumption. They suggest that house prices affect total expenditure through improved collateral, rather than directly through wealth. However, they also find that housing prices are stationary, which is not the case for Estonian data.

The central issue in analyzing wealth effects is timing. If the lag between a favourable shock to the household balance sheet and an increase in consumption spending takes many years to develop, then market fluctuations may have a limited impact on aggregate spending. If the link from net worth to consumption is powerful and immediate, then sharp changes in asset values may translate into sharp changes in consumer spending (Altissimo et al., 2005). Pachebo and Barata (2005) state that if consumption expenditures do not fully react immediately to changes in asset prices, then the wealth effect in the first periods will be smaller than the long-run effects. The error correction process will eventually bring actual spending into line with the long-run prediction of the life-cycle model. The paper by Aoki et al. (2002) shows that there is a strong co-movement between housing prices and consumption, and particularly between housing prices and consumer durables expenditure, and that house prices are not a source of fundamental shocks but are part of the transmission mechanism by which changes in short-term interest rates affect consumption, the output gap and inflation.

To give an overview of some studies that assess empirically the impact of housing wealth on consumption, I start with the study by Benjamin et al. (2004), which relies on US aggregate consumption and wealth data. The variables in their model are consumption, income other than from transfer payments, real estate wealth and financial wealth. Their results suggest that the MPC from housing wealth (0.08) is four times the size of that from financial assets (0.02). Case et al. (2005) rely on a panel of 14 countries observed annually and a panel of US states observed quarterly. The variables in their model are consumption, housing market wealth, stock market wealth and income. They find a statistically significant and rather large effect of housing wealth on household consumption. The housing wealth effect may be especially important in the preceding decades of the study of Case et al. (2005), as institutional innovations like second mortgages have made it simpler to extract cash from housing equity.

Altissimo et al. (2005) state that the variation in MPC of wealth estimates across studies, countries and estimation methods is disconcertingly large. It is quite difficult to compare different countries because their cultural and historical backgrounds may also have a significant influence on MPC.

Pachebo and Barata (2005) include in their model for nine EU countries the quarterly aggregate data for household consumption, disposable income, an equity price index, a residential price index, the unemployment rate, the short-term interest rate and the inflation rate. The selected time period differed across countries and the authors faced a serious lack of data and a short range of the time series for some of the smaller EU countries. In relation to residential prices, their results are significant for all the countries considered, with long-run consumption elasticity of residential prices between 6 and 21 per cent.

Gradner and Gstach (2006) study different papers on the effect of real estate wealth on consumption and find that in a series of recent papers, the real wealth effect on consumption from changes in housing prices has been found to be statistically significant and, for European countries, to exceed the stock market wealth effect considerably. Pachebo and Barata (2005) also state that the housing market appears to be more important than the stock market in influencing consumption. They report corresponding consumption elasticity of between 0.1 and 0.2 for various European countries, roughly the range of figures reported by Case et al. (2005). Greenspan (1999) gives a figure of roughly 0.05 for the USA and similar figures are reported for the G7 countries by Girouard and Blöndal (2001). Gan (2014) calculates the elasticity of consumption with respect to a change in housing value to be 0.17, based on data from Hong Kong. In some studies, self-reported housing values are used as a proxy for housing value, while in others only some areas of a country are covered, which makes it difficult to compare the results of different studies.

Muellbauer and Murphy (1994), using UK regional consumption data, find a negative effect from house prices. Girouard and Blöndal (2001) examine the impact of housing wealth on household consumption for France, Italy and the UK. Their results show that housing wealth exerts a positive influence on household consumption in France and in the UK in the short and long run, while Italy presents a negative relationship between those variables. Considering seven countries (Finland, Germany, Ireland, Italy, the Netherlands, Sweden and the UK), Henley and Morley (2001) suggest a significant degree of diversity in consumption functions and impacts from housing wealth. Ludwig and Slok (2004) evidence a significant long-run impact from stock market and housing market wealth to consumption. Aoki et al. (2002) say that house prices matter because houses can be used as loan collateral against which households borrow to finance housing investment or consumption (Pachebo and Barata, 2005).

The paper by Carroll et al. (2010) presents a simple new method for measuring wealth effects on aggregate consumption. This method exploits the stickiness of consumption growth to distinguish between immediate and eventual wealth effects. Using US data, they estimate that the immediate (next-quarter) MPC from a change in housing wealth is about 0.02, with a final eventual effect of around 0.09.

Altissimo et al. (2005) show estimates of MPC and elasticities of wealth components for different countries and by different authors. Their results concerning the topic of the current paper are shown in Table A1 in Appendix A. Unfortunately, these results cannot be directly compared because their data is based on a different foundation and the essence of the calculation of MPC and/or elasticity also differs across studies. For example, Ludwig and

Slok (2002) use panel data for 16 OECD countries and use a pooled estimation technique, which fixes the elasticity of wealth components to be equal within groups (countries have been divided into two groups: market-based economies and bank-based economies). Ludwig and Slok (2002) have not calculated MPC of housing wealth. Altissimo et al. (2005), however, have calculated MPC of housing wealth based on Ludwig and Slok (2002) and data on total household wealth decomposition into financial and housing wealth, and reach the conclusion that the corresponding MPCs of housing wealth would be around 0.075 euro cents for Germany and around 0.05 euro cents for France and Italy. In market-based economies, the estimates would be smaller; for example, 0.013 cents per euro for the Netherlands, 0.117 for the UK and 0.02 for the USA.

Sierminska and Tahktamanova (2007) have also gathered several authors' estimates of MPC and elasticities of different wealth components (see Table A2 in Appendix A). As can be seen, some authors have calculated MPCs of housing wealth, others have calculated elasticities. Some have used aggregate data, others household level data. The elasticities of consumption with respect to changes in housing wealth range from 0.017 up to 0.17. The MPCs of housing wealth are in the range of 0.015 up to 0.14. The elasticities and MPCs calculated for the same country differ also from author to author.

The relationship between real estate value and private consumption in Estonia has been previously investigated in three papers. Paabut and Kattai (2007) use a consumption function similar to the consumption function in the macro model EMMA of the Bank of Estonia. They use data up to 2006 and, using the Granger-Engle two-step methodology, find that the elasticity of private consumption with respect to a change in real estate value is 0.011. The present paper includes the years of the GFC and some years after that, having a much longer time series (73 quarters). Also, the method, used in the current paper is different. In 2012, Šonje et al. used data for four European post-transition countries, including Estonia. They use three variables (private consumption, residential real estate value and net wages) and the elasticity of private consumption with respect to a change in real estate value is found to be 0.04. They use data up to the first quarter of 2010 and two methods are applied: threshold error correction and vector error correction models. In my paper, GDP is used as a proxy for income and household debt is used as an additional variable. Although it is not customary for Estonians to take, for example, a second mortgage on their residential space, we can expect that there is a relationship between housing value and private consumption in Estonia. When the market value of their property rises, the homeowner may reduce precautionary savings. Estonian banks also offer a mortgage loan that does not necessarily need to be used to buy real estate. According to Aben et al. (2012), whose paper focuses on housing equity withdrawal in Estonia, the secured loan stock rose far more than the household net investment in housing during the years 2004–2008.

3. Methodology

The methodology of this paper follows the vector autoregression modelling approach (Sims, 1980). First, tests for the order of the integration of the variables are carried out to describe the characteristics of the data. If all the time series appear to be I(1), cointegration techniques are used. In the current paper, the Ng–Perron unit root test and Kwiatkowski–Phillips–Schmidt–Shin (hereinafter KPSS) stationarity test are used. A maximum of two lags is used. The Ng–

Perron test takes account of the fact that the bias in the autoregressive coefficients is highly dependent on the lag and adapts to the type of deterministic components present (Ng and Perron, 2001). This test uses a modified information criterion which selects the lag length according to the sample size (as opposed to the Akaike information criterion and Scwartz or Bayesian information criterion, which tend to choose lag values that are generally too small for unit root tests to have good size). Kwiatowski et al. (1992) argue that the way in which classical hypothesis testing is carried out makes it more likely that the null hypothesis is not rejected unless there is strong evidence against it. They propose that an alternative explanation for the common failure to reject a unit root is simply that most economic time series are not very informative about whether or not there is a unit root. Kwiatowski et al. (1992) provide a test of the null hypothesis of stationarity against the alternative of a unit root.

As will be shown, the variables are I(1), hence the Johansen cointegration approach (trace and maximum eigenvalue tests) is used. The choice of lag length is based on the Akaike information criterion. If the results of the trace and maximum eigenvalue tests show that there is cointegration between the series of the model, cointegration relation is highlighted using a VECM, which is used because it allows the capture of the long-term equilibrium and short-term relationships between variables and study of how deviations from the long-run equilibrium are corrected. Also, the multicollinearity effect is reduced in the error-correction form of the VAR model, because the first differences of the variables tend to be less correlated than levels.

Since this paper is focused on how the real estate variable affects the private consumption variable, it is useful to normalize the cointegration relationship to private consumption by setting the coefficient to unity. First a long-run relationship is shown and then short-run adjustments are calculated (private consumption adjusts to its long-run relationship with lags). The equation for the VECM (following e.g. Lütkepohl and Krätzig, 2004) is as follows (matrix notation in bold):

$$\Delta X_t = \Phi X_{t-1} + \mu + \sum_{j=1}^{k} \alpha_{t-j} \Delta X_{t-j} + \varepsilon_t \qquad (2)$$

where X_t is the vector of variables in the VECM ($\ln PC_t$, $\ln H_t$, $\ln GDP_t$, $\ln HD_t$) Φ is the vector of the adjustment coefficients of the long-term equilibrium equation, μ is the vector of intercepts of the short-run error correction model and α_{t-j} are the shortrun coefficients (j is the number of lagged periods). The variable *PC* is seasonally adjusted private consumption, *H* is seasonally adjusted residential real estate stock value, *GDP* is seasonally adjusted *GDP* and *HD* is seasonally adjusted household debt. I expect the signs of the long-term coefficients of all these variables to be positive. The variables are characterized in more detail in the data section of the paper.

Weak exogeneity of each variable for the long-run parameters is also tested using VECM restrictions on adjustment coefficients. If an adjustment coefficient appears not to be significant, the respective restriction is kept in the model and the variable is considered to be a pushing force in the system but not being pushed by the system. In order to see how responsive each variable in the VECM is to shocks to each of the variables, impulse responses are calculated and shown graphically. Based on the impulse response functions, the long-term elasticity of private consumption with respect to a change in real estate stock value is estimated.

The MPCs of real estate value, based on the elasticity from the long-term relationship, are calculated, varying each period with the variables PC_t and H_t. The MPCs show by how many euros private consumption changes if the real estate stock value index changes by one unit.

4. Data

Based on equation (1) shown in the literature review section, I construct a model with four quarterly variables: private consumption, residential real estate stock value, GDP and household debt. The data for private consumption and GDP is from Statistics Estonia and the data for household debt (the balance of housing loans and consumer loans) is taken from the statistics of the Bank of Estonia. The data for real estate value per square metre is from the Estonian Land Board and the number of square metres of existing residential real estate is calculated based on the population censuses of 2000 and 2011. All of the time series are in real values (nominal data is deflated, using the consumer price index from Statistics Estonia) and in natural logarithms, since using elasticities makes it easier to include households with real property of different value levels in the variable. As the main variable of interest of this paper is real estate stock value, this variable is explained more thoroughly below.

In order to do any calculations based on the value of real estate, it is necessary to specify how the value is estimated. Value is an estimation showing how much a property is worth. Real estate value in everyday appraisal practice (for residential real estate) is usually estimated using the sales comparison approach, meaning that the market value of residential real estate is based on the (historical) sales prices of comparable real properties; that is, value is estimated based on the magnitude of transactions. There is a full database of the sales transactions of real estate in Estonia: the transaction register of the Estonian Land Board. The data for the total value of residential real estate transactions is taken from that register, which is a database of all real property transactions in Estonia, including, for example, the address, the land size and the percentage of built area, but not including information about the condition of housing. The data is electronically sent to the Land Board by notaries. The transaction register was founded in 1997 and is part of the land cadastre. Only real estate appraisers who have a special licence issued by the Land Board can apply for data from the transaction register. Most of the larger real estate companies have a licenced appraiser and therefore they can buy data from the transaction register and in practice use it in their everyday real estate appraisal work. For this paper, public aggregated and systemized data has been used, while data for transactions has not been used. Quarterly data from 1997Q1 to 2015Q1 was available, so 73 periods are used in the analysis of this paper for all the time series, based on the availability of the transaction data starting from 1997.

From 1 January 2002 until 19 July 2003, only data for transactions with the state's pre-emptive purchase right of real properties situated or reaching the building prohibition zone or reserve area was collected. For that reason, the number of transactions in the years 2002 and 2003 is very low and cannot be used in the analysis in its original form. Starting from the second half of 2003, data on transactions of apartment properties was also included in the transaction register. Only apartment properties marked as residential space are included in the apartment properties data (not all apartment properties are actually residential space). Data about apartment properties where the seller was the municipal government was also excluded.

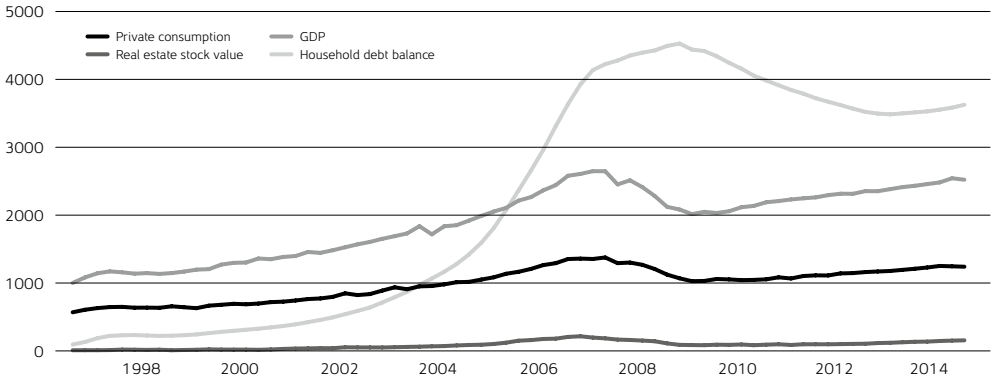
In order to calculate estimates of the value of residential real estate, first the residential housing space (in square metres) in Estonia had to be found, and to that end the results of the population censuses of 2000 and 2011 were used. To obtain data for the other years, the number of square metres was linearly interpolated and extrapolated. It appears that from 1997 approximately 33–41% of the residential housing space has been in Harju County

(including Tallinn). So the proportion was one-third but it has been growing steadily by around 3% a year. This is not surprising, since most of the new residential space has been built around Tallinn. In Tartu (the second-largest city in Estonia) and surroundings, the growth has been a bit less than 3% a year, being approximately 10% in 1997 and approximately 11.5% in 2014. The proportion of the total value of real estate transactions involving apartment properties is approximately 75% in Harju County (including Tallinn). The average proportion from all time periods of the value of other residential housing transactions is approximately 64% in Harju County (including Tallinn). So I make the assumption that 70% of the value of residential real estate housing stock is in Harju County (including Tallinn).

Based on the total value of residential real estate transactions, the transaction value per square metre was calculated for each county, separately for apartment properties and other residential housing properties. Unfortunately, the data for each county for the whole period was not available as there have been less than five transactions. For Harju County, data for total value of transactions and value of transactions per square meter were available from 2003Q3 until 2014Q4. From 1997Q1 until 2001Q4, information about transactions in regard to apartments was not available (because apartment properties were only included in the register from 2003), but it was available for other residential housing for that period. The values per square metre for the latter are available through the public search engine of the Land Board. It must be noted that the value of other residential housing besides apartment properties in the transaction register data reflects the value per square metre of the land on which the house is situated, not the number of square metres of the building. Therefore, the value per square metre is likely to be heavily underestimated.

In order to obtain the value of residential real estate (housing) stock, the respective index was multiplied by the respective type of space (in square metres). The seasonally adjusted value of the stock of residential real estate (value per square metre indexed, 1997=1, and multiplied by total number of square metres of residential real estate) is shown in Figure 1.

Figure 1. Seasonally adjusted data (EUR in millions, index for real estate stock value)



Note: All the variables, except real estate stock value, are in millions of euros; real estate value per square metre is indexed and multiplied by the number of square metres of existing residential real estate stock.

Since, as mentioned above, the data for total real estate value from 2002Q1 to 2003Q2 was not comparable with that of other periods, linear interpolation was used to recalculate the data for those six periods. Because of the nature and frequency of the data, seasonality was an issue; therefore, adjustments for seasonality were made using the X12 approach in Eviews.

To find the order of integration of the time series, I carried out unit root tests. The results of Ng–Perron unit root tests show that with the intercept as the deterministic component (Table 1), the null hypothesis of a unit root for all the time series cannot be rejected. I did the test also in first differences and, since in that case the null hypothesis of a unit root was rejected, I concluded that the time series are I(1). For robustness, I also tested for unit roots with the intercept and trend as the deterministic components, and these results (see Appendix B, Table B1) also suggest that all the time series are non-stationary.

Table 1. Ng–Perron unit root tests

	MZa	MZt	MSB	MPT
ln*PC*	0.284	0.231	0.814	41.981
ln*H*	2.696	1.639	0.608	38.750
ln*GDP*	0.459	0.392	0.853	47.200
ln*HD*	0.710	1.272	1.791	195.574

Notes: Intercept is taken as the deterministic component. The asymptotic critical values for 1%, 5% and 10% levels for MZa are −13.800, −8.100 and −5.700 respectively, for MZt are −2.580, −1.980 and −1.620 respectively, for MSB 0.174, 0.233 and 0.275 respectively and for MPT 1.780, 3.170 and 4.450 respectively (Ng and Perron, 2001).

For robustness analysis, an alternative test, the KPSS stationarity test, was also used. The results are shown in Table 2. The null hypothesis of the KPSS test is that the observable time series is stationary around a deterministic trend (Kwiatowski et al., 1992). The null hypothesis is rejected for all the time series with the intercept as the deterministic component, which means that all the time series are non-stationary. For robustness, the KPSS test was done also with the intercept and trend as the deterministic components (see Appendix B, Table B2), and these results also suggest that the time series are non-stationary.

Table 2. KPSS stationarity tests

	LM-statistic
ln*PC*	0.921
ln*H*	0.881
ln*GDP*	0.993
ln*HD*	1.040

Notes: Intercept is set as the deterministic component. The asymptotic critical values for 1%, 5% and 10% levels are with intercept and trend 0.216, 0.146 and 0.119 respectively and with intercept 0.739, 0.463 and 0.347 respectively (Kwiatowski et al., 1992).

Since the null hypothesis of a unit root was not rejected according to the Ng–Perron tests and the null hypothesis of stationarity was rejected according to the KPSS tests, and the time series appeared not to be of an order higher than one, then the time series can be considered to be integrated on the order one. Hence it is justified to test for cointegration and use the VECM to find the long-term equilibrium relation and the short-term adjustment.

5. Results

After finding unit roots for all the time series, cointegration was tested in order to find whether a VECM could be used and hence the long-term and short-term relationships between the variables could be modelled. The number of lags was chosen according to the lag length that minimizes the Akaike information criterion in the VAR model in levels minus one, because since VECM is in differences, the number of lags is smaller by one than in VAR in levels. The criterion suggested two lags; a lag length of two is in many cases sufficient to describe a rich dynamic structure (Juselius, 2006). According to the trace test and maximum eigenvalue test of the Johansen cointegration approach, one cointegration relationship with two lags is found using the intercept and no trend in the cointegration equation (Appendix B, Table B3). This means that, although all the time series appear to be non-stationary, there is a linear combination of the variables that is stationary, i.e. the residuals of this relationship are stationary.

After detecting cointegration between the time series, the relation between the four variables was modelled using a VECM. First the long-term estimates are presented and then short-term error correction estimates are shown. Earlier work on the relationship between residential real estate and private consumption in Estonia (Paabut and Kattai, 2007) has found cointegration in the long-term relationship; that particular paper was written before the GFC.

Equation (3) shows the long-term equilibrium relationship of the cointegrating variables. The coefficients of all the variables are statistically significant and the signs are positive, as expected based on theory and the results of most former studies on the topic. The positive coefficients of real estate stock value and GDP could already be expected based on the life-cycle model, which assumes a positive relationship between wealth, income and consumption. The positive coefficient of household debt can be explained by the increasing importance of consumer loans and, to some extent, also housing equity withdrawal.

$$\ln PC_{t-1} = 5.755 + \underset{(0.032)}{0.068}\ln H_{t-1} + \underset{(0.092)}{0.602}\ln GDP_{t-1} + \underset{(0.014)}{0.041}\ln HD_{t-1} \quad (3)$$

The short-term VECM estimates are shown in Table 3. According to the row of adjustment coefficients, 30.8% of the disequilibrium is corrected each quarter by changes in private consumption, 13.3% of the disequilibrium is corrected each quarter by changes in real estate stock value, 18.5% is corrected by changes in GDP and 45.6% by changes in household debt. The adjustment coefficients for real estate value and GDP are not statistically significant.

Table 3. VECM short-term estimates

	$\Delta \ln H$	$\Delta \ln GDP$	$\Delta \ln HD$	$\Delta \ln PC$
Adjustment coefficient	−0.133 (0.902)	−0.185 (0.142)	0.456 (0.106)	−0.308 (0.141)
$\Delta \ln PC_{t-1}$	−0.271 (1.121)	0.139 (0.176)	−0.363 (0.132)	−0.083 (0.175)
$\Delta \ln PC_{t-2}$	0.354 (0.996)	0.421 (0.157)	−0.175 (0.117)	−0.072 (0.155)
$\Delta \ln H_{t-1}$	0.031 (0.271)	0.054 (0.043)	0.115 (0.032)	0.046 (0.042)
$\Delta \ln H_{t-2}$	0.046 (0.260)	0.034 (0.041)	0.039 (0.031)	0.091 (0.041)
$\Delta \ln GDP_{t-1}$	1.661 (1.044)	−0.284 (0.164)	0.237 (0.123)	0.179 (0.163)
$\Delta \ln GDP_{t-2}$	0.344 (1.000)	0.041 (0.157)	0.179 (0.117)	0.130 (0.156)
$\Delta \ln HD_{t-1}$	−0.060 (0.851)	0.402 (0.134)	1.050 (0.100)	0.314 (0.133)
$\Delta \ln HD_{t-2}$	−0.073 (0.703)	−0.309 (0.111)	−0.463 (0.083)	−0.158 (0.110
Constant	0.016 (0.024)	0.003 (0.004)	0.013 (0.003)	−0.002 (0.004)

Notes: Intercept is set as the deterministic component, no VEC restrictions. Standard errors in parentheses.

Tests for weak exogeneity were performed next. If the null hypothesis for an adjustment coefficient being zero is not rejected, the variable in question can be considered a driving variable in the system: it pushes the system, but is not being pushed by it (Juselius & MacDonald, 2000). The results of the exogeneity test are shown in Table 4. As the null hypothesis of the adjustment coefficients for real estate stock value and GDP being zero together cannot be rejected, these adjustment coefficients are not significant at the 0.05 confidence level (together as well as separately). This means that, according to the test, those two variables influence the long-term stochastic paths of other variables in the system, but they are not influenced by the other variables. These restrictions are kept in the model, and real estate stock value and GDP are considered weakly exogenous for the long-term parameter. It could be argued that in an actual economy GDP depends, among other things, also on private consumption, but in this paper I consider GDP weakly exogenous, based on the statistical reasoning.

Table 4. Results for the weak exogeneity test of VECM variables

Restriction	Chi-square	Probability
$A(\Delta \ln PC) = 0$	4.475	0.034
$A(\Delta \ln H) = 0$	0.021	0.885
$A(\Delta \ln GDP) = 0$	1.514	0.218
$A(\Delta \ln HD) = 0$	15.186	0.000
$A(\Delta \ln H) = 0, A(\Delta \ln GDP) = 0$	1.633	0.442

Note: A is the adjustment coefficient of the respective variable.

After putting the restrictions of the adjustment coefficients into the model, the long-term equation becomes:

$$\ln PC_{t-1} = \underset{}{5.000} + \underset{(0.030)}{0.061}\ln H_{t-1} + \underset{(0.092)}{0.652}\ln GDP_{t-1} + \underset{(0014)}{0.033}\ln HD_{t-1} \quad (4)$$

The coefficients' magnitudes and signs have not changed much compared to the model with restrictions of adjustment coefficients and the coefficients are still statistically significant. The short-term VECM estimates of the adjusted model are shown in Table 5, where it can be seen that the adjustment coefficients of real estate stock value and GDP are set to zero, as the previous test suggested that they are weakly exogenous.

Table 5. VECM short-term estimates

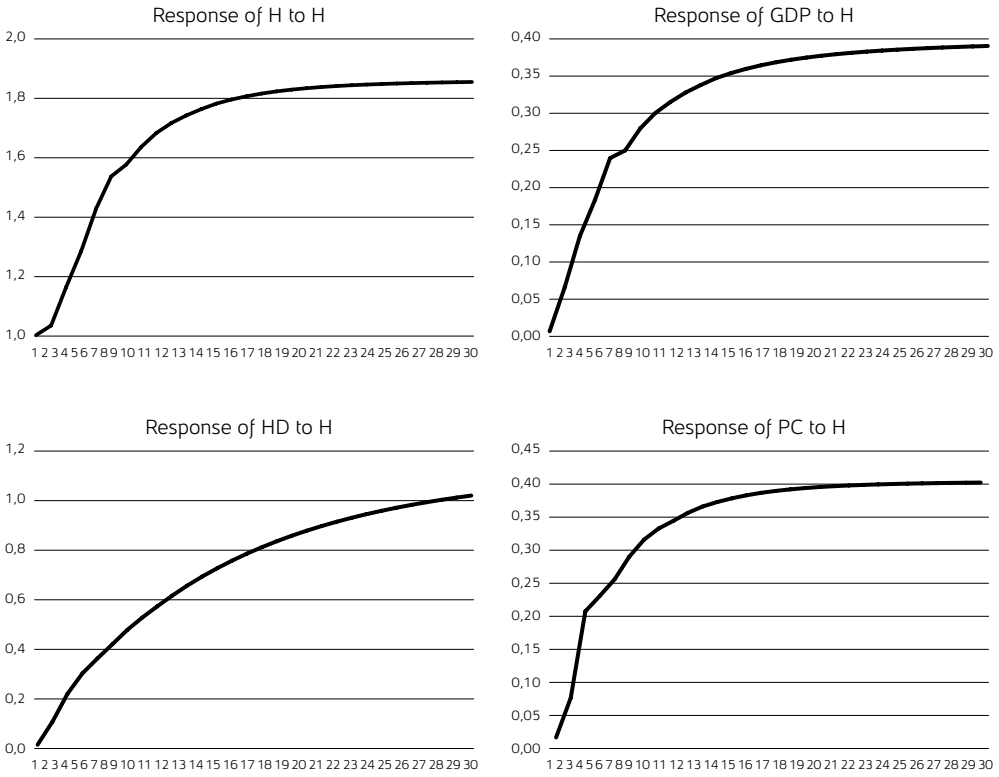
	$\Delta\ln H$	$\Delta\ln GDP$	$\Delta\ln HD$	$\Delta\ln PC$
Adjustment coefficient	0 –	0 –	0.504 (0.102)	−0.225 (0.114)
$\Delta\ln PC_{t-1}$	−0.169 (1.122)	0.119 (0.177)	−0.361 (0.132)	−0.078 (0.175)
$\Delta\ln PC_{t-2}$	0.410 (0.992)	0.407 (0.157)	−0.168 (0.117)	−0.074 (0.155)
$\Delta\ln H_{t-1}$	0.026 (0.271)	0.055 (0.043)	0.113 (0.032)	0.046 (0.042)
$\Delta\ln H_{t-2}$	0.042 (0.260)	0.035 (0.041)	0.039 (0.031)	0.090 (0.041)
$\Delta\ln GDP_{t-1}$	1.542 (1.065)	−0.270 (0.168)	0.254 (0.125)	0.162 (0.166)
$\Delta\ln GDP_{t-2}$	0.266 (1.007)	0.052 (0.159)	0.187 (0.119)	0.120 (0.157)
$\Delta\ln HD_{t-1}$	−0.011 (0.851)	0.393 (0.135)	1.051 (0.100)	0.316 (0.133)
$\Delta\ln HD_{t-2}$	−0.056 (0.704)	−0.309 (0.111)	−0.468 (0.083)	−0.153 (0.110)
Constant	0.014 (0.024)	0.003 (0.004)	0.013 (0.003)	−0.002 (0.004)

Notes: Intercept is set as the deterministic component. VEC restrictions: the adjustment coefficients of real estate stock value and GDP are set to zero. Standard errors in parentheses.

In order to distinguish between permanent and transitory shocks to the system and to see the estimated responses of all the variables to an unexpected 1% increase in real estate stock value, impulse response functions were calculated. I ordered the variables according to decreasing exogeneity, starting with real estate stock value, since I assume that in the first period real estate stock value is affected only by its own shock, not by shocks in other variables. The next variables are GDP, household debt and private consumption. I assume that if real estate value grows, people feel that they are wealthier and immediately start consuming more. They take out, for example, a consumer loan. Additionally, if their income grows, they also immediately start consuming more. For robustness, I also tried other orders (placing real estate stock value as the second, third or fourth variable), but the results do not change very much. Hence the model is not very sensitive to the ordering of variables.

The impulse responses in Figure 2 show that a 1% shock in real estate stock value leads to a permanent increase in private consumption (see the bottom right graph), GDP and household debt (and on itself).

Figure 2. VECM impulse responses to 1% shock in real estate stock value

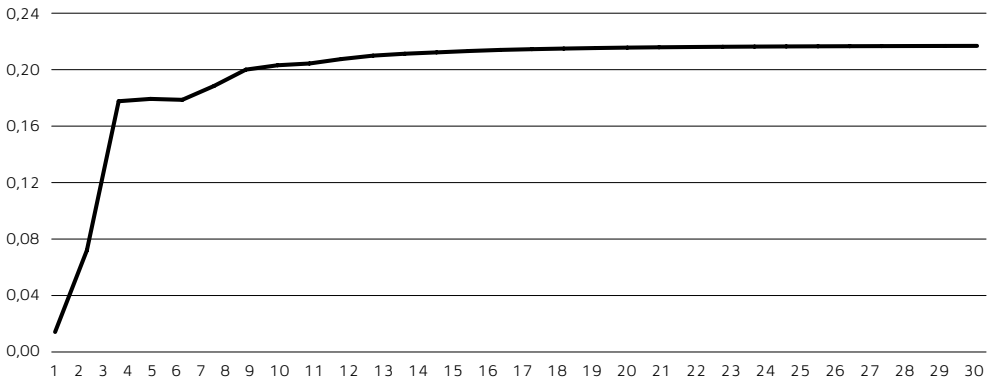


Notes: Intercept is set as the deterministic component. VEC restrictions: the adjustment coefficients of real estate stock value and GDP are set to zero.

The top left graph shows the response of real estate stock value and the other three graphs depict the responses of GDP, household debt and private consumption to an autoregressive shock in real estate stock value. The bottom right graph shows the response of private consumption to a 1% shock in real estate stock value. There is a small immediate response and a large response in the second half of the first year. After that the effect becomes smaller, until from the end of the fourth year it remains at the attained level. Hence, the effect of a shock in real estate stock value on private consumption is permanent.

The long-term elasticity of private consumption with respect to a change in real estate stock value was also calculated based on the IRFs and is plotted on Figure 3. It appears that a 1% change in real estate stock value leads to a 0.217% increase in private consumption in the long run. The initial effect is small, but it builds up in two years.

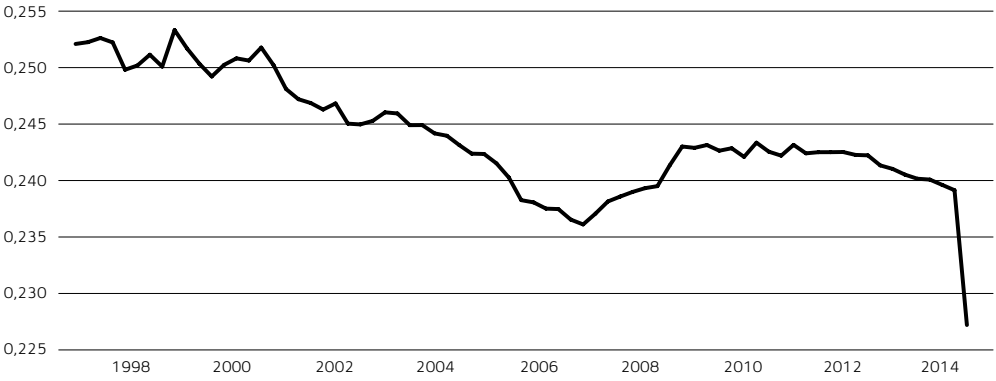
Figure 3. Long-term elasticity of private consumption with respect to real estate stock value



The estimated elasticity is large, compared to the findings of most previous authors. It may be assumed that one reason for this could be the high homeownership rate in Estonia. Another reason is likely the popularity of consumer loans in Estonia.

Quarterly MPCs were also calculated and they show how much private consumption changes when real estate stock value changes by one unit. The results are shown in Figure 4. It can be seen that the MPC fell remarkably during the GFC. This could show that people in Estonia were very cautious about spending money during the GFC – private consumption decreased more than the value of real estate stock.

Figure 4. The long-term marginal propensity to consume



For a robustness check, I replaced GDP with net personal wages as a proxy for income. Cointegration was detected between the variables of that alternative model also, but the long-term elasticities of private consumption with respect to a change in real estate stock value were not statistically significant when the trend and intercept were used as deterministic components. When only the intercept was used as a deterministic component, the elasticity was larger than one and hence is not economically reasonable; an elasticity of private consumption with respect to a change in real estate stock value larger than one would mean that if real estate value rose by 1%, then private consumption would rise by more than that. This would mean that a change in housing value would have an enormous effect on

consumption and also that the MPC would be larger than one. This is not supported by theory, nor has it been detected by previous authors. The results of the robustness check are available from the author upon request.

6. Conclusions

This paper has focused on the effect on private consumption of changes in residential real estate stock value in Estonia, a post-Soviet country with a very high rate of homeownership. Four variables have been used: private consumption, real estate stock value, GDP and household debt; as VECM has been used, all these variables were considered to be endogenous. One cointegration relationship was found and hence there is a linear combination of the time series that is stationary.

Long-term and short-term cointegrating equations were estimated simultaneously and weak exogeneity tests were run to see which variables were pushing the system but were not pushed by it. Impulse responses were also studied to distinguish between permanent and transitory shocks, to examine the response of each variable to shocks in other variables and to show the long-term elasticity of private consumption with respect to a change in real estate stock value.

According to the impulse response functions, a shock in the real estate stock value has a positive and permanent impact on private consumption; the long-term elasticity, based on impulse response functions, appeared to be 0.217. This high value of elasticity may be at least partly due to the high homeownership rate and the popularity of consumer loans in Estonia. Research in Estonia conducted before the GFC concluded that the elasticity was 0.010 (Paabut and Kattai, 2007). By comparison, the period under observation in this study is much longer and the GFC and the years following it included. In the current paper, all variables were assumed to be endogenous, as opposed to Paabut and Kattai (2007), who used the Granger-Engle two-step methodology, which assumes one variable is endogenous and the others are exogenous. Another innovation of this paper is the estimation and presentation of impulse response functions, which allows distinguishing between and showing the persistence of the impact of a shock to housing stock value in relation to other variables.

A broader implication of the current paper is that, in a country where housing wealth is a very important part of household wealth, it is important to follow developments in the real estate market, since it has a significant impact on consumption. Also, the analysis is based on data from a country which witnessed a very large price increase before the GFC and a marked price fall during the GFC. This means that policymakers should keep a close eye on the dynamics of the housing market to identify and possibly prevent large booms and busts. The results of this paper could also be applicable to other Central and Eastern European countries that have experienced large booms and busts.

Further research could include a similar study for the three Baltic countries (separately as well as a panel regression) and possibly for other post-Socialist countries as well. There could also be a quantitative study on the influence of the owner-occupancy rate on private consumption.

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Appendix A
Estimates of MPCs and Elasticities in Different Countries

Table A1. Estimates collected by Altissimo et al. (2005)

	BE	FR	DE	IT	NL	PT	SP	SE	UK	US	CA	JP
Ludwig and Slok (2002)												
MPC_W		1.4	2.0	3.0					4.9	4.0	4.0	4.0
e_H		0.11	0.11	0.11	0.03			0.03	0.03	0.03	0.03	0.11
Bertraut (2002)												
MPC_W									4.3	5.4	8.3	
MPC_H										9.7		
e_H									0.09	0.14	0.16	
Labhard, Sterne and Young (2005)												
MPC_W	0.7	0.8	7.8	2.8	1.3	-1.0	3.6		5.6	3.7	7.8	4.2
e_W	0.03	0.10	0.13	0.08	0.06	-0.02	0.07		0.16	0.12	0.19	0.16

Source: Altissimo et al. (2005).
Note: If the author(s) have calculated MPC and/or elasticity of housing wealth separately from total wealth, the indicator has been shown here. Otherwise MPC/elasticity of total wealth has been shown.

Table A2. Estimates collected by Sierminska and Takhtamanova (2007)

	MPC_H	e_H	Country
Davis & Palumbo (2001)	0.08		U.S.A.
Pichette & Tremblay (2003)	0.06		Canada
Carroll (2004)	0.09		U.S.A.
Case, Quigley & Schiller (2005) (aggregate data)		0.11–0.17	Panel of 14 developed countries
Dvornak & Kohler (2003)	0.03		Australia
Case, Quigley & Schiller (2005) (state-level data)		0.05–0.09	U.S.A.
Disney, Henley & Jevons (2003)	0.09–0.14		U.K.
Campbell & Cocco (2005)		0.017	U.K.
Grant & Pelton (2005)	0.014		Italy
Lehnert (2004)		0.04–0.05	U.S.A.
Bostic, Gabriel & Painter (2005)		0.06	U.S.A.
Bover (2005)	0.015		Spain

Source: Sierminska and Takhtamanova (2007).

Appendix B

Table B1. Ng–Perron unit root tests (intercept and trend)

	MZa	MZt	MSB	MPT
ln*PC*	−2.393	−1.007	0.421	34.468
ln*H*	−4.229	−1.265	0.299	19.786
ln*PW*	−5.360	−1.571	0.293	16.787
ln*GDP*	−5.123	−1.523	0.297	17.457
ln*HD*	0.148	0.088	0.592	79.693

Note: The asymptotic critical values for 1%, 5% and 10% levels for MZa are −23.800, −17.300 and −14.200 respectively, for MZt are −3.420, −2.910 and −2.620 respectively, for MSB are 0.143, 0.168 and 0.185 respectively and for MPT 4.030, 5.480 and 6.670 respectively (Ng and Perron, 2001).

Table B2. KPSS stationarity tests

	LM-statistic (intercept & trend)
ln*PC*	0.221
ln*H*	0.192
ln*PW*	0.225
ln*GDP*	0.227
ln*HD*	0.247

Note: The asymptotic critical values for 1%, 5% and 10% levels are with intercept and trend 0.216, 0.146 and 0.119 respectively and with intercept 0.739, 0.463 and 0.347 respectively (Kwiatowski et al., 1992).

Table B3. Johansen test

Trace test

Hypothesized no. of cointegrating equations	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.416	64.359	47.856	0.001
At most 1	0.243	26.681	29.797	0.110
At most 2	0.091	7.192	15.495	0.555

Maximum eigenvalue test

Hypothesised no. of cointegrating equations	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.**
None *	0.416	37.678	27.584	0.002
At most 1	0.243	19.488	21.132	0.084
At most 2	0.091	6.666	14.265	0.529

* denotes rejection of the hypothesis at the 0.05 level ** MacKinnon-Haug-Michelis (1999) p-values