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Housing Wealth Effect on Personal Consumption: Empirical Evidence from European Post-Transition Economies

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Abstract

Since housing is the most important component of non-financial personal wealth in most countries and in European post-transition economies in particular, this study estimates the impact of changes in the housing wealth effect on personal consumption in the long-run and the short-run in light of the permanent income hypothesis. In order to asses this relationship empirically, a pooled mean group estimator of dynamic heterogeneous panel data on a sample of six European post-transition economies, namely Bulgaria, Croatia, the Czech Republic, Estonia, Lithuania and Slovenia, was used. The panel was unbalanced with the longest time span for Estonia, ranging from the first quarter of 1997 to the third quarter of 2012. The main result of the analysis is the statistically significant long-run and short-run housing wealth effect in the analyzed economies, with the latter being less pronounced than the former, though these results are somewhat sensitive to the choice of estimation method.

1. Introduction

The conventional macroeconomic literature links personal consumption with income and wealth. This simple consumption function model with household income and wealth as only endogenous variables is motivated by several well-known theories, including the permanent income hypothesis (Friedman, 1957) and life-cycle hypothesis (Ando and Modigliani, 1963). Simply stated, personal consumption is determined by income and asset wealth, which implies real estate and stock ownership. However, in this paper only the housing wealth effect is explored, since very few recent studies provide evidence of a significant housing wealth effect in European post-transition economies (Seč and Zemčík, 2007; Aben et al., 2012; Ahec Šonje et al., 2012). Also, evaluating the importance of the macroeconomic impact of the housing wealth effect on European post-transition economies is especially interesting since the trend in real house prices changed rapidly after the financial and real estate crisis in late 2008, with the largest declines recorded in the countries which had previously reached the highest peaks (Ciarlone, 2011).

Within this framework it is very important to understand the fact that households both own housing assets and consume housing services resulting from those assets. Therefore, if there is an increase in house prices, homeowners may feel wealthier through both the realized and unrealized wealth effect. In other words, it is possible for homeowners to take out equity in the form of selling a house or mortgage withdrawal, or they can spend more today due to the higher discounted value of housing wealth. Furthermore, the increase in house prices might also lead to a rise in the value of housing services, thus generating a budget constraint effect on both homeowners and renters, which work in opposite directions with respect to the realized and unrealized wealth effect. Another characteristic of the housing market that has to be taken into account is illiquidity. It is relatively costly to convert increases in housing wealth to money that can be directly spent. More precisely, personal consumption would respond to a house price shock only if the accumulated price movement is larger than the transaction costs linked to adjusting that shock.

Nevertheless, interest in the housing wealth effect has recently revived as a result of developments in housing markets. Although many empirical studies exploring wealth effects have been published in recent decades, most of them refer to developed countries (for example Attanasio et al., 2009; Campbell and Cocco, 2007; Disney et al., 2010) and studies of European post-transition economies are still considerably rare. Therefore, the main goal of this study is to measure the housing wealth effect in the European post-transition countries. Furthermore, as the body of literature on the impact of the housing wealth effect on personal consumption in European post-transition countries is rather limited, this study also complements the earlier findings for those countries. Also, unlike other studies pertaining to the emerging markets, this study differentiates between the short-run and long-run housing wealth effect, and thus includes the latest housing prices in the analysis.

The rest of the paper is organized as follows: Section 2 briefly summarizes the existing empirical literature on the housing wealth effect on consumption. Section 3 presents research data, while Section 4 reviews the research methods. In Section 5 estimation results and robustness check results are given. The final section provides an overview of the main findings of the study.

2. Brief Housing Wealth Literature Review

Scientific literature that examines the wealth effect on personal consumption can be broadly divided in two categories: papers that model the wealth effect based on aggregated macroeconomic data and papers that examine the wealth effect on the basis of microeconomic data. Furthermore, in both of the above-mentioned groups of papers, three sub-groups of papers can be distinguished: those that model only the financial wealth effect on personal consumption, papers that model only the housing wealth effect on personal consumption and, finally, those that deal with both the housing and financial wealth effect on personal consumption (Paiella, 2009).

Empirical studies on the impact of housing wealth on personal consumption are mainly focused on advanced economies. Empirical analyses dating from the late 1990s and the 2000s find a small but statistically significant effect of housing wealth on consumption in the US and the UK (see, for example, Attanasio et al., 2009; Campbell and Cocco, 2007; Disney et al., 2010; Engelhardt, 1996; Skiner, 1996). However, there is still no consensus regarding the actual magnitude of the housing wealth effect, which is probably due to differences in data collection methodology, sampling periods or economic conditions.

The impact of housing wealth on consumption is still insufficiently explored in the emerging countries in general, particularly in the European post-transition countries, which is mostly due to a lack of data availability that restrains complete end effective empirical analysis. Very few recent studies provide evidence of significant housing wealth effects in European post-transition countries. Namely, Seč and

Zemčík (2007) find that housing price increases in the Czech Republic led to consumption growth for homeowners, but not for renters. Aben et al. (2012) provides some evidence of a close relationship between housing equity withdrawals and consumption in Estonia, while Ahec Sonje et al. (2012) find evidence of significant housing wealth effects in Bulgaria, Croatia, the Czech Republic and Estonia that are comparable in magnitude with advanced countries.

In a recent study using a pooled mean group estimator, Ciarlone (2011) estimated the impact of changes in real and financial wealth on private consumption for a panel of 17 emerging economies consisting of Asian and Central and Eastern European post-transition economies. He reached a dual conclusion. Namely, he found that the elasticity of consumption with respect to housing prices is larger than that for stock market prices and that the elasticity of housing wealth for CEE countries is larger than that for Asian countries in the sample.

In this paper, I will make use of Ciarlones study and try to present new estimates of the impact of only the housing wealth effect for a sample of six European post-transition economies, using the pooled mean group estimator of Pesaran et al. (1999). Therefore, taking into account the most recent data set available, capturing the time span before and after the financial crisis of 2008 (Q1 1997–Q3 2012) I will give long-run and short-run estimates of the housing wealth effect on personal consumption in selected European post-transition economies, which to the best of my knowledge was not previously studied on this sample of countries, this time span and using the methodology for non-stationary heterogeneous panel data.

3. Research Data

The dataset used in this research consists of quarterly indices for real estate prices, personal consumption, disposable income and wages for the sample of six European post-transition economies comprising Bulgaria, Croatia, the Czech Republic, Estonia, Lithuania and Slovenia, which were selected on the basis of availability of data for the variables of interest.

As far as the data sources are concerned, the real estate price indices in the empirical analysis are taken from the Property Price Statistics database compiled by the Bank for International Settlements. Personal consumption, disposable income and wages are from the International Financial Statistics, WIIW and Eurostat databases. Also, the housing wealth series, wages and disposable income series are given in real terms, while personal consumption is given in constant prices. Relevant information concerning data sources and the time period for each of the six countries under analysis, forming an unbalanced panel, are given in *Table 1* and descriptive statistics of the analyzed variables are given in *Table A1* in the *Appendix*.

Considering the broad coverage of this study, there are a number of data limitations. Firstly, data on housing wealth are not available for all of the countries in the panel, so real estate price indices are used as proxy variables for housing wealth. In a number of other studies regarding wealth effects such as Ludwig and Sløk (2004), Labhard et al. (2005), Case et al. (2005) and Carrol et al. (2006), to name only a few, price indices were also used as proxy variables for housing wealth. Secondly, data are given for the total aggregate consumption, so no distinction is made between consumption of durable and non-durable goods. Even though conventional consump-

Table 1 Data Description and Sources

Country	Data range	Real estate price	Personal consumption	Wage	Disposable income
Bulgaria	2000Q4– –2012Q4	Flats, existing, big cities, BIS	Constant prices, IFS	WIIW database	IFS database
Croatia	1998Q1– –2012Q3	All types of dwellings, new and existing, Croatian National Bank	Constant prices, IFS	WIIW database	IFS database
Czech Republic	1998Q1– –2012Q3	Single family houses and flats, BIS	Constant prices, IFS	WIIW database	IFS database
Estonia	1997Q1– –2012Q3	All types of dwellings, new and existing, BIS	Constant prices, IFS	WIIW database	IFS database
Lithuania	2000Q1– –2012Q4	All types of dwellings, new and existing, BIS	Constant prices, IFS	WIIW database	IFS Database
Slovenia	2003Q1– –2012Q3	All types of dwellings, new and existing, BIS	Constant prices, IFS	WIIW database	IFS Database

tion theories apply the flow of consumption, durable consumption can be considered as a replacement and addition to capital stock, so the approach in some studies is to use only non-durable consumption (Lettau and Ludwingson, 2004). However, a drawback of this approach might be that total (aggregate) consumption also includes expenditures on housing services, even though durable consumption goods are primarily spent on mortgage refinancing. Furthermore, in order to check the robustness of the results, two proxies for income are used: real net wage and total real disposable income.

Also, the longest possible data range for each country is used in order to capture as many asset price cycles as possible. The longest data range is available for Estonia; more precisely data are available from the first quarter of 1997 to the third quarter of 2012. The shortest data range is available for Slovenia, specifically from the first quarter of 2003 to the third quarter of 2012. The data ranges for the emerging economies are generally much shorter than for developed economies, as most of the data for the European emerging economies are not available at all prior to the 1990s. Finally, all of the variables are expressed in logarithms. In other words, the estimated coefficients can be interpreted as the elasticity of consumption to changes of individual regressors.

4. Research Methodology

Recent dynamic panel data literature emphasizes unit root and cointegration properties of variables observed over a relatively long time period and a large number of cross-section units (Pesaran et al., 1999). One of the central findings of the literature that deals with data sets with reasonably large T is that the assumption of homogeneity of slope parameters is often inappropriate.[1] Furthermore, with the increase in time observations in such panels, non-stationarity is also a concern. So, in the manner of Pesaran et al. (1999), relatively new techniques for estimation of non-stationary dynamic panels in which the parameters are heterogeneous across groups are employed in the empirical analysis of the impact of the housing wealth effect on personal

[1] For a discussion on this subject, see chapter 12 in Baltagi (2010).

consumption. Specifically, a mean group (MG) estimator which is based on estimating N time-series regressions and averaging of the coefficients (Pesaran and Smith, 1995), a dynamic fixed effects (DFE) estimator which pools the time series of all cross-sections and allows only intercepts to differ across groups and a PMG estimator which is a combination of pooling and averaging of coefficients (Pesaran et al. 1999) are used. The PMG estimator allows the intercept, short-run coefficients and error variances to differ across groups, but constrains the long-run coefficients to being equal across groups, which is convenient since the consumption function is estimated in this paper in light of the permanent income hypothesis (Friedman, 1957). Taking into account that the six analyzed post-transition economies are different with respect to their economic policies, the three above-mentioned dynamic panel models are estimated.[2]

However, before any econometric modeling,[3] all variables were tested for stationarity. According to the literature, panel-based unit root tests have higher power than unit root tests based on individual time series. With that in mind, a battery of panel unit root tests is conducted. More precisely, tests with common unit root processes are conducted: LLC (Levin et al., 2002), as are tests with individual unit root processes: IPS (Im et al., 2003) and the Fisher ADF test (Maddala and Wu, 1999; Choi, 2001). *Table A2* in the *Appendix* summarizes the panel unit root test results for three variables of interest: personal consumption ($C_{t,i}$), housing wealth ($w^h_{t,i}$) and household income ($Y_{t,i}$), and *Table A3* in the *Appendix* summarizes the unit root test results for all of the variables in the first differences. On the basis of the panel unit root tests presented in *Table A2*, it can be concluded that all of the series of interest are integrated of order one or difference stationary (*Table A3* in the Appendix).

However, the focus of this research is on the long-run relationship between personal consumption, housing wealth and income, which cannot be consistently estimated if all single variables have a unit root unless they are cointegrated in the long- run. For that reason, the next step of the empirical analysis was to perform panel cointegration tests. Since, according to Banerjee et al. (1998) and Kremers et al. (1992), residual-based cointegration tests can cause a significant loss of power due to the common-factor restriction, new panel cointegration tests developed by Westerlund (2007) are used. These tests are based on the structural rather than the residual dynamic and test the null hypothesis of no cointegration by inferring whether the error-correction term in a conditional panel error-correction model is equal to zero. All of these new tests are normally distributed, but they are also general enough to take into account the country-specific short-run dynamic, country-specific trend and slope parameters as well as cross-sectional dependence. More specifically, two of the tests are designed to test the alternative hypothesis that the whole panel is cointegrated, while the other two test the alternative hypothesis that at least one unit is cointegrated. The results of the performed Westerlund panel cointegration tests are summarized in *Table A4* in the *Appendix*. Accordingly, it can be concluded that

[2] OLS estimators are super-consistent in the case of co-integrated variables, but they are based on strong homogeneity assumptions among countries by imposing a single slope coefficient in the pooled estimation, which is inappropriate for this study with regard to potential country heterogeneity. This is the reason for using a PMG estimator instead of traditional panel techniques.

[3] All econometric analyses performed in this paper were done using Stata 12 and EViews 7 statistical software.

personal consumption and housing wealth are indeed cointegrated in the long-run. The null hypothesis of no cointegration is strongly rejected in all performed panel cointegration tests, with statistical significance of 1% and 5%.

Since the analysis has shown that all of the variables of interest have a unit root and are cointegrated in the long-run, the next step of the empirical analysis in this paper was to estimate the following simplified personal consumption equation:

$$C_{t,i} = \gamma_{0i} + \gamma_{1i} w^h_{t,i} + \gamma_{2i} Y_{t,i} + \varepsilon_{t,i}, \; i = 1,2,...,N, t = 1,2,...,T \tag{1}$$

where C is the logarithm of real personal consumption, w^h is the logarithm of real house prices and Y is the logarithm of real disposable income. The error term capturing the effects of unexpected shocks to personal consumption is denoted by $\varepsilon_{t,i}$. The subscripts i and t denote the country and time, respectively. Deviations from the long-run relationship given in equation (1) are possible in the short-run. Clearly there are various reasons for such deviations including adjustment cost, habit persistence and liquidity constraints (Mehra, 2001; Poterba, 2000; Campbell and Mankiw, 1991). Also, in this framework it is assumed that personal consumption differs across countries in the short-run. This assumption is implemented herein by using conventional statistical criteria and determining the lag length of each variable. Even though equation (1) can be generalized by introducing deterministic terms as well as an autoregressive lag polynomial for the dependent variable and complicatedly distributed lag schemes for explanatory variables, for the purpose of simplification it is assumed here (but relaxed afterwards) that only the first lag of each variable is important for determining personal consumption in each country. Thus, the model given in equation (1) can be written as an autoregressive distributed lag-ARDL (1,1,1) model:

$$C_{t,i} = \delta_{0,1} + \gamma_i C_{t-1,i} + \beta_{10,i} w^h_{t,i} + \beta_{11,i} w^h_{t-1,i} + \beta_{20,i} Y_{t,i} + \beta_{21,i} Y_{t-1,i} + \varepsilon_{t,i} \tag{2}$$

Since, in this case, all the variables under analysis are I(1) and cointegrated, the error term ($\varepsilon_{t,i}$) is an I(0) process for all of the countries in the sample (i). Statistically speaking, cointegrated variables show great responsiveness to any deviation from the long-run equilibrium, so an error-correction reparametrization can be employed:

$$\Delta C_{t,i} = \alpha_{0i} + \varphi_i \left(C_{t-1,i} - \alpha_{1,i} w^h_{t-1,i} - \alpha_{2,i} Y_{t-1,i} \right) + \beta_{10i} \Delta w^h_{t,i} + \beta_{20i} \Delta Y_{t,i} + \varepsilon_{t,i} \tag{3}$$

where:

$$\varphi_i = -\left(1 - \gamma_i\right), \alpha_{1i} = \frac{\beta_{10,i} + \beta_{11,i}}{1 - \gamma_i}, \; \alpha_{2,i} = \frac{\beta_{20,i} + \beta_{21,i}}{1 - \gamma_i} \tag{4}$$

The error-correcting speed of adjustment term is denoted by φ_i and it is expected to be statistically significant and negative. According to Engle and Granger (1987) there is a clear connection between cointegration and the error-correction mechanism.

As mentioned earlier, three alternative non-stationary dynamic models for heterogeneous panels are estimated, namely PMG, MG and DFE. Of these, the PMG estimator (Pesaran et al., 1999) is especially attractive because it assumes homogeneous long-run coefficients, allowing the short-run dynamic specifications to differ from country to country; thus the intercepts, short-run coefficients and error variances differ across the groups, but the long-run coefficients are constrained to being equal across the groups. Furthermore, the PMG estimator can simultaneously solve common econometric problems that occur when estimating the consumption function. More specifically, the serial autocorrelation problem and the problem of endogenous regressors are handled by choosing the appropriate lag structure for dependent and independent variables. Also, when *N* is rather small, as in this case, the PMG estimator is less sensitive to outliers (Pesaran et al., 1999). Since an important issue that needs to be handled in this empirical analysis is the dynamic structure of the consumption function model, assuming that certain economic aspects in each country prevent immediate adjustment of consumption to changes in housing wealth and household income, the panel autoregressive distributed lag model (ARDL) needs to be used.

In that sense, following the PMG procedure, the necessary first step was to choose the lag order of the ARDL model by applying the Schwarz information criterion, the results of which are shown in *Table A5* in the *Appendix*.

According to *Table A5,* there is no clear evidence of a most common representation. Even so, after choosing a country specific lag order of the ARDL model by applying the SBC information criterion, the preferred specification for the whole sample of analyzed countries was the ARDL (1,0,0):

$$C_{t,i} = \delta_{0,1} + \gamma_i C_{t-1,i} + \beta_{10,i} w_{t,i}^h + \beta_{20,i} Y_{t,i} + \varepsilon_{t,i} \tag{5}$$

That is, real personal consumption is lagged once, whereas real disposable income and real house prices are given in levels. So, equation (5) may be reparameterized as follows:

$$\Delta C_{t,i} = \alpha_{0i} + \varphi_i \left(C_{t-1,i} - \alpha_{1,i} w_{t,i}^h - \alpha_{2,i} Y_{t,i} \right) + \beta_{10i} \Delta w_{t,i}^h + \beta_{20i} \Delta Y_{t,i} + \varepsilon_{t,i} \tag{6}$$

and represents the preferred specification to be estimated using the PMG estimator.

5. Estimation Results

Table 2 presents the results of the baseline model of personal consumption specified by equation (6). Furthermore, the Hausman test of long-run homogeneity of coefficients is employed in order to determine which estimator is more appropriate (MG, DFE or PMG). According to Pesaran et al. (1999), the MG estimator provides consistent estimates of the mean of long-run coefficients, but these are inefficient if the slope homogeneity assumption holds. However, if the slope coefficients are indeed homogeneous, then the PMG and DFE estimators are consistent and efficient. According to *Table 2*, homogeneity restriction is not rejected by the data, implying that the PMG and DFE estimators are efficient under the null hypothesis and are preferred over the MG estimator. However, the PMG estimator is preferred over the DFE estimator because it allows for short-run coefficient heterogeneity.

According to *Table 2*, the adjustment coefficient for the analyzed panel has the correct negative sign and is statistically significant at the 1% significance level, which implies that an error-correction mechanism is in place. The average value of

Table 2 PMG and DFE Estimation Results

	PMG	DFE
speed of adjustment ϕ_i	-0.02632*** (0.0604)	-0.2207*** (0.0335)
	Long-run coefficients	
income $\alpha_{2,i}$	0.6579*** (0.0519)[a]	0.6158*** (0.0829)
housing wealth $\alpha_{1,i}$	0.1579*** (0.0242)	0.1469*** (0.0374)
	Short-run coefficients	
housing wealth β_{10i}	0.0635* (0.0367)	0.0767*** (0.0179)
income β_{20i}	0.1246*** (0.0282)	0.1474*** (0.0461)
constant	0.1048*** (0.0196)	0.1149*** (0.0232)
number of observations	291	291
number of countries	6	6
Hausman test for poolability of countries	0.2412	0.9011

Notes: Estimations are performed using the PMG and DFE estimators of Pesaran et al. (1999); all equations include a constant term; astandard errors are in brackets; ***, **, * denote significance at the 1%, 5% and 10% significance level, respectively.

Source: Author's calculations.

the error-correction coefficient (according to the PMG estimator) is –0.03, implying that equilibrium is reached in about 30 quarters. Also, the estimates suggest the presence of the long-run housing wealth effect with a properly signed coefficient that is statistically significant at the 1% significance level, with elasticity of consumption to changes in housing wealth of 0.15, which is in line with Ciarlone`s (2011) research results. The estimated long-run elasticity of personal consumption to changes in income is around one (0.66), as suggested by the Permanent Income Hypothesis.[4] It can be concluded that in the long run personal consumption is responsive to changes in both housing wealth and income, with the latter being more pronounced than the former.

In the short run, there is also evidence of the housing wealth effect in analyzed countries with a somewhat smaller but still statistically significant coefficient (0.06). The elasticity of consumption to changes in income is also statistically significant in the short run with the coefficient of 0.12. Since the PMG procedure allows for short-run heterogeneity, it is possible to estimate separate short-run coefficients for each country in the panel (see *Table 3*).

The estimates of the short-run, country-specific error-correction models also provide evidence of the housing wealth effect, with statistically significant coefficients for Bulgaria and Estonia, whereas that effect is more pronounced in Bulgaria. For the other countries in the sample (Slovenia, Croatia, the Czech Republic and

[4] However, the data do not support the unit elasticity hypothesis, since the corresponding value of $\chi^2(1)$ statistics of 43.41 leads to rejection of the null hypothesis of unit income elasticity.

Table 3 Short-Run, Country-Specific Estimates of Personal Consumption Model

Country	φ_i	ΔY	Δw^h	constant
Slovenia	-0.1231** (0.0598)[a]	0.2545* (0.1352)	-0.0198 (0.0770)	0.0687** (0.0333)
Estonia	-0.4175*** (0.0949)	0.0585 (0.0853)	0.0449* (0.0245)	0.1525*** (0.0353)
Croatia	-0.2035*** (0.0503)	0.1348 (0.0960)	0.0079 (0.0359)	0.1024*** (0.0244)
Czech Republic	-0.0857*** (0.0300)	0.0845* (0.0484)	0.0553 (0.0382)	0.0339*** (0.0129)
Bulgaria	-0.3197*** (0.1132)	0.0956 (0.1741)	0.2364*** (0.0838)	0.1135** (0.0459)
Lithuania	-0.4298*** (0.1222)	0.1195 (0.1350)	0.0565 (0.0571)	0.1580*** (0.0499)

Notes: *,**,*** indicates significance at 10%, 5% and 1% significance level, respectively; numbers in the brackets are standard errors for full PMG.

Source: Author's calculations.

Lithuania), there is a lack of short-run personal consumption reaction to changes in housing wealth. This can be due to underdeveloped financial markets and higher transaction costs that can prevent conversion of increases in housing wealth to money that can be directly spent. Although the error-correction term is statistically significant and correctly signed in all of the analyzed counties, the size of the coefficient is considerably larger for the Baltic countries (Estonia and Lithuania) compared to the SEE countries (Croatia, Bulgaria and Slovenia), with the smallest coefficient recorded for the Czech Republic. Also, consumption reacts to short-run changes in household income, with statistically significant coefficients recorded for Slovenia and the Czech Republic. However, in the other analyzed countries there is no statistically significant short-run response of consumption to changes in income.

5.1 Robustness Check Results

With the aim of investigating the robustness of the presented results, disposable income from the baseline model was replaced with another income proxy, wages, and a new model with the PMG and DFE estimators was evaluated. The result of this robustness check is given in *Table 4*.

The results of the robustness check clearly confirm the importance of housing wealth as one of the drivers of both short-run and long-run personal consumption changes in European post-transition economies. It can be concluded that the long-run housing wealth effect is more pronounced in the model where wages are used as a proxy variable for income[5] according to both the PMG and DFE estimators compared to the baseline model with income. Interestingly, the long-run coefficients of wages for both estimators (PMG and DFE) are much lower than in the case of using household income in the baseline model. Conversely, the short-run housing wealth coefficients in the model with income (baseline model) are lower than in the models where wages are used. Furthermore, it can be concluded that the adjustment coef-

[5] That might be due to the introduction of a new source of heterogeneity since the methodology of collecting data on wages differs significantly from country to country.

Table 4 Robustness Check Results (Wage)

	PMG	DFE
speed of adjustment φ_i	-0.1802** (0.0696)	-0.0697*** (0.0209)
	Long-run coefficients	
wage	0.2021*** (0.0492)[a]	0.2471 (0.2158)[a]
housing wealth	0.2437*** (0.0241)	0.2073** (0.1002)
	Short-run coefficients	
housing wealth	0.0909** (0.0446)	0.1067*** (0.2154)
wage	0.2569 (0.1566)	0.4227*** (0.1154)
constant	0.2002*** (0.0754)	0.0775*** (0.0226)
number of observations	301	301
number of countries	6	6
Hausman test for poolability of countries	0.1917	1.000

Notes: The estimates are performed using the PMG and DFE estimators of Pesaran et al. (1999); panel ARDL (1, 0, 0) model; equations include a constant term; standard errors are in brackets; *** denotes significance at the 1% significance level;** denotes significance at 5% significance level.

Source: Author's calculations.

ficient for both analyzed models presented in *Table 4* have the correct negative sign and are statistically significant at the 1% and 5% significance levels, which implies that an error-correction mechanism is in place. Also, according to the Hausman test for poolability, the homogeneity restriction is not rejected by the data, implying that the PMG and DFE estimators are efficient under the null hypothesis.

Finally, the long-run housing wealth effect in the baseline and the alternative model specification is in line with the previous research of Ciarlone (2011). Specifically, the results of his research showed that, according to different estimation procedures, the long-run elasticity of consumption to changes in house prices ranges from 0.06 to 0.20, with a mid-point of 0.13, and in this research the long-run elasticity of consumption to changes in house prices ranged from 0.14 to 0.24. As in Ciarlones research, the analysis conducted in this paper showed a significant short-run adjustment of income and house prices on consumption and also consumption adjusted to its long-run equilibrium with lags. In this research the estimated coefficients of housing wealth are somewhat higher than in Ciarlones research, but that might imply that countries in Central and Eastern Europe are vulnerable to developments in the housing sector, as he also concluded in his research in 2011.

6. Concluding Remarks

There are several conclusions that can be drawn from the analysis presented in this paper. Firstly, as suggested by the results of the cointegration tests and the PMG procedure, it is evident that personal consumption, housing wealth and income do form a long-run equilibrium relationship in the European post-transition economies.

Furthermore, the error-correction model estimates for the analyzed countries indicate that when the equilibrium relationship is disturbed, the resulting discrepancy is corrected by personal consumption. One can think of a number of reasons why the speed of adjustment of consumption reacts to changes in its fundamental determinants, including the presence of habit persistence and expectations, adjustment costs and liquidity constraints.

Secondly, according to the estimates from the baseline model, there is evidence of a statistically significant long-run and short-run housing wealth effect, with the latter being less pronounced than the former.

Thirdly, regardless of the model specification, the importance of housing wealth as one of the drivers of both short-run and long-run personal consumption changes in European post-transition economies is evident.

Finally, the long-run housing wealth effect in the baseline and the alternative model specification is in line with the previous research of Ciarlone (2011), suggesting that countries in Central and Eastern Europe are vulnerable to developments in the housing sector and that economic policymakers should take this into account in order to prevent adverse effects on consumption that changes of house prices might induce.

APPENDIX

Table A1 Descriptive Statistics of the Variables under Analysis

	Consumption (*C*)	Disposable income (*Y*)	Housing wealth (w^h)
Mean	2.029672	1.966513	2.079362
Median	2.044440	1.989572	2.068830
Maximum	2.212391	2.195499	2.491804
Minimum	1.810820	1.703493	1.678613
Std. Dev.	0.074666	0.133005	0.174529
Skewness	-0.417401	-0.308555	0.038825
Kurtosis	3.328945	1.884130	2.711424
Jarque-Bera	9.963124	20.12164	1.105154
Probability	0.006863	0.000043	0.575465
Sum	602.8127	584.0542	617.5705
Sum Sq. Dev.	1.650218	5.236353	9.016267
Observations	297	297	297

Source: Author's calculations.

Table A2 Panel Unit Root Tests (of Variables of Interest Given in Levels)

Variables		Method	Prob.*	Obs.
$C_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.9373	298
		Im, Pesaran and Shin W-stat	0.9995	298
		ADF-Fisher Chi-square	0.9970	298
		PP-Fisher Chi-square	0.9988	301
$w^h_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.6864	297
		Im, Pesaran and Shin W-stat	0.9994	297
		ADF-Fisher Chi-square	0.9979	297
		PP-Fisher Chi-square	1.0000	301
$Y_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.5979	322
		Im, Pesaran and Shin W-stat	0.9955	322
		ADF-Fisher Chi-square	0.9869	322
		PP-Fisher Chi-square	0.9914	323

Notes * The probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Im, Pesaran and Shin, ADF-Fisher and PP-fisher test—Null Hypothesis: Unit Root (Individual Unit Root process), Levin, Lin & Chu Test-null Hypothesis: Unit Root (common unit root process). Automatic lag length selection based on the Schwarz Criterion and Barlett Kernel.

Source: Author's calculations.

Table A3 Panel Unit Root Tests (of the Variables of Interest Given in First Differences)

Variables		Method	Prob.*	Obs.
$C_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.0000	295
		Im, Pesaran and Shin W-stat	0.0000	295
		ADF-Fisher Chi-square	0.0000	295
		PP-Fisher Chi-square	0.0000	295
$w^h_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.0000	293
		Im, Pesaran and Shin W-stat	0.0000	293
		ADF-Fisher Chi-square	0.0000	293
		PP-Fisher Chi-square	0.0000	295
$Y_{t,i}$	Constant and trend	Levin, Lin & Chu t	0.0000	310
		Im, Pesaran and Shin W-stat	0.0000	310
		ADF-Fisher Chi-square	0.0000	310
		PP-Fisher Chi-square	0.0000	317

Notes: *The probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Im, Pesaran and Shin, ADF-Fisher and PP-fisher test—Null Hypothesis: Unit Root (Individual Unit Root process), Levin, Lin & Chu Test-null Hypothesis: Unit Root (common unit root process). Automatic lag length selection based on Schwarz Criterion and Barlett Kernel.

Source: Author's calculations.

Table A4 Panel Cointegration Tests Results: Consumption and Property Price

Test	Null hypothesis	Alternative hypothesis	Name of the statistics	*p*-values
Westerlund	No ECM	All panels contain ECM	Gt	0.00
			Ga	0.02
		Some panels contain ECM	Pt	0.00
			Pa	0.00

Source: Author's calculations.

Table A5 Auto-Regressive Distributed Lag Specification

Country	Real house prices	Real personal consumption	Real disposable income
Bulgaria	0	0	0
Croatia	0	3	0
Czech Republic	1	0	1
Estonia	1	0	0
Lithuania	1	0	0
Slovenia	0	0	0

Note: Orders of lags in the ARDL model selected using the Schwarz Information Criterion.

Source: Author's calculations.

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