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Regional Housing Prices in the USA: An Empirical Investigation of Nonlinearity

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Abstract Existing literature on housing prices is predominantly in a linear framework, and an important question that has not been addressed is whether housing prices exhibit nonlinearity. We examine Smooth Transition Autoregressive (STAR) model based nonlinear properties of housing prices over the 1969–2004 period for the entire US and the four regions. Our main findings are (1) housing price for the entire US and all regions except for the Midwest show non-linearity, (2) the dynamic properties implied by the nonlinear estimation explain the typical patterns that have characterized each housing market, and (3) results of Granger causality tests look more plausible in the nonlinear framework where we find stronger evidence of Granger causality from housing price to employment and also from mortgage rates to housing price.

Keywords Housing market · STAR · Granger causality · Dynamic property

JEL Classification R10 · R21 · C12 · C13 · C32 · G10

Introduction

Housing is a substantial component of wealth for the typical household (Campbell and Cocco 2006; Poterba 1991). The influence of housing on the economy extends beyond its direct contribution because housing prices affect the level of consumer spending (Belsky and Prakken 2004). Most existing empirical work on modeling

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home prices is conducted in a linear framework. Nevertheless, it is clearly plausible that market behavior differs across expansion and contraction phases of the swings that characterize the real estate market. Therefore, it is important to first test if housing prices are nonlinear and then examine how they may be influenced by key variables in the nonlinear framework.

Within a linear framework the impact of economic fundamentals on housing prices is not entirely clear. Baffoe-Bonnie (1998) and Englund and Ioannides (1997) find a negative influence of mortgage rates on home prices using linear Granger Causality tests. However, McGibany and Nourzad (2004) find no such evidence even though they can establish a long-run cointegrating relationship between housing prices and real mortgage rates. When they take this cointegrating relationship into account in vector error correction models, they find that there is no Granger causality from mortgage rates to home prices.

Results from cointegration tests with panel data on metropolitan statistical area (MSA) level home prices and income from authors such as Gallin (2003) cast doubt on the very existence of a cointegration relationship between these variables. A “looser interpretation” of Gallin’s results is that even if long-run cointegration relations exist, they are impossible to verify and accurately estimate. He concludes by saying that this does not mean that fundamentals do not affect housing prices, and suggests that the level of housing prices does not appear to be tied with the level of fundamentals. If this is the case, the levels regressions found in the literature are most likely spurious, and the corresponding error correction models may be inappropriate.

The above literature indicates a lack of consensus on the appropriate framework in which housing prices are examined. An important fundamental question that has not been adequately addressed in the empirical literature on housing prices is whether home prices are nonlinear. There is plenty of theoretical discussion and empirical evidence on the testing for nonlinearity in other economic and financial variables. In the case of stock market returns, examples are, Sarantis (2001), Brock and Hommes (1998), Cecchetti et al. (1990), and Hsieh (1991). Evidence on the nonlinearity of macroeconomic variables such as GDP and unemployment rate using the Smooth Transition Autoregression (STAR) framework is presented by authors such as Skalin and Teräsvirta (1999, 2002) and McHugh et al. (2004).[1]

Although the nonlinear behavior in housing prices has been discussed and documented by authors such as Genesove and Mayer (2001) and Engelhardt (2001), formal empirical tests of nonlinearity have not been performed. Seslen (2004) argues that households exhibit rational responses to returns on the upside of the market but do not respond symmetrically to downturns.[2] On an upswing of the housing cycle, households exhibit forward looking behavior and are more likely to trade up, with equity constraints playing a minor role. On the other hand, households are less likely to trade when prices are on the decline causing stickiness on the downside of the housing market cycle. With developments in the field of behavioral economics and

[1] McHugh et al. (2004). “A Smooth-Transition Model of the Australian Unemployment Rate.” Retrieved from the world wide web at http://www.svt.ntnu.no/iso/WP/2002/10ausu.pdf.

[2] Seslen, T. (2004). “Housing Price Dynamics and Household Mobility Decisions.” Retrieved from the world wide web at http://www.usc.edu/schools/sppd/lusk/research/pdf/wp_2005.

finance, economists have explained the reluctance to move in down markets through loss aversion as opposed to equity constraints.[3] The loss itself rather than the tightening of equity constraints due to falling prices reduces mobility.

The presence of lumpy transactions costs in the housing market can result in important nonlinearities or threshold effects in the aggregate demand for housing. To quote Muellbauer and Murphy (1997, p 1721), "This arises from the extensive margin of housing demand: the greater is the appreciation of housing prices, actual and prospective, the more households are pulled over the transactions cost hurdle to engage in trade. At these times of heightened 'activity or frenzy', sharply increased demand feeds back into higher prices." Muellbauer and Murphy state that these spikes in the data can be successfully modeled with nonlinearity in the predicted rate of return. They also find evidence that sharp falls in the rate of return make households more cautious about entering the housing market.

In this paper we contribute to existing literature on housing prices in three ways. First, we empirically test for the possibility that housing price growth rates are non-linear using Smooth Transition Autoregression (STAR) model based tests. Second, we employ STAR estimations to specify dynamic properties of housing price growth rates. Third, we check for the presence of pairwise non-linear Granger causality between housing price growth rates and two key determinants of housing price: employment and mortgage rates.

Our focus is on housing markets at the regional level: Northeast, Midwest, West, and South. Since home prices are more responsive to regional economic and demographic shocks rather than to national shocks, a focus at the regional level enables us to compare the dynamics in housing markets across regions. From a policy point of view, developments in the housing markets have the strongest impact on the local economy.

A look at Fig. 1 below suggests that housing markets in the four regions of the US appear to have non-similar cycles. The ups and downs of housing price in the West and Northeast are more pronounced and have sharper swings than the rest of the US, especially the Midwest. In the case of the West and the Northeast, expansions are also noticeably steeper than contractions. For example, positive shocks to technology fueled an economic expansion in the Northeast from 1977–1988, referred to as the "Massachusetts Miracle" (Coulson and Steven 1995), leading to an annual appreciation rate of 11.2% over this period. This was followed by a fall in regional competitiveness (Grobar 1996) that led to contractionary phase when home prices fell from 1988 to 1995 at an annual percentage rate of 1.68%.[4] In the case of the 1985–1991 upswing in the West, nominal home prices increased at an annual rate of 5.96%, and then fell during the subsequent downturn at an annual rate of 1.41% until 1993. In contrast, the Midwest and to some extent the South show a steady increase in nominal home prices with the cycles not being very pronounced. This is not to deny the existence of swings in home prices in parts of these regions.

[3] For example, Tversky and Kahneman (1991).

[4] This steeper rise and milder fall also holds in real terms (Abraham and Hendershcott 1996; Angell and Williams 2005, please see footnote 5).

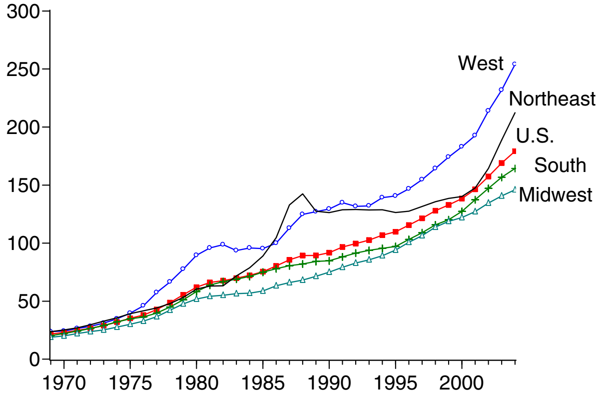

Fig. 1 Median Home Price in the USA and the Four Regions, 1969:1–2004:7

Angell and Williams (2005) identify pronounced cycles in the oil patch states, many of which fall in the South region as indicated in the Appendix of our paper.[5] Abraham and Hendershcott (1996) identify pronounced cycles in the Chicago MSA which falls in the Midwest region. Since our data is aggregated at the regional level, the pronounced cycles in individual MSAs do not reflect adequately in regional level fluctuations in the Midwest region and (to some extent) in the South as well. However, the West and Northeast are dominated by MSAs with wild swings in home prices that reflect at the regional level as well. Given the presence of pronounced swings in these markets, as opposed to the steady growth in the Midwest, the purpose of this paper is to empirically investigate if cyclical movements in housing prices exhibit asymmetric behavior.

The remainder of the paper is organized as follows. We first discuss the basic theoretical framework of non-linear STAR models. We then present the empirical results on the tests for linearity, the selection of the appropriate nonlinear STAR model, and the comparison of the results of the linear and nonlinear models. This is followed by an explanation of the nonlinear dynamic properties of housing price growth rates. Next, we perform bivariate nonlinear Granger causality tests between housing price and employment and between housing price and mortgage rate. We conclude and provide brief policy implications.

The Model

We focus on modeling home price growth rates as a non-linear and state-dependent variable, and we consider a model that allows regime switches to describe the dynamics of long-horizon housing price growth rates. Nonlinear models that allow for regime change can be the Threshold Autoregressive Model (TAR), developed by Tsay (1989), the Markov Switching Model developed by Hamilton (1989), or the smooth transition autoregressive (STAR) model developed by Luukkonen et al.

[5] Angell and Williams (2005). "Home Prices: Does Bust Always Follow Boom?" Retrieved from the world wide web at http://www.fdic.gov/bank/analytical/fyi/2005/050205fyi_table1.pdf.

(1988). While TAR and Markov switching models specify a sudden transition between regimes with a discrete jump, the dynamics of the STAR model allows a smooth transition between regimes. We employ the smooth transition autoregressive (STAR) model because we believe that home price growth rates are better characterized by the STAR model rather than by TAR or Markov regime switching models. The low speeds of transition obtained in the empirical results below validate the choice of the STAR model.

The main feature of the STAR model is to allow the dynamics of home price growth rates to evolve with a smooth transition between regimes that depends on the sign and magnitude of past realization of home price growth rates. For housing price growth rate, r_t, we specify the following STAR model of order p to capture the nonlinearities characterized by asymmetries in price growth dynamics.[6]

$$\begin{aligned} r_t &= \left[\phi_0 + \sum_{i=1}^{p} \phi_i r_{t-i}\right] + \left[\rho_0 + \sum_{i=1}^{p} \rho_i r_{t-i}\right] \cdot F(r_{t-d}) + \varepsilon_t \\ &= [\phi_0 + \phi(L) r_t] + [\rho_0 + \rho(L) r_t] \cdot F(r_{t-d}) + \varepsilon_t, \end{aligned} \quad (1)$$

where $F(r_{t-d})$, the transition function that controls the regime-shift mechanism, is a smooth and continuous function of past realized growth rates in housing price. In our STAR model, nonlinearities arise through conditioning the autoregressive coefficients $\rho(L)$ to change smoothly along with lagged housing price growth, so that past realized housing price growth rate r_{t-d} is the transition variable and d is the delay parameter which shows the number of periods that the transition variable leads the switch in dynamics.

STAR models could take the form of either the logistic smooth transition autoregressive (LSTAR) model or the exponential smooth transition autoregressive (ESTAR) model. In the LSTAR model the transition function, $F(\cdot)$ is given by the logistic function

$$F(r_{t-d}) = [1 + exp\{-\gamma(r_{t-d} - c)\}]^{-1} \quad (2)$$

In the ESTAR model, the transition function is given by the exponential function,

$$F(r_{t-d}) = 1 - exp\left\{-\gamma(r_{t-d} - c)^2\right\} \quad (3)$$

where $\gamma>0$ is the speed of transition between regimes and c indicates the halfway point between regimes. The adjustment parameter γ in both models governs the speed of transition between the two regimes, with greater values of γ implying faster transition between the regimes. In the limit, as the value of γ approaches infinity, the model degenerates into the conventional threshold autoregressive (TAR) model as in Tsay (1989). Alternatively, if γ approaches zero the model degenerates to the linear

[6] We use the growth rate of housing price so that r_t will be stationary.

AR model. For the choice between LSTAR and ESTAR, we follow the procedure suggested by Teräsvirta and Anderson (1992), which is described in "The Choice Between ESTAR vs LSTAR" below.

Empirical Results

In this section we first describe the data and the conclusions drawn from unit root tests for stationarity. Next, we perform the Lagrange Multiplier Smooth Transition (LM-STR) test for linearity of housing price growth rates. Then we conduct hypothesis tests to select the appropriate STAR model (LSTAR vs ESTAR). This section concludes with the estimation of ESTAR models and a discussion of the superiority of the performance of ESTAR models compared to linear AR models in terms of explaining house price movements in the case of the entire USA, the West, the Northeast, and the South.

Data

We use the nominal monthly seasonally adjusted median sales price of existing single-family homes over the period 1969:1–2004:7, published by the National Association of Realtors. This provides us with a reasonably large sample for a test of nonlinearity. We use the nominal home price in accordance with the findings of authors such as Genesove and Mayer (2001) and Engelhardt (2001) who show that sellers are averse to realizing nominal losses (not real losses) and therefore it is nominal house price changes that cause asymmetric effects on mobility and on the housing market.[7]

The employment data for each region is the sum of the monthly non-farm employment for each of the states that constitute that region (Appendix), and is obtained from the Bureau of Labor Statistics. The mortgage rate is the 30 year conventional mortgage rate (in nominal terms), and is obtained from Fannie Mae. The monthly data on economic time series are known to be noisy. Therefore we use the annual growth rate of housing price and employment both of which are measured as the growth rate with respect to the same month in the previous year. In the case of the mortgage rate variable, we use the annual first difference i.e., the difference with respect to the same month in previous year.[8] All variables are from 1969:1 to 2004:7, except the mortgage rate which begins in 1972:4. Results of the unit root test indicate that the null hypothesis of non-stationarity of growth rate of housing price, growth rate of employment, and first difference of mortgage rate is rejected in each case.[9]

[7] Engelhardt (2001) shows that households can lever nominal capital gains to purchase larger homes, but they became constrained by nominal capital losses.

[8] See Teräsvirta and Anderson (1992), Sarantis (2001) and the references therein for a similar transformation of various macro time series.

[9] Results of the unit root tests are not reported for the sake of brevity and may be obtained from the first author.

Lagrange Multiplier Smooth Transition (LM-STR) Test for Linearity

Following Teräsvirta and Anderson (1992), we estimate the following auxiliary regression

$$r_t = \phi_0 + \sum_{i=1}^{p} \phi_{1,i} \cdot r_{t-i} + \sum_{i=1}^{p} \phi_{2,i} \cdot r_{t-i} r_{t-d} + \sum_{i=1}^{p} \phi_{3,i} \cdot r_{t-i} r_{t-d}^2 + \sum_{i=1}^{p} \phi_{4,i} \cdot r_{t-i} r_{t-d}^3 + \varepsilon_t, \quad (4)$$

where the null hypothesis is $H_{01} : \phi_{2i} = \phi_{3i} = \phi_{4i} = 0$, for all i.

Accepting the null hypothesis implies that the appropriate model is a linear Autoregressive (AR) model against a non-linear STAR alternative.

The selection of the optimal lag, p, was made using the Akaike Information Criterion (AIC) over a range of lags from 1 through 20. The entire US, and each of the regions had an autoregressive lag of either 13 or 14. The delay lag (d) varies over the range 1≤d≤7. The estimate of the optimal delay lag (d) reported in Table 1 is chosen on the basis of the lowest p value (or highest F-statistic) associated with the test of the null hypothesis $H_{01} : \phi_{2i} = \phi_{3i} = \phi_{4i} = 0$, in Eq. 4 above. Table 1 indicates that in the case of the West and the Northeast we can reject the null hypothesis of linearity at the 5% level of significance or less. In the case of the South, the higher p value of 0.09 indicates a weaker rejection of the null of linearity. However, in the case of the Midwest the p value of 0.15 leads to a non-rejection of the null hypothesis of linearity at conventional levels of significance (5 or 10%). These results suggest the presence of linearity of housing prices in the Midwest and that there is a stronger case for nonlinearity in the West and Northeast than in the South.

The Choice Between ESTAR vs LSTAR

Given that linearity is rejected for all the regions except the Midwest, we now specify an appropriate STAR model (where the choice is between ESTAR and LSTAR models) to capture the nonlinear dynamics of regional housing markets. As suggested by Teräsvirta and Anderson (1992), the test of linearity as specified in Eq. 4 above can be used again to provide a sequence of nested hypothesis tests H_{04},

Table 1 LM-STR test for linearity: Housing Price Growth Rate

	Entire US p*=13	West p*=14	Northeast p*=13	Midwest p*=13	South p*=14
Optimal delay *d*	5 (0.0456)	2 (0.0520)	5 (0.0030)	3 (0.1515)	1 (0.0927)

The selection of the optimal lag, p^*, was made using the AIC statistic. The numbers in parentheses refer to the lowest p value of $H_{01} : \phi_{2i} = \phi_{3i} = \phi_{4i} = 0$ in Eq. 4 with the corresponding, d. For the Midwest region, the p values of H_{01} for each value of d where, $1 \leq d \leq 7$ are 0.2655, 0.3924, 0.1515, 0.5615, 0.4085, 0.2696, and 0.2954 respectively. Therefore the hypothesis of linearity is *not* rejected

H_{03}, and H_{02} for the choice between LSTAR and ESTAR alternatives. The sequence of nested tests for the coefficients in Eq. 4 above implies:

$$\begin{array}{llll} H_{04}: \phi_{4i} = 0, & & & i = 1, \ldots, p \\ H_{03}: \phi_{3i} = 0 & given & \phi_{4i} = 0, & i = 1, \ldots, p \\ H_{02}: \phi_{2i} = 0 & given & \phi_{3i} = \phi_{4i} = 0, & i = 1, \ldots, p \end{array} \tag{5}$$

We follow the standard procedure in the selection of the LSTAR vs ESTAR model as discussed in Teräsvirta and Anderson (1992), and implemented by several authors, including Sarantis (2001). Accordingly, we have three possible sequential outcomes, given the optimal delay lag (d) established in the previous test of linearity in Eq. 4, reported in Table 1.

First, rejection of $H_{04}: \phi_{4i} = 0$ in Eq. 5 implies selecting the LSTAR model. If, however, $H_{04}: \phi_{4i} = 0$ is not rejected, we move to the second part of the sequential test which tests if $H_{03}: \phi_{3i} = 0\, given\, \phi_{4i} = 0$. Rejection of H_{03} implies the selection of the ESTAR model. However, if H_{03} is not rejected, we move to the last part of the sequential test which tests: $H_{02}: \phi_{2i} = 0\, given\, \phi_{3i} = \phi_{4i} = 0$. Rejection of H_{02} implies the selection of the LSTAR model.[10] Table 2 reports the p values associated with the test of each hypothesis, H_{04}, H_{03}, and H_{02}. The fourth column which indicates the test of the null hypothesis H_{03} reports the lowest p values indicated with an asterisk. In choosing between the ESTAR and LSTAR models, we follow the practical recommendation of previous authors. According to Sarantis (2001, p. 461), "Granger and Teräsvirta (1993), Teräsvirta (1994), and Eitrheim and Teräsvirta (1996) argue that strict application of the sequential test may lead to wrong conclusions, because the higher order terms of the Taylor expansion used in deriving these tests are disregarded." These authors recommend that one should compute the p values of the F tests of each of the null hypothesis in Eq. 5 above and make the choice of the STAR model on the basis of the lowest p value. Accordingly, we choose an ESTAR model for all cases (the entire USA, the West, the Northeast, and the South) because the lowest p value occurs for the rejection of H_{03}. Since the LM-STR test of linearity (Table 1) does not reject linearity for the case of the Midwest, we treat home prices in the Midwest as linear.

Empirical Results of the STAR Models

In this section, we provide additional evidence of non-linearity by comparing the results of the estimation of the nonlinear ESTAR model in Eq. 6 below with the linear AR model in Eq. 7. Equation 6 is obtained by combining our basic nonlinear STAR model specification in Eq. 1 with the transition function of an ESTAR model given by Eq. 3.

$$r_t = \left[\phi_0 + \sum_{i=1}^{p} \phi_i r_{t-i}\right] + \left[\rho_0 + \sum_{i=1}^{p} \rho_i r_{t-i}\right] \cdot \left[1 - exp\left\{-\frac{\gamma}{\sigma^2(r_t)} \cdot (r_{t-d} - c)^2\right\}\right] + \varepsilon_t \tag{6}$$

[10] It should be noted that if $H_{02}: \phi_{2i} = 0$ *given* $\phi_{3i} = \phi_{4i} = 0$ is accepted then it implies that the hypothesis of linearity is accepted, implying that the linear AR model is appropriate.

Table 2 Specification of the Nonlinear Model for Housing Price Growth Rate

Region	Optimal delay lag d	$H_{04}: \phi_{4i}=0$, $i=1,\cdots,p$	$H_{03}: \phi_{3i}=0$ *given* $\phi_{4i}=0$	$H_{02}: \phi_{2i}=0$ *given* $\phi_{3i}=\phi_{4i}=0$	Selection of model
Entire USA	5	0.1278	0.0546*	0.3737	ESTAR
West	2	0.6068	0.0320*	0.1019	ESTAR
Northeast	5	0.0216	0.0210*	0.2166	ESTAR
South	1	0.3370	0.0814*	0.2443	ESTAR

The values for the nested tests H_{04}, H_{03}, H_{02} are probability p values. An asterisk indicates the lowest p value for the three tests

In Eq. 6 above we follow Teräsvirta (1994) and we standardize the exponent of the transition function $F(\cdot)$ to make the parameter γ scale-free by dividing the exponent of F by the variance, σ^2 (r_t), of r_t, the growth rate of housing price. If γ is statistically insignificant then Eq. 6 becomes the linear AR model of Eq. 7 below.

$$r_t = \left[\phi_0 + \sum_{i=1}^{p} \phi_i r_{t-i}\right] + \varepsilon_t. \tag{7}$$

Tables 3 and 4 report the estimated coefficients of the linear AR model and the nonlinear ESTAR model respectively, where the ESTAR estimation of Eq. 6 is conducted with nonlinear least-squares (NLS).

The mechanism that generates endogenous nonlinearity in the ESTAR model of Eq. 6 is the inclusion of the exponential transition function which conditions the autoregressive parameters, the ρ_{is}, to change smoothly along with past realized changes in growth rate of housing price. When either the adjustment parameter, γ, or the coefficients, ρ_i, are statistically insignificant in Eq. 6, we have the simple linear AR model of Eq. 7 which does not allow for the generation of nonlinearity.

Comparing the results of the two competing models across Tables 3 and 4, several features point to the superiority of the nonlinear estimation. First, the improvements in adjusted R^2 in the nonlinear ESTAR estimation compared to the AR estimation indicate that a substantial portion of variations in the growth of housing prices in the long-horizon is associated with nonlinear dynamics. Second, the measures of the standard error of the regression and the log likelihood value of the regression are both significantly improved in the nonlinear ESTAR estimations. Third, all the reported estimates ρ_i, i.e. coefficients of the nonlinear portion of Eq. 6 are statistically significant. Fourth, the value of γ, the parameter which indicates the speed of transition between regimes, is always positive as anticipated and is statistically significant at the 10% level (or less). The statistical significance of γ again confirms the presence of non-linearity as specified by Eq. 6 above. These results together provide strong support that the STAR models capture the inherent nonlinearity in long-horizon housing price growth rate dynamics in the case of the entire USA, and the regions of the West, Northeast, and South.[11] In these cases, a

[11] We perform Ljung Box and ARCH-LM tests to check for misspecification. These results, which can be available from the authors on request, indicate no evidence of misspecification.

Table 3 Estimation of the linear AR model in Eq. 7

Estimation
Entire USA: Adjusted R^2=0.7622, SER=1.6570, LLV=−831.31
$r_t = 0.3118(0.0451) + 0.7409\, r_{t-1}(0.0000) + 0.2117\, r_{t-2}(0.0000) - 0.4265\, r_{t-12}(0.0000) + 0.4261\, r_{t-13}(0.0000)$
West: Adjusted R^2=0.7365, SER=3.0183, LLV=−1115.50
$r_t = 0.3962(0.0523) + 0.6074\, r_{t-1}(0.0000) + 0.1828\, r_{t-2}(0.0032) + 0.1797\, r_{t-3}(0.0001) - 0.4289\, r_{t-12}(0.0000) + 0.2302\, r_{t-13}(0.0001) + 0.1812\, r_{t-14}(0.0006)$
Northeast: Adjusted R^2=0.6814, SER=3.8674, LLV=−1231.46
$r_t = 0.6031(0.0135) + 0.7012\, r_{t-1}(0.0000) + 0.2245\, r_{t-2}(0.0000) - 0.3876\, r_{t-12}(0.0000) + 0.3745\, r_{t-13}(0.0000)$
South: Adjusted R^2=0.6687, SER=2.2526, LLV=−979.37
$r_t = 0.4015(0.0455) + 0.5254\, r_{t-1}(0.0000) + 0.3504\, r_{t-2}(0.0000) - 0.3929\, r_{t-12}(0.0001) + 0.3422\, r_{t-13}(0.0000) + 0.1081\, r_{t-14}(0.0001)$
Midwest: Adjusted R^2=0.6676, SER=1.7787, LLV=−822.84
$r_t = 0.5367(0.0051) + 0.6523\, r_{t-1}(0.0000) + 0.1751\, r_{t-2}(0.0008) + 0.1523\, r_{t-5}(0.0001) - 0.4351\, r_{t-12}(0.0000) + 0.3692\, r_{t-13}(0.0000)$

The probability value (*p* value) is reported in parenthesis. In accordance with the practice followed by authors such as Teräsvirta (1994, p. 216) and Sarantis (2001, p 468) we include only significant lags
SER, standard error of regression; *LLV*, log likelihood value

linear model that ignores the nonlinear dynamics would be an incorrect specification because it does not allow the dynamics of home price growth rates to evolve with a smooth transition between regimes depending on the sign and magnitude of past realization of home price growth rates. Further evidence of nonlinearity of housing price growth rates, is provided by the results of the Ramsey model specification test (Table 5). The null hypothesis that the correct specification is a linear AR model (against a nonlinear ESTAR alternative) is rejected at the 1% significance level for all cases.

Two other features are worth noting from the estimation of the ESTAR models in Eq. 6, reported in Table 4. First, the parameter γ takes the values of 8.02 for the entire USA, 17.87 for the West, 17.98 for the Northeast, and 9.11 for the South. The higher γ for the West and Northeast regions indicates a sharper transition from one regime to another in these regions as compared to the South and the entire US. This is compatible with Fig. 1 where we see more pronounced fluctuations of home price in the West and Northeast regions. Also, the relatively small estimates of γ (17.98 and less) suggest a slower transition from one regime to the other, contrary to the TAR or Markov regime switching models, where γ is infinity and there is a sudden switch between regimes. This supports the choice of the STAR model rather than the TAR or Markov regime switching nonlinear models.

Table 4 Estimation of the ESTAR model in Eq. 6

Estimation

Entire USA (ESTAR): Adjusted R^2=0.8679, SER=1.3286, LLV=−697.45

$$\begin{aligned} r_t = {} & 2.0381(0.0055) + 0.8063\, r_{t-1}(0.0000) - 0.3579\, r_{t-3}(0.1112) - 1.4325\, r_{t-5}(0.0239) \\ & + 0.6383\, r_{t-6}(0.0015) - 0.4020\, r_{t-12}(0.0030) + 0.4236\, r_{t-13}(0.0000) + (-1.9038(0.0083) \\ & + 0.4534\, r_{t-3}(0.0472) + 1.4940\, r_{t-5}(0.0170) - 0.6241\, r_{t-6}(0.0040) + 0.1327\, r_{t-11}(0.0070)) \\ & \times \left[1 - \exp\left\{-8.2071(r_{t-d}(0.0103) - 0.6088(0.0186))^2\right\}\right] \end{aligned}$$

West (ESTAR): Adjusted R^2=0.8153, SER=2.8890, LLV=−1020.99

$$\begin{aligned} r_t = {} & 1.5378(0.1057) + 0.5751\, r_{t-1}(0.0000) - 0.8357\, r_{t-5}(0.0235) + 0.4531\, r_{t-6}(0.0453) \\ & +(-1.0892(0.2799) + 1.0367\, r_{t-5}(0.0057) - 0.4553\, r_{t-6}(0.0573)) \\ & \times \left[1 - \exp\left\{-17.8711(r_{t-d}(0.0992) - 0.3233(0.0779))^2\right\}\right] \end{aligned}$$

Northeast (ESTAR): Adjusted R^2=0.7882, SER=3.5117, LLV=−1096.16

$$\begin{aligned} r_t = {} & -0.2137(0.8043) + 0.7367\, r_{t-1}(0.0000) - 0.5406 r_{t-8}(0.0000) + 0.6746 r_{t-9}(0.0005) \\ & + 0.2359\, r_{t-11}(0.1204) - 0.4579\, r_{t-12}(0.0000) + 0.2945\, r_{t-13}(0.0014) + (1.2342(0.2426) \\ & + 0.2224\, r_{t-4}(0.0912) + 0.6829 r_{t-8}(0.0000) - 0.7012 r_{t-9}(0.0006)) \\ & \times \left[1 - \exp\left\{-17.9881(r_{t-d}(0.0318) + 1.4639(0.0000))^2\right\}\right] \end{aligned}$$

South (ESTAR): Adjusted R^2=0.7613, SER=1.9836, LLV=−859.76

$$\begin{aligned} r_t = {} & 1.5530(0.3673) + 0.6604\, r_{t-1}(0.0000) - 0.7677\, r_{t-5}(0.0993) + 5.9738\, r_{t-6}(0.1092) \\ & - 2.2288 r_{t-7}(0.0929) - 0.4243\, r_{t-12}(0.0000) + (-1.1372(0.5238) + 0.8702\, r_{t-5}(0.0714) \\ & - 5.9678\, r_{t-6}(0.1088) + 2.2817\, r_{t-7}(0.0854) + 0.4140\, r_{t-13}(0.0000)) \\ & \times \left[1 - \exp\left\{-9.1143(r_{t-d}(0.0096) + 2.9669(0.0000))^2\right\}\right] \end{aligned}$$

The probability value (p value) is reported in parenthesis.
SER, standard error of regression; *LLV*, log likelihood value

Second, the parameter c indicates the halfway point between the expansion and contraction phases of the housing markets. In all cases, the estimated value of c is statistically significant at 10% level. The estimated value of c is positive for the entire US and West markets, while it is negative for Northeast and South markets. This implies that for each market a different value of housing price growth rate shock triggers a shift in regimes.

Table 5 Ramsey Model Specification Test: Linear AR vs Nonlinear STAR Models

	Entire USA	West	Northeast	South
F statistics	55.0770	35.6669	41.7518	26.7674

Notes: The null hypothesis is that the model is linear. The F statistic is defined as $F = \frac{\left(R^2_{nonlinear} - R^2_{linear}\right)/m}{\left(1 - R^2_{nonlinear}\right)/(n-k)}$, where R^2_{linear} and $R^2_{nonlinear}$ denote R^2 s of linear AR and nonlinear STAR models respectively, m denotes the number of restrictions in the linear AR model, and k denotes the dnumber of parameters in the nonlinear STAR model

Dynamic Behavior

We investigate the dynamic behavior of the STAR models by examining the characteristic roots of the models derived from estimation of Eq. 6. The characteristic roots reported in Table 6 are computed from the following characteristic polynomial

$$\lambda^k - \sum_{j=1}^{k}\left(\phi_j + \rho_j F\right)\lambda^{k-j} = 0, \tag{8}$$

where the k^{th} order characteristic roots are denoted by a vector $\Lambda = (\lambda_1, \cdots, \lambda_k)$. We calculate the roots for the regime with F=0, which corresponds to the middle regime in the ESTAR model. Then, we calculate the roots for the regime with F=1, which describes the outer (either expansion or contraction) regime in the ESTAR model. First, we find that both regimes (outer and middle regimes in ESTAR model) include pairs of complex roots in all cases. This indicates that housing markets are characterized by cyclical movements during both the expansion and contraction phases, and the STAR models aptly describe asymmetric behavior in regime shifting

Table 6 Characteristic Roots and Modulus in Each Regime

Country	Regime	Most prominent roots	Modulus
Entire US (ESTAR)	Middle regime (F=0)	−0.9745±0.0980i	0.9794
		−0.2387±0.9362i	0.9661
		0.7798±0.6881i	1.0399
	Outer regime (F=1)	−0.6751±0.6728i	0.9531
		0.2498±0.9190i	0.9523
		0.9789	0.9789
West Region (ESTAR)	Middle regime (F=0)	−1.6131	1.6131
		−0.8465±0.4636i	0.9651
		0.4114±0.9751i	1.0583
		0.9281±0.2413i	0.9589
		1.0783±0.6689i	1.2689
	Outer regime (F=1)	−0.6708±0.6981i	0.9681
		0.9774	0.9774
Northeast Region (ESTAR)	Middle regime (F=0)	−0.9802	0.9802
		−0.8496±0.4424i	0.9578
		0.9863	0.9863
	Outer regime (F=1)	−0.9575	0.9575
		−0.6065±0.7652i	0.9764
		0.2722±0.9273i	0.9664
		0.7404±0.6439i	0.9812
		0.9628±0.2022i	0.9838
South Region (ESTAR)	Middle regime (F=0)	−1.1155	1.1155
		−0.3472±0.9293i	0.9920
		0.6411±0.7893i	1.0168
		1.2084	1.2084
	Outer regime (F=1)	0.9864	0.9864

Only roots with modulus ≥0.95 are reported and cyclical movements in housing markets of the entire US, West, Northeast, and South. Second, the middle regime in all cases except for the Northeast includes explosive roots, indicating that the growth rate of housing prices passes through the middle regime rapidly on its way up or down.[12] The middle regime in the case of the Northeast has roots with modulus values that are high, but less than 1. This reflects the fact that after a sharp drop (consistent with a high γ) from 1988 to 1990 in the Northeast, home prices remained fairly stagnant thereafter (i.e., in the middle regime), falling in nominal terms at an annual rate of 0.4% until 1995. Third, an important feature of the outer regime (which denotes either expansion or contraction) for all regions is that it does not include any explosive roots. This indicates that once housing markets are in the outer regime, they are more likely to stay there for prolonged periods.[13] For example, in the West home prices fell from 1982 to 1985 as a consequence of the national recession, followed by a prolonged increase until 1991 triggered by the boom in the defense industry. The cutback in defense spending caused a period of contraction in housing prices that lasted until the mid 1990s in many portions of the West. In the case of the Northeast, the surge in home prices lasted from 1977 to 1988, followed by a prolonged decline in home prices until 1995. This evidence supports our ESTAR-based dynamics which indicates that US housing markets stay for prolonged periods in the outer regime.

Housing prices have been increasing from the mid-1990s until 2006 in all regions of the US. The growth in home prices has been steady in the mid west, where our results indicate the presence of linearity. However, the growth has been dramatic in the West, Northeast, and some parts of the South, marking a prolonged and steep expansionary phase that ended by 2006 in all regions. Home prices have been falling since then and our study suggests that housing prices in these regions are now likely to reside in a contractionary phase for a prolonged period, without staying in the middle regime (or staying flat) for long. There is support for this analysis among housing analysts who project housing prices in the Midwest to be flat in the near term, but expect downturns in the markets that rose dramatically in recent years—the East Coast, the West Coast, and some portions of the South.[14]

Nonlinear Granger Causality

An important issue regarding regional housing markets is whether regional housing prices are related to the mortgage rate and to underlying cyclical movements in regional variables such as local employment. In this section we use pairwise non-linear Granger causality tests to test the causal relationship between (1) regional housing price growth rate and employment and (2) regional housing price growth rate and the mortgage rate. Skalin and Teräsvirta (1999) have developed Granger

[12] In the case of explosive roots, the modulus of the root is greater than 1.

[13] The outer regime in all cases includes roots with modulus values that are as low as 0.8, but are not reported in Table 7 for the sake of brevity.

[14] Housing at the Tipping Point (2006), Moody's Economy.com.

non-causality tests based on an additive STAR model. Skalin and Teräsvirta (1999) uses the following approximation:

$$r_t = \left[\phi_0 + \sum_{i=1}^{p} \phi_i r_{t-i}\right] + \left[\rho_0 + \sum_{i=1}^{p} \rho_i r_{t-i}\right] \cdot F(r_{t-d}) + \sum_{i=1}^{q} \delta_i y_{t-i}$$
$$+ \sum_{i=1}^{q}\sum_{j=1}^{q} \gamma_{ij} y_{t-i} y_{t-j} + \sum_{i=1}^{q} \psi_i y_{t-i}^3 + \varepsilon_t \qquad (9)$$

The null hypothesis that variable y does not Granger cause variable r is, H_0: $\delta_i' = \gamma_{ij}' = \psi_i' = 0$.[15]

Nonlinear Granger Causality Between Regional Housing Price and Employment

For purposes of comparison, the results for both linear Granger causality and STAR based nonlinear Granger causality tests are shown in Table 7 where the numbers in the cells represent the p values of the F test for the null hypothesis of Granger non-causality. Employment in all regions was found to be nonlinear.[16]

The nonlinear Granger causality tests in all cases—for Entire USA, West, Northeast, and South—indicate a strong rejection of the null of non Granger causality (i.e., indicate presence of significant Granger causality) from housing price to employment. A boom in the housing sector stimulates consumption and employment (Belsky and Prakken 2004) and has also been noted to lessen the severity of recessions by sustaining employment in the real estate sector.[17] This evidence of a short-run positive net effect of housing price on employment has stronger support in the non-linear framework as opposed to the linear framework, where Granger causality is absent in the case of the Northeast.[18]

In the case of testing for Granger causality from employment to housing price, in the Northeast, West, and to a smaller extent in the South, we find that housing prices are less dictated by employment in the nonlinear case.[19] This indicates that home prices in the Northeast and West and some portions of the South are more prone to bubbles, i.e. the momentum in home price is less related with fundamentals such as employment and more related with past values of housing prices itself. Abraham and

[15] Under the null, the test statistic has an F-distribution with degrees of freedom $q(q+1)/2 + 2q$ in the numerator and $T - n - q(q+1)/2 - 2q$ in the denominator, where T is the number of observatiuns and n is the dimension of the gradient vector, y_{t-i}, $y_{t-i}y_{t-j}$, and y_{t-i}^3, .

[16] The nonlinear estimation results on regional employment are available on request from the first author.

[17] We were not able to break down the testing of the Ganger causality results into sub periods of economic booms and recessions because the method of testing for nonlinearity calls for large data sets.

[18] When the variable y Granger causes variable r, the sum of significant coefficients (the θs below) of the "causing" variable (y) shows 'net effect of y on r'. In the following equation for the Granger causality test, the sum of θs represents the size of net effect from y to r. $r_t = \left[\phi_0 + \sum_{i=1}^{p} \phi_i r_{t-i}\right] + \left[\theta_0 + \sum_{i=1}^{q} \theta_i y_{t-i}\right] + \varepsilon_t$ Details of the Granger causality estimation can be obtained from the first author on request.

[19] This does not imply that employment is not an important determinant of home prices, but only that this relationship may not be evident in the short run.

Hendershott (1993) find price movements in the stable upper Midwest and Southeast are amenable to explanation within the basic Capozza-Helsley urban model, but they failed to explain the sharp, prolonged, cycles in the Northeast and West. They note that the West and Northeast show sustained serially correlated deviations in home prices. The results of the inability of employment to Granger cause regional home prices in the nonlinear case are in accordance with the findings of widespread discussion in the literature of the likely presence of several overvalued markets in the West and Northeast.

Nonlinear Granger Causality between Regional Housing Price and Mortgage Rate

The results for both linear Granger causality and STAR based nonlinear Granger causality tests between housing price and mortgage rate are shown in Table 7 for the period 1972:4 to 2004:6.[20] Since we do not see large differences in mortgage rates across regions we use only one national mortgage rate. In the case of tests of Granger causality from mortgage rate to home price, there appears a stronger rejection of the null hypothesis of non Granger causality (i.e. acceptance of Granger causality) in the nonlinear model as opposed to the linear model. Whereas the linear model suggests the lack of Granger causality from mortgage rate to home price in the West, the nonlinear model provides strong support for such short run effects of mortgage rate on home price in all cases. The negative impact of mortgage rates on home prices is to be expected, but is not always supported by empirical results of linear models. Our results suggest that a nonlinear framework is appropriate because it allows the impact of the mortgage rate to be asymmetric in housing market upswings versus downswings. For example, in the Northeast, home price increased by 125% as the mortgage rate fell from 16.63 to 10.33% in the 1981–1988 portion of the regional economic upswing. However the mortgage rate continued to fall from 1988 onwards, but this had little overall impact on home price (in the Northeast) that fell during the economic downswing until 1995. Similarly in the case of the West, the impact of falling mortgage rates was not large enough to cause a rise in home prices when the regional economy was in a slump from 1991 to 1993. Since 1995, the falling mortgage rate has had a strong, positive impact on home prices during the economic upswing caused by the “dot-com boom.” Home prices resided in the expansionary phase as mortgage rates continued to fall further and as the economy recovered from a brief recession in 2001.

Conclusions

Our paper is the first at empirically testing for the presence of nonlinearity in home price in the four regions of the USA, using the STAR framework. The baseline test of linearity is rejected in the case of the entire USA and the regions of the Northeast, West, and South, but is not rejected in the case of the Midwest. The strongest case for nonlinearity appears in the Northeast and the West. STAR model specification

[20] The mortgage rate data is available from April 1972 onwards only.

Table 7 Nonlinear (STAR-based) and Linear Granger Causality Tests

Region	Causing Variable: Housing Price Caused Variable: Employment	Causing Variable: Employment Caused Variable: Housing Price	Causing Variable: Mortgage Rate Caused Variable: Housing Price	Causing Variable: Housing Price Caused Variable: Mortgage Rate
Nonlinear Granger Causality Test				
Entire US	0.0013*	0.0455*	0.0000*	0.0027*
West	0.0233*	0.6949	0.0000*	–
North East	0.0001*	0.4106	0.0003*	–
South	0.0010*	0.1681	0.0000*	–
Linear Granger Causality Test				
Entire US	0.0152*	0.0174*	0.0000*	0.0006*
West	0.0006*	0.4500	0.1722	–
North East	0.1524	0.0765*	0.0181*	–
Midwest	0.6694	0.0293*	0.0000*	–
South	0.0028*	0.1656	0.0072*	–

Notes:
a. Housing price and employment are in annual growth rates. Mortgage rate is in annual first difference.
b. Numbers are probability p values of F test for the null hypothesis of Non Granger causality.
c. "*" implies significant at 10% level

tests indicate that the ESTAR model rather than the LSTAR model provides a better fit of the nonlinearity in the data.

The nonlinear ESTAR model outperforms the linear AR model on several fronts. The dynamic properties that are implied by the nonlinear estimation in each regional market match with the typical patterns of each housing market. The West and Northeast regions are characterized by a high speed of transition (γ) between regimes which explains the presence of more pronounced cycles (sharper swings) in these regions. In past booms in the West and Northeast, home price appreciation has been large and has lasted for several years. The sharp rise in home prices is followed by a prolonged period of falling prices, with the fall in price tending to be more gradual. A similar cycle occurs in the South, but since the South contains fewer bubbles prone areas, these swings appear smaller. In contrast, the Midwest, where we cannot reject the hypothesis of linearity, shows a steady upward climb in housing prices with very few and mild deviations from a linear trend.

When the asymmetry in the effect of mortgage rates—that mortgage rates have stronger impact on home prices when the housing market is in an upswing rather than in a downswing—is taken into account, we find strong support for Granger causality from mortgage rate to home price (in the nonlinear case). The nonlinear Granger causality tests also appear more plausible in the light of explanations of how the housing sector positively stimulates employment in all regions of the US.

Understanding the nonlinear nature of home prices alerts policy makers to the possibility that a contraction in home prices is likely to occur for a prolonged period. Since the downturn in housing prices has adverse effects on employment, falling home prices could be a blow to the local economy. While the STAR models we have fitted to the data provide several insights on the typical patterns of housing prices, future research needs to incorporate a more comprehensive set of factors that drive housing prices at the MSA level in the framework of multivariate STAR (nonlinear) Granger causality tests.

Appendix

Table 8 Definitions of Regions

Northeast	Midwest	South	West
Connecticut	Illinois	Alabama	Alaska
Maine	Indiana	Arkansas	Arizona
Massachusetts	Iowa	Delaware	California
New Hampshire	Kansas	District of Columbia	Colorado
New Jersey	Michigan	Florida	Hawaii
New York	Minnesota	Georgia	Idaho
Pennsylvania	Missouri	Kentucky	Montana
Rhode Island	Nebraska	Louisiana	Nevada
Vermont	North Dakota	Maryland	New Mexico
	Ohio	Mississippi	Oregon
	South Dakota	North Carolina	Utah
	Wisconsin	Oklahoma	Washington
		South Carolina	Wyoming
		Tennessee	
		Texas	
		Virginia	
		West Virginia	

Source: http://www.realtor.org

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