Identifying the Erratic Phase in the Behavior of the US Housing Prices

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Contents

I. Introduction

Recently, many industrialized economies have experienced significant and persistent increases in asset prices on a number of occasions over the past two decades, followed by sharp downward corrections. Such large swings in asset prices and housing prices in particular, have renewed a fundamental to central banks: whether the monetary authority should respond to asset prices in order to stabilize output and inflation variability

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or not.

For monetary authorities to answer the question above, it is important to understand the bi-directional relationship between monetary policy and asset prices. First, for central banks trying to stabilize the macro economy by affecting short-term interest rates, the causal effects of asset price movements on the short-run interest rates should be understood in advance (e.g., Rigobon and Sack (2003). Second, to the extent that the movements of asset prices, like stock returns or housing price inflation, have sizable effects on the macro economy, understanding how asset prices react to changes in monetary policy is another challenge to central banks (e.g., Rigobon and Sack (2004) and Ehrmann and Fratzcher (2004)).

This paper delves into the second relation posited above for the US economy. More specifically, we are interested in and delve into the reaction of housing price inflation to the interest rate policy of Federal Reserve Board. In so doing, we allow the possibility that the relation between housing price and key interest rates such as federal funds rate and Treasury Bill rate, like any other economic relations, is subject to sudden changes or abrupt shifts. There can be found a few pieces of work in the literature. Muellbauer and Murphy (1997) argue that there have been substantial shifts in housing price dynamics for the UK economy, and attribute such shifts to changing interactions between variables such as wealth, income, interest rates and housing prices.[1] Over the period spanning 1994 Jan to 2003 Dec, Davig and Gerlach (2006) identify a separate period from 1998 Sep to 2002 Sep, during which the volatility of the S&P 500 index responses to the action of the Fed is more than ten times greater than in the remaining periods. Others, Crawford and Fratantoni (2003) and Miles (2008) to name a few, investigate the

1) They also point out the possibility that large transaction costs associated with housing purchases are what causes the non-linearities in housing price adjustment.

forecasting performance of regime-switching models of the US housing price.

Motivated by those previous researches, we specify the evolution of the US housing price inflation as a two-state Markov switching model. Intuitively, the regime-switching model is very appealing for a real estate application, since real estate markets have historically been prone to booms and busts. The basic idea underlying this model is that housing markets perform differently in different economic environments and this change in behavior results in discrete changes in the time series properties of home price indices.

Using a two-state Markov switching model, we estimate the regime-specific behavior of the housing price inflation and the probabilities that the US economy is in respective regime. In terms of the estimated responses of the housing price inflation, one regime turns out to be 'atypical', in that the housing price inflation is positively related with real interest rate. Furthermore, compared with the other 'typical' regime in which housing price inflation responds negatively to real interest rate, any disturbance to the housing price inflation lingers for almost two years if the economy stays in the atypical regime. Also, data features in the atypical regime suggest that such a regime may be a mixture of the formation and collapsing of housing price bubbles.

To the extent that our results give a reasonable description of the US housing market, a daunting challenge to the Fed emerges: depending on the state of the economy, the appropriate policy responses of the Fed will be generally different. Therefore, another prerequisite of the Fed aiming at business cycle stabilization is to figure out what regime the economy is, and will be following its policy action to the current economic conditions.

The plan of the paper is as follows. Section II presents the regime switching model to use in the following analysis. In section III, we perform a preliminary investigation of data that motivates the use of regime switching approach. Section IV reports and interprets the results of applying the regime switching model to the US housing price. Section V briefly addresses the issue of optimal regime-specific monetary policy rules of the Fed. Section VI concludes.

Ⅱ. The Model

Aforementioned, we employ a Markov regime-switching model to allow for intrinsically different economic relations across the different regime of the economy.

The underlying intuition behind a regime-switching model is that the time series of interest behaves differently in different economic environments, characterized by a latent variable, denoting the state or regime of the economy. The characteristics in each regime are different, and there is a probability in each period that the economy will switch from one regime to the other. Potentially, we could use any model to describe the within-regime dynamics. However, the estimating parameters for each regime including the estimation of transition probabilities more than double the total number of parameters relative to single-regime models. In need of a parsimonious model, therefore, we consider the following equation for the housing price inflation

$$\pi_t^h = \gamma_{1,s_t}\pi_{t-1}^h + \gamma_{2,s_t}\pi_{t-2}^h + \gamma_{3,s_t}(y_{t-1} + y_{t-2}) + \gamma_{4,s_t}(i_{t-1} - \pi_{t-1}) + \gamma_{5,s_t}(e_{t-1} - e_{t-2}) + \varepsilon_t \quad (1)$$

, where π_t^h denotes the housing price inflation, y_t is the output gap (i.e., the excess of actual output over the potential output), i_t is the nominal interest rate, π_t is the rate of inflation, and e_t is the real exchange rate.[2] According to equation (1), the rate of housing price inflation depends on its two lags, the average of the output gap in the two recently previous periods, real interest rate, and the rate of appreciation of the dollar against the Euro.

In equation (1), we model the state variable $s_t = [1, 2]$ as a first order Markov−switching process. More specifically, the transition probabilities between the states $s_t = [1, 2]$ are specified as follows:

$$lP[s_t = 1 \mid s_{t-1} = 1] = p \ , \quad and \quad lP[s_t = 2 \mid s_{t-1} = 2] = q \qquad (2),$$

where the transition probabilities in equation (2) are assumed to be time−invariant.

Accordingly, the two sets $\Theta = (\Theta_1, \Theta_2)$ of regime−specific parameters are given by

$$\Theta_1 = (\gamma_{1,1}, \gamma_{2,1}, \cdots, \gamma_{5,1}, \sigma_1^2)$$
$$\Theta_1 = (\gamma_{1,2}, \gamma_{2,2}, \cdots, \gamma_{5,2}, \sigma_2^2),$$

where the indices 1 and 2 refer to different regimes, respectively. Note that we allow the volatilities of the disturbance ε_t to be regime−specific.

Statistical inferences in the context of the model (1)−(2) proceed via an iterative process of prediction-updating, described in Hamilton (1994) and Kim and Nelson (1999). This gives as a by-product the sample log-likelihood function, the maximization of yields in which the maximum likelihood estimates of the regime specific coefficients and variances are

2) This specification is reached after a series of preliminary ordinary least square regressions guided by the $\overline{R}^2$ criterion.

obtained. Furthermore, inferences about the unobserved regimes $\{s_t\}$ can be made, yielding the estimates of the probabilities that the economy is in each state in few of the whole data set.

Ⅲ. Data and Preliminary Results

1. Data

The data series used to estimate the model (1)−(2) span 1986:Q1 to 2008:Q2. The housing price series used is the quarterly house price index for the US, available from the Office of Federal Housing Enterprise Oversight. To circumvent the problem of seasonality, year-on-year rates of change are constructed for the housing price inflation. The nominal interest rate series is proxied by the federal funds rates and the 3 month Treasury Bill rates (both annualized), available from the FRB database. Output gap is constructed as the log-deviation of seasonally adjusted real GDP from the potential GDP, both available from the FRB database. The price inflation series is the year-on-year rate of change in the GDP deflator available from the FRB database. The real exchange rate series is constructed using the nominal exchange rate between the Euro and the US dollar and the GDP deflator in the two economies. More specifically, we construct the real exchange rate as $\frac{P_{US} \cdot NER}{P_{Euro}}$, where the nominal exchange rate (NER) corresponds to the units of Euro (or their equivalents for 1985:Q1−1998:Q4) per US dollar.[3)]

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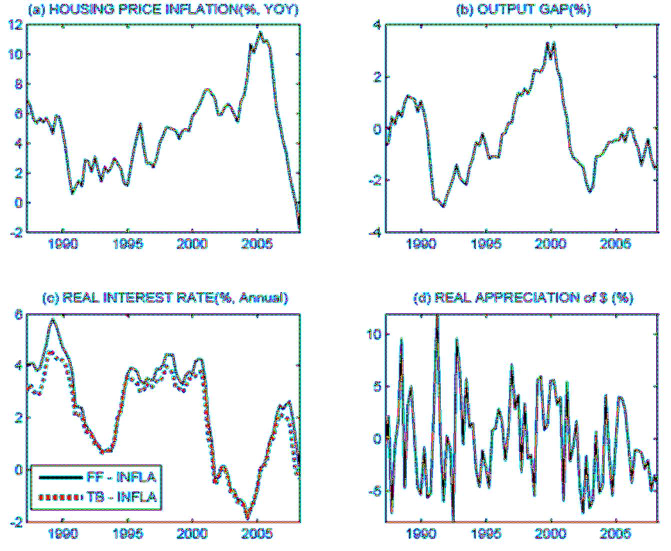
3) Therefore, higher real exchange rates thus constructed correspond to real appreciation of the US dollar.

[Figure 1]: Plots of Data Series



The plots of the data series on housing price inflation, output gap, real interest rate, and price inflation are shown in Figure 1. In the panel (a), the housing price inflation shows a prolonged upward movement over 1990-2005, followed by a sharp decline. Also, the period of upward trend has a few swings of smaller scale. Especially, the periods of 1995-1999 and 2000-2005 exhibit peak–trough features typical of business cycle fluctuations. The plots of output gap and real interest rate in panels (b) and (c) respectively show considerable degree of co–movement over the sample period, reflecting the action of the Federal Reserve Board for Business Cycle stabilization. Finally, the real appreciation of the US Dollar in panel (d) exhibits volatile fluctuations around a weakly downward trend, while the size of volatility has tended to decrease.

2. Preliminary Estimation

Before formally estimating the model (1)-(2), we perform a preliminary exercise to check for the existence of two distinct regimes. We consider a rolling regression, in which samples in four−year rolling windows are used to estimate the equation (1) via simple OLS (ordinary least square method). Here, we expect that the plots of rolling regression coefficients may exhibit abrupt shifts, provided that there are regime specific laws of evolution for the housing price inflation in response to other variables.

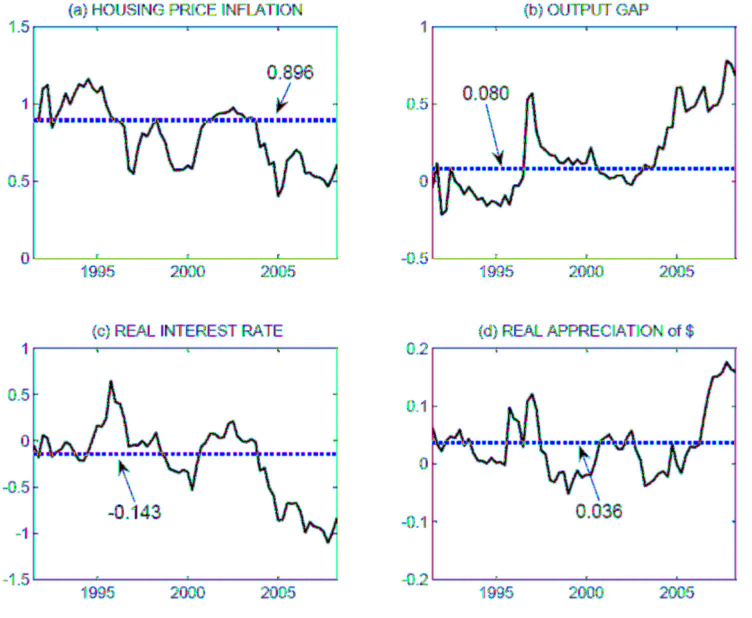
Figures 2A and 2B plot the coefficients from the rolling regression along with their counterparts from the full sample ordinary least square regression plotted as straight lines.[4] Figure 2A is for the specification in which the federal funds rates are used as the nominal interest rate, while Figure 2B corresponds to the case of using the 3 month Treasury Bill rates instead.

A few interesting features are found from the Figure 2A. First, there are episodes in which the autoregressive coefficients on the lagged housing inflation from the rolling regressions are considerably lower than the full sample estimates, roughly when the rolling windows span 1996−2001 and 2004-2008. The second feature is related to coefficients on real interest rate, which is traditionally expected to exert main effects on the housing price inflation. As evident in the panel (c), the coefficients from the rolling regressions exhibit frequent changes in sign, while its full sample estimate is small negative around −0.14. Our interpretation of this feature is that running a (single−regime) regression, without allowing for two distinct regimes, may lead to wrong policy prescriptions: when the coefficient on the real interest rate is lower than the single−regime estimate, the central

4) The coefficients on housing price inflation in the panal (a) are the sum of the two coefficients on π^h_{t-1} *and* π^h_{t-2}.

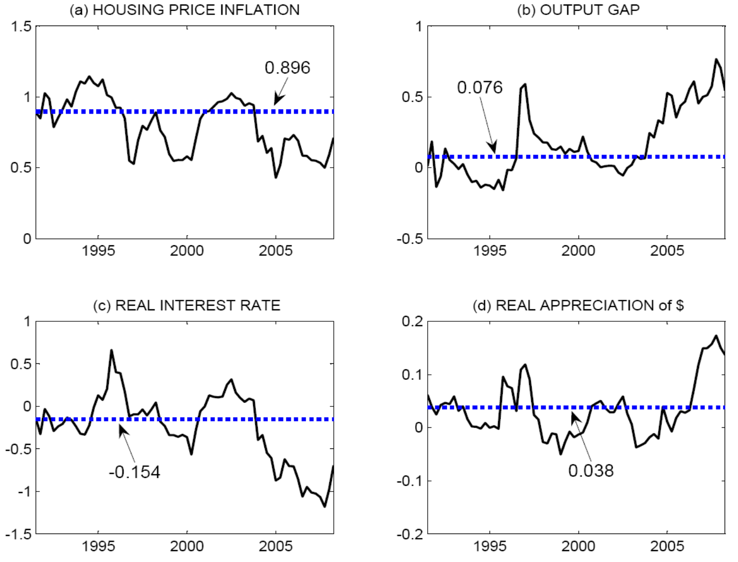
bank may underestimate the effect of the increases in the federal funds rate on the housing price inflation, leading to the possibility of a policy overkill. Furthermore, if the economy is currently in the atypical regime in which the housing price inflation is positively related to the real interest rate, the Fed unaware of the current regime may adjust the federal funds rate in the opposite direction to tame the hikes in the housing inflation.

[Figure 2A]: Coefficients from Rolling Regressions (Federal Funds Rate)



The coefficient plots in Figure 2B show similar features. The autoregressive coefficients on lagged housing price inflation exhibit period-dependent ups and downs, and the signs of the real interest rate change frequently.

[Figure 2B]: Coefficients from Rolling Regressions (Treasury Bill Rate)



In summary, the results of rolling regressions support the presence of two distinctive regimes for the evolution of the housing price inflation, especially with respect to its responses to the real interest rate. We will develop this idea more formally in the next section.

Ⅳ. Estimates from Regime-Switching Regression

Table 1 reports parameter estimates for the Markov-Switching model (1)-(2). Estimation procedure identifies two regimes. Table 1A is for the specification in which the federal funds rate is used as the nominal interest rate, and Table 1B corresponds to the case of using the 3 month Treasury Bill rate instead.

Table 1A: Estimation Results with the Federal Funds Rate

	Probabilities		Coefficients					σ^2
	P_{i1}	P_{i2}	γ_1	γ_2	γ_3	γ_4	γ_5	
Regime 1 $(i=1)$	0.490 (2.975)	0.510	1.665 (7.170)	−0.693 (−6.882)	0.201 (3.478)	0.249 (4.280)	0.056 (2.932)	0.003 (3.804)
Regime 2 $(i=2)$	0.111	0.889 (7.104)	1.126 (4.489)	−0.314 (−7.434)	0.133 (3.901)	−0.293 (−4.882)	0.024 (1.575)	0.007 (15.367)

According to the results in Table 1A, the model appears to be well identified: the estimated coefficients for each regime are sharply estimated, except that the response of the housing price inflation to the rate of real appreciation of the US dollar is not significant in the regime 2.[5)]

There are a few noticeable differences across the two regimes identified in Table 1A. In terms of the volatilities of the disturbance ε_t, the two regimes are quite different: the estimated standard deviation of ε_t for the regime 1 is less than 40% of that for the regime 2. The coefficient estimates are also different across the two regimes. First, the long-run AR coefficients on the lagged housing inflation are 0.972 and 0.812, for the regime 1 and regime 2, respectively. This means that, other things being constant, shocks of the same magnitude would have much longer effects on the housing price inflation in the regime 1. Second, the coefficient on the average of lagged output gaps in the regime 1 is 54% higher than that in the regime 2. Combined with the features of the long-run AR coefficient discussed above, the effect of one percent higher output gap in the current quarter on housing price inflation dies out to half after 24 quarters, if the economy stays in the regime 1. The half life reduces to merely 4 quarters if the economy remains in the regime 2.

5) We also check that the standardized residuals exhibit no signs of linear or non-linear dependence.

Third, the coefficients on the rate of real appreciation of the US dollar are clearly distinguished across the two regimes. In the regime 1, higher rate of appreciation leads to accelerations in the increase of housing price, while the effect in the other regime is not prominent in terms of either its size or significance. The most conspicuous difference between the two regimes, however, is captured by the estimated coefficients on the real interest rate. In the regime 2, housing price inflation decreases following the increase in the real interest rate, as suggested by standard asset pricing theory. To the contrary, the correlation between the housing price inflation and real interest rate is positive in the regime 1: unlike in the other regime, housing price inflation cannot be tamed by raising the federal funds rate if the economy is in the regime 1.

The results in Table 1B confirm that the above features, especially the sharp contrast in the coefficients on real interest rate across the two regimes, are to be found when the Treasury Bill rate is used as well.

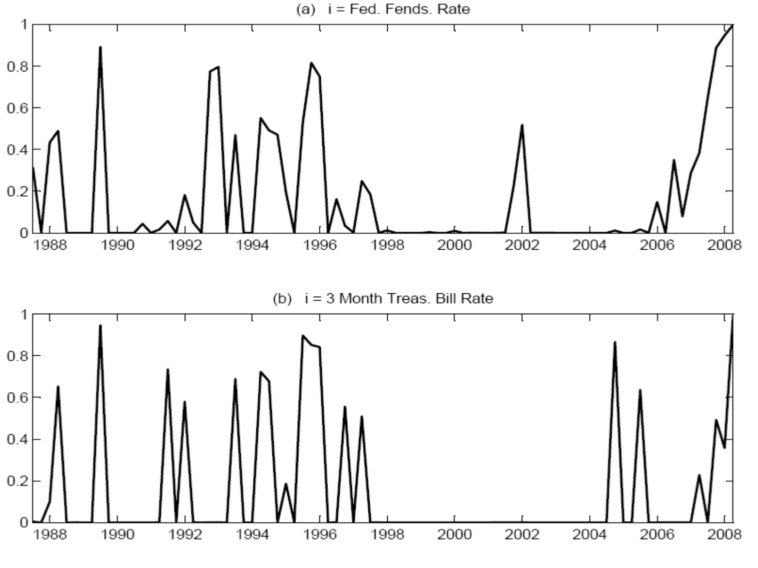
Table 1B: Estimation Results with the Treasury Bill Rate

	Probabilities		Coefficients					σ^2
	P_{i1}	P_{i2}	γ_1	γ_2	γ_3	γ_4	γ_5	
Regime 1 $(i=1)$	0.235 (2.380)	0.765	1.553 (4.200)	−0.603 (−7.537)	0.199 (14.186)	0.396 (12.277)	0.009 (0.713)	0.0007 (0.811)
Regime 2 $(i=2)$	0.137	0.863 (2.023)	1.226 (1.584)	−0.375 (−2.273)	0.120 (1.366)	−0.307 (−1.985)	0.040 (0.818)	0.007 (12.976)

We now turn to the identification of periods in each regime. To do so, we calculate the series of smoothed probabilities that the US economy is in regime 1, based on the estimation results of the Markov-switching model. Figure 3 plots the smoothed probabilities, where the upper panel is based on the estimation results in Table 1A and the lower panel concerns the estimation results with the Treasury Bill rate in Table 1B.

If we regard quarters with the smoothed probability of regime 1 higher than 50% as in the regime 1, the upper panel identifies {1989:Q3, 1992:Q4, 1993:Q1, 1994:Q2, 1995:Q3–Q4, 1996:Q1, 2002:Q1, 2007:Q3 –2008:Q2} as the regime 1 quarters. When the Treasury Bill rate is used as the nominal interest rate, the regime 1 probability is above 50% for {1988:Q2, 1989:Q3, 1991:Q3, 1992:Q1, 1993:Q3, 1994:Q2–Q3, 1995:Q3–Q4, 1996:Q1, 1996:Q4, 1997:Q2, 2004:Q4, 2005:Q3, 2008:Q2}, 15 quarters total. If we use the average of the two probabilities, the regime 1 quarters are {1988.Q2, 1989:Q3, 1993:Q2, 1994:Q2, 1994:Q5, 1995:Q3, 1995:Q4, 1996:Q1, 2007:Q4, 2008:Q1, 2008:2}. The proportion of the regime 1 quarters thus identified is 13.1% of the whole sample period, not much different from the estimates 17.9% and 15.1% of the steady state probability of regime 1, by the results in Table 1A and 1B, respectively.

[Figure 3]: Smoothed Probabilities of Regime 1



Having identified the periods in which the housing price inflation rates show the 'atypical' behavior, we now turn to the question of what economic conditions characterize such an atypical regime. To do so, we use the average of two smoothed probabilities in Figure 3 to identify periods in each regime, and compare the statistical properties of the data series. The results are reported in Table 2, where the regime-specific means of $\left(\pi_t^h, y_t, i_t - \pi_t, e_t - e_{t-1}\right)$ are reported.

Table 2: Comparison of the Two Regimes

	π_t^h	y_t	$FF - \pi_t$	$TB - \pi_t$	$e_t - e_{t-1}$
Regime 1	−0.0139	−0.0060	−0.0031	−0.0032	0.0047
Regime 2	0.0039	0.0022	0.0006	0.0008	−0.0043

Note: The regime−specific means are relative to the whole−sample mean.

The results in Table 2 suggest the following important characteristics: in terms of the regime-specific means, regime 1 is more likely when i) the housing price inflation and output gap are low, ii) real interest rates are low, and iii) the rates of real appreciation of Dollar are high, relative to the whole sample counterparts. We note that, among the three features above, the second and third one are usually considered to be what causes 'bubbles' in the housing market, while the first one is more akin to what happens when bubbles disappear or collapse. Our interpretation of the features in Table 2, therefore, is that the regime 1 identified by the two state Markov switching model is a mixture of the two atypical situations, i.e., the formation and collapsing of housing price bubbles.

V. Regime-Specific Optimal Responses of the Fed

In this section, we consider the optimal monetary policy rule of the Fed in each regime of the US economy identified in the previous section. To do so, we combine each version of the US housing price inflation equation in Table 1A with those describing the other sector of the US economy, and derive the optimal interest rate rule in each state for stabilizing the output and inflation variabilities with equal weights.[6] The results are given in Table 3.

Table 3: Optimal Responses of the Fed

	coefficients										
	π_t	π_{t-1}	π_{t-2}	y_t	y_{t-1}	i_{t-1}	i_{t-2}	π_t^h	π_{t-1}^h	e_t	e_{t-1}
Regime 1	−5.183	−0.129	0	−6.461	−1.750	0.129	0	−3.351	2.455	−0.024	0.132
Regime 2	14.735	−0.323	0	7.873	7.466	0.323	0	4.335	−2.932	−0.026	0.041

The upper panel of Table 3 shows the optimal response coefficients for the fixed regime 1, while those for the fixed regime 2 are in the lower panel. The comparison of response coefficients reveals an important difference across the two fixed regimes: the optimal policy under regime 2 requires the Fed to counteract the inflationary pressures from output, inflation, and housing price inflation, while the Fed is required to accommodate such inflationary pressure. For example, the long-run response coefficients of the Federal Funds rate with respect to (π, y, π^h) are all negative under the fixed regime 1. On the contrary, the corresponding long-run coefficients are all positive under the fixed regime 2.

6) Therefore, we assume that the US economy is in either regime forever. The details of the exercise are available from authors upon request.

We have raised the possibility that the regime 1 may correspond to the situation in which the housing price bubbles collapse. The coefficients on in Table 3 imply that the optimal action of the FRB in the bubble regime is **not** to prick the bubble. In this case, the FRB is required to keep the interest rate lower to forestall the harmful effects of the crash.

Ⅵ. Conclusion

The aim of this paper is to provide empirical evidence for the presence of two distinct regimes for the evolution of the US housing price inflation. We apply a two-state Markov switching model to estimate the regime-specific behavior of the housing price inflation and the probabilities that the US economy is in respective regime. One of the two regimes appears ‘typical’, in that the rate of housing price inflation negatively responds to higher real interest rate and that the effect of a shock to the housing inflation dies out in a reasonable period of time. The other regime, however, is ‘atypical’, in that the housing price inflation is positively related with real interest rate and that any disturbance to the housing price inflation lingers for almost two years. Data features of the unusual regime suggest that such a regime is a mixture of the formation and collapsing of housing price bubbles.

We also take an initial step toward the issue of optimal monetary policy in the presence of two distinctive regimes for the US housing price inflation. It is shown that, depending on which state the economy stays in, the appropriate policy responses of the central bank are qualitatively different. We plan to address the issue of optimal monetary policy further, especially taking into account the possibility that the US housing sector can alternate between the two regimes identified in the current paper.

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Abstract

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In this paper, we have provided empirical evidence on the presence of two distinct regimes for the evolution of the US housing price inflation. We apply a two-state Markov switching model to estimate the regime-specific behavior of the housing price inflation and the probabilities that the US economy is in respective regime. One of the two regimes appears 'typical', in that the rate of housing price inflation negatively responds to higher real interest rate and that the effect of a shock to the housing inflation dies out in a reasonable period of time. The other regime, however, is 'atypical', in that the housing price inflation is positively related with real interest rate and that any disturbance to the housing price inflation lingers for almost two years. Data features of the atypical regime suggest that such a regime is a mixture of the formation and collapsing of housing price bubbles. It is also shown that, depending on which state the economy is in, the appropriate policy responses of the central bank are qualitatively different.

Key words : Housing prices, Regime switching, US monetary policy JEL Classification: E21, D12, D91

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