Review of Economics & Finance
Submitted on 15/Aug./2013
Article ID: 1923-7529-2013-04-19-14

The Housing Markets in Spain and Portugal: Evidence of Persistence

Carlos P. Barros
Instituto Superior de Economia e Gestão; Technical University of Lisbon
Portugal and UECE (Research Unit on Complexity and Economics)
E-mail: cbarros@iseg.utl.pt

Luis A. Gil-Alana (Correspondence author)
Faculty of Economics, University of Navarra
E-31080 Pamplona, Pamplona, SPAIN
E-mail: alana@unav.es

Abstract: This paper analyses house prices in Spain and Portugal over the last twenty years. In recent decades, housing prices have increased dramatically in the two countries; However, in 2007 the US subprime mortgage crisis broke out, which had worldwide influences, including Spain and Portugal. The purpose of this research is to study the time series persistence and the potential presence of breaks in these two countries. From this viewpoint, it is interesting to see how housing policy makers design reforms to adjust the real estate market, an issue of real concern for both citizens and institutions.

JEL Classifications: C22, C52; R29

Keywords: Housing prices; Spain; Portugal; Persistence

1. Introduction

Housing prices in Spain and Portugal have registered one of the highest cumulative growth rates within the OECD (Organization for Economic Cooperation and Development) countries, since the mid 1990s until 2007. This behaviour, common to other countries as well, is the reason why national house prices have constituted a key area of research in recent years. The issue of the subprime mortgage crisis in the United States paralised all the transactions in the housing market influencing, therefore, the supply and demand for houses which brought down the price of property assets. The crisis soon spread across the world and put the world economy into difficulties, which resulted in the global intervention of governments. Since then, the control of housing prices is considered of great importance in the management of the economic crisis (Himmelberg, Mayer and Sinai, 2005).

Spain, as a member of the OECD which has witnessed (excluding Germany and Japan) the highest increase in housing prices, forms the core study in this research. So, the contribution of this paper to the existing literature relies mainly on two key factors. First, due to the important role that housing wealth has on consumption, changes in house prices and in the degree of persistence have important implications both for economic activity and policy (see Chen, Kuan and Lin, 2007). Secondly, house intermediaries rely on the price series to manage their activity (see García and Raya, 2011), therefore it is very important to investigate the statistical properties of housing prices

data. Through the study of the stability of the series we can reach a conclusive view of the degree of persistence, since the stability tests serve as robustness tests on the persistence analysis. Housing prices in Portugal in 2009 also fell relative to 2008 reflecting the impact of the subprime mortgage crisis and accompanying Spain in this fall. However the Portuguese house market was less damaged than Spain because the country was not in the middle of such a severe housing market crisis.

The focus of this paper is the analysis of the degree of persistence of house prices in Spain and Portugal, as it reflects the stability of the macroeconomic variables in the countries under study, (Holmes and Grimes, 2008). Are house prices in Spain and Portugal persistent (MacDonald and Taylor, 1993; Malpezzi, 1999), and if they are, do they present breaks and when have these breaks occurred? Some authors give evidence about the importance of persistence analysis, especially in terms of its direct impact on policy decisions, Alexander and Barrow (1994), Cook and Vougas (2009) and Gupta and Miller (2012a). In fact, when housing businesses have prior knowledge of the persistence behaviour of house sales, they can benefit from the positive effects, or avoid the disadvantages of a negative effect. The degree of persistence in a series characterises the different policy measures to be adopted.

The paper is organized as follows. Section 2 provides the background literature about persistence in housing prices. A description of the hypotheses to be tested is presented in Section 3. The data and the empirical results are given in Section 4, followed by the conclusions in Section 5.

2. Literature Review

There are several works dealing with the housing price indexes in different countries, e.g., De Vries, De Haan, van der Wal and Marien (2009) referring to the construction of a house price index in the Netherlands. The case of Spain is referred to by Ayuso and Restoy (2007), who analyse the fundamentals of Spanish housing prices, studying the relationship between prices and income with VAR and GMM methods. They conclude that the increase in the housing price/income ratio in the late nineties, can be seen as a return to a long-run equilibrium relationship. Later, the ratio increased above its long-run equilibrium level. Another contribution to the explanation of the changes in house prices in Spain by means of an econometric model is Rodriguez-Lopez and Fellinger (2007). These authors argue that the cost of housing and the increasing stock of houses are the main variables that affect the fluctuation of average housing prices by the square metre in Spain.

Englund, Quigley and Redfearn, (1998) analysed the evolution of the Swedish house prices. In contrast to other countries, there is a lot of literature on the British market, supporting the notion of a causal link from house prices in the South East of England to other regions. A range of studies have looked for regional-national house price convergence, Holmans (1990), MacDonald and Taylor (1993), Alexander and Barrow (1994), Drake (1995), Ashworth and Parker (1997), Meen (1996), Peterson *et al.* (2002), Holmes and Grimes (2008); etc.. They offer different conclusions in support of the ripple effect. Many of these studies employ Engle and Granger (1987) or Johansen (1988, 1991) likelihood ratio tests of cointegration. On this basis, MacDonald and Taylor (1993) and Alexander and Barrow (1994), investigate cointegrating relationships between the UK regional house prices, examining whether or not the UK prices are tied together in the long-run. They found a significant number of cointegrating relations, which provided evidence to support the interrelated housing market in the UK.

The existence of the long-run equilibrium relationship between the UK regional house prices is supported by authors such as Drake (1995), Meen (1996), Cook (2003, 2005), Cook and Thomas (2003), Holmes and Grimes (2008), among many others. In Cook (2005) the UK regional house price linkages are investigated from an aspect of asymmetric adjustment, and his findings revealed

there was an increase in the number of long-run relations. On the other hand, Holmes and Grimes (2008) employed a new test which combined the analysis of principal components with unit root testing. The conclusion in that work was that the UK prices were driven by a single common stochastic trend, which is regarded as evidence of strong convergence in the long-run.

Contrary to most of the above studies, we focus on univariate analysis and consider the cases of Spain and Portugal by applying unit root tests and, more generally, fractionally integrated techniques.

3. Hypotheses

In line with the aim of this study, i.e., to evaluate the degree of persistence and fractional integration in the housing prices of Spain and Portugal, we suppose that the series display long range dependence. The tests undertaken enable the definition of the following research assumptions or hypotheses:

***Hypothesis 1*:** The house prices of Spain and Portugal are persistent. Persistence is a measure of the extent to which short term shocks in current market conditions lead to permanent future changes. By a shock we mean an event which takes place at a particular point in the series, and it is not confined to the point at which it occurs. A shock is known to have a temporary or short term effect, if the series returns back to its original performance level after a number of periods. On the other hand, a shock is known to have a persistent or long term impact if its short run impact is carried over forward to set a new trend in performance (MacDonald and Taylor, 1993; Alexander and Barrow, 1994; Drake, 1995; Cook, 2003; Ashworth and Parker, 1997; Larraz-Iribas and Alfaro-Navarro, 2008; Chien, 2010). For instance, in the case of a unit root, shocks will be permanent and the series will be very persistent. On the other hand, if the series is I(0) or I(d) with $d < 1$, the series will be mean reverting and lower the value of d is, the lower the degree of persistence.

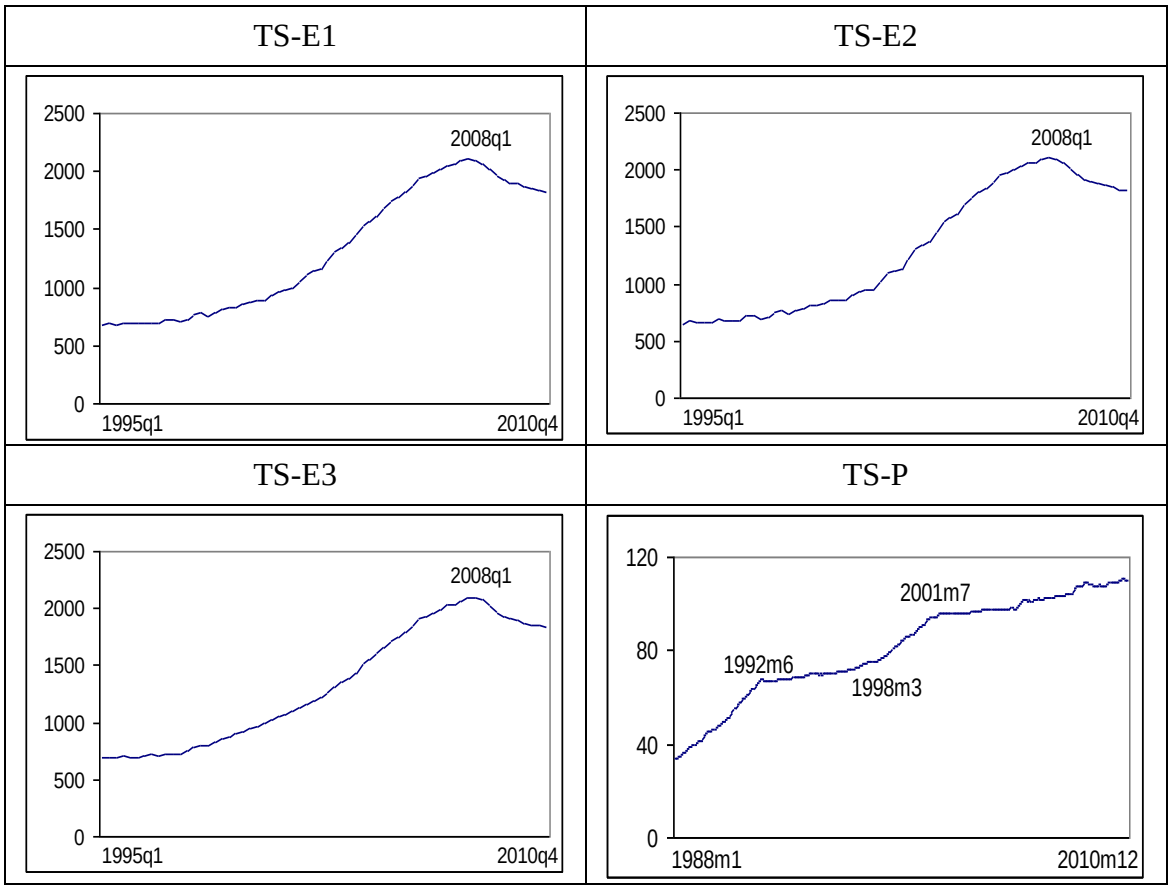
Hypothesis 2: The Spanish and Portuguese house prices present structural breaks. The existence of breaks is another important feature that may be present in house prices data, which may be attributed to fluctuation changes in the economic policy, and country-specific socio-economic events, among others. Indeed, if the house price series is stationary I(0), shocks will be transitory and following major structural breaks in the housing market, house prices will return to their original equilibrium with the disruptions in the housing price only having a temporary impact on economic activity. However, if the house price contains a unit root (i.e., if it is non-stationary I(1)), shocks due to breaks will have persistent effects on the market, thereby having a permanent impact on the economic activity (Coleman, LaCour-Little and Vandell, 2008; Chien, 2010; Gupta and Miller, 2012a,b).

These two hypotheses are tested below using fractional integration or I(d) methods, which are more general than the standard approaches based on I(0)/I(1) models.

4. Data and Empirical Results

Three series of housing price indices are available in the case of Spain. They are obtained from the Bank of Spain, quarterly, seasonally unadjusted, for the time period 1995Q1 – 2010Q4 and correspond to "*Residential property prices, all dwellings, per square meter*" (TS-E1); "*Residential property prices, existing dwellings, per square meter*" (TS-E2); and "*Residential property prices, new dwellings, per square meter*" (TS-E3), For Portugal, there is only one single series available, monthly, from 1988M1 to 2010M12, and referring to "*Residential property prices, all dwellings,*

per square meter" (TS-P). The data are collected by a private company named Confidencial Imobiliário Index (http://www.ci-iberica.com/?q=node/17). This series describes accurately the dynamics of the Portuguese house market and it is used by the Bank of Portugal to describe the house market in Portugal. The data are posted on the BIS - Bank of International Settlements website.



Note: The indices are measures at different units in the cases of Spain and Portugal.

Figure 1. Plots of time series

Plots of the four series are displayed in Figure 1. It is observed that the series grows throughout the sample period and the market reflects the crisis around the end of 2007.

We consider a linear regression model of form:

$$y_t = \beta^T z_t + x_t, \qquad t = 1, 2, \ldots, \tag{1}$$

where y_t is the time series we observe, β is a (kx1) vector of unknown coefficients and z_t is a set of deterministic terms that might include an intercept (i.e., $z_t = 1$), an intercept with a linear time trend ($z_t = (1, t)^T$), or any other type of deterministic processes such as dummy variables to examine the potential presence of breaks. The regression errors x_t above may be auto-correlated, and they can be I(0) (displaying, for instance, stationary ARMA structures) or I(1) in case of non-stationary unit roots. However, we will also allow for fractional degrees of differentiation. Thus, x_t in (1) can be described as

$$(1 - L)^d x_t = u_t, \qquad t = 0, \pm 1, \ldots, \tag{2}$$

where d can be any real value, L is the lag-operator ($Lx_t = x_{t-1}$) and u_t is an I(0) process, defined in this context as a covariance stationary process with spectral density that is positive and finite at any frequency. Thus, it includes weak autocorrelation of the ARMA-form, such that, if u_t is ARMA(p, q), x_t is then said to be a fractionally ARIMA, ARFIMA(p, d, q) process.

In the first part of this section we suppose that $z_t = (1,t)^T$ in (1), and we present the results for the three standard cases of no regressors, an intercept, and an intercept with a linear time trend, testing the null hypothesis $d = d_o$ in (2) for any real value d_o, We use here a parametric approach in the frequency domain developed by Robinson (1994). The functional form of this method can be found in any of the numerous empirical applications using this procedure (Gil-Alana and Robinson, 1997, Gil-Alana, 2000; Gil-Alana and Henry, 2003; etc.). The main advantages of Robinson (1994) compared with other methods are the following: first, it allows us to test any real value d, encompassing thus stationary ($d < 0.5$) and non-stationary ($d \geq 0.5$) hypotheses; second, the limit distribution is standard normal and this limiting behavior holds independently of the type of deterministic terms included in the model and of the way of modeling the I(0) error term u_t; finally, this method is the most efficient one in the Pitman sense against local departures from the null.[1]

In Table 1 we display the estimates of d (and the 95% confidence band using Robinson's (1994) approach) under the assumption that the error term u_t is white noise. As earlier mentioned we consider the three cases of: i) no regressors (i.e., $z_t = 0$ in (1)), ii) an intercept ($z_t = 1$), and iii) an intercept with a linear time trend ($z_t = (1, t)^T$). We first observe that if we do not include regressors, the estimated value of d is below 1 for the Spanish series, and it is slightly above 1 for the Portuguese one, however, the unit root null cannot be rejected in any of the four series. Including an intercept or an intercept with a linear trend, we observe that the unit root null cannot be rejected in case of TS-E1, and this hypothesis is decisively rejected in favour of $d > 1$ for the remaining three series (TS-E2, TS-E3 and TS-P)

Table 1. Estimated values of *d* under the assumption of white noise errors

Series	No regressors	An intercept	A linear time trend
TS-E1	0.964 (0.793, 1.213)	**0.988** **(0.819, 1.235)**	0.897 (0.719, 1.151)
TS-E2	0.988 (0.819, 1.235)	**1.481** **(1.363, 1.672)**	1.481 (1.364, 1.671)
TS-E3	0.897 (0.719, 1.151)	**1.554** **(1.455, 1.693)**	1.553 (1.455, 1.693)
TS-P	1.004 (0.936, 1.090)	1.284 (1.225, 1.360)	**1.258** **(1.199, 1.333)**

Notes: 1. Reported in bold, the selected models for each series.
2. In parenthesis, the 95% confidence band of the non-rejection values of *d*.

Table 2 displays the estimates of the coefficients in the selected models according to the specification of the deterministic terms. It is observed that the time trend is statistically insignificant in the three series corresponding to the Spanish housing prices with only the intercept being required. However, in the case of the Portuguese data, the time trend is statistically significant at

[1] Using other approaches in the time domain (Sowell, 1992) or in the frequency domain (Dahlhaus, 1989) lead essentially very similar results.

conventional statistical levels.[2] As earlier mentioned the unit root cannot be rejected in the case of the TS-E1 series being rejected in favour of d > 1 in the remaining three series.

Table 2. Estimated coefficients in the selected models

Series	*d*	An intercept	A linear time trend
TS-E1	0.988 (0.819, 1.235)	665.237 (35.716)	---
TS-E2	1.481 (1.363, 1.672)	634.203 (26.809)	---
TS-E3	1.554 (1.455, 1.693)	693.068 (42.371)	---
TS-P	1.258 (1.199, 1.333)	33.544 (88.281)	0.3028 (3.444)

Notes: In parenthesis in the second column, the 95% confidence band of the non-rejection values of *d*; In the third and fourth columns, t-values.

The results presented so far, however, might be biased due to the lack of autocorrelation for the error term. In what follows we suppose that u_t in (2) is weakly autocorrelated. However, instead of imposing an ARMA structure, we approximate that with the approach of Bloomfield (1973). This is a non-parametric model that produces autocorrelations decaying exponentially as in the AR(MA) case and that accommodates extremely well in the context of Robinson's (1994) tests.[3] The results based on this approach are given in Tables 3 and 4.

Table 3. Estimated values of *d* under the assumption of auto-correlated errors

Series	No regressors	An intercept	A linear time trend
TS-E1	0.741 (0.582, 1.221)	**1.388** **(1.173, 1.631)**	1.392 (1.172, 1.635)
TS-E2	0.771 (0.597, 1.210)	**1.284** **(1.063, 1.519)**	1.278 (1.044, 1.523)
TS-E3	0.695 (0.551, 1.223)	**1.791** **(1.579, 2.059)**	1.790 (1.579, 2.061)
TS-P	1.004 (0.900, 1.151)	1.352 (1.248, 1.484)	**1.312** **(1.212, 1.430)**

Notes: 1. Reported in bold, the selected models for each series.
2. In parenthesis the 95% confidence band of the non-rejection values of *d*.

The results here are very similar to the case of white noise disturbances. Thus, if we do not include regressors, the unit root null hypothesis cannot be rejected at the 5% level, however, including deterministic terms this hypothesis is rejected in favour of d > 1 now even for TS-E1. The estimated values of d based on the selected models are now 1.388, 1.284, 1.791 and 1.312 respectively for TS-E1, TS-E2, TS-E3 and TS-P. (See Table 2B).

[2] Note that the t-values are valid in the context of the model given by (1) and (2) noting that the estimates of the deterministic terms are based on the d_o-differenced processes, which are supposed to be I(0) under the null.

[3] See Gil-Alana (2004) for a paper dealing with the model of Bloomfield (1973) using Robinson's (1994) tests.

Table 4. Estimated coefficients in the selected models

Series	*d*	An intercept	A linear time trend
TS-E1	1.388 (1.173, 1.631)	665.826 (33.797)	---
TS-E2	1.284 (1.063, 1.519)	636.177 (25.540)	---
TS-E3	1.791 (1.579, 2.059)	693.897 (48.151)	---
TS-P	1.312 (1.212, 1.430)	33.540 (89.437)	0.3158 (2.784)

Notes: In parenthesis in the second column, the 95% confidence band of the non-rejection values of *d*; In the third and fourth columns, t-values.

Finally, and given the seasonal nature of the series examined (quarterly in the case of the Spanish data, and monthly for Portugal), we also investigate a model based on seasonal AR(1) disturbances.[4] The results are displayed in Tables 5 and 6.

Table 5. Estimated values of d under the assumption of seasonal errors

Series	No regressors	An intercept	A linear time trend
TS-E1	0.939 (0.713, 1.210)	**1.653** **(1.452, 1.995)**	1.652 (1.449, 1.998)
TS-E2	0.955 (0.731, 1.229)	**1.559** **(1.379, 1.858)**	1.554 (1.372, 1.862)
TS-E3	0.884 (0.659, 1.148)	**1.485** **(1.344, 1.654)**	1.485 (1.345, 1.653)
TS-P	0.999 (0.924, 1.089)	1.278 (1.215, 1.356)	**1.253** **(1.193, 1.331)**

Notes: 1. Reported in bold, the selected models for each series.
2. In parenthesis the 95% confidence band of the non-rejection values of *d*.

Table 6. Estimated coefficients in the selected models

Series	*d*	An intercept	A linear time trend
TS-E1	1.653 (1.452, 1.995)	664.886 (52.569)	---
TS-E2	1.559 (1.379, 1.858)	633.431 (42.864)	---
TS-E3	1.485 (1.344, 1.654)	692.716 (42.687)	---
TS-P	1.253 (1.193, 1.331)	33.544 (86.216)	0.3010 (3.437)

Notes: In parenthesis in the second column, the 95% confidence band of the non-rejection values of *d*; In the third and fourth columns, t-values.

[4] Here we use the model $u_t = \rho u_{t-4} + \varepsilon_t$ for the case of the Spanish series, and $u_t = \rho u_{t-12} + \varepsilon_t$ for the Portuguese series. Higher seasonal AR orders produced very similar results.

Once more the results are consistent with the previous cases though the estimated values of d are now even higher, and based on the selected specifications (with one intercept for Spain, and with a linear time trend for Portugal) the values of d are 1.653, 1.559 and 1.485 for the three Spanish series, and 1.253 for the Portuguese data. (Table 6). These results clearly indicate that, if we do not take structural breaks into account, the four series examined are non-stationary and even the first differenced series present a component of long memory behavior.

As mentioned above, the results presented so far do not allow the existence of breaks in the data. Next we take this possibility into account. First, we suppose that the break date is known. For the case of Spain, we consider a single break at 2008q1, and we allow for a change in the intercept and the slope of the time trend. The considered model is then:

$$y_t = \alpha_1 I(t \le T^*) + \alpha_2 I(t > T^*) + \beta_1 t I(t \le T^*) + \beta_2 t I((t - T^*) > T^*) + x_t ; \quad (1 - L)^d x_t = u_t, \tag{3}$$

with T^* = 2008q1. In the case of Portugal, three potential breaks are examined: 1992m6, 1998m3 and 2001m7, however, only the first break was found to be significant at conventional statistical levels. Therefore, we consider the same model as in (3) with T^* = 1992m6. The results based on this approach are displayed in Tables 7 - 9.

Table 7. Estimates in the context of a dummy variable for the break and white noise errors

Series	*d*	α_1	β_1	α_2	β_2
TS-E1	1.357 (1.249, 1.534)	654.260 (34.208)	22.299 (2.399)	1863.372 (3.804)	-31.265 (-2.228)
TS-E2	1.301 (1.199, 1.486)	622.714 (26.170)	24.797 (2.583)	1968.796 (3.889)	-31.089 (-2.063)
TS-E3	1.307 (1.217, 1.432)	679.372 (41.767)	22.650 (3.382)	1899.445 (5.378)	-30.492 (-2.914)
TS-P	1.179 (1.113, 1.264)	33.335 (88.413)	0.639 (6.608)	67.458 (12.899)	0.174 (2.757)

Notes: In parenthesis, in the second column, the 95% confidence interval for the values of *d*;
In the remaining columns, they are t-values.

Table 8. Estimates in the context of a Dummy variable for the break and autocorrelated (Bloomfield) errors

Series	*d*	α_1	β_1	α_2	β_2
TS-E1	1,182 (0.969, 1.420)	651.002 (33.204)	25.893 (4.981)	2048.801 (7.456)	-28.482 (-3.104)
TS-E2	1.092 (0.873, 1.326)	619.557 (25.480)	27.501 (5.912)	2104.307 (8.543)	-28.583 (-3.089)
TS-E3	1.462 (1.272, 1.756)	684.025 (44.095)	16.767 (1.921)	1591.630 (2.924)	-35.968 (-2.418)
TS-P	1.199 (1.091, 1.348)	33.341 (88.743)	0.638 (6.174)	67.406 (12.072)	0.171 (2.445)

Table 9. Estimates in the context of a Dummy variable for the break and seasonal AR(1) errors

Series	d	α_1	β_1	α_2	β_2
TS-E1	1.424 (1.273, 1.647)	655.249 (52.043)	20.369 (2.704)	1762.402 (4.444)	-32.826 (-2.989)
TS-E2	1.385 (1.232, 1.600)	623.170 (42.660)	23.145 (2.982)	1882.857 (4.605)	-32.655 (2-833)
TS-E3	1.234 (1.098, 1.389)	677.129 (42.934)	24.416 (4.847)	1990.857 (7.484)	-28.806 (-3.429)
TS-P	1.177 (1.110, 1.264)	33.335 (88.507)	0.639 (6.660)	67.449 (13.092)	0.174 (2.795)

We notice here first that all the deterministic terms result in significant values. If the disturbances are white noise (Table 7) or seasonal AR (Table 9) the estimated values of d are above 1 in all cases. If we suppose that u_t follows the exponential spectral model of Bloomfield (1973) the unit root cannot be rejected in the cases of TS-E1 and TS-E2 being rejected in favour of d > 1 in the other two series (Table 8). As expected, the time trend coefficient is negative after the break in the case of Spain, and positive though small in the second subsample in the case of Portugal.

Finally, we allow the break dates to be endogenously determined by the model. Here we employ a procedure developed by Gil-Alana (2008), which is based on minimizing the residuals sum squares across the different subsamples. We consider the following model,

$$y_t = \beta_i{}^T z_t + x_t; \quad (1-L)^{d_i} x_t = u_t, \quad t = 1,\ldots,T_b^i, \quad i = 1,\ldots,nb, \qquad (4)$$

where nb is the number of breaks (i.e., nb = 0, 1, 2, 3), y_t is the observed time series, the β_i's are the coefficients corresponding to the deterministic terms; the d_i's are the orders of integration for each subsample, and the T_b^i's correspond to the times of the unknown breaks. The method is based on minimizing the residuals sum squares for a grid of values of the fractional differencing parameters and the time breaks.[5]

The results using this approach indicate that there are no endogenous breaks in the case of the Spanish series and that there is at most one break in the Portuguese series, and it occurs at 1992m6 for all types of disturbances, probably related with the increase in oil prices due to the invasion of Kuwait by Iraq in 1990. The estimated coefficients at each subsample are displayed in Table 10. [6]

[5] The number of breaks is also determined by the model itself through the Schwarz (1978) criterion. See Gil-Alana (2008).

[6] The fact that no breaks are found in any of the series at the time of the subprime crisis may be related with the fact that the break may have taken place at a time that is very close to the end of the sample, implying then that there are very few observations in the second subsamples to justify long memory results.

Table 10. Estimates of the parameters in the context of a single break at an unknown period of time for the case of Portugal

u_t	Break date	1st sub-sample			2nd sub-sample		
		d_1	α_1	β_1	d_2	α_1	β_1
White noise	1992 m6	0.981 (0.834, 1.214)	33.250 (76.831)	0.6384 (11.55)	1.220 (1.148, 1.315)	67.228 (188.79)	0.1694 2.348)
Bloomfield	1992 m6	0.929 (0.667, 1.413)	33.210 (76.872)	0.6398 (13.97)	1.239 (1.123, 1.407)	67.2352 (189.54)	0.1657 (2.100)
Seasonal AR(1)	1992 m6	0.984 (0.833, 1.231)	33.251 (41.434)	0.6383 (6.18)	1.221 (1.147, 1.319)	67.2284 (186.547)	0.1692 (2.305)

Notes: In parenthesis in the 3rd and 6th columns, the 95% confidence band of the non-rejection values of *d*; In the remaining columns, t-values.

We notice here (in Table 10) that there is a substantial increase in the degree of integration after the break. Thus, for the first subsample, the estimated *d*'s are 0.982, 0.929 and 0.984 respectively for white noise, Bloomfield and seasonal AR(1) errors, and the unit root cannot be rejected in any of the three cases. However, for the subsample after the break, the estimated values of d are 1.220, 1.239 and 1.221 and the unit root is now decisively rejected in the three cases. That means that the degree of persistence has substantially increased in the housing prices in Portugal after 1992.

5. Discussion and Conclusion

This study examines the degree of persistence in Spain and Portugal housing prices by using fractional integration. We employ three housing prices series for Spain, quarterly, from 1995Q1 to 2010Q4, and one monthly series for Portugal (1988M1 – 2010M12). It is found in all the series that housing prices are better explained in terms of a long memory model that incorporates persistence and autoregressive components. If we do not include structural breaks, the results indicate strong degrees of persistence in all cases, with values of d above 1 in the majority of the cases. That means that according to this specification the first differenced processes still present a component of long memory behaviour. In fact, though not reported, performing standard unit root testing methods (Dickey and Fuller, ADF, 1979; Phillips and Perron, PP, 1988; Elliot et al., ERS, 1996; and Ng and Perron, 2001) the results supported the unit root hypothesis in practically all cases, which is not surprising given that these methods have very low power in the context of fractional models.[7] If we incorporate an exogenous structural break in the Spanish series at the time of the subprime crisis, the results still support the hypothesis of persistence, with orders of integration about 1. For the Portuguese data, an endogenous break is found around 1992m6, a significant increase being observed in the degree of persistence after the break.

[7] Diebold and Rudebusch (1991), Hassler and Wolters (1994) and Lee and Schmidt (1996) among many others found that standard unit root methods (ADF, PP, etc) have very low power in the context of fractional integration.

The results presented in this work strongly support hypothesis 1 indicating that the housing price series in Spain and Portugal are non-stationary and highly persistent, with orders of integration which are equal to or higher than 1. That means that shocks in the series will have permanent effects, and strong measures must be adopted to recover the original trends. Also, hypothesis 2 seems to be satisfied in the case of Portugal, finding evidence of a break at 1996 and a significant increase in the degree of persistence after such a break.

In summary, as house prices are driven by market demand and supply forces, they are persistent in Spain and Portugal, which is an endogenous characteristic of the housing prices in these two countries, explained by increase in costs and increasing demand. Moreover, it is found that this persistence is present when we take structural breaks into account. Furthermore, housing prices may behave differently during economic crisis and normal times based on the evidence of breaks, signifying that prices are persistent all the time, but during economic crises they break with a small immediate change in the growth trajectory, but later they continue to show persistent growth. Finally, although prices behave with specific breaks in these two neighbouring countries, they are similarly persistent, which results from market dynamics.

In terms of the literature and industry contribution of our research, this study first contributes to the literature by providing more accurate insights into the persistence of housing prices in the Iberian Peninsula, while most previous studies have ignored the persistence on the short and long term dependence of housing prices. Long memory models have also not been implemented previously on housing prices despite the fact that they include as particular cases the standard AR(I)MA models widely employed in the literature. This paper was also one of the first to adopt fractional integration, while most previous papers adopted a traditional integrated I(1) approach (Giussani and Hadjimatheou, 1991, 1992; MacDonald and Taylor, 1993; Alexander and Barrow, 1994; Meen, 1996; Ashworth and Parker, 1997; Cook, 2003; Drake, 1995; Larraz-Iribas and Alfaro-Navarro, 2008; Holmes and Grimes, 2008; Zohrabyan, Leatham and Bessler, 2008; Chien, 2010; Gupta and Miller, 2012a,b). Models based on fractional integration are more general than the classical models based on integer degrees of differentiation and thus allow for a much richer degree of flexibility in the dynamic specification of the series. Note that an added contribution of this paper is that it provides evidence from Spain and Portugal, while most previous studies have focused on one single country, usually the US and/or the UK.

The findings can also directly assist policy makers in the Spanish and Portuguese housing industry. In fact, when realtor authorities have a *priori* knowledge of the persistence on housing prices, they can design appropriate housing strategies to adjust persistence in house prices. Finally, we plan to consider in future papers the relationship between the house price series in the two countries through fractional cointegration techniques. Since Spain and Portugal share a common currency with free labour mobility, it could be of great interest to examine issues related with fractional cointegration, providing thus a contribution on integration in the Eurozone. As we provide evidence from Spain and Portugal, the results may offer generalised indications for other countries. More research is needed to confirm the present results.

Acknowledgments: Luis A. Gil-Alana acknowledges financial support from the Ministry of Education of Spain (ECO2011-2014 ECON Y FINANZAS, Spain).

References

[1] Alexander, C. and Barrow, M. (1994), "Seasonality and Cointegration of Regional House Prices in the UK", *Urban Studies*, 31(10): 1667-689.

[2] Ashworth J. and Parker, S.C. (1997), "Modeling Regional House Prices in the UK", *Scottish Journal of Political Economy*, 44(3): 225-246.

[3] Ayuso, J. and Restoy, F. (2007), "House prices and rents in Spain: does the discount factor matter?", *Journal of Housing*, 16(3-4): 291-308.

[4] Bloomfield, P. (1973), "An exponential model in the spectrum of a scalar time series", *Biometrika*, 60(2), 217-226.

[5] Chen, C.-L., Kuan, C.-M. and Lin, C.-C. (2007), "Saving and housing of Taiwanese households: New evidence from quantile regression analyses", *Journal of Housing Economics*, 16(2): 102-126.

[6] Chien, M. S. (2010), "Structural breaks and the convergence of regional house prices", *Journal of Real Estate Finance and Economics*, 40(1): 77 - 88.

[7] Coleman, M., Lacour-Little, M. and Vandell, K. (2008), "Subprime Lending and the Housing Bubble: Tail Wags Dog?", *Journal of Housing Economics*, 17(4): 272-290.

[8] Cook, S. (2003), "The Convergence of Regional House Prices in the UK", *Urban Studies*, 40(11): 2285-2294.

[9] Cook, S. (2005), "Detecting Long-Run Relationships in Regional House Prices in the UK", *International Review of Applied Economics*, 19(1): 107-118.

[10] Cook, S., and Thomas, C. (2003), "An Alternative Approach to Examining the Ripple Effect in the UK Housing Market", *Applied Economics Letters*, 10(13): 849-851.

[11] Cook, S., and Vougas, D. (2009), "Unit Root Testing against an ST-MTAR Alternative: Finite-Sample Properties and an Application to the UK Housing Market", *Applied Economics* 41(11): 1397-1404.

[12] Dahlhaus, R. (1989), "Efficient parameter estimation for self-similar process", *Annals of Statistics*, 17(4): 1749-1766.

[13] De Vries, P., De Haan J., Van Der Wal E. and G. Marien (2009), "A house price index based on the SPAR method", *Journal of Housing Economics*, 18(3): 214-223.

[14] Dickey, D. and W.A. Fuller (1979), "Distributions of the estimators for autoregressive time series with a unit root", *Journal of the American Statistical Association*, 74(336): 427-431.

[15] Diebold, F.S. and G. Rudebusch (1991), "On the power of Dickey-Fuller tests against fractional alternatives", *Economic Letters*, 35(2): 155-160.

[16] Drake, L. (1995), "Testing for Convergence between UK Regional House Prices", *Regional Studies*, 29(4): 357-366

[17] Elliot, G., Rothenberg, T. and J. Stock (1996), "Efficient tests for an autoregressive unit root", *Econometrica*, 64(4): 813-836.

[18] Engle, R.F. and C.W.J. Granger (1987), "Cointegration and error correction model, Representation, estimation and testing", *Econometrica*, 55(2): 251-276.

[19] Englund, P., J.M. Quigley and C.L. Redfearn (1998), "Improved price indexes for real estates: Measuring the course of Swedish housing prices", *Journal of Urban Economics*, 44(2): 171-196.

[20] Garcia, J. and Raya, J. M. (2011), "Price and Income Elasticities of Demand for Housing: Characteristics in the City of Barcelona", *Regional Studies*, 45(5): 597-608.

[21] Gil-Alana (2000), "Mean reversion in real exchange rates", *Economics Letters*, 69(3): 285-288.

[22] Gil-Alana, L.A. (2004), "The use of the model of Bloomfield (1973) as an approximation to ARMA processes in the context of fractional integration", *Mathematical and Computer Modelling*, 39(4-5): 429-436.

[23] Gil-Alana, L.A. (2008), "Fractional Integration and Structural Breaks at Unknown Periods of Time", *Journal of Time Series Analysis*, 29(1): 163-185.

[24] Gil-Alana, L.A. and B. Henry (2003), "Fractional integration and the dynamics of UK unemployment", *Oxford Bulletin of Economics and Statistics*, 65(2): 221-240.

[25] Gil-Alana, L.A. and Robinson, P.M. (1997), "Testing of unit roots and other non-stationary hypotheses in macroeconomic time series", *Journal of Econometrics*, 80(2): 241-268.

[26] Giussani, B. and Hadjimatheou, G. (1991), "Modeling Regional House Prices in the UK", *Papers in Regional Science*, 70(2): 201-219.

[27] Giussani, B. and Hadjimatheou, G. (1992), "House prices: an econometric model for UK", Netherland *Journal of Housing and the Built Environment*, 7(1): 31-58.

[28] Gupta, R., and Miller S. M. (2012a), "The Time-Series Properties on Housing Prices: A Case Study of the Southern California Market", *The Journal of Real Estate Finance and Economics*, 44(3): 339-361.

[29] Gupta, R. and S.M. Miller (2012b), "'Ripple Effects' and Forecasting Home Prices in Los Angeles, Las Vegas, and Phoenix", *Annals of Regional Science*, 48(3): 763-782.

[30] Hassler, U. and J. Wolters (1994), "On the power of unit root tests against fractional alternatives", *Economics Letters*, 45(1): 1-5.

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

[32] Holmans, A. (1990), "House price: changes through time at national and sub-national level", Working Paper 110, Government Economic Service Working Papers.

[33] Holmes, J.M. and A. Grimes (2008), "Is there long-run convergence among regional house prices in the UK?", *Urban Studies*, 45(8): 1531-1544.

[34] Johansen, S. (1988), "Statistical analysis of cointegrating vectors", *Journal of Economics Dynamics and Control*, 12(2-3): 231-254.

[35] Johansen, S. (1991), "Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models", *Econometrica*, 59(6): 1551-1580.

[36] Larraz-Iribis, B. and J.-L. Alfaro-Navarro (2008), "Asymmetric Behaviour of Spanish Regional House Prices", *International Advances in Economic Research*, 14(4): 407-421.

[37] Lee, D. and P. Schmidt (1996), "On the power of the KPSS test of stationarity against fractionally integrated alternatives", *Journal of Econometrics*, 73(1): 285-302.

[38] MacDonald R. and M.P. Taylor (1993), "Regional House Prices in Britain: Long-Run Relationships and Short-Run Dynamics", *Scottish Journal of Political Economy*, 40(1): 43-55.

[39] Malpezzi, S. (1999), "A Simple Error Correction Model of House Prices", *Journal of Housing Economics*, 8(1): 27-62.

[40] Meen, G. (1996), "Spatial aggregation, spatial dependence and predictability in the UK housing market", *Housing Studies*, 11(3): 345-372.

[41] Ng, S., and P. Perron (2001), "Lag length selection and the construction of unit root tests with good size and power", *Econometrica*, 69(6): 1519-1554.

[42] Peterson, W., Holly, S. and Gaudoin, P. (2002), "Further Work on an Economic Model of the Demand for Social Housing", Report to the Department for the Environment, Transport and the Regions, London.

[43] Phillips, P.C.B. and P. Perron (1988), "Testing for a unit root in time series regression", *Biometrika*, 75(2): 335-346.

[44] Robinson, P.M. (1994), "Efficient tests of nonstationary hypotheses", *Journal of the American Statistical Association*, 89(428): 1420-1437.

[45] Rodriguez-Lopez, J. and Fellinger J.E. (2007), "El mercado de la vivienda en España. Previsiones 2007-2009", *Universidad de Alcalá*. OMPV(In Spainish).

[46] Schwarz, G. (1978), "Estimating the dimension of a model", *Annals of Statistics*, 6(2): 461-464.

[47] Sowell, F. (1992), "Maximum likelihood estimation of stationary univariate fractionally integrated time series models", *Journal of Econometrics*, 53(1-3): 165-188.

[48] Zohrabyan, T., D. Leatham and D. Bessler (2008), "Cointegration analysis of regional house prices in US", Proceedings: 2007 Agricultural and Rural Finance Markets in Transition, Regional Research Committee NC-1014 St Louis, Missouri, Oct.4-5, 2007.