What drives housing price dynamics in Greece:
New evidence from asymmetric ARDL cointegration

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Abstract

In this paper, we study the dynamics between house prices and selected macroeconomic fundamentals in Greece. The empirical analysis applies the asymmetric ARDL cointegration methodology proposed by Shin, Yu and Greenwood-Nimmo (2011) over the period January 1999 to May 2011. The evidence suggests that ignoring the intrinsic non-linearities may lead to misleading inference. In particular, the results reveal significant differences in the response of house prices to positive or negative changes of the explanatory variables in both the long- and short-run time horizon. The obtained evidence of asymmetry could be of major importance for more efficient policymaking and forecasting in the Greek house market.

Keywords: Asymmetric cointegration, nonlinear ARDL (NARDL), asymmetric dynamic multipliers, house prices.

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1. Introduction

In Greece, residences are considered to be the most significant wealth component of households. In fact, housing constitutes the 80-90% of the total household wealth, with homeownership reaching 78% (Simigiannis and Hondrogiannis, 2009). The housing sector has always performed a major role in the overall growth of the Greek economy. Between 2000 and 2008, investments in the housing sector had accounted for 7.5% of GDP (Hardouvelis, 2009) reaching 9.1% in 2006. Furthermore, over the period 2003-2008, reduction in interest rates made residences a more attractive wealth component, which in conjunction with the increase in house prices between 1993 and 2007 (mean growth rate of 4.9%) resulted in housing loans to account for 75% of the investment in residences and 31% of the private sector loans' portfolio of the Greek banking sector. Over the last two decades, rising house prices undoubtedly have produced a substantial wealth effect for homeowners.

The international literature on housing market suggests as major determinants of the behaviour of house prices, consumer prices, income, interest rates, constructions costs and housing credit, with the first two being the most cited. More particularly, Malpezzi (1990 and 1999) investigated the relationship between house prices and income posing a long-run equilibrium ratio between typical house prices, and income. Theory based on the permanent income hypothesis argues that in any particular period housing consumption is a stable function of the average income over the current cycle. However, this permanent income hypothesis proves insufficient to explain the fluctuations of prices in the housing market. This could be attributed to the fact that housing is a multidimensional commodity being regarded both as a durable consumer good and as an asset for investment (Chen et al. 2007).

With respect to the impact of inflation on the housing sector, there exist different views (Feldstein, 1992; Poterba, 1992). Feldstein (1992) supported that increasing inflation discourages people to invest in real estate, which in turn lowers housing demand. On the other hand, it can be argued that inflation pushes nominal housing payments and construction costs upwards, which implies a lower housing demand.

The majority of the research on house price modelling has been conducted in a linear framework. However, many macroeconomic variables incorporate nonlinear properties, especially in the area of business cycles (Neftci, 1984; Falk, 1986). As house prices are driven by the economic activity they could also be expected to exhibit nonlinearities. This possibly implies that linear models may not be appropriate to explore the determinants of house prices and could provide misleading evidence. More specifically, in the presence of nonlinearities, the response of the house market to positive shocks in the economy's price level may be different from the response to negative shocks.

Furthermore, the two aforementioned key determinants of house prices, inflation and income, have been found by the relevant empirical literature to asymmetrically affect a wide range of macroeconomic fundamentals (Holmes, 2000; Wang et al., 2008; Kamalian et al.,2010). Hence, inflation and income are possible candidates for causing asymmetric impacts on the house market.

To investigate the potential asymmetric relationship between house prices and their determinants, this paper employs the asymmetric cointegration methodology making further steps regarding the nonlinear framework. In particular, the asymmetric Autoregressive Distributed Lag (ARDL) cointegration technique is deployed, to test for a possible nonlinear relationship between Greek house prices and a set of macroeconomic variables which includes consumer prices and income proxied by the industrial production index. To our knowledge, this represents an enrichment of the existing literature on house prices modelling, and for the Greek house market in particular.

The rest of the paper is structured as follows: the second section presents the relevant literature on house price modeling, the third section focuses upon the methodology applied, while the fourth section presents the model and the empirical results. Finally, the last section provides a short summary and conclusions.

2. Literature review

Current research on the factors that affect house prices uses a main body of determinants that mostly includes inflation, interest rates, and GDP as a proxy of income. This set is usually completed by other more specific determinants, regarding the particular aim of the research, such as construction costs, housing credit, stock market index, money supply, employment, demographic factors and others. The relevant studies employ either vector autoregression models and error correction models, or panel data models, assuming a linear framework.

Baffoe-Bonnie (1998), employed the VAR methodology and concluded that mortgage rates, inflation, employment and money supply have significant effects on house prices and the stock of houses sold, while Sutton (2002), confirmed the importance of income, interest rates and stock prices for several developed countries.

Tsatsaronis and Zhu (2004), estimated a structural vector autoregressive (SVAR) model of house prices for 17 countries using five endogenous variables: the growth rate of GDP, inflation, the real short-term interest rate, the growth rate of inflation-adjusted bank credit and the term-spread calculated as the yield difference between a long-maturity government bond and the short-term interest rate. Their findings suggest that inflation accounts for more than half of the total variation in house prices in a five-year horizon, while in the short-run the impact is even greater. Furthermore, variables concerning the financial sector, such as the bank credit, the short-term interest rate and the term-spread, explain almost one third of the house prices variance in the long-run, while the growth rate of GDP has very little impact on house prices variation.

Égert and Mihaljek (2007), confirmed that house prices in 19 OECD and 8 Central and Eastern European countries (CEE) are driven, to a large extent, by the conventional variables such as GDP per capita, real interest rates, housing credit and demographic factors. They also add transition-specific factors such as the institutional development of housing markets and housing finance and quality effects.

Furthermore, McQuinn and O'Reilly (2008), estimated a model for the Irish housing market, where the demand for housing is determined by the amount of loans that can be offered to individuals, based on their disposable income and interest rates. Their results indicated the existence of a long-run relationship between house prices and the money amount borrowed by individuals.

In a recent study, Adams and Füss (2010) applied panel cointegration methodology to study the long and short-term dynamics of construction costs, long-term interest rates and economic activity on house prices of 15 countries. They described economic activity with a set of variables which included real money supply, real consumption, real industrial production, real GDP and employment. According to their results, an increase in economic activity has a positive effect on house prices while long-term interest rates and construction costs have a negative and positive effect respectively.

The relevant literature for the Greek housing market is rather limited. Apergis and Rezitis (2003) studied the effects of specific macroeconomic factors on house prices in Greece, i.e. inflation, employment, money supply and the mortgage interest rate. Their results, derived by means of variance decomposition analysis, suggested that all the variables under consideration affect house prices, with the mortgage interest rate having the largest explanatory power.

Merikas et al. (2009), employed an error correction model of house prices in Greece considering inflation, unemployment, the long-run interest rate, the production index and the Athens Stock exchange general index as explanatory variables. Their empirical results indicated inflation as the most important determinant of house prices.

Simigiannis and Hondrogiannis (2009) studied the recent experience in the Greek housing market by estimating a model where house prices were determined by the amount of loans that is offered to individuals based on their income and the mortgage interest rates, as well as by a parameter that depends on the demand and supply house price elasticity.

Finally, Brissimis and Vlassopoulos (2008) analyzed the interaction between house prices and mortgage lending. They found that, in the long-run the causation does not run from mortgage lending rate to house prices, while, in the short-run, a bi-directional dependence is acceptable.

All the studies mentioned above assumed that house prices and their determinants are linearly related. In fact, economic theory is mostly concerned with the set of the appropriate determinants but not the functional form of the relationship between them. The imposition of a linear long-run relationship is a rather restrictive assumption and may not be appropriate when the cointegration relationship is in fact nonlinear. A limited number of studies employing nonlinear models are available. Kim and Bhattacharya (2009), have employed a smooth transition autoregressive (STAR) model to examine the non-linearity properties of house prices in the US housing market while Zhou (2010), used the ACE algorithm and applied non-linear methodology to examine the relationship between house prices and macroeconomic fundamentals for the US housing market. Tsai et al. (2011), investigated the relationship between the US housing and stock markets applying the momentum-threshold autoregressive (M-TAR) model to test for non-linear cointegration. The results provided evidence of asymmetric wealth effects in the examined markets. All these research efforts confirmed the existence of non-linearities between house prices and the examined fundamentals.

3. Methodology and model structure of the asymmetric ARDL cointegration

The existing literature concerning asymmetry is dominated by three regime-switching models. First, the threshold ECM (Balke and Fomby 1997), where regime shifts are triggered by the level of observed variables in relation to an unobserved threshold. Second, the Markov-switching ECM (Psaradakis et al. 2004) in which the regime shifts evolve according to a Markov chain. And third, the smooth transition regression ECM (Kapetanios et al. 2006) which considers the threshold ECM as a special case by allowing the transition from one regime to another as a smooth function.

Granger and Yoon (2002) introduced the term “hidden cointegration”. According to them, two time series have hidden cointegration if their positive and negative components are cointegrated each other. They showed that, standard linear (symmetric) cointegration is a special case of hidden cointegration and hidden cointegration is simple case of nonlinear cointegration. Schorderet (2003), proposed a bivariate asymmetric cointegrating regression to analyze hidden cointegration where only one component of each series appears in the cointegrating relationship.

In a more recent paper, Shin et al. (2011) developed an asymmetric ARDL cointegration methodology, which uses positive and negative partial sum decompositions, allowing for the detection of asymmetric effects both in the long- and the short-run. Actually, the specification of the asymmetric ARDL allows the joint analysis of the issues of non-stationarity and non-linearity in the context of an unrestricted error correction model.

In the context of cointegration, if house prices and the examined macroeconomic fundamentals are found cointegrated, means that, although they may temporarily drift apart from each other, in the long-run they tend to return to equilibrium. We can discriminate between three possible cases; the existence of linear cointegration, the existence of nonlinear cointegration, and lack of cointegration.

The conventional cointegration approach initially employed in this paper, is based on the linear ARDL model proposed by Pesaran and Shin (1998) which performs better for determining cointegrating relationships in small samples (Pesaran et al. 1996, Pesaran et al. 2001 and Romilly et al. 2001). It also maintains the additional advantage that it can be applied irrespective of the regressors' order of integration, I(0) or I(1), thus allowing for statistical inferences on long-run estimates, which are not possible under alternative cointegration techniques. However, the linear ARDL cointegration technique is not valid in the presence of I(2) variables.

The general form of the ARDL model (Pesaran and Shin, 1998) is defined as:

$$\Phi(L)y_t = \alpha_0 + \alpha_1 w_t + \beta^{'}(L)x_{it} + u_t , \quad (1)$$

where: $\Phi(L) = 1 - \sum_{i-1}^{\infty} \Phi_i L^i$ and $\beta(L) = \sum_{j=1}^{\infty} \beta_j L^j$, with (L) being the lag operator and (w_t) being a vector of deterministic variables such as the intercept, seasonal dummies, time trends or other exogenous variables (with fixed lags).

The recently developed asymmetric ARDL model, applied in this paper, is a new technique for detecting non-linearities focusing on the long and short-run asymmetries among economic variables. The technique was advanced by Shin et al. (2011) and is an asymmetric expansion of the above mentioned linear ARDL model.

Following Pesaran and Shin (1998), Pesaran et al. (2001), Schorderet (2003) and Shin et al. (2011), we consider the following nonlinear asymmetric cointegrating regression:

$$y_t = \beta^+ x_t^+ + \beta^- x_t^- + u_t , \qquad (2)$$

where β^+ and β^- are the associated long-run parameters and x_t is a k×1 vector of regressors decomposed as:

$$x_t = x_0 + x_t^+ + x_t^- , \qquad (3)$$

where, x_t^+ and x_t^- are partial sum processes of positive and negative changes in x_t :

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \max(\Delta x_j , 0) ,\ x_t^- = \sum_{j=1}^{t} \Delta x_i^- = \sum_{j=1}^{t} \min(\Delta x_j , 0) , \qquad (4)$$

By associating (2) to the ARDL(p, q) case, we obtain the following asymmetric error correction model (AECM)[1]:

$$\Delta y_t = \rho y_{t-1} + \theta^+ x_{t-1}^+ + \theta^- x_{t-1}^- + \sum_{j=1}^{p-1} \varphi_j \Delta y_{t-j} + \sum_{j=0}^{q} (\pi_j^+ \Delta x_{t-j}^+ + \pi_j^- \Delta x_{t-j}^-) + e_t \ \text{for}\ j = 1,...,q \qquad (5)$$

where $\theta^+ = -\rho\beta^+$ and $\theta^- = -\rho\beta^-$.

This empirical analysis follows four steps; namely, step one concerns the estimation of model (5) which can be estimated by standard OLS. Step two is the establishment of the long-run relationship between the levels of the variables y_t , x_t^+ , x_t^- , by means of a modified F-test, while using the bounds-testing procedure advanced by Pesaran et al. (2001) and Shin et al. (2011), which refers to the joint null, $\rho = \theta^+ = \theta^- = 0$ in (5). In step three, using the Wald test, we examine for: long-run symmetry where, $\theta = \theta^+ = \theta^-$, and short-run symmetry which can take one of the following forms (i) $\pi_i^+ = \pi_i^-$ for all $i = 1,...,q$ or (ii) $\sum_{i=0}^{q} \pi_i^+ = \sum_{i=0}^{q} \pi_i^-$. Finally, in step four we use the asymmetric ARDL model (5) to derive the asymmetric cumulative dynamic multiplier effects of a unit change in x_t^+ and x_t^- , respectively, on y_t :

$$m_h^+ = \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial x_t^+} , m_h^- = \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial x_t^-} , h = 0,1,2... , \qquad (6)$$

Note that as $h \to \infty$, then $m_h^+ \to \beta^+$ and $m_h^- \to \beta^-$, where β^+ and β^- are the asymmetric long-run coefficients calculated as $\beta^+ = -\theta^+ / \rho$ and $\beta^- = -\theta^- / \rho$ respectively.

[1] For a more extensive derivation of the model see Shin et al. (2011)

4. Model and empirical results

The data used in the empirical analysis are monthly, covering the period from January 1999 to May 2011, and are collected from the databases of the Hellenic Statistical Authority of Greece (ELSTAT) and the International Monetary Fund (IMF). More specifically, the variables employed are the following: the Greek housing price index (HPI)[2], which is a mixed index composed from weighted partial indices for rentals, materials, services, water supply and fuel, as a proxy of house prices[3]; the consumer price index (CPI) and the industrial production index (IP) as a proxy of GDP. All variables are expressed in logarithmic form.

The analysis is performed on the following general empirical model:

$$lnHPI_t = f(lnCPI_t^+, lnCPI_t^-, lnIP_t^+, lnIP_t^-) \qquad (7)$$

where, $lnCPI^+$, $lnCPI^-$, $lnIP^+$and $lnIP^-$ are partial sums of positive and negative changes in lnCPI and lnIP respectively.

With respect to the application of the ARDL method although it can be applied irrespective of the regressors' order of integration, it is necessary to initially test the integration properties of the involved variable to ensure that the series used are not I(2). In such a case the computed F-statistics turn invalid (Ouattara 2004). Consequently, we apply the Augmented Dickey-Fuller (1979) unit root test. The findings presented in Table 1, suggest that all examined variables are nonstationary in levels while, they turn stationary in first differences and thus we can proceed with testing for cointegration in the ARDL framework.

[Table 1 here]

In the first step, we test for cointegration, using the unrestricted error correction model derived from specification of the form (1).

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[2] The use of the housing price index (HPI) which is provided by (ELSTAT) on a monthly basis and not house prices (stock variable) may raise some comments, since it is a component of CPI. However, the data set provided by the Bank of Greece is limited covering the period from the first quarter of 1997 to the second quarter of 2011. Such a small data set limits significantly the power of the error correction especially when the sample size is smaller than 100 observations (Shin et al. 2011).

[3] The HPI and house prices variables are found highly correlated (r=0.89). Besides, we found evidence in favor of the existence of a very strong long-run relationship between them. Actually, the results, based on the Johansen estimation technique provided evidence of cointegration with the estimated vector being [-1, 0.99083]. Further, we imposed the restriction [-1, 1] and the LR χ^2 value was calculated 0.6216 with a p-value 0.43 indicating acceptance of the null hypothesis.

$$\Delta lnHPI_t = cons + \sum_{i=1}^{p} b_i \Delta lnHPI_{t-i} + \sum_{i=0}^{q} c_i \Delta lnCPI_{t-i} + \sum_{i=0}^{q} d_i \Delta lnIP_{t-i} + \delta_1 lnHPI_{t-1} + \delta_2 lnCPI_{t-1} + \delta_3 lnIP_{t-1} + e_t \quad (8)$$

The optimal lag structure of the unrestricted error correction model is chosen based on the Akaike Information Criterion. The estimates, presented in Table 2, provide evidence in favour of the non-rejection of the null hypothesis of no cointegration (F-value=1.769 and smaller than the lower bound critical value). A possible reason for the non-detection of a causal long-run relationship might be the existence of nonlinearities among the variables.

[Table 2 here]

To test for this possibility, we proceed with the estimation of the nonlinear error correction model[4] of the form (5) as below.

$$\Delta lnHPI_t = cons + \rho lnHPI_{t-1} + \theta_1^+ lnCPI_{t-1}^+ + \theta_1^- lnCPI_{t-1}^- + \theta_2^+ lnIP_{t-1}^+ + \theta_2^- lnIP_{t-1}^- + \sum_{i=1}^{p-1} \varphi_i \Delta lnHPI_{t-i} + \sum_{i=0}^{q} \pi_{1,i}^+ \Delta lnCPI_{t-i}^+ + \sum_{i=0}^{q} \pi_{1,i}^- \Delta lnCPI_{t-i}^- + \sum_{i=0}^{q} \pi_{2,i}^+ \Delta lnIP_{t-i}^+ + \sum_{i=0}^{q} \pi_{2,i}^- \Delta lnIP_{t-i}^- + e_t \quad (9)$$

The cointegration test applied on the unrestricted model is an F-test on the joint hypothesis that the coefficients of the lagged level variables are jointly equal to zero. The results reveal statistically significant evidence in favour of the existence of a long-run cointegrating relationship between the examined variables (Table 2). Indeed, the F-statistic for the joint significance of the parameters of the lagged level variables is found 4.572 and exceeds the upper bound critical value[5].

The estimates of the unrestricted asymmetric ARDL regression of the form (5) are presented in Table 3. It should be noted that in order to select the final ARDL specification, we followed the general-to-specific approach. The preferred specification, is chosen by starting with max p = max q = 12 and dropping all insignificant stationary regressors. The inclusion of insignificant lags, in practice, is likely to lead to inaccuracies in the estimation and may introduce noise into the dynamic multipliers.

[Table 3 here]

[4] In the context of the ARDL methodology, the appropriate modification of the orders of the ARDL(p,q) model is sufficient to simultaneously correct for the residual serial correlation and the problem of endogenous regressors (Pesaran and Shin, 1999). Regarding the asymmetric ARDL model, if the decomposed series x_t^+ and x_t^- are not I(1) then the degree to which any endogeneity is corrected will depend on the degree of persistence - e.g. if they are I(d) then the correction is better for values of d closer to 1.

[5] We adopted the conservative approach to the selection of critical values as recommended by Shin et al. (2011) and select k=3.

In order to verify the appropriateness of an asymmetric model, we applied the Wald test for both long- (W_{LR}) and short-run (W_{SR}) symmetry. Regarding the long-run time horizon, the results are reported in the lower panel of Table 3 and suggest the rejection of the null hypothesis of long-run symmetry between the positive and negative components of each one of the examined variables. More specifically, for the CPI components the Wald test is found 7.476 (p-value=0.006), while for the IP components is found 14.293 (p-value=0.000). These findings further support that a linear model for the behaviour of house prices in Greece would be probably misspecified.

Before we examine the magnitude of these long-run asymmetric effects, we proceed with the analysis of the short-run dynamics. The null hypotheses of symmetry in the short-run impacts against the alternative of asymmetry are tested using the Wald statistic with null: $H_{SR,lnCPI}$: $\sum_{i=0}^{q}\pi_{1,i}^{+}=\sum_{i=0}^{q}\pi_{1,i}^{-}$ and $H_{SR,lnIPI}$: $\sum_{i=0}^{q}\pi_{2,i}^{+}=\sum_{i=0}^{q}\pi_{2,i}^{-}$, respectively. The results (Table 3, lower panel), suggest the rejection of the null hypothesis of a weak form summative symmetric adjustment (W_{SR}) for all included variables. More specifically, for the CPI components the Wald test is found 4.675 (p-value=0.031), while for the IP components is found 9.759 (p-value=0.002).

Next, we turn to the analysis of the long-run dynamics presented in Table 3. Focusing on the estimated long-run coefficients of the asymmetric ARDL model, we note that for the consumer price index, significance is confirmed for both positive (L_{CPI}-) and negative (L_{CPI}-) long-run coefficients, with the signs being positive and in line with the reported literature. The estimated long-run coefficients on $lnCPI^{+}$ and $lnCPI^{-}$ are 1.20 and 3.38 respectively. Therefore, we may conclude that a 1% increase in the consumer price index results in a 1.20% rise in house prices. Similarly, a 1% decrease in the consumer price index leads to a 3.38% decrease in house prices. Hence, our results indicate that the greater effect is sourcing from the negative changes.

Regarding the industrial production, a statistically significant long-run impact is detected only from the positive component (L_{IP}+). Analytically, the long-run coefficient on $lnIP^{+}$ is 0.23 indicating that, a positive change in the industrial production of 1%, results in an increase of 0.23% in house prices. The size of the positive long-run coefficient is rather small, raising some doubts regarding its accuracy. In contrast, the response of house prices to a negative change in the industrial production is distinctly not significant (L_{IP}- =0.0033) and is statistically insignificant at the 5% level of significance.

The analysis of the dynamic effects between the involved variables can be further enriched by studying the dynamic multipliers, considering the fully asymmetric case of equation (5). Figure 1, plots the dynamic effects of positive and negative changes in consumer prices where we observe that house prices respond more rapidly to increases in the consumer price index than to decreases, achieving equilibrium nearly after a 12-month time horizon. The response to decreases in CPI is obviously more gradual with equilibrium correction achieved in the beginning of the 2nd year. More particularly, in the first 6-month period the absolute effect of an increase in CPI is larger than that of a decrease while a decrease results in a greater in absolute terms effect than that of an increase only after the 12th month.

[Figure 1 here]

The overall impression is that when inflation rises, the housing market in Greece reacts immediately and more strongly in the short-run. However, the gap in magnitude between positive and negative shocks in the CPI diminishes very rapidly and disappears after nearly 6-months. Thereafter, the gap turns opposite revealing that, when considering a longer time horizon, the effect of downwards changes in consumer prices significantly dominates upwards changes.

These findings and in particular the behaviour of house prices in the case of a decreasing CPI, are indicative of a relative short-run rigidity in the Greek housing market. On the other hand, in cases of inflationary pressures, house prices are pushed upwards, further confirming that house ownership is always an important alternative for Greek investors.

Regarding the dynamic impacts of output changes on house prices, the study of the dynamic multipliers presented in Figure 2, reveals that it is the positive changes mainly that cause house prices to respond. More particularly, the housing market responds more rapidly and significantly stronger in positive changes in the economic activity, as proxied by the industrial production index, with full adjustment to equilibrium occurring around the middle of the 2nd year.

[Figure 2 here]

As it concerns the negative changes, we notice a very short-living reaction of house prices at the end of the first year while thereafter the magnitude diminishes rapidly being statistically zero towards the end of the second year. Simply put, the results suggest that the Greek housing market is more sensitive and reacts faster in periods of increasing economic activity than in recessions.

Comparing our results with the findings of other previous studies on Greek house prices, we are in line with Apergis and Rezitis (2003), confirming a positive explanatory power of the consumer price index. Such finding is justified in Kearl (1979) who argued that inflation increases housing prices and may eventually reduce housing demand. However, our results indicate that the greater effect on house prices sources from the negative changes.

Regarding the industrial production, our findings reveal the existence of a positive and significant effect on house prices in the long-run, sourcing from positive only changes in industrial production. However, our findings are in contrast to those reported in Merikas et al. (2009), who found evidence of a negative association between house prices and GDP in the Greek case. Overall, there is clear evidence of asymmetric adjustment in the housing market in output shocks with positive changes prevailing.

5. Summary and conclusions

The present paper investigated the dynamics between house prices and selected macroeconomic fundamentals, using Greek monthly data. Our analysis contributes to the literature by using a non-linear cointegration methodology and more specifically the asymmetric Autoregressive Distributed Lag cointegration technique, which permits the exploration of possible asymmetric effects in both the long- and short-run time horizon. The employed data sample covers the period from January 1999 to May 2011 and includes the housing price index, the consumer price index and the industrial production index.

Our results indicated the presence of asymmetric long-run effects, from the consumer price index and the industrial production index, towards house prices. Regarding the short-run time horizon, we found statistically significant asymmetric effects, running from all the examined variables towards house prices. However, there seem to exist important differences, in the response of house prices to positive or negative changes of the explanatory variables.

Overall, we conclude that the imposition of a linear symmetric model could be misleading in the case of the Greek housing market. The use of the asymmetric ARDL model for house prices contributes to the understanding of the non-linear dynamics among house prices and specific macroeconomic fundamentals, thus leading to more efficient policymaking and forecasting.

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TABLES

Table 1: ADF unit root tests

	Include an intercept, but not a trend			Include an intercept and a trend		
Series in levels	Test Statistic	*k*	Critical Value	Test Statistic	*k*	Critical Value
lnHPI	0.263	2	-2.824	-2.899	1	-3.456
lnIP	1.481	10	-2.807	-0.206	9	-3.419
lnCPI	0.381	11	-2.841	-1.696	11	-3.391
Series in first difference	Test Statistic	*k*	Critical Value	Test Statistic	*k*	Critical Value
Δ*lnHPI*	-8.293	1	-2.882	-8.310	1	-3.402
Δ*lnIP*	-8.884	7	-2.881	-6.824	10	-3.386
Δ*lnCPI*	-7.106	10	-2.791	-7.090	10	-3.349

Notes: The optimal lag structure of the ADF test is chosen based on the Akaike Information Criterion, while k denotes lag order. The critical values are 95% simulated critical values using 135 obs. and 1000 replications.

Table 2: Bounds test for cointegration in the linear and the nonlinear specifications

Dependent variable ΔlnHPI	F-statistic	95% lower bound	95% upper bound	Outcome
Linear ARDL(3,6,10) model	$F_{PSS\text{-}Linear}=1.769$	3.856	4.923	No Cointegration
Asymmetric ARDL model*	$F_{PSS\text{-}Nonlinear}=4.572$	3.219	4.378	Cointegration

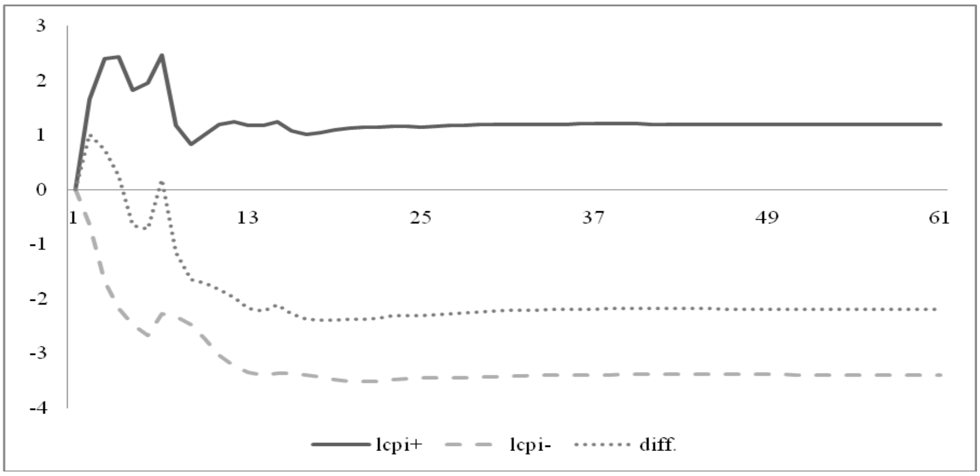
Notes: $F_{PSS\text{-}Linear}$ and $F_{PSS\text{-}Nonlinear}$ denote the PSS F-statistic testing the null hypothesis $\rho=\theta=0$ and $\rho=\theta^{+}=\theta^{-}=0$ respectively. * The exact specification of the asymmetric ARDL model is presented analytically in Table 3.

Table 3: Dynamic asymmetric estimation of house prices adjustments

Dependent Variable: ΔlnHPI			
Variable	Coefficient	Standard Error	T-Ratio[Prob]
Constant	0.5468	0.1223	4.468 [0.000]
lnHPI(-1)	-0.1337	0.0298	-4.4853 [0.000]
$\ln CPI^{+}(-1)$	0.1607	0.0838	1.9154 [0.058]
$\ln CPI^{-}(-1)$	0.4534	0.1835	2.4701 [.015]
$\ln IP^{+}(-1)$	0.0316	0.0124	2.5365 [0.013]
$\ln IP^{-}$ (-1)	0.000449	0.0107	0.0415 [0.967]
ΔlnHPI(-1)	0.2658	0.072	3.6892 [0.000]
ΔlnHPI(-8)	0.1372	0.0757	1.8124 [0.072]
$\Delta\ln CPI^{+}$	1.6558	0.2314	7.1545 [0.000]
$\Delta\ln CPI^{+}$ (-1)	0.3663	0.159	2.3030 [0.023]
$\Delta\ln CPI^{+}$ (-3)	-0.4572	0.2131	-2.1454 [0.034]
$\Delta\ln CPI^{+}$ (-4)	0.378	0.1593	2.3725 [0.019]
$\Delta\ln CPI^{+}$ (-5)	0.5844	0.1443	4.0501 [0.000]
$\Delta\ln CPI^{+}$ (-6)	-1.267	0.2353	-5.3844 [0.000]
$\Delta\ln CPI^{+}$ (-9)	0.6405	0.2205	2.9042 [0.004]
$\Delta\ln CPI^{-}$	0.4959	0.2214	2.2392 [0.027]
$\Delta\ln CPI^{-}$ (-4)	-0.542	0.2334	-2.3224 [0.022]
$\Delta\ln IP^{+}$	0.0449	0.0227	1.9766 [0.050]
$\Delta\ln IP^{+}$ (-2)	-0.04019	0.0238	-1.6837 [0.095]
$\Delta\ln IP^{+}$ (-4)	0.097	0.0194	4.9823 [0.000]
$\Delta\ln IP^{-}$ (-8)	0.03564	0.0196	1.8178 [0.072]
$L_{\ln CPI}^{+}$	1.2011 *	$L_{\ln IP}^{+}$	0.2364 *
$L_{\ln CPI}^{-}$	3.3895 *	$L_{\ln IP}^{-}$	0.0033
R^2	0.5141	R-bar^2	0.4318
X^2_{SC}	13.273 [0.349]	X^2_{FF}	3.142 [0.076]
X^2_{NORM}	33.093 [0.000]	X^2_{HET}	7.052 [0.008]
$W_{LR,\ \ln CPI}$	7.476 [0.006]	$W_{LR,\ \ln IP}$	14.293 [0.000]
$W_{SR,\ \ln CPI}$	4.675 [0.031]	$W_{SR,\ \ln IP}$	9.759 [0.002]

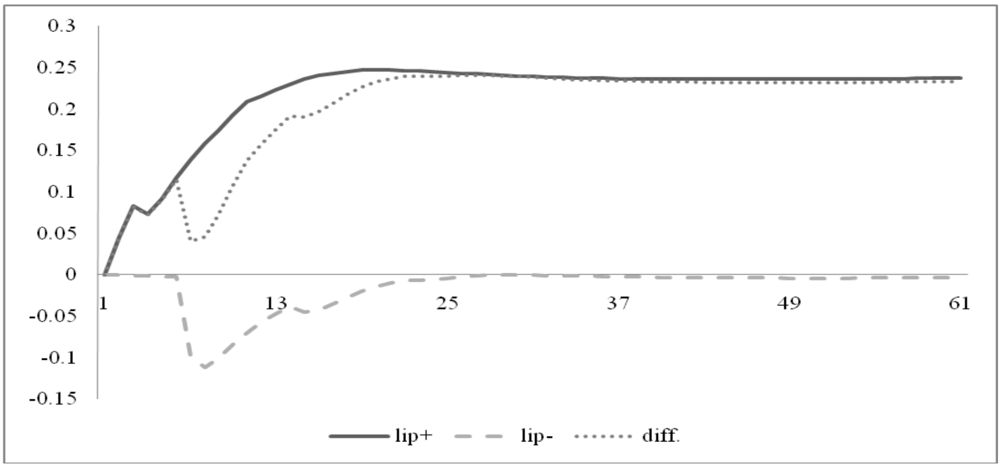
Notes: The superscripts "+" and "-" denote positive and negative partial sums, respectively. L^{+} and L^{-} are the estimated long-run coefficients associated with positive and negative changes, respectively, defined by $\hat{\beta} = -\hat{\theta}/\hat{\rho}$. X^2_{SC}, X^2_{NORM}, X^2_{FF}, X^2_{HET} denote LM tests for serial correlation, normality, functional form and heteroscedasticity, respectively. $W_{LR,\ \ln CPI}$ and $W_{LR,\ \ln IP}$ refer to the Wald test for the null of long-run symmetry defined by $-\hat{\theta}_1^{+}/\hat{\rho} = -\hat{\theta}_1^{-}/\hat{\rho}$ and $-\hat{\theta}_2^{+}/\hat{\rho} = -\hat{\theta}_2^{-}/\hat{\rho}$, respectively. $W_{SR,\ \ln CPI}$ and $W_{SR,\ \ln IP}$ refer to the Wald test for the null of the additive short-run symmetry condition defined by $\sum_{i=0}^{q} \pi_{1,i}^{+} = \sum_{i=0}^{q} \pi_{1,i}^{-}$ and $\sum_{i=0}^{q} \pi_{2,i}^{+} = \sum_{i=0}^{q} \pi_{2,i}^{-}$, respectively. * denotes the 5% significance level.

FIGURES



LR and SR asymmetry

Figure 1: Greek house prices-consumer price index dynamic multipliers



LR and SR asymmetry

Figure 2: Greek house prices-industrial production index dynamic multipliers