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# Contents

# The Nonlinear House Price Adjustment Process in Developed and Transition Countries

**Abstract:**
We use a nonlinear framework in order to explore house price determinants and their 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). All eight countries experienced an intensive increase in house prices during the 1990s and the first half of this decade. In addition to testing for nonlinearities, we focus on house price determinants in these four transition countries of Central and Eastern Europe. An asymmetric house price adjustment is present in all transition countries and the U.S., while no threshold effects are detected in developed European countries. In a threshold error correction framework, house prices are aligned with the 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 U.S.

**Key words:** house prices, threshold cointegration, asymmetric adjustment, transition
**JEL classification:** C22, R21, R31

# Nelinearnost procesa prilagodbe cijena nekretnina u razvijenim i tranzicijskim zemljama

**Sažetak:**
U članku se koristi nelinearna metodologija pomoću koje se istražuju determinante cijena nekretnina i karakteristike njihove kratkoročne prilagodbe. Testira se kointegracija s uključenim pragom na uzorku od četiri razvijene zemlje (SAD, Velika Britanija, Španjolska i Irska) i četiri tranzicijske zemlje (Bugarska, Hrvatska, Češka i Estonija). Svih je osam zemalja zabilježilo intenzivan rast cijena nekretnina tijekom zadnjeg desetljeća prošlog stoljeća i prve polovine ovog desetljeća. Osim testiranja nelinearnosti, članak se fokusira i na utvrđivanje determinanti cijena nekretnina u četiri tranzicijske zemlje Srednje i Istočne Europe. Opaža se da asimetrična prilagodba cijena nekretnina postoji u svim tranzicijskim zemljama i u SAD-u. Model korekcije odstupanja s pragom sugerira da cijene nekretnina odražavaju kretanje makroekonomskih fundamentala, no perzistentnost cijena nekretnina te spora i asimetrična prilagodba mogli su pogodovati eksploziji cijena nekretnina u tranzicijskim zemljama i SAD-u.

**Ključne riječi:** cijena nekretnina, kointegracija s uključenim pragom, asimetrična prilagodba, tranzicija
**JEL klasifikacija:** C22, R21, R31

# 1 Introduction*

Housing is an essential good, accounting for a large share of household expenditures and assets and a significant part of economic activity. By affecting the net wealth of households and their capacity to borrow and spend, as well as profitability and employment in the construction and real estate industries, developments in house prices have major economic implications. The importance of housing is reflected in the great number of papers on house price modeling. Thus far, the majority of empirical studies on house prices have been conducted using a linear framework for the data sample of developed countries. 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; Enders and Chumrusphonlert, 2004). Common sense would suggest that house prices also incorporate some nonlinear properties. Moreover, one of the few papers exploring house price nonlinearities (Kim and Bhattacharya, 2009: 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 an equity constraint plays 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 could be justified by asymmetric properties of house price determinants like GDP or interest rates (Neftci, 1984; Enders and Siklos, 2001). Threshold effects may also stem from high transaction costs inherent to the property transactions. As such, small deviations from the 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 housing markets tend to increase transaction costs.

The aim of this paper is to test for nonlinear house price properties, such as threshold cointegration and the asymmetric adjustment of house prices in relation to the long-run discrepancies proposed by Enders and Siklos (2001). We test the given methods on a

* This research was supported by a grant from the CERGE-EI Foundation under a program of the Global Development Network. All opinions expressed are those of the authors and have not been endorsed by CERGE-EI or the GDN. For the helpful suggestions and comments, we thank the participants of the GDN Regional Research Competition Conference held in Prague on August 9-10, 2009. We are also grateful to Petr Zemčík for his valuable comments and discussions.

sample that includes 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 uses this methodology on house prices and one of the few papers dealing with house price nonlinearities in general. By applying the threshold cointegration method, we want to explore whether house price nonlinearities have contributed to a house price boom. Furthermore, by incorporating Central and Eastern European countries in our sample, we explore house price properties and determinants in the region where house price appreciation has been more intensive when compared to developed countries that have experienced a house price boom. However, unlike in developed countries, housing markets in Central and Eastern European countries have not been intensively investigated, and this paper may shed more light on the subject and allow us to compare the characteristics and behavior of developed and transition housing markets.

The remainder of the paper is organized as follows. Section 2 is a review of the literature on house price modeling. 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 transition countries of Central and Eastern Europe. Section 3 presents the data and the applied methodology 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 given to house price modeling within a linear framework. In general, such studies use vector autoregression models, cointegration and error correction models, or panel data models in order to identify house price determinants. Some studies including Sutton (2002), McQuinn and O'Reilly (2008), Pagés and Maza (2003), 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. Égert and Mihaljek (2007) reached the same conclusion by examining a sample of developed and European transition economies.

Other studies like Gallin (2006), Shiller (2005), and Mikhed and Zemčík (2009) showed that changes in fundamentals did not explain the rapid growth of house prices in the U.S. 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 usual suspects such as income and interest rates, empirical studies also detected several other house price drivers. Abelson et al. (2005) showed that changes in 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 on Sweden (Hort, 1998), the U.S. (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 roots and cointegration have low power in the presence of an asymmetric adjustment. Hence, if house prices exhibit nonlinear properties as Kim 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 the best of our knowledge, there are only two papers dealing with nonlinear properties of house prices, i.e. Abelson et al. (2005) and Kim and Bhattacharya (2009). Abelson et al. (2005) estimate a cointegration and the asymmetric error correction model with 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 ($\alpha$) during boom periods has been somewhat greater when compared to non-boom periods (-0.21 and -0.14 respectively). However, one has to notice that the specification of an asymmetric error correction model does not rely on the statistical literature and, therefore, the power and test size properties for the asymmetric adjustment are not known. Moreover, the chosen model of asymmetric adjustment is not a generalization of any cointegration method, which in turn means that the cointegration test that the authors conducted might have been misspecified due to the presence of nonlinearities. Lastly, the estimates of two threshold adjustment parameters should have been tested for equality in order to make sure that the adjustment process indeed contains threshold effects. Since the difference between two adjustment parameters is very small, it is quite probable that, contrary to the conclusion of the study, there is no asymmetric adjustment of house prices in Australia.[1]

Kim and Bhattacharya (2009) determined that a nonlinear smooth transition autoregressive model is able to explain house price growth rates in three out of four U.S. regions much better than a linear autoregressive model. They also conducted the

[1] In the Enders and Siklos (2001) framework, one adjustment parameter has to be at least several times smaller or bigger than the other one in order to reject the null hypothesis of their equality and, thus, confirm the existence of threshold error correction.

asymmetric Granger non-causality test and concluded that in a nonlinear framework mortgage rates had a significant impact on house prices. Namely, mortgage rates had a stronger impact on house prices when the housing market was in an upswing rather 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 the fundamentals.

Unlike developed countries, house prices in European transition countries are far less explored. To the best of our knowledge, only a few papers on the subject have been published. Clearly, more studies are needed since house prices increased more intensively in transition countries while housing was in comparative terms less affordable (Vizek, 2009). Égert and Mihaljek (2007) estimated panels composed of eight transition and 19 developed OECD economies. Firstly, two panel variables were various proxies of income and interest rates while the third variable was varied. Using such a framework, Égert and Mihaljek 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 VAR methodology combined with a regression in order to analyze house price determinants in three EU-15 countries and three European transition countries. Their results suggest that in Croatia, Ireland, Poland, and Spain house price persistence was the most important determinant for explaining the variance of house prices. On the other hand, interest rates in the U.K. and Estonia 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. Supply side factors did not seem to play a role in short-run house price dynamics. Moreover, house prices in three EU-15 countries explained a significant fraction of GDP, construction activity, and interest rates variance.

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, the 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 the fundamentals.

# 3 Empirical Analysis

## 3.1 The 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 them, Hall et al. (1997), Hort (1998), Malpezzi (1999), Wang (2000), Meen (2002), Gallin (2006), Pagés and Maza (2003), McQuinn and O'Reilly (2008), and Mikhed and Zemčík (2009) applied cointegration in order to model house prices.

We take the cointegration approach to house price modeling one step further. While assuming that the long-run behavior of house prices and their determinants is symmetric, we allow for their asymmetric adjustment in the short-run. We use and examine an explicit test for cointegration with the asymmetric error correction developed by Enders and Siklos (2001) in order to examine a possible asymmetric adjustment toward a long-run cointegrating relationship. 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. More details about the basic TAR and M-TAR models can be found in Tong (1983), Caner and Hansen (1998), and Enders and Siklos (2001), for example.

As a starting point of our analysis, for each country we consider the following linear regression basis for cointegration tests in order to estimate the long-run equilibrium relationship:

$$x_{1t} = \beta_0 + \beta_1 x_{2t} + \beta_2 x_{3t} + \ldots + \beta_{k-1} x_{kt} + \mu_t, \tag{1}$$

where $x_{1t}$ is a house prices series, while $x_{2t},..,x_{kt}$ are house price determinants. All series are random variables integrated of degree 1. $\mu_t$ is the disturbance term that may be serially correlated, $k$ may vary from 2 to 4 depending on the established determinants of house prices for that country. A thorough explanation of the analyzed regression equations and the corresponding variables for each country is given in Appendix. The Granger representation theorem guarantees that in the presence of cointegration, Equation (1) implies the existence of an error-correction representation of the variables. The point is that these cointegration tests and their extensions are misspecified if adjustment is asymmetric. Therefore, we adopt the notation from Enders and Siklos (2001) and consider alternative specifications of the error-correction model, namely the TAR and M-TAR models given by:

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

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

$$I_{1t} = \begin{cases} 1 & \text{if} \quad \mu_{t-1} \geq \tau_1 \\ 0 & \text{if} \quad \mu_{t-1} < \tau_1 \end{cases} \tag{3}$$

in the TAR case, and

$$I_{2t} = \begin{cases} 1 & \text{if} \quad \Delta\mu_{t-1} \geq \tau_2 \\ 0 & \text{if} \quad \Delta\mu_{t-1} < \tau_2 \end{cases} \tag{4}$$

in the M-TAR case. $\tau_1$ and $\tau_2$ are the values of the threshold and $(\varepsilon_t)$ is a sequence of independent and identically distributed random variables with a zero mean and a constant variance, and the residuals from (1) are used to estimate (2). Furthermore, $\varepsilon_t$ is independent of $\mu_s$, for $s < t$.

Equations (1) and (2) are consistent with a wide variety of error-correction models, and the necessary and sufficient condition for the stationarity of $(\mu_t)$ is $\rho_1 < 0,\ \rho_2 < 0$ and $(1+\rho_1)(1+\rho_2) < 1$ for any value of the threshold $\tau$ (Petrucelli and Woolford, 1984); and the least squares estimates of $\rho_1$ and $\rho_2$ have an asymptotic multivariate normal distribution (Tong, 1983; 1990). Given the existence of a single cointegrating vector in the form of (1), the error-correcting model for any variable $x_{it}$ can be written in the form

$$\Delta x_{it} = \rho_{1,t} I_{jt}\mu_{t-1} + \rho_{2,i}(1 - I_{jt})\mu_{t-1} + \ldots + v_{i,t} \qquad j = 1,2, \tag{5}$$

where $\rho_{1,i}$ and $\rho_{2,i}$ are the speed of adjustment coefficients of $\Delta x_{it}$, and the latter can differ for each of the $\Delta x_{it}$.

In general, the value of the threshold $\tau$ is unknown and needs to be estimated along with the parameters $\rho_1$ and $\rho_2$. For both the models, we first set $\tau_{1,2} = 0$ in order for the cointegrating vector to coincide with the attractor and also in order to estimate the value of the threshold according to the algorithm specified in Enders and Siklos (2001) since there is no a priori reason to expect the threshold to coincide with the attractor.[2] In each of the cases, depending on the type of asymmetry under consideration $(I_{1t}$ or $I_{2t})$, a regression Equation (2) was estimated and both the null hypotheses $\rho_i = 0$ and $\rho_1 = \rho_2 = 0$ were tested using the larger of the *t*-statistics and the *F*-statistic respectively.[3] The sample statistics were then compared with the appropriate critical values from Enders and Siklos (2001). Also, if the alternative hypothesis of stationarity is accepted, it

[2] Estimates of the threshold and all test statistics related to threshold cointegration were obtained by using MATLAB.

[3] Furthermore, these statistics were denoted by Tmax and $\Phi$ both in the text and in the corresponding tables.

is possible to test for symmetric adjustment (i.e. $\rho_1 = \rho_2$), and this is done by performing the Wald test. Finally, diagnostic checking of the residuals are undertaken to ascertain whether the residual series $(\hat{\varepsilon}_t)$ satisfy the assumed properties of a white noise process. If the residuals were found to be correlated,[4] the model was re-estimated in the form of

$$\Delta\hat{\mu}_t = I_{jt}\rho_1\hat{\mu}_{t-1} + (1 - I_{jt})\rho_2\hat{\mu}_{t-1} + \sum_{k=1}^{p}\gamma_k\Delta\hat{\mu}_{t-k} + \varepsilon_t, \qquad j = 1,2 \tag{5}$$

where $(\hat{\mu}_t)$ is the residual series and $p$ is the lag length determined by an analysis of the regression residuals.

## 3.2 Data

We collected data for eight countries which experienced a prolonged increase in house prices in the last two decades. The data set includes four developed countries (the United States, the United Kingdom, Spain, and Ireland) and four transition countries (Bulgaria, Croatia, the Czech Republic, and Estonia). Table 1 displays house price developments in the analyzed countries, presents the cumulative increases of house prices recorded from 1998 to the point when house prices peaked, and the cumulative decreases of house prices recorded from the peak to the latest available data observation. We choose 1998 as a starting year because for some of the countries (Bulgaria and the Czech Republic) the data are not available before that year.

One can notice that there are substantial differences in both cumulative house price inflation and deflation among countries. The highest house price increase is recorded in Estonia where prices increased almost 400 percent in just nine years. A similar scenario is witnessed in Bulgaria and the Czech Republic where prices rose by 359 and 220 percent in approximately eleven years. One may speculate that astounding house price inflation in these three transition countries can be associated with some kind of "catching-up" process that has occurred due to a big gap in house price levels. On the other hand, an increase in house prices in the remaining transition country (Croatia) seems to be quite modest (89 percent). This is partially due to the fact that Croatia started its transition process with a somewhat higher house price level in comparison to other countries in the region. As opposed to the countries of Central and Eastern Europe (CEE), house price inflation in European countries seems to have been following a more coherent pattern. In all three countries, the prices have almost tripled in approximately eleven years. In the U.S., house prices measured by the Case-Schillier U.S. National Home Price Index rose 121 percent before reaching a peak in the second quarter of 2006.

[4] The Ljung-Box test was performed in order to test for autocorrelation of the residuals of the corresponding regression equation. The results are given in Appendix.

As far as house price deflation is concerned, the biggest cumulative drop has been recorded so far in the United Kingdom and the United States (40 and 30 percent respectively), followed by Estonia and Bulgaria (28 and 16 percent respectively). Lastly, one can notice that a house price peak across countries varies from the second quarter of 2006 in the case of the United States to the third quarter of 2008 in the case of Bulgaria. CEE countries exhibit more disparity when the dating of a turning point is in question, while house price cycles in European countries seem more synchronized.

Table 1 **House Price Developments, 1998-2009**

| | Spain | U.K. | Ireland | U.S. | Estonia | Bulgaria | Croatia | The Czech Rep.* |
|---|---|---|---|---|---|---|---|---|
| Cumulative increase | 193.4 | 178.5 | 205.2 | 121.6 | 399.2 | 359.5 | 89.0 | 220.9 |
| Cumulative decrease | -6.8 | -40.9 | -15.8 | -30.1 | -28.1 | -16.0 | -4.7 | - |
| House price peak point | 2008q1 | 2007q3 | 2007q2 | 2006q2 | 2007q1 | 2008q3 | 2007q4 | - |

*Note: *Data available until the second quarter of 2008.*
*Source: See Appendix.*

Aside from the house price series, the data set for each country is comprised of the 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. An exception to this rule is a house price series which is not fully comparable across countries due to methodological issues.

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

Data span for transition countries is somewhat shorter; i.e. 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. Series 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 series were tested for unit roots using the Ng-Perron test (Perron and Ng, 1996). The results suggest that all series are stationary in first differences. Due to space considerations, the results of the unit root test are not presented in this paper, but can be obtained upon request from the authors. All series except interest rates were transformed into logarithms. More details on all the series are available in Appendix.

## 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 a dependent variable, the long-run equation incorporates the following explanatory variables: the real GDP, the interest rate on a housing loan, the total amount of housing loans, employment, and construction activity. Residuals from the cointegration equation are then used to test for threshold cointegration. We tested for both TAR and M-TAR threshold cointegration, thereby using the following two thresholds: 0 and a consistent estimate of the threshold calculated by applying the Chan (1993) algorithm. If tests did not detect the presence of any threshold cointegration, we left out one explanatory variable, 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. 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. However, the results of the analysis suggested that more parsimonious models yield more plausible results; not only in terms of the presence of threshold cointegration, but also in terms of the sign and magnitude of the long-run coefficients. Namely, in almost all cases, the threshold cointegration was only confirmed in the most reduced trivariate case.

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 case of the Czech Republic; suggesting that the dispersion of the coefficients is larger for transition countries in comparison to developed countries. GDP coefficients for Ireland, and the U.K. (when the sample starts from 1995) are close to unity. While in the case of Spain, and the U.S. it is somewhat lower than unity. Égert and Mihaljek's (2007) findings also suggest that the dispersion of income coefficients is larger for transition countries in comparison to OECD countries.

Interest rate elasticities are rather high in some countries; in the U.S., Croatia, Estonia, Ireland, and Spain they exceed in absolute value GDP elasticities. The opposite is true in the U.K., the Czech Republic, and Bulgaria.

| Table 2 **Engle-Granger Cointegration Coefficients** | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| **Dependant variable:** *house price$_t$* | **Bulgaria** | **Croatia** | **Estonia** | **The Czech Rep.** | **Ireland** | **Spain** | **U.K. (1969)** | **U.K. (1995)** | **U.S. (1975)** | **U.S. (1995)** |
| Constant | - | 1.88 (4.41) | - | -7.104 (-12.6) | - | - | - | - | 1.354 (35.75) | 0.951 (7.892) |
| $gdp_t$ | 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 in parenthesis. *In order to obtain interest rates elasticities, one must multiply coefficients by 100.*
*Source: Authors' calculation.*

Table 3 summarizes the most important findings related to the threshold cointegration. It displays the results of the M-TAR tests with the 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 the threshold 0, M-TAR with the threshold 0, TAR with an unknown threshold, and M-TAR with an unknown threshold. The estimation results suggest that the M-TAR test with the unknown threshold was the most successful in detecting the threshold cointegration, which should not come as a surprise given the fact that the M-TAR has greater power when compared to the TAR test (Enders and Siklos, 2001). As suggested by the $\Phi$ statistic values, asymmetric adjustment of house prices to disequilibrium is present in all four transition countries.[5] The $\Phi$ statistic is also significant for the U.S. when tested on both samples; one dating back to 1975 and the other dating back to 1995, thus supporting Kim and Bhattacharya (2009) findings, that also suggest house prices in the U.S. have asymmetric properties. For all countries which exhibit threshold cointegration except the Czech Republic, the Wald test for the equality of $\rho_1$ and $\rho_2$ suggests that adjustment parameters are significantly different from each other.[6] 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 the unknown threshold also indicated the presence of threshold cointegration. In the case of Estonia, the M-TAR test with the unknown threshold also detected a threshold cointegration between house prices, GDP, the interest rate and construction activity (details are displayed in Appendix). On the other hand, in developed European countries no evidence of asymmetric adjustment was found. The results of threshold cointegration tests which did not detect the presence of threshold cointegration can be obtained upon request from the authors.

[5] If one would 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 US (shorter sample). However, Enders and Siklos (2001) showed that in the M-TAR framework $\Phi$ statistics has substantially more power than t-max statistics. Hence, when ambiguity regarding the existence of cointegration arises, $\Phi$ statistics should be consulted.

[6] One must note that M-TAR models for the US, the Czech Republic and Ireland were augmented with lagged changes of the residuals in order to account for autocorrelation. Parameters $\gamma_1$, $\gamma_2$, and $\gamma_3$ presented in Table 3 are estimated coefficients of the lagged values of the residual changes.

Table 3 **M-TAR Threshold Cointegration with the Unknown Threshold – Summary of Results**

| | Bulgaria | Croatia | Estonia | The Czech Rep. | Ireland | Spain | U.K. (1969) | U.K. (1995) | U.S. (1975) | U.S. (1995) |
|---|---|---|---|---|---|---|---|---|---|---|
| $\rho_1$ | -0.5437 | -0.4584 | -0.1723 | -0.0577 | -0.2427 | 0.00645 | -0.1174 | 0.12531 | -0.0024 | -0.3664 |
| $\rho_2$ | -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 |
| $\Phi$ $H_0: \rho_1 = \rho_2 = 0$ | 42.024* | 39.048* | 14.303* | 10.353** | 3.0358 | 3.9301 | 5.3729 | 1.1957 | 34.12* | 10.07** |
| $W$ $H_0: \rho_1 = \rho_2$ | 29.941* | 11.845* | 6.536** | 1.7445 | 0.9563 | 2.896*** | 2.0148 | 0.5427 | 16.4* | 4.54** |
| $\gamma_1$ | - | - | - | 0.5584 | 0.163 | - | - | - | 0.429 | - |
| $\gamma_2$ | - | - | - | - | - | - | - | - | 0.0504 | - |
| $\gamma_3$ | - | - | - | - | - | - | - | - | 0.252 | - |

*Notes: *Null hypothesis rejected at a 1 percent significance level. **Null hypothesis rejected at a 5 percent significance level. ***Null hypothesis rejected at a 10 percent significance level. $\rho_1$ and $\rho_2$ are adjustment parameters, parameters $\gamma_1$, $\gamma_2$, and $\gamma_3$ are estimated coefficients of the lagged values of residual changes.*
*Source: Authors' calculation.*

The threshold value is negative for four countries out of five, which exhibit an asymmetric adjustment of house prices to GDP and interest rates. Moreover, for those countries (Croatia, Estonia, the Czech Republic, and the U.S. (1975 sample)) the adjustment is faster when the disequilibrium is below the threshold, while the adjustment is slower when the disequilibrium is larger than the threshold. However, when the opposite is true, the adjustment is much slower (in the case of Croatia and Estonia 46 and 17 percent, respectively, of disequilibrium is adjusted in the next period) or barely takes place (as in the case of the U.S., and the Czech Republic where only 0.2 and 6 percent of disequilibrium is adjusted respectively). Results of the M-TAR exercise for Bulgaria indicate that house prices in Bulgaria adjust more strongly when the disequilibrium is larger than the threshold. The same is true for the U.S. if one should judge from the estimation on the shorter sample. One must, however, note that the TAR test with an unknown threshold for Bulgaria also yielded a negative value of the threshold, while $|\delta_2| > |\delta_1|$.

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. Estimated coefficients and respective 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 the equilibrium in all countries. One must, however, note that in all countries house prices adjust only if discrepancies are either larger or smaller than the threshold. In the case of the U.S. (1995 sample), and Bulgaria house prices adjust if disequilibrium is lower than the threshold. While in the U.S. (1975 sample), Estonia, Croatia, and the Czech Republic they adjust if disequilibrium is larger than the threshold, while the discrepancies smaller than the threshold persist. Expanding the U.S. sample thus reveals that the nature of the house price threshold adjustment in the U.S. has shifted during time.

Statistically significant adjustment parameters for all countries except Croatia are also quite small and range from -0.029 in the case of the U.S. (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. Namely, a country whose house prices do not fully adjust to disequilibrium also exhibit house price persistence. Namely, in Bulgaria, the Czech Republic, Estonia, and the U.S. past values of house price changes Granger cause present house price changes. In such a situation, one would expect that fundamentals take a longer time to kick in, which in turn prevents the adjustment to unfold 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 when compared to other countries. This in turn might explain why Croatia did not experience such a dramatic house price increase when compared to other countries.

Granger causality test results reveal that changes in GDP lead to house price changes in Estonia and the U.S. (both samples), while interest rate changes lead to house prices in Bulgaria, Croatia, and the U.S. (1995 sample). It is also quite interesting to note that the interest rates do not Granger cause house prices in the U.S. when 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 U.S. during the last decade of 20$^{th}$ century played an important role in house price developments. We can conclude that house prices were not entirely misaligned from the fundamentals in the observed period. However, a slow and asymmetric correction of disequilibrium coupled with house price persistence probably facilitated the emergence of the house price boom.

| Table 4 **Threshold EC Model – Summary of Results** | | | | | | |
|---|---|---|---|---|---|---|
| **Dependant variable:** $\Delta house\_price_t$ | **Bulgaria** | **Croatia** | **Estonia** | **The Czech Rep.** | **U.S. (1975)** | **U.S. (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] |
| $\rho_1$ | 0.036 [0.573] | -0.551 [0.011] | -0.1807 [0.03] | -0.091 [0.057] | -0.02997 [0.001] | -0.083 [0.251] |
| $\rho_2$ | -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: *Numbers represent F-statistics and the corresponding p-values of the Granger causality test for the respective variable. P-values are presented in brackets.*
*Source: Authors' calculation.*

# 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 the house prices is characterized by threshold effects. We show that the adjustment process of house prices in four transition countries in Europe (Croatia, Bulgaria, the Czech Republic, and Estonia) that experienced an intensive increase of house prices is asymmetric. The asymmetric adjustment of house prices is also present in the U.S. 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 U.S., past values of house price changes Granger cause present house price changes. Thus, house price persistence, which prevents fundamentals from kicking in and adjusting the disequilibrium, might provide some explanation for the fact that threshold adjustment parameters are small in magnitude. In addition to house price persistence, Granger causality test results also indicate that changes in GDP lead to house price changes in Estonia and the U.S.; while interest rate changes influence house prices in Bulgaria, Croatia, and the U.S. (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 supported by house price persistence coupled with a slow and asymmetric adjustment process.

# Appendix

## Data Description and Sources

| Country: **Croatia**<br>Data range: 1996:Q4 – 2009:Q1 | | |
|---|---|---|
| House price | Real Estate Exchange Database (Burza Nekretnina) | Average purchase-sale of all housing units (houses and apartments; old and used) consisting the database |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Central Bureau of Statistics | Volume of construction works undertaken by legal entities with 25 or more employees, 2000=100 |
| Number of employed persons | Central Bureau of Statistics | Total number of employed persons in legal entities, crafts and free lance activities, in 000 |
| Housing loans | Croatian National Bank | Housing loans series is available from July 1999, before July 1999 the series was reconstructed using growth rates of total loans to households, in millions EUR |
| Short term interest rate | Croatian National Bank | Overnight money market rate |
| Interest rate on housing loans | Croatian National Bank | Average annual interest rates to housing loans is available since January 2002, before 2002 average annual interest rate for long-term housing loans with currency clause series was mean adjusted and used |
| CPI | International Financial Statistics | Consumer price inflation, base index. Before 1998, the retail price index was used |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices |

| Country: **Bulgaria**<br>Data range: 1998:Q1 – 2009:Q1 | | |
|---|---|---|
| House price | National Statistical Institute | Average market prices of homes, quarterly |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Eurostat | Construction production index, 2005=100 |
| Number of employed persons | Eurostat | Total employment – national concept, in 000 |
| Housing loans | Bulgarian National Bank | Loans for house purchase, in 000 BGN |
| Short term interest rate | Eurostat | Overnight money market interest rate |
| Interest rate on housing loans | Bulgarian National Bank | Average interest rate on EUR loan for house purchase |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices |

| Country: **Estonia**<br>Data range: 1997:Q1 – 2009:Q1 | | |
|---|---|---|
| House price | Estonian Statistics | Average purchase-sale price per square meter of a two room and a kitchen dwellings of satisfactory condition in capital city (Tallin) intermediated by real estate agencies, in EUR<br><br>The series is highly correlated with average purchase-sale price series for entire Estonia which could not be used since it starts from 2002 |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Eurostat | Construction production index, 2005=100 |
| Number of employed persons | Eurostat | Total employment – national concept, in 000 |
| Housing loans | Bank of Estonia | Total housing loans, in millions EUR |
| Short term interest rate | Bank of Estonia | 1 month TALIBID rate |
| Interest rate on housing loans | Bank of Estonia | Weighted average annual interest rate to housing loans granted to individuals |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices |

| Country: **The Czech Republic**<br>Data range: 1998:Q1 – 2008:Q2 | | |
|---|---|---|
| House price | Czech Statistical Office | Apartment price indices (2005=100) |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Eurostat | Construction production index, 2005=100 |
| Number of employed persons | Eurostat | Total employment – national concept, in 000 |
| Housing loans | National Bank of Czech Republic | Lending to households for long-term house purchase, in millions EUR |
| Short term interest rate | National Bank of Czech Republic | NBCRs´ refinancing rate |
| Interest rate on housing loans | International Financial Statistics | Interest rate charged on loans to households |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices, 2000=100 |

| | | |
|---|---|---|
| Country: **Spain**<br>Data range: 1995:Q1 – 2009:Q1 | | |
| House price | National Institute of Statistics | Average price pre square meter of a real, in EUR |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Eurostat | Construction production index, 2005=100 |
| Number of employed persons | Eurostat | Total employment – national concept, in 000 |
| Housing loans | Bank of Spain | Total housing loans, in millions EUR |
| Short term interest rate | Bank of Spain | Interbank overnight rate |
| Interest rate on housing loans | Eurostat;<br>Bank of Spain | For the period from 1995 Q1 – 2003 Q1 average annual interest rate on housing loans for households, from 2003 Q2 onwards average interest rate on housing loans over 5 years maturity, outstanding amount |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices, 2000=100 |

| | | |
|---|---|---|
| Country: **United Kingdom**<br>Data range: 1995:Q1 – 2009:Q1 and 1969 – 2008 (annual frequency) | | |
| House price | Department for Communities and Local Government | Average sale prices of new and old house, in EUR |
| Gross domestic product | Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates) |
| Construction production index | Eurostat | Construction production index, 2005=100 |
| Number of employed persons | Eurostat | Total employment – national concept, in 000 |
| Housing loans | Bank of England | Total secured sterling lending to individuals and house associations, outstanding amount, in millions EUR |
| Interest rate on housing loans | Bank of England | Average standard variable mortgage rate to households |
| Short term interest rate interest rate | Bank of England | BoEs´ official interest rate |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices, 2000=100 |

| Country: **Ireland**<br>Data range: 1995:Q1 – 2008:Q4 | | |
|---|---|---|
| House price | Department for environment, heritage and local government | Average national new house price, in EUR |
| Gross domestic product | Irish Statistical Office Eurostat | Gross domestic product, millions EUR, chain-linked volumes, reference year 2000 (at 2000 exchange rates)<br><br>The data for period 1995:Q1 – 1996:Q4 were reconstructed using quarterly growth rates of industrial production volume from Irish statistical office |
| House completion index | Irish Statistical Office | Calculated using the quarterly series of house completion number in all local authorities, 2000=100 |
| Number of employed persons | Irish Statistical Office | Persons aged 15 years and over in employment , in 000 |
| Housing loans | Department for environment, heritage and local government | Total housing loan payments, banks and building societies, in millions EUR |
| Interest rate on housing loans | Department for environment, heritage and local government | Average annual building society mortgage interest rate |
| Short term interest rate interest rate | Bloomberg | EONIA |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices, 2000=100 |

| Country: **United States**<br>Data range: 1975:Q1 – 2009:Q2 | | |
|---|---|---|
| House price | Federal Housing Finance Agency | FHFA house price index - all transactions<br>(for the analysis of 1975-2009 period) |
| Gross domestic product | International Financial Statistics | Gross domestic product, millions of US$, chain-linked volumes, reference year 2005 |
| Number of employed persons | International Financial Statistics | Total employment, in 000 |
| Housing loans | Federal Reserve Board | Total real estate loans – all commercial banks |
| Interest rate on housing loans | International Financial Statistics | Mortgage rate |
| Short term interest rate interest rate | International Financial Statistics | FED discount rate |
| CPI | International Financial Statistics | Consumer price inflation, base index |
| CPI deflator | International Financial Statistics | Calculated by using quarterly base index of consumer prices, 2000=100 |

## Results of Threshold Cointegration

| Table A1 **Bulgaria - Unknown Threshold** | | | | |
|---|---|---|---|---|
| **Threshold TAR =-0.1129** | | | **TAR** | |
| | | | **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 | $\Phi(\rho_1 = \rho_2 = 0) =$ | 9.5394* |
| | | | $W(\rho_1 = \rho_2) =$ | 2.493 |
| | | | Residuals | No autocorrelation |

*Notes: *Null hypothesis rejected at a 1 percent significance level. The Box-Ljung test for the autocorrelation of the residuals is applied.*
*Source: Authors' calculation.*

| Table A2 **Estonia - Unknown Threshold** | | | | |
|---|---|---|---|---|
| **Threshold M-TAR = -0.04531** | | | **M-TAR** | |
| | | | **Parameters and tests** | **Values** |
| Engle – Granger cointegration | | | $\rho_1 =$ | -0.45534 |
| Variables | β coefficients | t -values | $\rho_2 =$ | -1.00901 |
| GDP | 0.258 | 4.30 | Tmax | -2.94186 |
| Interest rate on a housing loan | 0.00013 | 0.0035 | $\Phi(\rho_1 = \rho_2 = 0) =$ | 30.5609* |
| Construction | 0.965 | 9.48 | $W(\rho_1 = \rho_2) =$ | 4.4315** |
| | | | Residuals | No autocorrelation |

*Notes: *Null hypothesis rejected at a 1 percent significance level. **Null hypothesis rejected at a 5 percent significance level. The Box-Ljung test for the autocorrelation of the residuals is applied.*
*Source: Authors' calculation.*

| Table A3 **The Czech Republic - Unknown Threshold** | | | | |
|---|---|---|---|---|
| **Threshold TAR = -0.0392** | | | **TAR** | |
| | | | **Parameters and tests** | **Values** |
| Engle-Granger cointegration | | | $\rho_1 =$ | -0.0453 |
| Variables | β coefficients | t -values | $\rho_2 =$ | -0.1848 |
| Constant | -7.104 | -12.6 | Tmax | -0.7805 |
| GDP | 2.113 | 12.0 | $\gamma_1 =$ | 0.5466 |
| Interest rate on a housing loan | 0.0106 | 2.78 | $\Phi(\rho_1 = \rho_2 = 0) =$ | 9.7114** |
| | | | $W(\rho_1 = \rho_2) =$ | 2.1814 |
| | | | Residuals | No autocorrelation |

*Notes: **Null hypothesis rejected at a 5 percent significance level. The Box-Ljung test for the autocorrelation of the residuals is applied.*
*Source: Authors' calculation.*

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