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Are House Prices Characterized by Threshold Effects?
Evidence from Developed and Post-Transition Countries*

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Abstract
We use a nonlinear framework in order to explore house price determinants and adjustment properties. We test for threshold cointegration using a sample of four developed countries (the United States, the United Kingdom, Spain, and Ireland) and four transition countries (Bulgaria, Croatia, the Czech Republic, and Estonia). In addition to testing for nonlinearities, we explore house price determinants in these four transition countries of Central and Eastern Europe. Asymmetric house price adjustment is present in all transition countries and the USA, while no threshold effects are detected in developed European countries. In a threshold error correction framework, house prices are aligned with fundamentals, but house price persistence coupled with a slow and asymmetric house price adjustment process might have facilitated the house price boom in transition countries and the USA.

1. Introduction

Housing is an essential good, accounting for a large share of household expenditure and assets and a significant part of economic activity. When modeling house price behavior, it is useful to differentiate between demand and supply factors. The former factors include household income, shifts in a country's demographic structure, changes in the tax system promoting a higher owner-occupancy ratio, the interest rate level, and the amount of housing loans granted. The availability and cost of land, the cost of construction, investment in improving the quality of the existing housing stock, and housing stock changes are considered the most important housing supply determinants. One must also note that the feedback effects of house price changes can exhibit just as strong an impact on the economy as the effect that economic fundamentals exert on house prices. In that sense, housing market developments can emphasize business cycle fluctuations or even contribute to business cycle changes. The most extensively researched house price feedback mechanism is undoubtedly the wealth effect, which influences aggregate consumption spending by affecting the net wealth of households and their capacity to borrow and spend. House prices can also affect investment spending as well as profitability and employment in the construction and real estate industries. The impact of house prices on investment activity is particularly pronounced in countries that have ample residential investments.

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The importance of housing is reflected in the large number of papers about house price modeling. Thus far, the majority of empirical studies on house prices have been conducted using a linear framework for the developed countries data sample. However, if house prices are characterized by nonlinear properties, this in turn implies that linear house price models are not an appropriate tool for such an analysis.

Judging from the literature, many other economic series and phenomena, such as stock market returns, purchasing power parities, GDP, industrial production, and unemployment rates, incorporate nonlinear properties (Neftci, 1984; Falk, 1986; Bradley and Jansen, 1997; Sarantis, 2001). Common sense would suggest that house prices also incorporate some nonlinear properties. Moreover, one of the few papers exploring house price nonlinearities (Sei-Wan and Bhattacharya, 2009, p.444) states "it is clearly plausible that market behavior differs across expansion and contraction phases of the swings that characterize the real estate market". Abelson et al. (2005) suggest that households are keener to get into the housing market when prices are on the rise. This is partly due to a fear that a delay would result in paying even higher prices. Hence, when prices are on the rise, households exhibit forward-looking behavior, while equity constraints play only a minor role. On the other hand, households are less keen to buy or sell a house when prices are on the decline due to loss aversion and more pronounced equity constraints, causing stickiness on the downside of the housing market cycle. The threshold adjustment of house prices may be justified by asymmetric properties of house price determinants such as GDP or interest rates (Neftci, 1984; Enders and Siklos, 2001). Threshold effects may also stem from the high transaction costs inherent in property transactions. As such, small deviations from equilibrium will not be corrected, while larger discrepancies are expected to be mean-reverting such that speed of adjustment is an increasing function of the size of the discrepancy. However, in this case threshold effects should be more pronounced in transition countries because lower property rights standards, underdeveloped financial markets, and less liquid property markets tend to increase transaction costs.

The aim of this paper is to test for nonlinear house price properties such as threshold cointegration and asymmetric adjustment of house prices in relation to long-run discrepancies as proposed by Enders and Siklos (2001). We test the given methods on a sample of four developed countries (Ireland, Spain, the United States, and the United Kingdom) and four transition countries (Bulgaria, Croatia, the Czech Republic, and Estonia). To the best of our knowledge this is the first paper that applies this methodology on a sample of developed and transition countries' house prices and one of the few papers dealing with house price nonlinearities in general.[1] By applying the threshold cointegration method, we want to explore whether house price nonlinearities have in part contributed to house price booms. Furthermore, by including Central and Eastern European countries in our sample, we explore house price properties and determinants in a region where house price appreciation has been more intensive than in developed countries that have experienced house price

[1] We must note that comparing any European country with the United States should be done extremely cautiously, since it involves significant qualitative and scale differences. Specifically, the empirical analysis in this paper was performed without controlling for several socio-economical features of the United States, such as its specific migration patterns, its distinctive labor market and financial system characteristics, and its settlement structure. All these features might have influenced the empirical results had they been included in the analysis.

booms. However, the housing markets in Central and Eastern European countries have been less intensively researched than those in the developed countries, and this paper might shed more light on the subject.

The remainder of the paper is organized as follows. Section 2 is a review of the literature on house price modeling. The results of studies undertaken in the linear and nonlinear framework are summarized, with special attention being given to empirical studies dealing with house price modeling in the transition countries of Central and Eastern Europe. Section 3 presents the data and the methodology applied and includes a detailed description of the results of the empirical analysis. Section 4 concludes the paper.

2. Literature Review

In developed countries, a lot of attention has been devoted to house price modeling within a linear framework. In general, such studies apply vector autoregression models, cointegration and error correction models, or panel data models in order to identify house price determinants (see *Table 1*). Some studies, including Sutton (2002), McQuinn and O'Reilly (2008), Pagés and Maza (2007), Schnure (2005), Abelson et al. (2005), and Meen (2002), confirmed the importance of income and interest rates as house price drivers in several developed economies. Egert and Mihaljek (2007) reached the same conclusion by examining a sample of developed and European transition economies.

Other studies, such as Gallin (2006) and Mikhed and Zemčík (2009), showed that changes in fundamentals did not explain the rapid growth of house prices in the USA during the period prior to the house price correction that started in 2006. Tsatsaronis and Zhu (2004) also concluded that GDP in 17 developed countries had very little explanatory power over house price movements. Annett (2005) suggested that real income per capita was not a major determinant of short-run house price dynamics in the panel of the EU-15 countries and was significant only in some countries (Germany, Ireland, and Finland).

In addition to the obvious suspects such as income and interest rates, empirical studies detected several other house price drivers. Abelson et al. (2005) showed that changes in the housing stock and equity prices explained house prices in Australia. Sutton (2002) also stressed the importance of equity prices as a house price determinant in developed countries, while Hort (1998) suggested that changes in both construction and user cost have affected house prices in Sweden. Tsatsaronis and Zhu (2004) concluded that inflation and variables related to mortgage finance have been the most important drivers of house prices in developed countries. Furthermore, empirical studies done for Sweden (Hort, 1998), the USA (Lamont and Stein, 1999), the EU-15 (Annett, 2005), and a sample of Central and Eastern European and EU-15 countries (Posedel and Vizek, 2009) concluded that the growth of real house prices has been very persistent, i.e., that there would be a strong tendency for real house prices to rise tomorrow if they rose today.

All the above-mentioned studies assume that house prices behave in a linear fashion. If house prices, however, do incorporate nonlinear properties or threshold effects, then a linear empirical framework is not appropriate. For example, Balke and Fomby (1997) and Enders and Siklos (2001) showed that conventional tests for unit

Table 1 Selected Recent Empirical Studies on the Determinants of House Prices

Authors and country	Methodology	House price determinants		
		Income	Interest rates	Other
6 developed countries Sutton (2002)	VAR	Yes	Yes	Equity prices
Ireland McQuinn and O´Reilly (2009)	Cointegration and ECM	Yes	Yes	-
Spain Pagez and Maza (2007)	Cointegration and ECM	Yes	Yes	Equity prices, housing loans
The USA Schunure (2005)	Panel regression	Yes	No	Unemployment
The UK and the USA Meen (2002)	Cointegration and ECM	Yes	Yes	Housing stock
The USA Gallinn (2006)	Panel cointegration	No	-	-
The USA Mikhed and Zemčík (2009)	Panel dana unit root and cointegration test	No	No	No cointegration found for any set of explanatory variables
17 developed countries Tsatsaronis and Zhu (2004)	VAR	No	Yes	Inflation
Euro area Annett (2005)	Panel regression	Yes	Yes	Real credit, real money
Australia Abelson et al. (2005)	Threshold ECM	Yes	Yes	Unempolyment, CPI, equity prices, housing stock
Sweeden Hort (1998)	Cointegration and ECM	Yes	-	User cost, construction cost
27 OECD and CEE countries Egert and Mihaljek (2007)	Panel regression	Yes	Yes	Real credit, population, and construction cost
3 CEE and 3 developed countries Posedel and Vizek (2009)	VAR	Yes	Yes	Housing loans
The USA Sei-Wan and Bhattacharya (2009)	STAR	-	Yes	-

Source: compiled by the authors

roots and cointegration have low power in the presence of asymmetric adjustment. Hence, if house prices exhibit nonlinear properties, as Sei-Wan and Bhattacharya (2009) claim, then nonlinear methods have to be applied if one wishes to examine how house prices may be influenced by the key variables.

To our knowledge there are only three papers dealing with nonlinear properties of house prices, namely, Abelson et al. (2005), Cook (2005), and Sei-Wan and Bhattacharya (2009). Abelson et al. (2005) estimate a cointegration and asymmetric error correction model for Australia. They use the Heaviside indicator function, which defines boom observations as observations for which the real price growth over the past year has been over two percent. These results suggest that the speed of adjustment (α) during boom periods has been somewhat greater than during non--boom periods (-0.21 and -0.14, respectively). Cook (2005) applies threshold cointegration on house price data for the UK regions and concludes that while conventional methods fail to detect the long-run relationships between house prices in the different regions of the UK, the threshold cointegration approach reveals the existence of a large number of cointegrating relationships. Moreover, the adjustment of discrepancies is asymmetric, with reversion to equilibrium occurring more rapidly in some regions.

Sei-Wan and Bhattacharya (2009) determine that a nonlinear smooth transition autoregressive model is able to explain house price growth rates in three out of four US regions much better than a linear autoregressive model. They also conduct the asymmetric Granger non-causality test and conclude that in a nonlinear framework mortgage rates had a significant impact on house prices. Specifically, mortgage rates had a stronger impact on house prices when the housing market was in an upswing than in a downswing. In the same framework, house prices explained employment, while the opposite was not true, which in turn indicated that house prices were not aligned with fundamentals.

Compared to developed countries, house prices in European transition countries are far less explored. Egert and Mihaljek (2007) used a panel data model composed of eight transition and 19 developed OECD economies and concluded that GDP and interest rates are the most important determinants of house prices, with their elasticities with respect to house prices being higher for transition countries, which exhibited a more intensive house price increase. The results of the analysis also suggested that growth of credit, population changes, and changes in construction costs also explained changes in house prices.

Posedel and Vizek (2009) applied the SVAR model in order to analyze house price determinants in three EU-15 countries and three European transition countries. In Croatia, Ireland, Poland, and Spain house price persistence was the most important determinant for explaining the variance of house prices. In the UK and Estonia, on the other hand, interest rates explain the biggest portion of the house price variance. Besides house price persistence and interest rates, GDP and housing loans were also important for explaining the variance of house prices, but to a lesser degree than house price persistence.

Zemčík (2009) tested the relationship between house prices and rents in the Czech Republic using panel data stationary techniques with the aim of determining whether there was a bubble in the Czech housing market. The results suggest that housing in the Czech Republic was somewhat overpriced. However, the degree of overpricing seems small, which in turn means that a large house price correction is not expected. Finally, according to that study, changes in rents in the capital city predicted changes in prices and vice versa, which indicates that house prices in the Czech Republic are aligned with fundamentals.

3. Empirical Analysis

3.1 Methodology

The analysis of non-stationary series for assets was first introduced by Campbell and Shiller (1987), who tested the present value model for bonds and stocks using cointegration. Following in their footsteps, many authors, including Hall et al. (1997), Hort (1998), Malpezzi (1999), Wang (2000), Meen (2002), Gallin (2006), Pagés and Maza (2007), McQuinn and O'Reilly (2008), and Mikhed and Zemčík (2009), have applied cointegration in order to model house prices.

In order to detect threshold effects in house price behavior, we take the cointegration approach to house price modeling one step further and use a threshold cointegration method developed by Enders and Siklos (2001). Unlike the Engle-Granger (1987) or Johansen (1996) methods, which assume linear behavior in the long and short run, this method allows for threshold adjustment in the short run while maintaining linearity in the long run. Threshold cointegration in essence allows for the discrete adjustment process observed in many economic situations (in particular those that involve transaction costs): movement toward long-run equilibrium may occur in some instances, but not in others. Therefore, threshold cointegration divides the available data span into two distinct regimes. In one regime the adjustment process is present and the strength of the error-correction effect depends, in part, on how far away from equilibrium the variable is, while in the other regime unit root behavior may dominate if adjustment is not present or takes place at a different speed. Linear error-correction behavior is then obtained as the average behavior across the two regimes (Balke and Fomby, 1997).

Threshold adjustment of house prices may exist due to the presence of households' rational responses to returns on the upside of the housing market. On the other hand, a response symmetric to that during a market upswing does not usually take place during a downswing (Sei-Wan and Bhattacharya, 2009). This discrepancy occurs because households are more likely to trade up during a housing cycle upswing, partly also because of equity constraints playing a minor role during upswings. In turn, house prices may adjust the long-run disequilibrium error more quickly during upswings. In the same vein, households are less likely to trade when prices are on the decline, causing stickiness on the downside of the housing market cycle and slowing down or completely eliminating the house price adjustment process.

We use and examine an explicit test for cointegration with asymmetric error correction. In this class of models, the Enders and Granger (1998) threshold autoregressive (TAR) and momentum-TAR (M-TAR) tests for unit roots are generalized to a multivariate context. In principle, the TAR model allows the degree of autoregressive decay to depend on the state of the variable of interest, while the M-TAR model allows a variable to display differing amounts of autoregressive decay depending on whether it is increasing or decreasing (Tong, 1983; Caner and Hansen, 1998; Enders and Siklos, 2001).

Firstly, in order to estimate the long-run equilibrium relationship, for each country we consider the following linear regression basis for cointegration tests:

$$x_{1t} = \beta_0 + \beta_1 x_{2t} + \beta_3 x_{3t} + ... + \beta_k x_{kt} + \mu_t \qquad (1)$$

where x_{1t} is a house price series, while $x_{2t},..,x_{kt}$ are house price determinants. All series are random variables integrated of degree 1. μ_t is the disturbance term, which may be serially correlated. k may vary from 2 to 4 depending on the determinants of house prices established for that country. A thorough explanation of the analyzed regression equations and the corresponding variables for each country is given in the *Appendix* Data Sources available *on the web site* of this journal. Equation (1) implies the existence of an error-correction representation of the variables (Engle and Granger, 1987), but the main issue is that these cointegration tests and their extensions are misspecified if adjustment is asymmetric.

After the Engle-Granger model of the long-run behavior of house prices is estimated, we adopt the notation of Enders and Siklos (2001) and consider alternative specifications of the error-correction model – the TAR and M-TAR models. In these models the disturbance term (μ_t) is used to formulate the threshold cointegration model with the following specification:

$$\Delta\mu_t = I_{jt}\rho_1\mu_{t-1} + \left(1 - I_{jt}\right)\rho_2\mu_{t-1} + \varepsilon_t \qquad j = 1,2 \tag{2}$$

where I_{1t} and I_{2t} are the Heaviside indicator functions for the TAR model and M-TAR model, respectively, such that

$$I_{1t} = \begin{cases} 1 & \text{if} \quad \mu_{t-1} \geq \tau_1 \\ 0 & \text{if} \quad \mu_{t-1} < \tau_1 \end{cases}$$

in the TAR case, and

$$I_{2t} = \begin{cases} 1 & \text{if} \quad \Delta\mu_{t-1} \geq \tau_2 \\ 0 & \text{if} \quad \Delta\mu_{t-1} < \tau_2 \end{cases}$$

in the M-TAR case. τ_1 and τ_2 are the values of the threshold and $\left(\varepsilon_t\right)$ is a sequence of independent and identically distributed random variables with mean zero and a constant variance. The residuals from (1) are used to estimate (2) and ε_t is independent of μ_s for $s < t$.

Equations (1) and (2) are consistent with a wide variety of error-correction models.

Since the least squares estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution (Tong, 1983, 1990), and given the existence of a single cointegrating vector, the error-correcting model for any variable x_{it} can be written in the form:

$$\Delta x_{it} = \rho_{1,i}I_{jt}\mu_{t-1} + \rho_{2,i}\left(1 - I_{jt}\right)\mu_{t-1} + \ldots + v_{i,t} \qquad j = 1,2 \tag{3}$$

where $\rho_{1,i}$ and $\rho_{2,i}$ are the speed of adjustment coefficients of Δx_{it}.

In general, the value of the threshold τ is unknown and needs to be estimated along with the parameters ρ_1 and ρ_2. First, we test for threshold cointegration using the TAR and M-TAR models setting the value of the threshold τ to zero. Moreover, we also test for threshold cointegration using the TAR and M-TAR models with an unknown threshold. We estimated the threshold τ using the Chan (1993) algorithm. In each of the four cases, depending on the type of asymmetry under consideration $\left(I_{1t} \text{ or } I_{2t}\right)$, a regression equation (2) was estimated using the ordinary least squares method. The significance and magnitude of the asymmetry parameters, ρ_1 and ρ_2, are necessary in order to establish if positive or negative departures from long-run equilibrium are eventually eliminated and if one of the possible discrepancies persists for a different length of time than the other one. From the specified regression, both the null hypothesis $\rho_i = 0$ and the joint hypothesis $\rho_1 = \rho_2 = 0$ were tested using the larger in magnitude of the *t* statistics (Tmax) and the *F* statistic (Φ), respectively. Since the necessary conditions for convergence are $\rho_1<0$, $\rho_2<0$ and $\left(1+\rho_1\right)\left(1+\rho_2\right) < 1$ for any value of the threshold τ (Petrucelli and Woolford, 1984), both the tests are direct tests of the existence of cointegration. The empirical F statistic Φ is compared to the critical values tabulated by Wade, Gilbert, and Dibooglu (2004). Finally, we test the null hypothesis $\rho_1 = \rho_2$ using the Wald test in order to determine whether the cointegration relationship is characterized by threshold effects in the short run.[2]

3.2 Data

The data set includes four developed countries (the United States, the United Kingdom, Spain, and Ireland) and four post-transition European countries (Bulgaria, Croatia, the Czech Republic, and Estonia). Aside from the house price series, the data set for each country comprises real GDP, the interest rate on a housing loan, total housing loans, employment, and construction activity. Since we adopted a comparative approach, we collected series that are as similar as possible across countries. The exception to this rule is the house price series, which is not fully comparable across countries due to methodological issues.

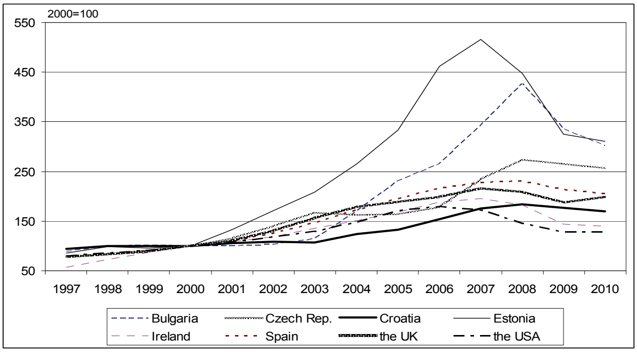
When modeling house prices for post-transition and developed countries one has to bear in mind that post-transition economies are characterized by many features related to housing markets which are not present in developed economies, but which heavily influence house price dynamics in post-transition countries. These features often cannot be proxied in applied econometric studies, but they can influence the study results. Thus, one has to take them into consideration when comparing the results for developed and post-transition countries. These specific post-transition features include the initial undershooting of house prices in the early 1990s, the poor quality of the existing housing stock, the limited supply of new residential units, weak housing market institutions, the initial absence and subsequent rapid development of housing finance, and external demand for housing. The same disclaimer has to be noted when comparing any European country to the United States due to the specific features of the US economy, such as its distinctive labor market and financial system characteristics, its migration patterns, and its settlement structure. A possible solution to these qualitative and scale differences between the US and any individual European country would be to compare the United States with the EU as a whole. However, the lack of a homogeneous EU-wide house price series precludes modeling the EU housing market as one observational unit.

The data range differs somewhat across countries, which is a consequence of the availability of house price series. The data for the developed countries start in the first quarter of 1995. The last observation available for Ireland is for the last quarter of 2008. For Spain and the UK, the data extend to the first quarter of 2009, while in the case of the USA, data are available up to the second quarter of 2009 (we used the Federal Housing Finance Agency house price index). As cointegration is a long-run phenomenon, we also tested for asymmetric adjustment in the USA and the UK, the two developed countries in our sample that have longer house price series. In the case of the USA we used quarterly data starting from 1975, while in the case of the UK we used annual data available from 1969. (See *Figure 1*.)

The data span for transition countries is somewhat shorter. The starting observation for Croatia is the fourth quarter of 1996, for Estonia it is the first quarter of 1997, and for Bulgaria and the Czech Republic it is the first quarter of 1998. The se-

[2] The null hypothesis assumes linearity, while the alternative assumes threshold behavior. Test statistics are denoted by *W* both in the text and in the corresponding tables.

Figure 1 House Price Developments



Source: Various sources, for details see in the *Appendix* Data Sources available *on the web site* of this journal.

ries for all transition countries end in the first quarter of 2009, except for the Czech Republic, where house price data are available until the second quarter of 2008. Series expressed in nominal terms, such as house prices, interest rates, and housing loans, were deflated using the consumer price index.

All the series were tested for unit roots using the Ng-Perron test (Perron and Ng, 1996). The results suggest that all the series are stationary in first differences. The results of the unit root test are presented here in the *Appendix*. All series except interest rates were transformed to logarithms. More details on all the series are available in in the *Appendix* Data Sources available *on the web* site of this journal.

3.3 Results

At the beginning of the empirical analysis, the Engle-Granger cointegration equation is estimated for each country. Aside from house prices being the dependent variable, the long-run equation incorporates the following explanatory variables: real GDP, the interest rate on a housing loan, the total amount of housing loans, employment, and construction activity. The residuals from the cointegration equation are then used to test for threshold cointegration. We tested for both TAR and M-TAR threshold cointegration, using the following two thresholds: 0 and a consistent estimate of the threshold as explained in section 3.1. If the tests did not detect the presence of any threshold cointegration, we left the explanatory variable which appears the least in the literature as a house price determinant, re-estimated the cointegration equation, and tested for threshold cointegration among the reduced number of variables. This procedure was repeated until the tests confirmed the existence of threshold cointegration among a given set of variables or until the cointegration equation was reduced to only three variables: house prices, the interest rate on a housing loan, and GDP. The first variable to be excluded from the model is housing loans, then employment, and finally construction activity. We decided to pursue this general-to-specific approach because we wanted to make sure that none of the potentially important house price determinants was omitted from the analysis. The results of this exercise suggested that threshold effects are present only in more parsimonious models. Specifically, in almost all cases, threshold cointegration was only confirmed in the most reduced trivariate case.

Table 2 Engle-Granger Cointegration Coefficients

Dependant variable: *house price$_t$*	BL	CRO	EE	CZ	IR	E	UK (1969)	UK (1995)	USA (1975)	USA (1995)
gdp_t	0.649 (130.0)	0.303 (2.78)	0.827 (181.0)	2.113 (12.0)	1.18 (933.0)	0.589 (213.0)	1.479 (18.9)	0.936 (222.0)	0.303 (14.87)	0.533 (9.204)
ir_t*	-0.0047 (-2.79)	-0.0099 (-2.96)	-0.0268 (-7.34)	0.0106 (2.78)	-0.0137 (7.36)	-0.0336 (-7.79)	-0.0087 (3.89)	-0.0058 (-6.52)	-0.0037 (-3.57)	-0.0097 (-3.87)

Notes: *t*-values presented in parenthesis.
* in order to obtain the interest rate elasticities, one must multiply the coefficients by 100.

Source: authors' calculation

Table 3 M-TAR Threshold Cointegration with an Unknown Threshold Summary of Estimation Results

	BL	CRO	EE	CZ	IR	E	UK (1969)	UK (1995)	USA (1975)	USA (1995)
ρ_1	-0.5437	-0.4584	-0.1723	-0.0577	-0.2427	0.00645	-0.1174	0.12531	-0.0024	-0.3664
ρ_2	-0.0039	-1.6528	-0.8835	-0.2029	-0.091	-0.20946	-0.508	-0.0545	-0.1504	-0.0627
Threshold value	0.0218	-0.0273	-0.051	-0.0139	0.00745	-0.0084	-0.0244	0.0322	-0.0042	0.00609
Tmax	-0.0793	-3.461	-1.6866	-1.0789	-1.3083	0.1196	-1.014	0.533	-2.002	-1.013
Φ $H0{:}\rho_1 = \rho_2 = 0$	42.024*	39.048*	14.303*	10.353**	3.0358	3.9301	5.3729	1.1957	34.12*	10.07**
W $H0{:}\rho_1 = \rho_2$	29.941*	11.845*	6.536**	1.7445	0.9563	2.896***	2.0148	0.5427	16.4*	4.54**

Notes: * null hypothesis rejected at 1 percent level of significance; ** null hypothesis rejected at 5 percent level of significance, *** null hypothesis rejected at 10 percent level of significance.

Source: authors' calculation

The Engle-Granger cointegration relationship coefficients for the trivariate case are displayed in *Table 2*. One can notice that all coefficients, except the interest rate coefficient for the Czech Republic, have the expected sign. The magnitude of the GDP coefficient ranges from 0.3 in the case of Croatia to 2.1 in the case of the Czech Republic, suggesting that the dispersion of the coefficients is larger for transition countries than for developed countries. The GDP coefficients for Ireland and the UK (when the sample starts from 1995) are close to unity, while in the case of Spain and the USA they are somewhat lower than unity. Egert and Mihaljek's (2007) findings also suggest that the dispersion of income coefficients is larger for transition countries than for OECD countries. The interest rate elasticities are rather high in some countries. In the USA, Croatia, Estonia, Ireland, and Spain they exceed the GDP elasticities in absolute value. The opposite is true in the UK, the Czech Republic, and Bulgaria.

Table 3 summarizes the most important findings relating to threshold cointegration. It displays the results of the M-TAR tests with an unknown threshold for the long-run equation consisting of three variables: house prices, the interest rate on a housing loan, and GDP. As was already stated, four different cases of threshold cointegration were tested: TAR with threshold 0, M-TAR with threshold 0, TAR with an unknown threshold, and M-TAR with an unknown threshold. The estimation results suggest that the M-TAR test with an unknown threshold was the most successful in detecting threshold cointegration. This should not come as a surprise given that the M-TAR test has greater power than the TAR test (Enders and Siklos, 2001). As suggested by the Φ statistic values, cointegration is confirmed in all four transition countries.[3] Comparing the value of the Φ statistic to the critical values tabulated in Wade, Gilbert, and Dibooglu (2004) also reveals that threshold cointegration is present in the USA in both samples, the one dating back to 1975 and the other dating back to 1995, thus supporting Sei-Wan and Bhattacharya's (2009) findings, which also suggest that house prices in the USA have asymmetric properties. For all countries which exhibit threshold cointegration except the Czech Republic, the Wald test for the equality of ρ_1 and ρ_2 suggests that the adjustment parameters are significantly different from each other.[4] For the Czech Republic, the equality of adjustment parameters is marginally accepted. Moreover, in the case of Bulgaria and the Czech Republic, the TAR test with an unknown threshold also indicated the presence of threshold cointegration. In the case of Estonia, the M-TAR test with an unknown threshold also detected threshold cointegration between house prices, GDP, the interest rate, and construction activity (details are displayed in the *Appendix* Data Sources available *on the web site* of this journal). On the other hand, in developed European countries no evidence of asymmetric adjustment was found. The results of the threshold cointegration tests which did not detect the presence of threshold cointegration can be obtained upon request from the authors.

After testing for threshold cointegration and for the equality of adjustment parameters, we proceeded by formulating a threshold error correction model of house prices for countries exhibiting threshold cointegration. The estimated coefficients and corresponding *p*-values of the adjustment parameters, the Granger causality test for lagged changes of house prices, GDP and interest rates, and diagnostic tests are presented in *Table 4*. One can notice that house prices are not weakly exogenous, i.e., they react to discrepancies from equilibrium in all countries. One must, however, note that in all countries house price adjustment occurs only during one regime, when the discrepancies are either larger or smaller than the threshold, while during the other regime unit root behavior persists. In the case of the USA (1995 sample) and Bulgaria, house prices adjust if the disequilibrium is smaller than the threshold, whereas in the USA (1975 sample), Estonia, Croatia, and the Czech Republic they adjust if the disequilibrium is larger than the threshold, while discrepancies smaller than the threshold persist. Expanding the USA sample thus reveals that the nature of house price threshold adjustment in the USA has shifted over time.

The statistically significant adjustment parameters for all countries except Croatia are also quite small and range from -0.029 in the case of the USA (1975 sample) to -0.181 in the case of Estonia. Even the adjustment parameter for Croatia (-0.55) is not large enough to correct all discrepancies in one period. One possible explanation for the lack of adjustment can be traced back to the results of the Granger causality tests for lagged values of house prices. Past house price changes in all countries except Croatia Granger cause contemporaneous house changes, which in turn suggests

[3] If one were to judge only on the basis of *t*-max statistics, the null hypothesis of no cointegration would not be rejected in the case of Bulgaria, the Czech Republic, and the USA (shorter sample). However, Enders and Siklos (2001) showed that in the M-TAR framework Φ statistics have substantially more power than *t*-max statistics. Hence, when ambiguity arises regarding the existence of cointegration, Φ statistics should be consulted.
[4] One must note that the M-TAR models for the USA, the Czech Republic, and Ireland were augmented with lagged changes of the residuals in order to account for autocorrelation.

Table 4 Threshold EC model – Summary of Estimation Results

Dependant variable: $\Delta house\_price_t$	BL	CRO	EE	CZ	USA (1975)	USA (1995)
Constant	-0.009 [0.05]	0.008 [0.241]	0.00072 [0.930]	0.006 [0.178]	-0.00039 [0.931]	-0.00034 [0.748]
ρ_1	0.036 [0.573]	-0.551 [0.011]	-0.1807 [0.03]	-0.091 [0.057]	-0.02997 [0.001]	-0.083 [0.251]
ρ_2	-0.064 [0.003]	-0.121 [0.73]	0.369602 [0.095]	-0.022 [0.855]	-0.02997 [0.245]	-0.071 [0.053]
$A_1(L)\Delta house\_price_{t-1}$*	30.118 [0.0000]	0.67127 [0.5758]	8.7720 [0.005]	9.25 [0.0002]	53.207 [0.0000]	24.092 [0.0000]
$A_2(L)\Delta gdp_{t-1}$*	1.3280 [0.2788]	1.3328 [0.2804]	13.783 [0.0006]	0.918 [0.47]	2.1893 [0.0743]	3.4427 [0.0105]
$A_3(L)\Delta ir_{t-1}$*	6.3324 [0.0047]	3.6345 [0.0227]	0.01302 [0.909]	0.539 [0.71]	0.43555 [0.7827]	3.7800 [0.0064]
R^2	0.75	0.52	0.39	0.75	0.71	0.88
Number of lags of explanatory variables	3	3	1	4	4	6
AR test	0.367 [0.777]	0.567 [0.688]	0.83 [0.518]	1.18 [0.34]	0.479 [0.79]	0.334 [0.85]
ARCH test	1.28 [0.30]	0.959 [0.447]	1.91 [0.134]	0.496 [0.69]	0.729 [0.57]	0.552 [0.70]

*Notes:** the numbers represent the *F* statistics and the corresponding *p*-values of the Granger causality test for the given variable; *p*-values are presented in brackets.

Source: authors' calculation

house price persistence is present. Consequently, because of the long memory of house prices, fundamentals take a longer time to kick in, which in turn prevents the house price adjustment from unfolding fully. Croatia is the only country where house price persistence does not seem to play a role and, consequently, its adjustment coefficient is much larger than those of other countries. This in turn might explain why Croatia did not experience such a dramatic house price increase when compared to the other countries.

The Granger causality test results reveal that changes in GDP lead to house price changes in Estonia and the USA (both samples), while interest rate changes lead to house price changes in Bulgaria, Croatia, and the USA (1995 sample). It is also quite interesting to note that interest rates do not Granger cause house prices in the USA when the threshold error correction model is estimated on the sample starting in 1975, while they do seem to matter from 1995 onwards. This suggests that financial liberalization in the USA during the last decade of the 20th century played an important role in house price developments. We can conclude that house prices were not entirely misaligned from fundamentals in the observed period. However, slow correction of disequilibrium in one regime coupled with house price persistence and unit root behavior in the other regime might have facilitated the emergence of the house price boom.

4. Concluding Remarks

The aim of this paper was to test whether house prices and their most important determinants are cointegrated in the long run, while the short-run adjustment of house prices is characterized by threshold effects. We show that the house price adjustment processes in four transition countries in Europe (Croatia, Bulgaria, the Czech Republic, and Estonia) are characterized by threshold effects. Threshold adjustment of house prices is also present in the USA. On the other hand, we find no evidence of threshold cointegration in three developed European countries that also witnessed strong house price appreciation. An asymmetric error correction model of house prices suggests that in Bulgaria, the Czech Republic, Estonia, and the USA, past values of house price changes Granger cause present house price changes. Thus, house price persistence, which prevents fundamentals from adjusting a disequilibrium, might provide some explanation for the fact that the threshold adjustment parameters are small in magnitude. In addition to house price persistence, the Granger causality test results indicate that changes in GDP lead to house price changes in Estonia and the USA, while interest rate changes influence house prices in Bulgaria, Croatia, and the USA (when tested on the shorter sample). This in turn suggests that house prices in the observed period were not completely detached from fundamentals. However, the emergence of the house price boom was perhaps supported by house price persistence coupled with either a slow adjustment process or a complete lack of adjustment.

APPENDIX

Results of Unit Root Test and Threshold Cointegration

Table 1 Ng-Perron Unit Root Test Results

Variable	House price		GDP		Interes rate		House loans		Employment		Construction activity	
	MZt test statistics											
Country	**Levels**	**1st diff.**	**Levels**	**1st diff.**	**Levels**	**1st diff.**	**Levels**	**1st diff.**	**Levels**	**1st diff.**	**Levels**	**1st diff.**
Bulgaria	0.004	-2.45**	2.72	2.12**	-1.12	2.32**	3.46	-1.87*	0.11	-3.16*	0.96	-3.78*
Croatia	-1.06	-1.92***	1.07	-3.28*	-0.89	-3.13*	2.73	-1.32	3.80	-2.83	-0.87	-1.46
Czech Rep.	1.74	-2.29**	1.43	-1.74***	-1.12	-3.03*	3.08	-2.53**	0.62	-1.60	0.59	-2.59*
Estonia	-0.34	-2.95*	0.46	-2.21**	-1.51	-2.04**	3.42	-2.32**	-1.25	-2.59*	-0.23	-3.58*
Ireland	0.77	-3.00*	1.88	-4.40*	-0.51	-3.66*	2.50	0.86	2.94	-1.02	-0.76	-3.71*
Spain	1.10	-1.63***	3.18	-3.93*	-0.97	-3.31*	5.29	-2.92*	1.42	0.91	-0.18	-3.09
UK	-0.52	-2.29**	2.15	-2.96	-1.34	-3.14*	0.40	-2.62*	3.77	-3.24	-0,71	-3.02*
USA	-1.08	-2.36**	3.41	-3.19*	-0.11	-2.24**	1.17	-1.96**	1.31	-3.12*	0.36	-1.52

Notes: Tests specification includes a constant. Bartlett kernel is used for estimating spectral density of the residuals. * null hypothesis rejected at 1 percent level of significance; ** null hypothesis rejected at 5 percent level of significance, *** null hypothesis rejected at 10 percent level of significance.

Source: calculation of the authors

Table 2 Bulgaria – Unknown Threshold

BULGARIA			TAR	
Threshold TAR =-0.1129			Parameters and tests	values
1 lag added			$\rho_1 =$	0.0033
Engle-Granger cointegration			$\rho_2 =$	-0.1033
variables	β coefficients	*t*-values	$\gamma_1 =$	0.3285
GDP	0.649	130.0	Tmax	0.0962
Interest rate on a housing loan	-0.0047	-2.79	Φ ($\rho_1 = \rho_2 = 0$) =	9.5394*
			W ($\rho_1 = \rho_2$) =	2.493
			Residuals	no autocorrelation

*Notes:** Null hypothesis rejected at 1 percent level of significance. ** null hypothesis rejected at 5 percent level of significance, Box-Ljung test for the autocorrelation of the residuals applied.

Source: calculation of the authors

Table 3 Estonia – Unknown Threshold

ESTONIA			M-TAR	
Threshold M-TAR =-0.04531			Parameters and tests	values
Engle-Granger cointegration			ρ_1 =	-0.45534
variables	β coefficients	*t*-values	ρ_2 =	-1.00901
GDP	0.258	4.30	Tmax	-2.94186
Interest rate on a housing loan	0.00013	0.0035	Φ ($\rho_1 = \rho_2 = 0$) =	30.5609*
Construction	0.965	9.48	W ($\rho_1 = \rho_2$) =	4.4315**
			Residuals	no autocorrelation

Notes:* Null hypothesis rejected at 1 percent level of significance. ** null hypothesis rejected at 5 percent level of significance, Box-Ljung test for the autocorrelation of the residuals applied.

Source: calculation of the authors

Table 4 Czech Republic – Unknown Threshold

CZECH REPUBLIC			TAR	
Threshold TAR =-0.0392			Parameters and tests	values
Engle-Granger cointegration			ρ_1 =	-0.0453
variables	β coefficients	*t*-values	ρ_2 =	-0.1848
Constant	-7.104	-12.6	Tmax	-0.7805
GDP	2.113	12.0	γ_1	0.5466
Interest rate on a housing loan	0.0106	0.0035	Φ ($\rho_1 = \rho_2 = 0$) =	9.7114**
			W ($\rho_1 = \rho_2$) =	2.1814
			Residuals	no autocorrelation

Notes:* Null hypothesis rejected at 1 percent level of significance. ** null hypothesis rejected at 5 percent level of significance, Box-Ljung test for the autocorrelation of the residuals applied.

Source: calculation of the authors

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