Silvo Dajčman [*]

JEL Classification D14, G21, R22
Preliminary statement
https://doi.org/10.32910/ep.70.6.1

HOUSE PRICES IN THE EURO AREA: THE FUNDAMENTALS AND THE ROLE OF FINANCIAL MARKET STRESS

The paper aims to assess whether financial market stress is associated with real house prices in the euro area. Building on the theory of house prices fundamentals, we first apply the second generation cointegration tests and reject a stable long-run relationship between house prices and the variables identified in the theory as their main determinants (fundamentals). Short-run panel data models are then estimated, relating real house prices to their fundamentals and the financial market stress. The results imply that the real GDP per capita growth rate and the loans to households for house purchase are the main determinants of real house prices growth in the short run. Financial market stress is significantly associated with real house prices changes only in some euro area countries. Different panel data estimators are used to show that heterogeneity and cross-section dependence needs to be accounted for to obtain robust estimates. The differences between two group of euro area countries (the PIIGS and the non-PIIGS euro area) are also compared.

Keywords: house prices, euro area, financial market stress

[*] S. Dajčman, PhD, associate professor, University of Maribor, Faculty of Economics and Business (e-mail: silvo.dajcman@um.si). The paper was received on September 4th, 2018. It was accepted for publication on March 5th, 2019.

1. Introduction

After almost ten years of a steady rise, real house prices in the euro area have reached their peak in the third quarter of 2007[1]. The trough of the boom-bust cycle of the real house prices in the euro area was reached in the last quarter of 2013[2], following financial market and economic stress episodes, including the subprime mortgage, global financial, and the euro area crises. Looking at the country level, between the third quarters of 2007 and 2016, the house prices cycle has not been synchronized across the euro area. One group of countries, including Austria, Belgium, Finland, Germany, and Luxembourg, experienced relatively modest drop in house prices. In countries more affected by the economic and financial market stress episodes, including Greece, Ireland, Italy, Portugal, and Spain, the drop in house prices has been more pronounced. The latter group of countries also experienced a sharper deterioration of macroeconomic activity[3] than the first group of countries and a "systemic financial stress that expanded beyond global financial crisis" (Duprey et al., 2015). Ample international empirical evidence (e.g. Reinhart and Rogoff, 2008; Quigley, 1999; Goodhart and Hofmann, 2008; Andrews, 2010; Holly et al., 2010; Corradin and Fontana, 2013) suggest that real house prices are positively related to measures of economic activity such as GDP per capita and identifies also several other factors (fundamentals) that determine house prices. The knowledge of house prices fundamentals and whether they are related to house prices in the long- or short-run only is important for economic policy from macroeconomic and financial stability perspectives (see e.g. ECB, 2015).

The theory (e.g. Poterba, 1984; Gallin, 2006; Holly et al. 2010) identifies several determinants (fundamentals) of house prices. The most frequently identified include disposable income (or gross domestic product (GDP) per capita), interest rate on loans for house purchases (or long-term interest rate), outstanding loans for house purchase, construction costs, and population. The theory also predicts a stable long-run equilibrium (or cointegrating in econometric terms) relationship between house prices and some fundamentals. Following a change in a specific fundamental variable, only short-run deviations from the equilibrium relationship

[1] See e.g. the Bank for International Settlements Residential property prices statistics database. Throughout the paper, we treat houses and residential property as synonymous.

[2] The Residential property prices statistics of the Bank for International Settlements (2017) shows that at the end of 2013 the real house price index in the euro area was approximately 16 %, while at the end of the third quarter of 2016 approximately 10% below the peak level.

[3] According to Eurostat data, at the end of the third quarter of 2016, the real GDP in the former group of countries was above the level of the third quarter of 2007; whereas in the latter group, the output was above the level at the start of the period only in Ireland, and it was at the same level as at the start of the period in Spain.

between the house prices and the fundamental variable should be observed (Gallin, 2006). The empirical studies – they mostly rely on panel data (panels of cities, regions or countries) – present mixed results: some confirm and some reject a stable long-run (cointegrating) relationship. Only a fraction of these studies (e.g. Holly et al., 2010; Clark and Coggin, 2011; Liu, 2015), none for the euro area, apply the second generation cointegration tests[4] and panel model estimators that take account of cross-section dependence between cross section units of the panel. As noted among others by Breitung and Pesaran (2005), Baltagi and Pesaran (2007), and Sarafidis and Wansbeek (2012), the failure to control for cross-section dependence can lead to misleading inferences.

Some empirical studies, explicitly concentrated on the short-run economic perspective, have recently shown that financial market uncertainty affects macroeconomic activity (e.g. Bloom, 2009; Bloom et al., 2012; Hirata et al., 2013), and also house prices (see Hirata et al., 2013). It is still not documented in the literature how a more broadly defined financial market uncertainty that reflects uncertainty in several segments of financial assets markets affects house prices in the euro area. Recently, Duprey et al. (2015) defined the Country-Level Index of Financial Stress (CLIFS); an indicator that measures financial market stress[5] in three segments of individual euro area countries' financial markets: the equity, bond, and foreign exchange markets.

The empirical evidence on determinants of house prices for the euro area as whole is limited (e.g. Annett, 2005; Corradin and Fontana, 2013; Kulikauskas, 2016; Zhu et al., 2017), and the samples in these studies comprise only a fraction of the euro area member states. The panel data studies for the euro area also do not explicitly account for heterogeneity and cross-section dependence between cross-section units (i.e. the euro area countries).

This study aims to fill the gaps in the literature by analyzing how the main fundamentals of house prices, identified in the theoretical and empirical literature, affect real house prices in the euro area as whole. We use quarterly data for 18 euro area countries and first test for a cointegrating relationship between real house prices and their fundamentals by explicitly accounting for cross-section dependence. Based on the results we then build and estimate an empirical panel data model that relates real house prices to their main fundamentals and the finan-

[4] Several studies simply assume a cointegrating relationship between real house prices and their fundamentals without providing statistical evidence (see Gallin (2006) for review of these studies).

[5] A similar indicator for the euro area as whole, called Composite indicator of Systemtic Stress, was developed by Holló et al. (2012), who note that the stress indicator not only proxies for changes in financial market uncertainty but also for changes in the risk aversion, information asymmetry, and disagreement among investors in the financial markets.

cial market stress indicator. Cross-section dependence is accounted for in the estimation of the empirical model by applying the common correlated effects mean group (CCEMG) estimators developed by Pesaran (2006).

2. Literature review

In the theory there are two general approaches to house prices modelling. The first[6] approach derives the house prices equation as a reduced-form of the structural housing market model (see e.g. Quigley, 1999; Égert and Mihaljek, 2007; Coleman et al. 2008; Clark and Coggin, 2011). Following these studies, the demand for housing can be expressed as a function of real[7] house prices, loan market characteristics, and real macroeconomic and demographic variables, while the supply of housing by real house prices and the real costs of construction. Also other demand-shifting or supply shifting factors can be included in the equations. It is assumed that supply in the short-run is given and that the market is in equilibrium. Under these conditions the reduced-form house prices equation is derived. The second approach to house prices modeling assumes that the user cost of housing should equal rents in the long-run[8] (see e.g. Poterba, 1984; Gallin, 2006; Mikhed and Zemčík, 2009; Holly et al., 2010). If the costs of owning the house exceed rental costs, renting the house becomes more attractive and the house prices are expected to fall in the long-run. The same logic with the opposite conclusion applies to the case when the house prices are below the costs of renting a house. It is common to both approaches to model the real house prices as a function of fundamentals such as real (disposable) income (or real GDP per capita), real interest rate on loans for house purchase (or long-term interest rate), and other demand-for-housing shifters such as changes in population, credit conditions, and real wealth. Under certain conditions, it can be shown that theoretically real house prices and certain fundamentals (foremost GDP per capita) are cointegrated (see e.g. Gallin et al. 2006; Holly, et al. 2010).

The empirical literature on modeling house prices is vast. One strand of empirical literature, related to our study, investigates whether a stable long-run rela-

[6] The inverted demand approach or the supply and demand approach (see Muellbauer, 2012).

[7] It must be noted that the theory is inconclusive whether variables in the house prices model should be real or nominal. We found it more common in theoretical and empirical papers that the variables are real (see e.g. Gerlach and Peng, 2006; Oikarinen, 2009; Holly et al. 2010), but there are papers in which variables are nominal (e.g. Coleman et al. (2008)) or a mixture of real and nominal (e.g. interest rate) (e.g. Gimeno and Martínez-Carrascal, 2006).

[8] Typically, certain house-ownership costs such as maintenance, taxes and depreciation are ignored (see above mentioned studies for details).

tionship between real house prices and their fundamentals exists. The results of these studies are mixed. Some studies find a cointegrating relationship between real house prices and their fundamentals (e.g. Gerlach and Peng, 2005; Gimeno and Martínez-Carrascal, 2006; Égert and Mihaljek, 2007; Oikarinen, 2009; Holly et al. 2010; Gattini and Hiebert, 2010; Corradin and Fontana, 2013; Liu, 2015; Turk, 2015; Kulikauskas, 2016), others do not (e.g. Malpezzi, 1999; Gallin, 2006; Zhou and Sornette, 2006; Mikhed and Zemčík, 2009; Clark and Coggin, 2011). Based on the finding of a stable long-run relationship, some of these studies then estimate long- and/or short-run relationships (e.g. Gerlach and Peng, 2005; Oikarinen, 2009; Corradin and Fontana, 2013; Liu, 2015; Kulikauskas, 2016).

Another strand of empirical literature (e.g. Goodhart and Hofmann, 2008; Calza et al., 2009; Hirata et al., 2013; Zhu et al., 2017) concentrates on the short-run economic perspective and is thus able to analyze the effect of changes in (the level) stationary economic determinants (e.g. monetary policy stance or financial market uncertainty) on the short-run house prices formation[9]. It has been recently proved by Hirata et al. (2013) that financial market uncertainty can affect current house prices. Hirata et al. (2013), building on the literature that investigates the effect of uncertainty shocks on the macroeconomy (e.g. Bloom et al. 2012; Stock and Watson, 2012), show that uncertainty in financial markets (proxied by realized equity price volatility) and real house prices in G7 countries are significantly positively related. Theoretically, a positive or a negative relationship between house prices and uncertainty in financial markets can be argued. A positive relationship can be expected if house equity is perceived by financial market investors as a safe haven in uncertain times (Hirata et al., 2013). Increased uncertainty in stock markets then stimulates portfolio rebalancing towards housing assets (ibidem). If this is not the case, the housing is perceived by consumers more like an ˝ordinary˝ durable consumption good. Increased uncertainty in stock markets in this case is expected to bear negatively on housing consumption and consequently on house prices.

Empirical evidence on house prices determinants for the euro area as whole is limited. Majority of the studies concentrate either on a single country (e.g. Gimeno and Martínez-Carrascal, 2006; Oikarinen, 2009; Turk, 2015), a panel limited to a fraction of euro area countries (Annett, 2005; European Commission, 2012; Kulikauskas, 2016; Zhu et al., 2017)[10] or analyze aggregated euro area time series (Gattini and Hiebert, 2010).

Our study is the most related to the studies of Annett (2005), Gattini and Hiebert (2010) and European Commission (2012). Annett (2005) analyses the interrelation-

[9] In the empirical literature house prices (index) are usually found to be a unit root process.

[10] In some studies some of the euro area countries are included in a larger panel sample (e.g. Hilbers, 2008; Iossifov et al. 2008; Andrews, 2010; Corradin and Fontana, 2013; Hirata, 2013).

ship between real house prices, inflation, real disposable income per capita, real long-term interest rate, and real credit (alternatively money) for a panel of eight euro area countries (Germany, France, Italy, Ireland, Spain, Netherlands, Belgium, and Finland) in the period 1970-2003. A ˝short-to-medium run˝ model is estimated with the LSDV (Least Squares Dummy Variable), pooled OLS (Ordinary Least Squares), and the Arellano-Bond GMM (Generalized Method of Moments) estimators, while the long-run model with the Engle-Granger two-step procedure. The results of the short-to-medium run model show that all variables have expected signs. Only the lagged house prices and real long-term interest rate statistically significantly explain the real house price dynamics. The long-run model includes three regressors: real disposable income per capita, real long-term interest rate, and real credit and/or money. All the regressors are statistically significant, whereby house prices in the long-run are negatively related to real long-term interest rate, and positively to real disposable income per capita and outstanding credit/money in the economy.

Gattini and Hiebert (2010) analyze and forecast real house prices dynamics in the euro area as whole on aggregated time series data for the period 1970-2009. They find that real house prices are positively related to real income and negatively to the housing investment and real interest rate on government bonds. A vector error-correction regression is used to provide the out-of-the sample forecasts that show that house prices were overvalued at the start of the global financial crisis.

The European Commission (2012) evaluates house price developments for an unbalanced panel of 11 euro area countries for the period 1972-2011. They apply price-to-income and price-to-rents indicators to assess if the real house prices changes were supported by fundamentals, including population, real disposable income, and long-term interest rate[11]. Additionally, they estimate a long-run relationship between real house prices and the specified fundamentals (but do not present the results) and then calculate out-of-sample forecast of real house prices to assess price misalignments for individual countries. Their results show unsynchronized house price cycles in the euro area.

There are other studies that concentrate on the euro area, but capture only some of the member states. Zhu et al. (2017) estimate how the monetary policy and the housing market regulation affect ˝non-fundamental˝ house prices. They first estimate changes in fundamental real house prices for each of 11 euro area countries for the period 1992Q1-2012Q4, i.e. the part of changes in real house prices that are

[11] There are several approaches to assess whether house prices debase from their fundamental value including relative simple indicators such as the price-to-rent ratio, the price-to-income ratio, which may be misleading, because they do not account for changing interest rates, expected inflation, expected house prices appreciation or taxes (Himmelberg et al., 2005), and income distribution of households (André, 2010). A more robust judgement demands cointegration analysis (Gallin, 2006).

explained by changes in population, building permits, real income per capita, real GDP, unemployment rate, CPI and mortgage rate. The results of this exercise show diversity in the relevance of the fundamental house prices determinants across the countries. They then calculate changes in ˝non-fundamental˝ house prices as residual changes in house prices, i.e. house prices changes that are unexplained by fundamentals, and apply a panel VAR model to assess how they are affected by negative shocks in the monetary policy stance. In another study, Kulikauskas (2016) investigates deviations from fundamental real house prices in three Baltic states, all of which are currently euro area members, for the period 2000–2014. He applies a panel error-correction model to model short-run real house prices by fundamentals, including real household income, population dynamics, real interest rates on loans for house purchase, loans for house purchase, and construction costs. The results show that construction costs, population, income and loans are positively, while interest rates negatively associated with real house prices. He finds actual house prices to deviate from the fundamental prices, confirmed also by the price-to-income and price-to-rents indicators.

3. Methodology

This research applies the inverted demand approach to modelling house prices (Quigley, 1999; Égert and Mihaljek, 2007; Coleman et al. 2008; Clark and Coggin, 2011; Kulikauskas, 2016). The approach considers simultaneously the demand for and supply of housing and then derives the reduced-form equation of house prices. Guided by the empirical literature, the price equation in this paper is modeled as a function of the following demand and supply fundamentals: the real GDP per capita, the real volume of loans to households for house purchase, the real interest rate on loans for house purchase, and the real construction costs. We additionally consider the country-level index of financial market stress as a possible determinant of house prices.

The first objective of the paper is to test for a stable long-run relationship between real house prices and their fundamentals in the euro area as whole. Following e.g. Holly and Pesaran (2010) and Banerjee and Silvestre (2017), a panel cointegration test is used for the purpose. We first apply the second-generation unit root test of Pesaran (2007) (CIPS test) that can account for cross section dependence[12].

[12] Baltagi and Pesaran (2007) note that the use of the first-generation panel unit root and cointegration tests – they do not account for cross-section dependence – could lead to significant size distortions. Breitung and Pesaran (2005) review the main issues and compare the performance of the first- and second-generation panel unit root and cointegration tests.

Having identified variables with unit root processes, we then test for cointegration by applying two second-generation panel cointegration tests that account for cross-section dependence between cross section units (or panel groups), i.e. countries of euro area in our case. The first is the cointegration test of Westerlund (2007) and the second is the CCEP-based[13] cointegration test, suggested by Holly et al. (2010) and formally developed by Banerjee and Carrion-i-Silvestre (2017).

The second objective of the paper is to assess how the main fundamentals, identified in the theoretical and empirical literature, affect real house prices in the euro area. Because our panel data sample can be defined as a macro-panel (large *T*, time dimension, and moderate-to-large *N*, cross section dimension), according to the macro panel (or panel-time series) literature (see e.g. Baltagi and Pesaran, 2007; Baltagi, 2013; Bonizzi, 2017) the empirical analysis has to address the issues of cross-section dependence and (cross-section) heterogeneity. The cross-section dependence across cross-section units of the panel may be present only between some cross-section units due to physical or economic proximity (a weak form of cross-section dependence) or it may be pervasive across all or most cross-section units due to some common shocks (e.g. macroeconomic shocks, oil price shocks, changes in technology) that affect all cross section units (see Holly et al., 2010; Pesaran and Tosetti, 2011; Banerjee and Carrion-i-Silvestre 2017). Ignoring cross-section dependence can have serious consequences for validity of inference (Pesaran and Tosetti, 2011) and the conventional panel data estimators (e.g. fixed effects) can even be inconsistent (Phillips and Sul, 2003; Pesaran and Tosetti, 2011; Chudik and Pesaran, 2013).

Heterogeneity is another issue to be dealt with in empirical analysis of panel data. Given heterogeneity in house prices and economic cycles of euro area countries witnessed in the statistical data and evidence by empirical studies (e.g. European Commission, 2012; Zhu et al., 2017) the conventional panel data estimators that allow only for intercept heterogeneity may produce biased results (see Pesaran and Smith, 1995). Consistent results may be obtained by applying the mean group estimators that allow for slope heterogeneity (ibidem).

The results of the cointegration tests (presented in continuation) show that there is no stable long-run relationship between the real house prices and the fundamentals. On this premise, and guided by the existent empirical studies, data availability for euro area countries, and the recent advancements in the macro-panel analysis literature, we propose to estimate the following short run (stationary)[14]

[13] CCEP is an abbreviation for the pooled common correlated effect estimator.

[14] Given that a cointegrating relationship between the house prices and their fundamentals does not exist for our sample, and because (non-)stationarity is also an issue in macro panels due to a large time dimension of the data (see e.g. Baltagi and Kao, 2001; Baltagi, 2013), a stationary panel model is proposed.

panel data model of real house prices that accounts for cross-section dependence and heterogeneity of the panel data structure (see e.g. Pesaran 2006, 2007):

$$\Delta hpr_{it} = \alpha_i + \boldsymbol{\beta}_i^{'}\boldsymbol{x}_{it} + \boldsymbol{\eta}_i^{'}\boldsymbol{f}_t + \varepsilon_{it}, \tag{1}$$

where $i = 1, ..., N$ denotes a set of countries, $t = 1, ..., T$ denotes time, Δhpr_{it} is the logarithmic growth[15] of the real house prices, α_i are country fixed effects, $\boldsymbol{x}_{it}$ is a 5x1 vector of observed exogenous stationary variables (house prices fundamentals) including: the logarithmic growth of the real GDP per capita ($\Delta gdppcr_{it}$), the logarithmic growth of the real volume of loans to households for house purchase ($\Delta loansr_{it}$), the real interest rate on loans for house purchase (r_{it}), the country-level index of financial market stress ($clifs_{it}$), and the logarithmic growth of real construction costs ($\Delta cocor_{it}$). $\boldsymbol{\beta}'_i$ is a 1x5 vector of slope coefficients. $\boldsymbol{\eta}'_i\boldsymbol{f}_t + \varepsilon_{it}$ is a multifactor structure of the error-term, where $\boldsymbol{f}_t$ denotes a *mx1* vector of unobserved common factors, driving the cross section dependence, $\boldsymbol{\eta}'_i$ is a *1xm* vector of factor loadings, and ε_{it} is the idiosyncratic error term, uncorrelated with the explanatory variables, but allowed to be weakly spatio- and temporally correlated (see Pesaran, 2006 or Chudik and Pesaran, 2013).

Pesaran (2006, 2007) shows that the unobserved common effects can be captured by the cross-section averages of the observed variables (dependent and explanatory), serving as proxies for $\boldsymbol{f}_t$, in the above specification. With this augmentation of the multifactor error structure of model (1), the unobserved common effects are asymptotically (as $N \rightarrow \infty$) eliminated (de Vos and Everaert, 2016). Against this background and assuming the number of unobserved common effects is equal to the number of observed variables[16] ($m=k$), model (1) can be rewritten as (see e.g. Holly et al. 2010; Chudik and Pesaran, 2013; Desbordes and Eberhardt, 2014; Banerjee and Carrion-i-Silvestre, 2017):

$$\Delta hpr_{it} = \alpha_{i,CCEMG} + \boldsymbol{\beta}_{i,CCEMG}^{'}\boldsymbol{x}_{it} + c_{i,CCEMG}\overline{\Delta hpr}_{it} + \boldsymbol{d}_{i,CCEMG}^{'}\overline{\boldsymbol{x}}_{it} + \varepsilon_{it,CCEMG}, \tag{2}$$

where $\overline{\Delta hpr}_{it}$ is the cross-section average of the dependent variable Δhpr_{it}, $c_{i,CCEMG}$ is a regression coefficient, $\overline{\boldsymbol{x}}_{it}$ is the 5x1 vector of cross-section averages of the explanatory variables (fundamentals) from model (1), and $\boldsymbol{d}'_{i,CCEMG}$ a 1x5 vector of parameter estimates of cross-section averages of the explanatory variables from

[15] Logarithmic growth of a variable in the paper is calculated as the first difference of the natural logarithm of a variable.

[16] This is a common practice in empirical applications (see Pesaran, 2006; Holly et al. 2010; Desbordes and Eberhardt, 2014; Banerjee and Carrion-i-Silvestre, 2017).

model (1). Model (2) assumes heterogeneity of slope coefficients, $\boldsymbol{\beta}'_{i,CCEMG}$, yielding the mean group common correlated effects panel data model (CCEMG).

Pesaran (2006) shows that consistent estimates of the regression (slope) coefficients can be obtained by the CCEMG panel data estimators without the need to determine the number of unobserved common factors, given the regressors are stationary and exogenous. Consistency of the estimators is robust to non-stationarity or common factors (Pesaran, 2006), cointegration between them (Kapitanios et al., 2011), or spatial correlation in the idiosyncratic errors (Pesaran and Tosetti, 2011)[17].

Model (2) nests also a model that does not account for common correlated effects – the mean group (MG) model (see Pesaran and Smith (1995)):

$$\Delta hpr_{it} = \alpha_{i,MG} + \boldsymbol{\beta}'_{i,MG}\boldsymbol{x}_{it} + \varepsilon_{it,MG}. \tag{3}$$

Models (2) and (3) are estimated by the Stata routine xtmg of Eberhardt (2012).

The results of model (3) will be compared to the results of model (2) to investigate the possible issue of cross-section dependence. The models will be estimated for the total euro area sample (euro area as whole), and separately for the PIIGS (Portugal, Ireland, Italy, Greece, Spain) and the non-PIIGS euro area sample, the later including the remaining countries in the sample.

4. Data and the empirical results

An unbalanced panel data for a group of 18 euro area countries is used, including Austria, Belgium, Cyprus, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Lithuania, Luxembourg, Malta, Netherlands, Portugal, Slovakia, Slovenia, and Spain[18]. Quarterly data is used and the observation period is not the same across countries. The variables that define the start of the data sample are either the real house prices index or the credit market variables (the real interest rate on loans for house purchase or the real value of loans for house purchase outstanding). Definition of the variables and their transformations are described in Table 1.

[17] Chudik and Pesaran (2015) show that when a lagged dependent variable enters the list of regressors (this would then be a dynamic panel data model) only the CCEMG estimator remains consistent under certain conditions.

[18] Euro area currently consists of 19 countries. Estonia is excluded from the sample because of relatively short observation period. The data on the country-level index of financial market stress for Estonia is available only until 2010Q4, whereas the data on credit market variables only from 2008Q1, leaving us with only 12 observations.

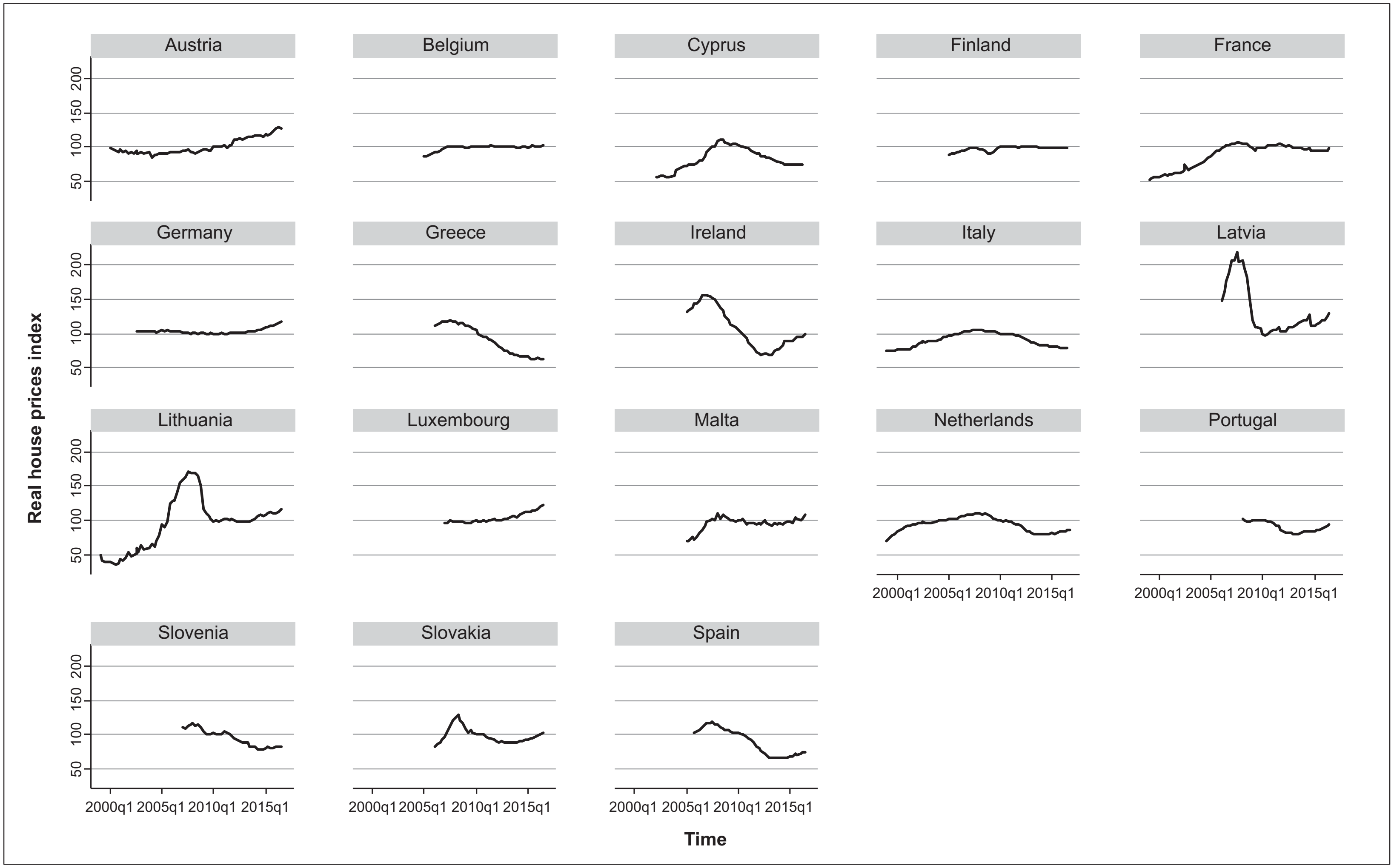
Table 1:

DESCRIPTION OF VARIABLES USED

Variable	Description of the primary data	Source
Natural logarithm of the quarterly real house prices index (*hpr_{it}*)	Quarterly real house prices index	Bank for International Settlements (2017)
Natural logarithm of the annual real GDP per capita index (*$gdppcr_{it}$*) on a quarterly frequency	Real annual GDP per capita was calculated as a sum of real GDP per capita in the last four quarters and converted to index	Eurostat
Natural logarithm of the quarterly real volume of loans for house purchase outstanding index (*$loansr_{it}$*)	Loans to households and non-profit institutions serving households (NPISH) for house purchase outstanding at the end of quarter, all maturities; nominal outstanding amount was converted to index and then transformed to real index by HICP	European Central bank and Eurostat (for HICP)
Quarterly real interest rate on loans for house purchase (r_{it})	Average nominal interest rate for loans to households and NPISH for house purchase outstanding (i_t) was the primary data. r_{it} was obtained by first calculating monthly real interest rate for individual countries by the Fisher equation: $r_t = (1 + i_t)/(1 + \pi_t) - 1$, where π_t is the annual inflation for month t calculated from the HICP. Finally, the quarterly real interest rate was calculated as the average of the monthly levels of the real interest rates	European Central Bank
Quarterly level of country-level index of financial market stress (*$clifs_{it}$*)	Quarterly country-level index of financial stress was calculated as a simple average of monthly levels of the country-level index of financial stress as developed and calculated by Duprey et al. (2015); this index is constructed to reflect volatility in three financial market segments – equity market, bond market and foreign exchange market (see Duprey et al. (2015) for a comprehensive description of how the index is constructed)	European Central bank
Natural logarithm of the quarterly real construction costs index (*$cocor_{it}$*)	Quarterly level of nominal construction cost index for residential buildings, except residences for communities; real index was calculated by deflating nominal index by HICP	Eurostat
Observation period for individual countries: Austria: 2003Q1-2016Q3; Belgium: 2005Q1-2016Q3; Cyprus: 2008Q1-2016Q2; Finland: 2005Q1.2016Q3; France. 2003Q1-2016Q3; Germany: 2003Q1-2016Q3; Greece: 2006Q1-2016Q3; Ireland: 2005Q1-2016Q3; Italy: 2003Q1-2016Q3; Latvia: 2010Q3-2016Q3; Lithuania: 2004Q2-2016Q3; Luxembourg: 2007Q1-2016Q3; Malta: 2005Q1-2016Q3; Netherlands: 2003Q1-2016Q4; Portugal: 2008Q1-2016Q3;Slovakia: 2006Q1-2016Q3; Slovenia: 2007Q1-2016Q3; Spain: 2005Q1-2016Q3		

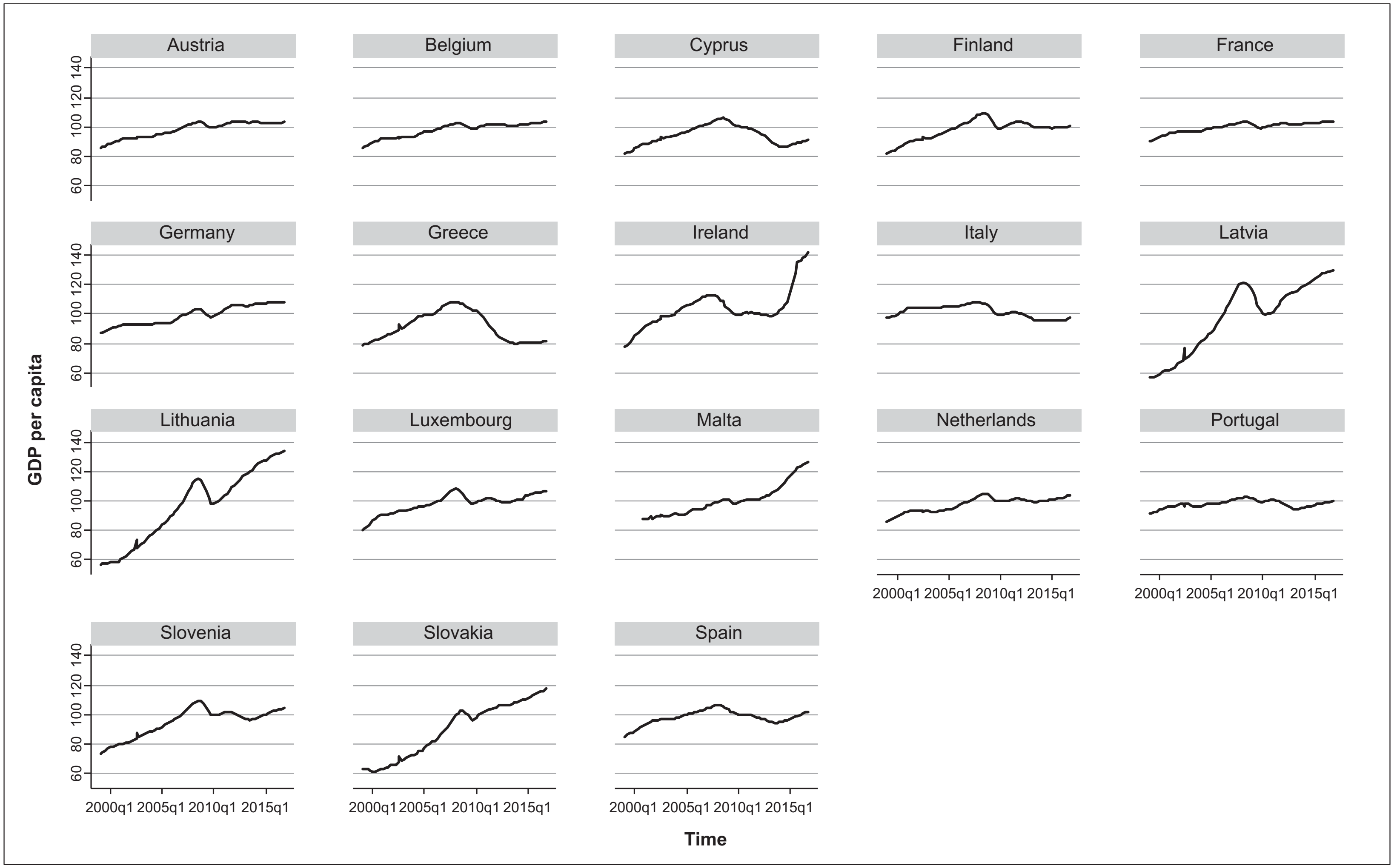
Figures 1 and 2 present the dynamics of real house prices and the GDP per capita in the sampled euro area countries.

Figure 1: THE REAL HOUSE PRICES IN THE EURO AREA



Notes: the y-axis presents the index (2010=100).

Figure 2: GDP PER CAPITA IN THE EURO AREA



Notes: The y-axis presents the index (2010=100).

Observing the figures, the heterogenous dynamics of the observed variables becomes evident. Application of regression estimation methods that account for heterogeneity is therefore well-suited.

The cross-section dependence in the log-level and growth transformed variables was tested by the Pesaran's (2004) cross-section dependence (CD) test. Pesaran (2004) shows that the test is robust to variety of panel data model characteristics, including non-stationarity, structural breaks, and time dimension of the panel data. The results, presented in the first column of Table A1 in Appendix show that for all variables the null hypothesis of no cross-section dependence can be rejected, implying that the unit-root and cointegration tests that account for the cross-section dependence must be applied.

The stationarity of variables was checked by the Pesaran's (2007) CIPS test. The results, presented in the third column of Table A1, indicate that for the log-level variables hpr_{it}, $gdppcr_{it}$, $loansr_{it}$, and $cocor_{it}$ the null hypothesis of unit root cannot be rejected at any conventional significance level (significance levels are indicated in the brackets under the statistics). The result is robust to alternative specification of deterministics (constant or constant plus trend) and the lags of the underlying CADF regression. First-differencing of these variables (this transformation is indicated by Δ before the variable notation) yields logarithmic growth of the variables and achieves their stationarity. The real interest rate on loans for house purchase (r_{it}) and the country-level index of financial market stress ($clifs_{it}$) are stationary in levels.

The results of the Westerlund's (2007) cointegration test are presented in Table A2 in Appendix, while the PCCE-based cointegration test, suggested by Holly et al. (2010) and formally developed by Banerjee and Carrion-i-Silvestre (2017) in Table A3 in Appendix.

The results of both tests unanimously do not reject the hypothesis of no cointegration between the variables with unit root process implying there is no stable long-run relationship between the house prices and the fundamentals (as specified above). The result is robust to alternative specifications. Our findings thus support the findings of a large body of empirical studies performed for other (group of) countries (including e.g. Malpezzi, 1999; Gallin, 2006; Zhou and Sornette, 2006; Mikhed and Zemčík, 2009; Clark and Coggin, 2011) that question the validity of the theoretical models predicting a stable long-run relationship between house prices and their fundamentals. Literature offers some explanations why this contradiction may exists – relationship between house prices and fundamentals may not be stable over time due to changes in regulatory environment (Gallin, 2006) or because house prices and the fundamentals are exposed to common permanent (e.g. changes in productivity) and transitory (e.g. economic policy) shocks (Fraser et. al., 2012). No cointegrating relationship can also result when deviations from the stable long-run relationship between house prices and their fundamentals are very persistent. It may last for several decades to restore long-run equilibrium – in this case the time period covered in the empirical analysis is too short to find a cointegration relationship by cointegration tests (see e.g. Ambrose et al., 2013). But as noted among others by Gallin (2006) or Holly et al. (2010), in order to assert a cointegrating relationship between house prices and their fundamentals an empirical investigation is needed, based on the available data and in this paper we use the data for the longest term available.

Several authors in the literature argue that the absence of a stable long-run relationship between house prices and their fundamentals is an indication of a house prices bubble formation (see e.g. Mikhed and Zemčík, 2009; Clark and Coggin, 2011). However, some authors are more cautious and state that the absence of cointegration just implies that the possibility of a bubble formation is increased (Case and Shiller, 2003; Holly et al., 2010).

Given that the cointegrating relationship between house prices and their fundamentals for our sample could not be established, stationary panel data models of real house prices proposed by equations (2)-(3) were estimated. The results of this exercise are presented in Table 2. The first three columns present the results of model (3) and the last three of model (2).

We present first the results for the euro area as whole (columns [1] and [4]). The diagnostics provided in table (RMSE and CD test results) indicate the CCEMG model (model (2)) as a better statistical fit to the data as the MG model, therefore this model is highlighted in explanation of the results.

The CCEMG model [4] indicates that a one percentage point increase in the growth of real GDP per capita ($\Delta gdppcr_{it}$) is associated with an increase in the growth of real house prices of 0.55 percentage points. The results also show a positive association between the outstanding loans for house purchase and real house prices: a one percentage point increase in the real volume of loans for house purchase outstanding ($\Delta loansr_{it}$) is associated with a 0.17 percentage point increase in the growth of real house prices. The significant positive association between the variables is expected and has already been reported in the empirical studies (e.g. Gerlach and Peng, 2005; Annett, 2005; Oikarinen, 2009; Kulikauskas, 2016). The real interest rate on loans for house purchase (r_{it})[19], the financial market stress ($\Delta clifs_{it}$), and the growth in real construction costs ($\Delta cocor_{it}$)[20] are not significantly related to the growth of real house prices.

[19] Interest rates are documented to be an insignificant real house price fundamental among others in Oikarinen (2009), Holly et al. (2010), and Kulikauskas (2016).

[20] Construction costs are a determinant of a housing market supply. As the housing supply is rigid in the short run (see Hilbers et al., 2008), in a static model as ours the construction costs may be insignificant variable of real house price changes. In the long-run models and in dynamic models, the variable may become significant (see e.g. Kulikauskas, 2016).

Table 2:

THE RESULTS OF HOUSE PRICE MODELS (2) AND (3)

Explanatory variable	**Model (3) (i.e. MG model)**			**Model (2) (i.e. CCEMG model)**		
	Euro area as whole [1]	PIIGS euro area [2]	Non-PIIGS euro area [3]	Euro area as whole [4]	PIIGS euro area [5]	Non-PIIGS euro area [6]
$\Delta gdppcr_{it}$	.50847*** (.19217)	.562090 (.37807)	.43425* (.23542)	.55041*** (.17988)	.21481 (.25387)	.40690** (.19479)
r_{it}	.24076** (.10178)	.32644** (.15030)	.09120 (.12558)	-.03309 (.07814)	-.23879** (.10013)	.09645 (.1115)
$\Delta loansr_{it}$	.25897*** (.07877)	.02494 (.19408)	.35836*** (.08839)	.16852*** (.06237)	-.02411 (.08048)	.26626** (.13319)
$clifs_{it}$	-.03160*** (.01120)	-.06103*** (.01156)	-.01515 (.01377)	.01679 (.02343)	-.00429 (.02567)	.00452 (.02664)
$\Delta cocor_{it}$	.12125 (.09441)	.30902** (.13554)	.03723 (.13197)	.15977 (.10162)	.32872** (.15900)	-.03840 (.14495)
Diagnostics:						
RMSE	0.0205	0.0149	0.0222	0.0157	0.0093	0.0166
CD test	4.75483 (p=0.000)	0.722 (0.470)	3.327 (0.001)	-2.64041 (p=.008)	-3.637 (0.000)	-2.702 (0.007)
Number of observations	774	205	569	774	205	569

Notes: The robust mean values of slope coefficients were calculated (see Eberhardt, 2012). The standard errors in these two models are robust, calculated as suggest by Pesaran and Smith (1995). The Stata routine xtmg of Eberhardt (2012) was used to estimate models (2) and (3). */**/*** denote the 10%/5%/1% significance levels for the rejection of the null hypothesis of the average of slope coefficient being equal to zero. Diagnostics that are usually reported in macro panel data studies are reported as well (see e.g. Eberhardt and Presbitero, 2015). RMSE is the root mean squared error, while CD test is a cross-section dependence test of Pesaran (2004). We report the CD statistics and the corresponding p-value for the rejection of the null of no cross-section dependence in the residuals in the brackets. The Stata routine xtcd2 was used for this purpose (see Ditzen, 2018). We also calculated (not reported in the table) the Hausman test to test the difference in the slope coefficients between the FE model and the random effects model (for the euro area as whole sample). The results showed that the consistent FE should be preferred over the random effects model.

We must note (see Eberhardt and Presbitero, 2015) that the insignificance of the slope coefficients for variables r_{it}, $\Delta cocor_{it}$ and $clifs_{it}$ does not necessarily imply that there is no association between these variables and the growth of real house prices for each euro area country. Insignificance of slope coefficients may simply be a consequence of heterogeneity of the relationship between house prices and fundamentals across the countries canceling out on the average. The regression estimates (not shown here) for individual countries show that the slope coefficient for the real interest rate on loans for house purchase (r_{it}) is significant, with a negative sign, only for one country, Ireland. The slope coefficient for the variable of growth in the real construction costs ($\Delta cocor_{it}$) is significantly positive at the 5% level for Belgium, Ireland and Italy, and at the 10% level for Spain. The slope coefficient for country-level index of financial market stress ($clifs_{it}$) is significantly positive at the 5% level for Austria and Finland and at the 10% level for Portugal; it is statistically negative at the 10% level for Greece.

As also stressed by some extant studies (e.g. European Commission, 2012; Zhu et al., 2017), the relationship between real house prices and fundamentals in the euro area is heterogeneous. The presented results indeed show heterogenous characteristics of house price determinants in the PIIGS and non-PIIGS euro area. For the non-PIIGS euro area, the CCEMG estimator (column [6]) provides a better fit to the data in regards of cross-section dependence and the RMSE, while for the PIIGS euro area we prefer to highlight the results of the MG model (column [2]), given the results of the CD test. In the non-PIIGS euro area, the real GDP per capita and the real growth in loans for house purchase are found to be the main determinants of real house prices dynamics, while in the PIIGS euro area the real interest rates and the financial market stress.

The robustness of the results of model (2) for euro area as whole was checked threefold. First, the empirical model (2) was modified by including another potentially important determinant of real house prices dynamics identified in the literature – population growth. We used the Eurostat's data on population and included the second difference of the natural logarithm of the number of population[21]. The results (not presented here to save space) showed that the variable is highly statistically insignificant and does not affect the significance of other variables. The second robustness check consisted of replacing the real interest rate on loans for house purchase with the nominal interest rate on loans for house purchase. As noted among others by Gimeno and Martínez-Carrascal (2006) or Martínez-Carrascal and Rio (2004), an increase in the nominal lending interest rate (leaving the real rate unchanged) increases the debt burden and reduces the availability of funds for consumption and/or investment in the housing. The results of this alternation of model (2) are presented in Table 3.

[21] We found the variable to be integrated of order 2 (I(2)).

Table 3:

ROBUSTNESS CHECK OF MODEL (2) – ALTERNATIVE INTEREST RATE SPECIFICATIONS

Explanatory variable	**Model (2): CCEMG model (euro area as whole)**
$\Delta gdppcr_{it}$	.34942* (.21081)
Δlni_{it}	-.11168*** (.03939)
$\Delta loansr_{it}$	.20451** (.07994)
$clifs_{it}$	.00662 (.01917)
$\Delta cocor_{it}$	.11227 (.08461)
Diagnostics:	
RMSE	0.0159
CD test	-2.66576 (p=.008)
Number of observations	771

Notes: Δlni_{it} is the first difference of the natural logarithm of the nominal interest rate on loans for house purchase. The first difference was taken to obtain a stationary variable. For other description, the notes of Table 2 apply.

The results show that the growth in the nominal interest rate on loans for house purchase is significantly negatively associated with the growth in real house prices. Substitution of the real interest rate with the nominal one reduces the size of the slope coefficient of the growth in the real GDP per capita. The slope coefficients for the country-level index of financial market stress and the growth in real construction costs remain statistically insignificant.

The third robustness check considered the role of house prices appreciation in the house prices formation. Following e.g. Di Pasquale and Wheaton (1994), Bar Nathan et al. (1995), Higgins and Osler (1998), and Annett (2005), a first-order autoregressive price process was considered. Since it is known that the CCEMG estimator in dynamic panel model is inconsistent, the price equation (1) was estimated by the bias-corrected least-squares least dummy variable estimator (LSDV) of Bruno (2005a). The results of this exercise are presented in Table A4 in Appendix. The results show persistence in house prices, implying adaptive expectations (Bar

Nathan et al. 1995). All slope coefficients are statistically significant. Caution in interpretation is warranted because the LSDV estimator is not robust to cross section dependence.

The results of our study have important implications for economic policy. They show that not only real income but credit availability as well is an important determinant of real house prices dynamics in the short-run. The result is not surprising as the inter-linkage between the credit, the housing market and the economic cycles has been documented in the literature (see, e.g., Mian and Sufi, 2011; Jordà et al., 2014). The economic policy (including primarily monetary and macroprudential policies, but also fiscal policy) can stimulate the demand for housing by reducing the costs of external financial funds for the consumers and increasing the willingness of banks to supply loans. Therefore they stimulate also macroeconomic activity mainly through the wealth and the collateral effects on consumption (ECB, 2015). House price cycles are inherent feature of the housing markets and the role of economic policy is not to generally prevent house prices to fall but to prevent the house prices cycles having detrimental effects on the financial stability and the macroeconomic activity (see e.g. Ambrose et al., 2013). The literature identifies the macroprudential policy as the best equipped for the job (see e.g. Crowe, 2011, 2013; ECB, 2015). Our results show that no stable long-run relationship between the house prices and their fundamentals can be identified. Against this background it may be difficult to determine the fundamental house prices level and to assess to what level the market house prices are misaligned. This raises the challenge facing economic policy to prevent a potential housing bubble to develop and/or to bust. Given the differences in the main determinants of short-run house prices dynamics, the optimal policy-mix very likely differs across individual euro area countries.

5. Conclusion

The paper applied the second generation cointegration tests that control for cross-section dependence and found that in the euro area as whole a stable long-run relationship between the real house prices and their main fundamentals, including real GDP per capita, the real volume of loans for house purchase outstanding, and the real construction costs, does not existent. A short-run panel data model is then proposed to analyse how real house prices growth is related to its main determinants identified in the literature and to the indicator (index) of country-level financial market stress. Controlling for the cross-section dependence and (cross-section) heterogeneity, our results show that the growth of real house prices is positively related to the growths in real GDP per capital and real volume of loans for house purchase outstanding. The real interest rate on loans for house purchase and the growth in real construction costs are not significantly associated with the growth of real house prices. A positive association between the country-level index of financial market stress and the growth of real house prices is established but is not statistically significant. We discuss several implications of the findings for the economic policy.

References

Ambrose B.W., Eichholtz P., & Lindenthal T. (2013). House Prices and Fundamentals: 355 Years of Evidence. *Journal of Money, Credit and Banking*, 45(2-3): 477-491.

André C. (2010). A Bird´s Eye View of OECD Housing Markets. *OECD Economics Department Working Papers* No. 746.

Andrews D. (2010). Real house Prices in OECD Countries. *OECD Economics Department Working Papers* No. 831.

Annett A. (2005): House prices and monetary policy in the euro area. *IMF Country Report* No.05/266.

Baltagi B.H. (2013). *Econometric Analysis of Panel Data, 5th Edition*. Wiley, Chichester.

Baltagi B.H., & Kao C. (2001). Nonstationary panels, cointegration in panels and dynamic panels: A survey. In Baltagi B.H., Fomby T.B., Hill R.C. (eds): *Nonstationary Panels, Panel Cointegration, and Dynamic Panels (Advances in Econometrics, Volume 15)*. Emerald Group Publishing Limited, 7 – 51.

Baltagi B.H., Pesaran M.H. (2007). Heterogeneity and cross section dependence in panel data models: theory and applications. Introduction. *Journal of Applied Econometrics*, 22(2): 229-232.

Banerjee A., & Carrion-i-Silvestre J.L. (2017). Testing for Panel Cointegration Using Common Correlated Effects Estimators. *J. Time Ser. Anal.*, 38(4): 610-636.

Bank for International Settlements. (2017). *Sources: National sources, BIS Residential Property Price database*. Available: www.bis.org/statistics/pp.htm.

Bar-Nathan M., Beenstock M., & Haitovsky Y. (1995). An econometric model of the Israeli housing market. *Bank of Israel Research Department Discussion Paper* No. 95.02.

Bloom N. (2009). The Impact of Uncertainty Shocks. Econometrica, *77*(3): 623–685.

Bloom N., Floetotto M., Jaimovich N., Saporta-Eksten I., & Terry S. (2012). Really Uncertain Business Cycles. *NBER Working Paper* No. 18245.

Bonizzi B. (2017). Institutional Investors Allocation to Emerging markets: a Panel Approach to Asset Demand. *Journal of International Financial Markets, Institutions and Money*, 47(C): 47-64.

Breitung J., & Pesaran M.H. (2005). *Unit Roots and Cointegration in Panels*. *Deutsche Bundesbank Discussion Paper Series 1: Economic Studies* No. 42/2005.

Bruno G.S.F. (2005a). Estimation and Inference in Dynamic Unbalanced Panel Data Models with a Small Number of Individuals. *Centro di Ricerca sui Processi di Innovazione e Internazionalizzazione Working Paper* 165.

Bruno G.S.F. (2005b). XTLSDVC: Stata module to estimate bias corrected LSDV dynamic panel data models. Statistical Software Components S450101, Boston College Department of Economics, revised 08 2005.

Calza A., Monacelli T., & Stracca L. (2009). *Housing finance and monetary policy*. *ECB Working Paper Series* No. 1069.

Case K.E., & Shiller R. (2003). Is there a bubble in the housing market? *Brookings Papers on Economic Activity*, 34(2): 299-342.

Chudik A., & Pesaran M.H. (2013). *Large Panel Data Models with Cross-Sectional Dependence: A Survey*. *Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute Working Paper* No. 153.

Chudik A., & Pesaran M.H. (2015). Common correlated effects estimation of heterogeneous dynamic panel data models with weakly exogenous regressors. *Journal of Econometrics*, 188(2): 393-420.

Clark S.P., & Coggin T.D. (2011). Was there a U.S. house price bubble? An econometric analysis using national and regional panel data. *The Quarterly Review of Economics and Finance*, 51(2): 189-200.

Coleman M.I.V., LaCour-Little M., & Vandell K.D. (2008). Subprime lending and the housing bubble: Tail wags dog? *Journal of Housing Economics*, 17(4): 272-290.

Corradin S., & Fontana A. (2013). House price cycles in Europe. *ECB Working Paper Series* No. 1613.

Crowe C., Dell'Ariccia G., Igan D., & Rabanal P. (2011). Policies for Macrofinancial Stability. Options to Deal with Real Estate Booms. *IMF Staff Discussion Note* SDN11/02.

Crowe C., Dell'Ariccia G., Igan D., & Rabanal P. (2013). How to deal with real estate booms: Lessons from country experiences. *Journal of Financial Stability*, 9(3): 300-319.

Desbordes R., & Eberhardt M. (2017). *Gravity models in the presence of slope heterogeneity and cross-section dependence*. Centre d´Études de Recherchers sur le Développement International working paper. Available: http://cerdi.org/uploads/sfCmsNews/html/2830/2014-01-27%20gravity%20-%20heterogeneity%20-%20csd.pdf.

De Vos I., & Everaert G. (2016). Bias-corrected Common Correlated Effects Pooled estimation in homogenous dynamic models. *Universiteit Gent Working Paper* No. 2016/920.

Di Pasquale D., & Wheaton W.C. (1995). Housing Market Dynamics and the Future of House Prices. Journal of Urban Economics, 35(1): 127, 1994.

Ditzen, J. (2018). Estimating dynamic common-correlated effects in Stata. Stata Journal 18(3): 585-617.

Duprey T., Klaus B., & Peltonen T. (2015). Dating systemic financial stress episodes in the EU countries. *ECB Working Paper Series* No. 1873.

Eberhardt, M. (2012). Estimating panel time-series models with heterogeneous slopes. *The Stata Journal*, 12(1): 61-71.

Eberhardt M. (2017). XTCD Stata routine. Available at: https://sites.google.com/site/medevecon/code#TOC-xtcd.

Eberhardt M., & Presbitero A.F. (2015). Public debt and growth: Heterogenity and nonlinearity. *Journal of International Economics*, 97(1): 45-58.

ECB. (2015). The state of the house price cycle in the euro area. *ECB Economic Bulletin*, Issue 6 / 2015.

Égert B., & Mihaljek D. (2007). Determinants of house prices in central and eastern Europe. *BIS Working Papers* No. 236.

European Commission. (2012). Assessing the dynamics of house prices in the euro area. *Quarterly Report on the Euro Area* 4/2012.

Fraser P., Hoesli M., & McAlevey L. (2012). House Prices, Disposable Income, and Permanent and Temporary Shocks. *Journal of European Real Estate Research*, 5(1): 5-28.

Gallin J. (2006). The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets. *Real Estate Economics*, 34(3): 417-438.

Gattini L., & Hiebert P. (2010). Forecasting and assessing euro area house prices through the lens of key fundamentals. *ECB Working Paper Series* No. 1249.

Gerlach S., & Peng W. (2005). Bank lending and property prices in Hong Kong. *Journal of Banking & Finance*, 29(2): 461-481.

Gimeno R., & Martínez-Carrascal C. (2006). The interaction between house prices and loans for house purchase. The Spanish case. *Banco de España Documentos de Trabaj* No. 0605.

Goodhart C, & Hofmann B. (2008). House prices, money, credit, and the macroeconomy. *Oxford Review of Economic Policy*, 24(1): 180-205.

Higgins M., & Osler C. (1998). Asset market hangovers and economic growth: U.S. housing markets. *Federal Reserve Bank of New York Research Pa*per 9801.

Hilbers P., Hoffmaister A., Nanerji A., & Shi H. (2008). House Price Developments in Europe: A Comparison. *IMF Working Paper* WP/08/211.

Himmelberg C., Mayer C., & Sinai T. (2005). Assessing High House Prices: Bubbles, Fundamentals, and Misperceptions. *Journal of Economic Perspectives*, 19(4): 67-92.

Hirata H., Kose M.A., Otrok C., & Terrones M.E. (2013). Global House Price Fluctuations: Synchronization and Determinants. *IMF Working Paper* WP/13/38.

Holló D., Kremer M., & Lo Duca M. (2012). *CISS – A composite indicator of systemic stress in the financial system*. ECB Working Paper Series No. 1426.

Holly S., Pesaran M.H., & Yamagata T. (2010). A spatio-temporal model of house prices in the USA. *Journal of Econometrics*, 158(1): 160-173.

Iossifov P., Čihák M., & Shangavi A. (2008). Interest Rate Elasticity of Residential Housing Prices. *IMF Working Paper* WP/08/247.

Kulikauskas D. (2016). Fundamental housing prices in the Baltic States. Empirical approach. *Baltic Journal of Economics*, 16(2): 53-80.

Lewandowski P. (2007). PESCADF: Stata module to perform Pesaran's CADF panel unit root test in presence of cross section dependence. Available: http://econpapers.repec.org/software/bocbocode/s456732.htm.

Liu X. (2015). The Income Elasticity of House Prices in New South Wales: A Panel Study. *ARC Centre of Excellence in Population Ageing Research Working Paper* 2015/19.

Malpezzi S. (1999). A simple error correction model of house prices. *Journal of Housing Economics*, 8(1): 27-62.

Martínez-Carrascal C., & Rio A. (2004). Household borrowing and consumption in Spain: a VECM approach. *Banco de España Documentos de Trabaj* No. 0421.

Mikhed V., & Zemčík P. (2009). Do house prices reflect fundamentals? Aggregate and panel data evidence. *Journal of Housing Economics*, 18(2): 140-149.

Muellbauer J. (2012). When is a Housing Market Overheated Enough to Threaten Stability? In Heath A, Packer F, Windsor C (eds): *Property Markets and Financial Stability. Proceedings of a Conference, held in Sydney August 20-21 2012*. Sydney: Reserve Bank of Australia, 73-108.

Oikarinen E. (2009). Interaction between housing prices and household borrowing: The Finnish case. *Journal of Banking & Finance*, 33(4): 747-756.

Persyn D., & Westerlund J. (2008). Error Correction Based cointegration Tests for Panel Data. *Stata Journal*, 8(2): 232-241.

Pesaran M.H. (2004). General Diagnostic Tests for Cross Section Dependence in Panels. *IZA Discussion Paper* No. 1240.

Pesaran M.H. (2006). Estimation and Inference in Large Heterogeneous panels with a multifactor error structure. *Econometrics*, 74(4): 967-2012.

Pesaran M.H. (2007). A Simple Panel Unit Root Test in the Presence of Cross Section Dependence. *Journal of Applied Econometrics*, 22(2): 265-312.

Pesaran M.H., & Smith R. (1995). Estimating long-run relationships from dynamic heterogeneous panels. *Journal of Econometrics*, 68(1): 79-113.

Pesaran M.H., & Tosetti E. (2011). Large panels with common factors and spatial correlation. *Journal of Econometrics*, 161(2): 182-202.

Phillips P., & Sul D. (2003). Dynamic panel estimation and homogeneity testing under cross section dependence. *The Econometrics Journal*, 6(1): 217-259.

Poterba J. (1984). Tax Subsidies to Owner-occupied Housing: An Asset-Market Approach. *Quarterly Journal of Economics* 99(4): 729–52.

Quigley J.M. (1999). Real Estate Prices and Economic Cycles. *International Real Estate Review*, 2(1): 1-20.

Reinhart C.M., & Rogoff K. (2008). Is the 2007 U.S. subprime financial crisis so different? An international historical comparison. *American Economic Review* 98(2): 39-44.

Sarafidis V., & Wansbeek T. (2012). Cross-sectional Depdendence in Panel Data Analysis. *Econometric Reviews*, 31(5): 483-531.

Stock J.H., & Watson M.W. (2012). Disentangling the Channels of the 2007-2009 Recession. *Brookings Papers on Economic Activity*, 44(1): 81-135.

Turk A.R. (2015). Housing Price and Household Debt Interactions in Sweden. *IMF Working Paper*, WP/15/276.

Westerlund J. (2007). Testing for Error Correction in Panel Data. *Oxford Bulletin of Economics and Statistics*, 69(6): 709-748.

Zhou W.X., & Sornette D. (2006). Is there a real-estate bubble in the USA? *Physica A*, 361: 297-308.

Zhu B., Betzinger M., & Sebastian S. (2017). Housing market stability, mortgage market structure, and monetary policy: Evidence from the euro area. *Journal of Housing Economics*, 37(September): 1-21.

Appendix

Table A1:

RESULTS OF THE CROSS-SECTION DEPENDENCE (CD) TEST AND THE CIPS TEST

Variable	[1] CD test statistics	[2] Averaged cross-correlation coefficient	[3] CIPS test statistics					
			CIPS (Lag 1)		CIPS (Lag 2)		CIPS (Lag 3)	
			Cons.	Cons.+ trend	Cons.	Cons.+ trend	Cons.	Cons.+ trend
hpr_{it}	17.23 (0.000)	0.208	1.816 (0.965)	-0.398 (0.345)	1.945 (0.974)	1.213 (0.887)	-0.046 (0.482)	-0.719 (0.236)
Δhpr_{it}	19.91 (0.000)	0.245	-11.144 (0.000)	-10.748 (0.000)	-5.904 (0.000)	-5.746 (0.000)	-4.413 (0.000)	-3.897 (0.000)
$gdppcr_{it}$	66.26 (0.000)	0.633	-3.003 (0.001)	-4.284 (0.000)	-2.359 (0.009)	-3.489 (0.000)	0.403 (0.657)	-0.637 (0.262)
$\Delta gdppcr_{it}$	55.07 (0.000)	0.529	-6.460 (0.000)	-4.168 (0.000)	-7.264 (0.000)	-5.033 (0.000)	-9.938 (0.000)	-8.269 (0.000)
r_{it}	37.69 (0.000)	0.456	-4.102 (0.000)	-2.842 (0.002)	-5.371 (0.000)	-4.494 (0.000)	-3.867 (0.000)	-2.971 (0.001)
$loansr_{it}$	34.14 (0.000)	0.379	2.889 (0.998)	1.856 (0.968)	2.703 (0.997)	2.841 (0.998)	2.053 (0.980)	0.818 (0.793)
$\Delta loansr_{it}$	26.48 (0.000)	0.312	-11.386 (0.000)	-11.189 (0.000)	-8.972 (0.000)	-8.852 (0.000)	-5.314 (0.000)	-4.163 (0.000)
$clifs_{it}$	51.47 (0.000)	0.512	-6.940 (0.000)	-5.945 (0.000)	-6.888 (0.000)	-6.007 (0.000)	-6.816 (0.000)	-6.280 (0.000)
$\Delta clifs_{it}$	32.90 (0.000)	0.319	-1.296 (0.098)	-0.702 (0.241)	-0.170 (0.432)	0.739 (0.770)	-0.574 (0.283)	0.785 (0.784)
$\Delta cocor_{it}$	15.19 (0.000)	0.149	-17.651 (0.000)	-16.986 (0.000)	-14.757 (0.000)	-13.774 (0.000)	-7.161 (0.000)	-5.208 (0.000)

Notes: Δ denotes the first-difference transformation, Δhpr_{it} thus denotes the logarithmic growth of the real house prices index, $\Delta gdppcr_{it}$ the logarithmic growth of the real GDP per capita (index), $\Delta loansr_{it}$ the logarithmic growth of the real volume of outstanding loans for house purchase (index), and $\Delta cocor_{it}$ the logarithmic growth of the real construction costs (index). CD test is the cross-section dependence test of Pesaran (2004) and was performed with the xtcd Stata routine of Eberhardt (2017). The CD tests statistics and the corresponding p-value for the rejection of the null of no cross-section dependence are presented in the brackets. The table also reports the averaged cross-correlation value, calculated as the arithmetic correlation of the absolute correlation of the variables among the panel groups (countries). CIPS test is based on Pesaran (2007). Codes of Lewandowski (2007) were used for the purpose. The standardized CIPS statistics (i.e. averaged CADF statistics calculated on individual cross section units), and the corresponding significance levels (in brackets) are reported. To account for a potential serial correlation, up to 3 lags were included in the CADF regression.

Table A2:

RESULTS OF THE WESTERLUND'S (2007) COINTEGRATION TEST

	Error correction model specification: Lags only (determined by AIC)		**Error correction model specification: 1 lag, 1 lead**	
Test statistic	Test statistic value	Robust p-value	Test statistic value	Robust p-value
G_τ	-1.652	0.943	-1.920	0.611
G_α	-5.655	0.979	-7.143	0.648
P_τ	-6.195	0.766	-7.537	0.461
P_α	-4.986	0.661	-6.126	0.386
Average lag selected by AIC	0.33		/	

Notes: The test assumes no cointegration and provides four test statistics based on the error-correction specification (see Westerlund (2007) for details). Two test statistics (P_τ and P_α) test the null hypothesis of no cointegration for the panel as whole, and two (G_τ and G_α) for the null of no cointegration for at least one cross-section group. τ and α denote that the standard errors in the error-correction specification are standard and the Newey and West, respectively. Two alternative specifications of the error-correction regression were used: In the first, only lags of the variables were specified, setting maximum lag order to 2 (i.e. maximally allowed considering the time dimension of the panel sample) and letting the Akaike information criteria (AIC) to decide the optimal lag used. The second error correction specification used for the test allowed for one lag and one lead of the variables in the error correction model. Including leads in the specification allows the variables to be only weakly exogenous (see Westerlund and Persyn, 2008). Cross-section robust p-values of the test statistics were obtained by bootstrapping the test statistics with 800 replications. The xtwest Stata routine of Westerlund and Persyn (2008) was used.

Table A3:

RESULTS OF COINTEGRATION TEST OF BANERJEE AND CARRION-I-SILVESTRE (2017)

Integrated variables (I(1)) included in the long-run specification in the first step of the test	$CADFC_p$ statistics Model with a constant only		$CADFC_p$ statistics Model with a constant and trend	
	Lag 1	Lag2	Lag 1	Lag2
hpr_{it}, $gdppc_{it}$, $loansr_{it}$, $cocor_{it}$	-1.977 (-2.89/-2.79)	-1.543 (-2.81/-2.70)	-1.567 (-3.26/-3.16)	-0.947 (-3.16/-3.05)
hpr_{it}, $gdppc_{it}$ *	-1.657 (-2.50/-2.40)	-0.958 (-2.46/-2.36)	-0.832 (-2.94/-2.85)	-0.439 (-2.91/-2.81)

Notes: The Westerlund's test is an error-correction based test that accounts for the cross-section dependence by bootstrapping the test statistics (Westerlund, 2007; Persyn and Westerlund, 2008). The test of Banerjee and Carrion-i-Silvestre (2017) is a multi-step test. In the first step, a long-run relationship between integrated variables is assumed and the long-run regression coefficients are calculated by the pooled common correlated effects estimator (PCCE) of Pesaran (2006, 2007). In the second stage, the ˝residuals˝ are calculated as $\hat{u}_i = y_i - \beta_{iPCCE}\boldsymbol{x}_i$, where $\hat{u}_i$ is a $(T\ x\ 1)$ vector of residuals, y_i is a $(T\ x\ 1)$ vector of the dependent variable, $\boldsymbol{x}_i$ is a $(T\ x\ k)$ matrix of explanatory variables (k is the number of explanatory variables), and β_i is a $(k\ x\ 1)$ vector of regression coefficients from the estimated long-run specification in the first stage. In the third step, the deterministics (the constant, or the constant plus trend) in the $\hat{u}_i$ is subtracted. In the fourth step the CIPS test of Pesaran (2007) is applied to the residuals obtained in stage three to test for a unit root. The rejection of the null hypothesis of unit root implies that the cointegration relationship between the variables in the long-run model exists, whereas the non-rejection implies no cointegration. Banerjee and Carrion-i-Silvestre (2017) in their paper calculate the critical values of the CCEP-based cointegration test statistic (they denote it $CADFC_p$). Both cointegration tests share the null hypothesis of no cointegration and the rejection of the hypothesis implies a stable long-run relationship between house prices and the fundamentals. The number of common factors is assumed to be the same as the number of observables (i.e. number of variables in the long-run specification in the first step of the test; see Banerjee and Carrion-i-Silvestre (2017) for advantage of such an agnostic specification). 2 lags are the maximum lag augmentation in the CIPS test in the fourth step of the test, because Banerjee and Carrion-i-Silvestre (2017) do not report test statistics for more than 2 lags. We first included four variables in the long-run specification of the first step of the test (i.e. hpr_{it}, $gdppcr_{it}$, $loansr_{it}$ and $cocor_{it}$; r_{it} is not included because it is a stationary process). The critical values of the $CADFC_p$ statistics for the rejection of the null of no cointegration for N=20, T=70 are taken from Banerjee and Carrion-i-Silvestre (2017) and are noted in the brackets (the first for the 5% and the second for the 10% significance level). *Alternatively, only two variables were included in the long-run specification of the first step of the test (hpr_{it} and $gdppc_{it}$), similar to the study of Holly et al. (2010). An intuitive explanation of the test procedure in Stata can be found also on the Eberhardt's website: https://sites.google.com/site/medevecon/code#TOC-Cointegration-Testing (see the xtbcis procedure).

Table A4:

ROBUSTNESS CHECK OF MODEL (1): THE ROLE OF HOUSE PRICES EXPECTATIONS (IN EURO AREA AS WHOLE)

Explanatory variable	**Parameter estimates**
$\Delta hpcr_{it-1}$	.13652 *** (.03407)
$\Delta gdppcr_{it}$	.66733 *** (.12382)
r_{it}	.24394 (.07106)
$\Delta loansr_{it}$	.17718*** (.03463)
$clifs_{it}$	-.02647 ** (.01073)
$\Delta cocor_{it}$	.18142*** (.06577)
Diagnostics:	
CD test	27.338 (p=.00)

Notes: The results are obtained by the LSDV estimator of Bruno (2005a), corrected for endogeneity bias. 1000 repetitions were performed to obtain bootstrapped standard errors. Stata code xtlsdvc (Bruno 2005b) was used.

CIJENE NEKRETNINA U EURO PODRUČJU: FUNDAMENTI I ULOGA NESTABILNOSTI FINANCIJSKOG TRŽIŠTA

Sažetak

Ovaj rad ima za cilj procijeniti povezanost nestabilnosti na financijskom tržištu s realnim cijenama nekretnina u euro području. Polazeći od teorije fundamenata cijena nekretnina, prvo primjenjujemo kointegracijske testove druge generacije i odbacujemo stabilan dugoročni odnos između cijena nekretnina i varijabli koje su u teoriji identificirane kao njihove glavne odrednice (fundamenti). Zatim se procjenjuju kratkoročni panel modeli, povezujući cijene nekretnina s njihovim osnovama i nestabilnošću na financijskom tržištu. Rezultati pokazuju da su rast realnog BDP-a po glavi stanovnika i zajmovi kućanstvima za kupnju nekretnina glavne odrednice rasta realnih cijena nekretnina u kratkom roku. Nestabilnost financijskog tržišta značajno je povezan s promjenama cijena nekretnina samo u nekim zemljama euro područja. Različiti procjenitelji panel podataka se primjenjuju kako bi se pokazalo da heterogenost i ovisnost prostornih podataka treba biti uračunata kod robusnih procjena. Također se uspoređuju razlike između dvije skupine zemalja euro područja (PIIGS i euro područje koje nije PIIGS).

Ključne riječi: cijene nekretnina, euro područje, nestabilnost financijskog tržišta