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Monetary Policy and the House Price Boom Across U.S. States

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Abstract

We use a dynamic factor model estimated via Bayesian methods to disentangle the relative importance of the common component in OFHEO house price movements from state- or region-specific shocks, estimated on quarterly state-level data from 1986 to 2005. We find that historically movements in house prices have mainly been driven by the local (state- or region-specific) component. The recent period (2001-2005) has been different, however: "Local bubbles" have been important in some states, but overall the increase in house prices is a national phenomenon. We then use a VAR to investigate the extent to which expansionary monetary policy is responsible for the common component in house price movements. We find the impact of policy shocks on house prices to be small in comparison with the magnitude of fluctuations in the recent period .

KEY WORDS: Housing, Monetary Policy, Bayesian Analysis

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"We don't perceive that there is a national [housing] bubble but it's hard not to see ... that there are a lot of local bubbles."

Chairman Alan Greenspan (Economic Club of New York, May 20, 2005; CNN Money)

1 Introduction

In some U.S. metropolitan areas house prices increased dramatically during the last few years. The increase in house prices is substantial even if one looks at the average state-level price, which smooths out the differences across local markets within each state. The dark bars in Figure 1 show the annualized average growth rates from the first quarter of 2001 to the last quarter of 2005 in the OFHEO (Office of Federal Housing Enterprise Oversight) house price indexes, deflated by the core PCE inflation, for the forty-eight contiguous U.S. states. In this five-year period house price indexes increased more then ten percent per year in several states on both the East and the West Coasts, notably California, Florida, Nevada, Maryland, Rhode Island, New Jersey and Virginia. The rise in house prices has been very uneven across the nation, with some states, like Texas and Ohio, growing at two percent per year. If we compare the growth in house prices in the last five years with the average growth since 1986, we find that states like Florida have grown two and half times their average, while other states, like Michigan, have grown twenty-five percent less than average.

From the perspective of the current debate, an important question is whether the widespread, but not homogeneous, increase in house prices reflects a national phenomenon or rather, in the words of Chairman Greenspan, a collection of "local bubbles." The answer to this question has important policy implications. "Local bubbles" are most likely attributable to local factors, i.e., circumstances that are specific to each geographic market, rather than to monetary policy, which is the same across the nation. On the contrary, if the boom in house prices is a national phenomenon, monetary policy may well be a likely suspect.

To address the issue of a potential national housing cycle we estimate a dynamic factor model in the spirit of Geweke (1977), Sargent and Sims (1977), and Stock and Watson (1989), on state-level OFHEO house price indexes from the mid-eighties to the end of 2005. We then use the factor model to disentangle the component of the increase in the value of housing that is common to all states from the component that is idiosyncratic, i.e. specific to each state. The latter component is meant to capture the "local bubbles" Chairman Greenspan refers to, while the former captures co-movement across all states, and therefore, potentially, what has been referred to as a "national bubble." We find that historically movements in house prices have mainly been driven by the local (state- or region-specific) component. Indeed, growth rates in OFHEO house price index are far less synchronized across states than are the growth rates in real per capita income, which are a measure of the business cycle at the state level.

However, the recent period has been different in this regard. While for a number of states local factors are still very important, for many states that experienced large increases in house prices a substantial fraction of these increases is attributable to the national factor. How can we reconcile this finding with the fact that increases in house prices have been uneven across states? Of course, part of the cross-state heterogeneity is due to local factors. But about sixty percent of the heterogeneity is due to the fact that states have different exposures to the common cycle: Some states, like Iowa, Nebraska, or Oklahoma, are barely affected by the common cycle, while others, for instance most states in the North-east, are strongly affected.

Since in the recent period the common component of the growth in house prices across states has been sizable, we ask to what extent monetary policy is behind this co-movement. Of course there are many other potential causes of the house price boom, such as mortgage market innovations for instance, but here we focus on one of them only: monetary policy. We follow Bernanke and Boivin (2003) and estimate a VAR where the common factor in house prices is one of the variables, while the other variables measure the stance of monetary policy (the federal funds interest rate, money supply), aggregate U.S. inflation and output, and the thirty-year mortgage rate. We identify monetary policy shocks using sign restrictions á la Uhlig (2005) and Canova and De Nicoló (2002). Perhaps not surprisingly, we find that monetary policy has been expansionary in the recent period, in the sense that more deviations from the implied policy rule have been on the side of "loose" rather than "tight" monetary policy. The analysis of impulse responses shows that expansionary monetary policy shocks lead to increase in the housing factor. We then perform the following counterfactual thought experiment: What would have been the alternative path of the housing factor had there not been any monetary policy shocks from the first quarter of 2001 onward? And, in turn, what would have been the counterfactual growth in house prices across states? The results from this counterfactual experiment indicate that the impact of monetary policy shocks on house prices is non-negligible, but overall fairly small in comparison with the magnitude of the price increase over the last five years. Therefore, our analysis suggests that expansionary monetary policy is not behind the recent boom in house prices. It is important to stress the following limitation of our findings. We do not conclude that the low interest rate environment experienced by the US economy is not responsible for the housing boom. Here, we only consider the component of the low interest rate that is attributable to policy shocks – that is, to the Fed deviating from its historical policy rule in an expansionary way. Had the Fed followed a different rule, the results might have been different. But this is a much more difficult question that goes beyond the scope of this paper.

While there are established literatures studying the effect of housing on asset pricing, portfolio choice, business cycles and consumption, the literature on the relationship between housing prices and monetary policy is fairly limited. Chirinko et. al. (2004) study the interrelationship between stock prices, house prices, and real activity in a thirteen country sample. Their primary focus is in determining the role asset prices play in formulating monetary policy. Iacoviello and Minetti (2005) document the role that the housing market plays in creating a credit channel for monetary policy. Their empirical analysis uses a sample of four countries that does not include the U.S. As in this paper, Iacoviello (2005) estimates a VAR to assess the impact of monetary policy shocks on housing prices. Iacoviello estimates a VAR in interest rates, inflation, and detrended output and house prices using quartely data from 1974 to 2003. He then identifies monetary policy shocks using a Choleski decomposition with the interest rate ordered first, and finds that policy shocks have a significant effect on house prices. Our VAR results complement those of Iacoviello, as we consider different datasets and identification approaches. In addition, we provide evidence on the impact of policy shocks at the state level, with particular emphasis on the recent boom.

Perhaps the closest study to ours is Fratantoni and Schuh (2003) who study the effects of monetary policy on regions in the U.S. from 1966-1998. They find that the response of housing investment to monetary policy varies by region. Our paper differs from the previous literature both in terms of methodology and of focus. In terms of methodology, we use a factor model to extract the common cycle in house price fluctuations. In terms of focus, like Fratantoni and Schuh – and unlike Chirinko et. al. (2004), Iacoviello and Minetti (2005), and Iacoviello (2005)– we are interested in the regional differences in the response of house prices to policy shocks. Differently from all these papers, we are particularly interested in the role of monetary policy in the latest housing boom.

The remainder of the paper is as follows. Section 2 describes the dynamic factor model;

section 3 describes the data, section 4 the empirical results, and section 5 concludes.

2 Model

Our approach consists of two steps. In the first step we use a purely statistical model to distinguish the common component of fluctuations from state or region-specific fluctuations. In this step we avoid making too many a-priori assumptions on the drivers of common fluctuations – that is, we let the common factor be latent instead of pre-specifying a number of regressors. Once we have obtained an estimate of the common factor from the statistical model, in the next step we investigate what lies behind it – and in particular, we focus on the role of monetary policy shocks.

Our statistical model is a dynamic factor model estimated via Bayesian methods, as in Kose, Otrok, and Whiteman (2005). The model is used to differentiate movements in house price levels that are common across all states from those that are region- or state-specific. The model postulates that the observable variables ($y_{n,t}$, $n = 1, .., N$, $t = 1, .., T$), the growth rates in state-level house price indexes, depend on a number of latent factors, which capture comovement at the national (f_t^0) or at the regional (f_t^r) level, as well as on state-specific shocks $\epsilon_{n,t}$. Specifically, the model is:

$$y_{n,t} = \mu_n \ + \ \beta_n^0 \ f_t^0 \ + \sum_{r=1}^{R} \ \beta_n^r \ f_t^r \ + \ \epsilon_{n,t}, \tag{1}$$

where μ_n is the average growth rate, which is allowed to differ across states, and β_n^r represents the exposure of state n to factor r. The estimates μ_n may be of interest in themselves, as they reflect historical trends in terms of state-level demographics, economic growth, *etc cetera*, but are not the focus of this paper. This paper focuses on the deviations from the historical mean, and in particular on the decomposition of these deviations in the recent period. We impose the natural restriction that $\beta_n^r = 0$ if state n does not belong to region r. Section 3 discusses the definition of regions.

The law of motions for the factors is given by an $AR(q)$ process:

$$f_t^r = \phi_1^r f_{t-1}^r + .. + \phi_q^r f_{t-q}^r + u_t^r, \quad u_t^r \rightsquigarrow N(0,1), \ r = 0, .., R., \tag{2}$$

where the variance of the innovations u_t^r is normalized to one. An alternative normalization assumption would be to set one of the β_n^0 loadings to one. We choose the former approach to avoid the arbitrariness with the latter. In particular, a bad choice for the normalizing state

n under the second approach (the state for which β_n^0 is set to one) may lead to very imprecise posterior estimates of the factor.[1] One drawback of the first approach is that the magnitude of the movements in the factor itself does not have a direct economic interpretation. For much of the paper this is not a problem, because the quantity of interest is $\beta_n^0 f_t^0$, which is directly interpretable in economic terms as the component of growth in house prices in state n attributable to the national factor, regardless of the normalization. The issue arises only when we report f_t^0 by itself. In order to address this issue, we exploit the fact that the formula used by OFHEO to compute the growth in the aggregate US index can be very well approximated by $y_{US,t} = \sum_n w_n y_{n,t}$, where the weights w_n are the fraction of detached single-family homes in state n. This implies from equation (1) that the impact of a one percent movement in the factor on the aggregate OFHEO index is approximately $\sum_n w_n \beta_n^0$, since:

$$\sum_n w_n y_{n,t} = \sum_n w_n \mu_n \; + \; \left(\sum_n w_n \beta_n^0 \right) \; f_t^0 \; + \sum_n w_n \left(\sum_{r=1}^{R} \beta_n^r \; f_t^r \; + \; \epsilon_{n,t} \right). \tag{3}$$

Therefore, in the remainder of the paper any time we show plots of f_t^0, or compute impulse responses, we multiply it by the posterior median of the quantity $\sum_n w_n \beta_n^0$, so that movements in the factor can be directly mapped into movements in the aggregate OFHEO price index.

The law of motions for the state-specific shocks is given by an $AR(p_n)$ process:

$$\epsilon_{n,t} = \phi_{n,1} \epsilon_{n,t-1} + .. + \phi_{n,p} \epsilon_{n,t-p} + u_{n,t}, \quad u_{n,t} \rightsquigarrow N(0, \sigma_n^2). \tag{4}$$

A key identification assumption which allows us to disentangle the factors from one another and from the state-specific shocks is that all innovations are mutually independent. Intuitively, the factor model disentangles "co-movement" (the f_ts) and "idiosyncratic fluctuations" (the $\epsilon_{n,t}$s) apart. In order for this decomposition to be meaningful the innovations to these different components need to be orthogonal:

$$I\!\!E[u_t^r, u_{n,t}] = 0 \text{ all } r, n, t; \;\; I\!\!E[u_{m,t}, u_{n,t}] = 0 \text{ all } m, n, t. \tag{5}$$

If the innovation to the idiosyncratic components were correlated across series, or with the factors, they would cease to be "idiosyncratic". The attempt to satisfy this assumption is one reason why we include regional factors in the analysis – that is, to explicitly capture

[1] There is here an analogy with the identified VAR literature. There, the choice is between setting the variance of the identified innovations to one, or setting the diagonal elements of the impact matrix to one. Waggoner and Zha (2000) argue in favor of the first approach when using Bayesian methods.

possible sources of comovements across states. Of course, to the extent that we are omitting other important sources of comovements our results may not be robust.[2] Likewise, we assume that the innovations across factors are also orthogonal:

$$\mathbb{E}[u_t^r, u_t^s] = 0 \text{ all } r, s, t, \tag{6}$$

otherwise the distinction between national and regional factors would again lose its meaning.

An important question for this paper is: What is the difference between the national factor and the U.S. OFHEO price index? To what extent does the latter already correctly capture co-movements in house prices across U.S. states, thereby making our approach redundant? Doesn't the fact that the U.S. price index grew so much in the last two years provide by itself evidence that the increase in house prices is a national phenomenon? Equation (3) shows that national factor and the aggregate OFHEO price index are perfectly correlated only under the assumption that the term $\sum_n w_n \left(\sum_{r=1}^R \beta_n^r f_t^r + \epsilon_{n,t} \right)$ is zero. This condition is met *ex ante* only if the weights on each single state or region are negligible, so that the law of large numbers applies and both idiosyncratic and regional shocks average out. In practice this condition is not met: California alone is 10 percent of the OFHEO index for instance. As a consequence, movements in the aggregate OFHEO price index could be driven by either movement in the common component (f_t^0) or by large movements in those states or regions that have large weight in the index. This makes it hard to solve the "local versus national factors" question just by staring at the U.S. OFHEO price index. The factor model is designed to address this identification problem. The model tries to extract the common component of fluctuations across states, without having any information on the relative weight of each state: *ex ante*, all states weight the same in the factor model.

The paper will decompose movements in $y_{n,t}$ into fluctuations due to each of the three components: national, regional, and state-specific component. The statistic $v_n(t_0, t_1)$ computes the variance of fluctuations due to the national factor as the fraction of the sum of the variance of all three components over the sub-sample (t_0, t_1):

$$v_n(t_0, t_1) = \frac{\sum_{t=t_0}^{t_1} (\beta_n^0 \; f_t^0)^2}{\sum_{t=t_0}^{t_1} (\beta_n^0 \; f_t^0)^2 + \sum_{t=t_0}^{t_1} (\beta_n^r \; f_t^r)^2 + \sum_{t=t_0}^{t_1} \epsilon_{n,t}^2}. \tag{7}$$

[2]The literature has also considered "approximate" factor models, that is, models where the idiosyncratic shocks can be cross-sectionally correlated. Doz, Giannone and Reichlin (2006) show that even in this situation maximum likelihood delivers consistent estimates of the factors. While this result is in principle important for us since we effectively use maximum likelihood techniques, in practice some of the conditions underlying their result are not met here. In particular, the size of the cross-section is far from infinity.

This variance decomposition is computed for each state n for the entire sample, as well as for sub-periods of interest, notably the last four years.

The Bayesian procedure used to obtain the posterior distribution of the parameters of interest, the Gibbs sampler, is straightforward for this model. We now give a brief description of it. The Gibbs sampler is a zig-zag procedure where a set of parameters is drawn conditional on another set of parameters, and vice-versa, exploiting the fact that the conditional posterior distributions have a known form, even if the joint posterior does not. The Gibbs sampler for this problem has two step. In the first step we condition on the factors and draw all other parameters. Conditional on the factors, each equation (1) is a regression model with AR(p_n) errors. The procedure developed by Chib and Greenberg (1994) makes it possible to draw μ_n, $\beta^r_{n,t}$, $\phi_{n,j}$, σ^2_n (see Otrok and Whiteman 1998). Since the errors are independent across n, the procedure can be applied equation by equation. The same procedure is applied to the parameters ϕ^r_j of the law of motion of the factors (2). In the second step we draw the factors, conditional on all other parameters. The model is already written in state-space form, equation (1) being the measurement equation, and equations (2) and (4) being the transition equation for the unobserved states, which include both the factors and the idiosyncratic shocks. We can then use the algorithm developed by Carter and Kohn (1994) (see Kim and Nelson 1999, Cogley and Sargent 2002, and Primiceri 2005) to obtain draws of the states. This approach leads to a curse of dimensionality however, as the number of states grows proportionally with the cross-sectional dimension N. A solution to this curse of dimensionality is to pre-whiten the data, that is, to pre-multiply each measurement equation by $1 - \sum_{j=1}^{p} \phi_{n,j} L^j$, so to get rid of the dynamics in the state-specific shocks (see Kim and Nelson 1999, and Quah and Sargent 1993), obtaining:

$$(1 - \sum_{j=1}^{p} \phi_{n,j} L^j) y_{n,t} = (1 - \sum_{j=1}^{p} \phi_{n,j}) \mu_n \ + \sum_{r=0}^{R} \ \beta^r_n \ (1 - \sum_{j=1}^{p} \phi_{n,j} L^j) f^r_t \ + \ u_{n,t}. \qquad (8)$$

To our knowledge, the literature that followed this path has in general conditioned on the initial p observations $(y_{n,1}, \ldots, y_{n,p})$. We use the full sample instead. The extra step required to do this amounts to computing the expectation of the first p realizations of the factors conditional on the initial observations.

The priors used in this paper are quite standard, and similar to those used in Kose, Otrok, and Whiteman (2005). Importantly, we use identical prior for both the house price and the real income data sets, obtaining very different results in terms of the national/local factor decomposition, as we will see. This is indirect evidence that the priors do not drive the main results of the paper. The prior for constant μ_n is normal with 2 and precision (the inverse of the variance) 1. The prior for the loadings β_n^r is fairly loose: it is Gaussian with zero mean and precision equal to 1/250. We have experimented with different degrees of tightness and found that the results of the paper are qualitatively unchanged. The prior for the idiosyncratic innovation variance σ_n^2 is an inverted gamma with parameters 4 and 0.1. The priors for the parameters of the AR polynomial are Normal with mean zero and precision equal to 1 for the first lag, and then increasing geometrically at rate .75 for the subsequent lags. We choose a lag length equal to $q = 3$ for the factors and $p = 2$ for the idiosyncratic shocks. All priors are mutually independent.

3 The Data

Most of the data were obtained from Haver Analytics (Haver mnemonics are in italics). The Housing Price Index (HPI; *HPI@REGIONAL*) is published by the Office of Federal Housing Enterprise Oversight (OFHEO), and captures changes in the value of single-family homes. The HPI is a weighted repeat sales index: It measures average price changes in repeat sales or refinancings on the same properties and weights them (see Calhoun, 1996, for an in-depth description of how the HPI is constructed). The price information is obtained from repeat mortgage transactions on single-family properties whose mortgages have been purchased or securitized by Fannie Mae or Freddie Mac since January 1975.[3] While the housing price data has been criticized for its construction, to our knowledge it is the best data available to the public at the state (or more disaggregated) level.[4] Additionally, we will be working with growth rates of the housing price data so issues related to bias in the level estimates are not relevant. Also, while we use state level data other levels of aggregation (e.g. Metropolitan Statistical Area) are available. We find that the state level

[3]An alternative measure that is available at the state level at quarterly frequency is the Conventional Mortgage Home Price Index (CMHPI), published by Freddie Mac. This measure is roughly based on the same data on which the OFHEO HPI is constructed. Indeed, we find that the correlation between the growth rates in the two price indexes is above .9 over the entire sample period.

[4]Some authors, notably Peach and McCarthy (2004), have emphasized the differences between the OFHEO house price and the constant quality house price index produced by the U.S. Bureau of the Census. They argue that home renovations and improvements lead to an overstatement of the average growth in the OFHEO house prices. The constant quality house price index is simply not available at the disaggregated level. We are aware that the potential mis-measurement of quality can lead to an upward bias in our estimated mean growth rate. However, the average growth rate in house prices is not the focus of the paper. The focus is on comovements in state-level house prices, especially during the recent boom. From this perspective, we think that taking home renovations and improvements into account makes little difference for our analysis.

data is disaggregated enough to establish our main conclusions, and yet the cross-section is small enough for our computational approach to be feasible. We compute growth rates using annualized log-differences, in percent. The HPI data are nominal. We deflate the data using core PCE inflation (*JCXFEBM@USECON*), which measures inflation in the personal consumption expenditure basket less food and energy.

While the HPI data are available from 1975, we use in our estimation only data beginning in the first quarter of 1986. In the working paper version (Del Negro and Otrok, 2005) we document that state-level HPI data are extremely noisy for a number of states before the mid-eighties, with sharp appreciations immediately followed by sharp depreciations. From the perspective of the dynamic factor model, the noise in the series is not necessarily a problem in terms of estimation, as it is captured by the idiosyncratic component. However, our methodology cannot deal with very large time variation in the importance of the noise component, particularly when the time variation is very large as it is for the HPI data. The noise abates considerably for most states after the mid-eighties. We choose the first quarter of 1986 as the starting date for our analysis. Large structural changes in the credit market, such as the end of regulation Q, provide another reason for leaving the first part of the sample out of the analysis. Additionally, these sample gives us a period with one monetary policy regime, which is convenient for the identification of monetary policy shocks. The sample ends in the last quarter of 2005.[5] In summary, we have 20 years (80 quarters) of data for the 48 contiguous U.S. states. We have checked for the robustness of our results to moving the start date to the first quarter of 1985, and found that the results to be robust. The real per capita personal income data (*YPPHQ@PIQR*) are computed by deflating the nominal per capita income data from the Bureau of Economic Analysis using PCE inflation.

The regional factors are defined by the geography. Our baseline specification includes five regions. The first three regions follow the Census definition. These are the North-East Region, which includes the New England and Middle Atlantic Divisions; the Mid-West Region, which includes the East- and West-North-Central Divisions (the former includes the Great Lakes regions, while the latter includes the Plains); the West Region, which includes the Mountains and the Pacific Divisions. We split the South Region, which includes the South Atlantic, the East-South-Central, and the West-South-Central Divisions, into two separate regions: South Atlantic and the East-South-Central (i.e., Alabama, Kentucky, Mississippi, and Tennessee) on one side, and the West-South-Central division (Arkansas, Louisiana, Oklahoma, Texas), which includes a number of oil-producing states, on the other.

[5]The working paper version shows that the results are robust to the exclusion of the last four quarters.

We have also tried a specification with nine regions, the nine Census Divisions, and obtained very similar results. The only difference is that for some of the Divisions with few states the regional factors were not well identified, hence we preferred the five regions specification.

The data used in the VAR include two measures of monetary policy, total reserves (in results not reported here we use non-borrowed reserves as a robustness check) and the federal funds rate, inflation as measured by the GDP deflator and real output growth as measured by the growth in real GDP. All data were taken from the Federal Reserve Bank of St Louis database (FRED). The FRED mnemonics are *BOGNONBR*, *TOTRESNS*, *FEDFUNDS*, *GDPDEF*, and *GDPC1*, respectively. The 30 year mortgage rate is obtained from Haver Analytics (FCM@USECON). All data are quarterly, and the time-period coincides with that used in the factor model.

4 Empirical Results

We have three sets of empirical results. The first set of results provides evidence on the relative importance of national versus regional or state-specific shocks in driving movements in house prices across US states over the past twenty years. We document that there is a large degree of heterogeneity across states in regard to relative importance of the national factors. Overall, however, we find that historically movements in housing prices are mainly driven by local factors, either regional or state-specific. The second set of results argues that the recent period has been different in this regard. While local factors have remained important, the increase in house prices that occurred in several states in the last four years is mainly driven by the national factor. Given the importance of the national factor, in the third sub-section we ask whether or not expansionary monetary policy lies behind the recent national housing boom. Specifically, we identify monetary policy shocks using a VAR with sign restriction, and investigate their impact on house prices. We find the impact of monetary policy on the housing boom to be non-negligible but small relative to the size of the housing boom.

4.1 House price fluctuations and business cycles across US states

We first want to establish the degree of comovement in housing prices across states. Figure 2 contains three charts. The top charts shows the data – the growth rate in house price indexes for the forty-eight contiguous states from QI-1986 to QIV-2005. The other two charts decompose movements in the growth rates into national and local factors. By national factor (middle chart) we mean the impact of national shocks on state n, that is, the component $\beta_n^0\ f_t^0$ of equation (1). All lines in the middle chart are perfectly correlated by construction, with the different amplitudes of fluctuations reflecting the size of the exposure β_n^0 to national shocks. By local factors (bottom chart) we mean the joint impact of regional and state specific shocks, that is, the $\beta_n^r\ f_t^r\ +\ \epsilon_{n,t}$ component of equation (1). The data show two periods of relatively high volatility in the growth rates. The first coincides with the early part of the sample – from 1986 to the early 1990s. The second coincides with the most recent period. A difference between these two episodes is in the role played by local factors, as shown by the bottom chart. In the first episode local factors are behind most of the volatility.[6] Simply eyeballing the data (top chart) it is very hard to tell whether there is a common component of house price movements in the late eighties and early nineties, for any co-movement is shadowed by the importance of local factors. In contrast, in the recent episode the common component is quite apparent. Except for two states – notably Nevada and California– the volatility in local factors in the 2001-2004 period is no higher than it had been in the previous four years. In the last two quarters of 2005, toward the very end of the boom, state-specific volatility rises slightly, yet not to levels comparable with the first part of the sample.

We have performed the same decomposition for state-level growth rate in real per capita income. Real per capita income can be seen as a proxy for state-level business cycles (output is not available at the quarterly level). The model used on the real income data is the same, including the specification of all the priors, as that used on the house price data.The data sets however are quite different, and so are the results of the decomposition, which we show in Figure 4 of the working paper version and briefly summarize here. First, state-level real income growth rates are less volatile than the house price data, with some exceptions for rural states like North Dakota. Second, state-level business cycles appear to move more in step than fluctuations in house prices. Local factors are important, but appear to be largely short frequency deviations from the common component. Finally, for the income data there is no evidence of the two high volatility episodes that characterize the house price data. In particular, not much is happening toward the end of the sample.

The top panel of Figure 3 quantifies the relative importance of the national factor for the house price and the income data over the entire sample. In particular, for each state the

[6] Some of the local factors are correlated across subsets of states, as they represent regional shocks to the South West and Mid West regions.

figure shows the magnitude $v_n(1, T)$ (see equation (7)), that is, the variance of fluctuations due to the national factor as a fraction of the variance of all components, for housing on the horizontal axis and for income on the vertical axis. For all states that are above the 45 degree line the common component of fluctuations is more important for income than it is for house prices. About seventy percent of states are above the 45 degree line. For the median state the national factor explains only a quarter of the variance of house price movements. This is in contrast with state-level business cycles. For the median state the national factor explains about fifty percent of the variance in real per-capita income growth. One should bear in mind that these figures include the housing boom years. If we drop the last five years, the fraction of states for which the national factor is more important for income than for housing rises to about ninety percent.

Recent literature (see for instance McCarthy and Peach, 2004) tries to explain the recent increase in house prices by means of an affordability index, which measures the extent to which current house prices are "affordable" given the level of income and of mortgage interest rates. The underlying idea is that as income grows and interest rates decline, households bid up house prices because they simply can afford them (that is, the mortgage payments remain constant as a fraction of income). This theory posits a tight relationship between income and house prices. The state-level evidence presented here presents a challenge to this theory, as it shows that local factors dominate fluctuations in house prices, at least in the first part of the sample, but not in per-capita income. An interesting hypothesis, which we do not investigate further in this paper, is that segmentation in the mortgage markets up until the mid-nineties lies in part behind the importance of the local factors.

4.2 The recent housing price boom

Figure 4 plots the posterior median of estimated national housing factor f_t^0 (black, scale on the left axis), the ninety percent bands (dotted lines), as well as the OFHEO U.S. price index (gray dash-and-dotted, right axis). The figure shows that the recent period, particularly the last two years, has been one of unprecedented volatility at the national level. The questions we try to address in this section are twofold. First, are shocks at the national level, or local factors, behind the recent increase in house prices across many U.S. states? Quantitatively, what is the relative importance of the two? Second, to what extent can different exposures to national shocks explain the heterogeneity across states in the house price increases?

Figure 4 also shows that the estimated national factor and the OFHEO U.S. price index are very correlated, particularly in the second part of the sample. As discussed in section 2, movements in the OFHEO U.S. price index could either be driven by movements in the common component (f_t^0) or by large shocks in states or regions that have large weight in the index. Indeed, a key issue in the debate is whether the increase in the national index reflects a national phenomenon, or local phenomena in a number of very highly populated (and therefore highly weighted) states, like Florida or California. The fact that the estimated national factor and the OFHEO U.S. price index move in sync in the last five years suggests that the recent housing boom is a national phenomenon.

In order to quantify the importance of national shocks relative to local factors in the recent period we plot for each state the the magnitude $v_n(t_0, t_1)$ for the sub-sample 2001-2005 on the vertical axis, and 1986-2000 on the horizontal axis in the bottom panel of Figure 3. Again, $v_n(.,.)$ represents the variance of house prices fluctuations in state n due to the national factor as a fraction of the variance of all components. For all states that are above the 45 degree line the common component of fluctuations is more important in the recent period than in the remainder of the sample. Figure 3 shows only the median of the posterior distribution of $v_n(\text{QI-1986,QIV-2000})$ and $v_n(\text{QI-2001,QIV-2005})$, but in general these magnitudes are tightly estimated: The difference $v_n(\text{QI-1986,QIV-2000}) - v_n(\text{QI-2001,QIV-2005})$ is almost always statistically significant. Between 2001 and 2005 the relative importance of national shocks has increased for all states. For the median state the explanatory power of the national factor in terms of the variance of house price movements has increased three-fold, from 11 to 34 percent. Overall the figure provides evidence that the recent period is different from the rest of the sample, in that the relative importance of national shocks has increased substantially. Even in the recent period, heterogeneity across states is large – for a number of South-western and Mid-western states the importance of the national factor remains negligible even in the last four years. For many states local factors are still the dominant source of fluctuations. However, we now proceed to show that for many of the states that witnessed large increases in house prices, the national factor lies behind such increases.

The gray bars in Figure 1 represent the component of the average annualized growth rates in the OFHEO house price index for the 2001-2005 period that can be attributed to the national factor. Specifically, for each state the figure shows the median of the posterior for the quantity $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}} (\mu_n + \beta_n^0 f_t^0)/20$.[7] The dark bars represent the total annualized

[7] To avoid cluttering the picture we do not show the ninety percent bands for the component attributable to the national factor: Taking estimation uncertainty into account does not alter the conclusions. The figure

growth rates, as discussed in the introduction. The figure shows that a substantial fraction of the cross-state heterogeneity in the recent housing price boom is attributable to the fact that states have different exposures to the national factor. Indeed, the cross-sectional variance in the component of house price growth explained by the common factor (gray bars) amounts to fifty-seven percent of the cross-sectional variance in average house price growth (dark bars). Heterogeneity in the mean growth rates over the whole sample (μ_n) plays virtually no role in driving this result: the fraction of the cross-sectional variance explained by different exposures to the national factor is still fifty-seven percent using demeaned data. Of course, local factors – shown by the distance between the dark and the light gray bars – are still important, especially for states like California and Nevada. But it is certainly not the case that all cross-state heterogeneity in the house price boom is driven by "local bubbles".[8]

4.3 Expansionary monetary policy and the housing price boom

We have so far made a case that the national factor has been important in the 2001-2005 period. What economic forces are behind the common factor? The answer must lie in a common set of shocks or changes to the housing market. It is beyond the scope of this paper to analyze all possible explanations for the house price boom. Here, we focus on one of the most likely potential culprits, monetary policy shocks. By most accounts, including the Fed's own FOMC statements, monetary policy has been expansionary in recent history. In the press, the boom in house prices has often been associated with loose monetary policy. This section uses a VAR to investigate the importance of monetary policy shocks in driving the common component of house price increases in the recent period.

The VAR includes standard macroeconomic data, as well as our estimates of the common house price factor. We use some newly developed VAR identification techniques that require minimal assumptions to identify and extract monetary shocks. We then use the identified shocks in a counterfactual experiment where we eliminate these shocks from QI-2001 on to create a counterfactual housing price factor. Next, we are able to identify the effects of

is available from the authors upon request.

[8] What state-specific characteristics drive the differences in exposures to the common factor is an interesting question which we leave for further research. We observe that the size of the light gray bars (that is, the magnitude of the exposures) has a clear geographical pattern: It is generally large for states on both the East and the West coasts – states where fluctuations in residential land prices are likely to be larger (see Davis and Heathcote 2004), possibly because of zoning and other land use controls (see Glaeser and Gyourko, 2002), and small for states in the center.

this shock on each state-level house price index via its factor loading or sensitivity to the national cycle. These calculations give us an idea on the magnitude of the importance of expansionary monetary policy on housing prices.

The virtue of our approach is that we can combine of a small number of national level variables and with a wide selection of regional variables capturing local economic conditions without loosing too many degrees of freedom (see Bernanke and Boivin 2003). The factor-augmented-VAR thus yields a parsimonious model that allows us to study the effects of national shocks on regional economies. A similar approach was developed by Otrok and Terrones (2004) to study the effects of US monetary policy shocks on global house prices.

The reduced form VAR is given by:

$$Y_t = A(L)Y_{t-1} + e_t,\ e_t \rightsquigarrow N(0, \Sigma), \tag{9}$$

where Y_t is an $m \times 1$ vector that includes observable data as well as the housing factor and $A(L)$ is a matrix lag polynomial. Ideally, one would estimate the VAR coefficients as part of the MCMC algorithm and allow the uncertainty in the estimates of the factors to be captured in the uncertainty in the VAR parameters and the impulse response functions (see Bernanke, Boivin, Eliasz 2005) . However, since estimating the dynamic factor model itself is computationally costly, combining this with estimating and identifying the VAR at each step of the MCMC algorithm would be computationally very costly.[9] Given how tightly the latent factor is estimated (see Figure 4), we believe this to be a relatively minor issue.

The identification of monetary shocks is controversial. We identify monetary shocks using a sign restrictions approach introduced by Faust (1998) and further developed by Uhlig (2005) (see also Canova and DeNicoló 2002). Under this approach identification is achieved by placing restrictions on the sign of the impulse responses with respect to the shock(s) of interest for some variables for a number of periods in the future. For example, after a contractionary monetary policy shock we restrict the impulse response function of reserves to be non-positive, and that of interest rates to be non-negative. One advantage of Uhlig's procedure is that it allows one to identify only the innovation of interest (here as in Uhlig's work, monetary policy shocks), without making assumptions on the remainder of the system. A second advantage is that by changing the set of variables that are subject to the sign restrictions one can analyze the outcome of a variety of different identification

[9] Our identification procedure, described below, is based on sign restrictions. This procedure requires many draws of potential impulse response functions and is also computationally intensive procedure. It is the combination of this procedure with the dynamic factor model that is infeasible.

approaches. We exploit this feature by exploring a number of them, which we describe below. We then pick the identification that gives the largest impact of monetary policy shocks on the house factor, thereby showing an upper bound on the importance of policy shocks.

Our VAR consists of six variables: the house factor, total reserves, CPI inflation, GDP growth, the thirty-year mortgage rate and the Federal Funds rate. The VAR is estimated using quarterly data from QI-1986 to QIV-2005 and has four lags. We use a Litterman prior for the VAR coefficients, taking into account that it applies to variables in differences rather than levels. [10] The first four variables are in growth rates while the fed funds rate and mortgage rate are first differenced.[11] We then identify the (contractionary) monetary shock as the shock that results in a set of impulse response functions consistent with 1) a increase in the Fed Funds rate, 2) a non-positive change in growth of total reserves, 3) a non-positive change in CPI growth, 4) a non-positive change in GDP growth. These restrictions hold for three quarters after the shock. We leave the responses to the house factor and the mortgage rate unrestricted. The restriction on the response of CPI inflation and GDP growth to be non-positive in response to a contractionary monetary shock is not without controversy. Indeed, the former is at the center of the 'price-puzzle' debate while the latter has generated controversy in the literature (Uhlig 2005). Among the alternatives identification schemes we study are: i) leaving CPI and GDP unrestricted, ii) constraining the impact on the mortgage rate to be non-negative, and iii) changing the number of quarters for which the restrictions have to hold to two. We also considered alternative model specifications, where we include non-borrowed reserves in place of total reserves, or alternatively include both measures of reserves. In another specification we use both the Fed Funds rate and mortgage rates in levels. Finally, we increased the lag lengths for the VAR to five and six lags. For none of these alternative identification schemes/model specifications is the impact of policy shocks on the house factor is larger than reported.

Figure 5 displays the impulse response function of all of the variables to the monetary policy shock, along with the 68 percent posterior coverage intervals.[12] The house

[10]This implies that the prior on the first lag is centered at zero rather than one.

[11]From a Bayesian perspective it would be natural to estimate the model in levels (e.g. Uhlig 1994). For space consideration we do not report the results for a model in levels but the affects of monetary policy are smaller in that specification. We choose to focus on the results with first differences since the factor model requires stationarity and hence the factor is the common factor in the growth rate of housing prices. To maintain symmetry in the model we also work with differences or growth rates of the other variables

[12]The coverage interval reflects the uncertainty in the posterior distribution of both the VAR rotation matrix and the VAR parameters.

factor shows a significant and persistent drop following the contractionary monetary shock. This result confirms that monetary policy shocks can impact housing prices, as found in Iacoviello (2005). The mortgage rate increases after the shock, although the uncertainty surrounding the size of the impact is large, and the posterior coverage intervals include zero. For all other variables the responses are constrained, at least for the first three period after impact, so little interpretation need be given to the sign of these responses. From a quantitative standpoint, the effect of the monetary shock on housing prices is considerably larger then the impact on the inflation rate, consistent with Iacoviello's findings. The size of both impulse responses are also roughly in line with his. Two remarks are in order. First, in our sample period house prices have been much more volatile than inflation. Hence it is not too surprising to find that the impulse responses to monetary shocks are larger for housing prices than for inflation. Second, we deliberately search among all possible "reasonable" identification procedures, and VAR specifications, for the one that deliver the largest impact of monetary policy on house prices, in order to find an upper bound. The reader should bear in mind that there are other reasonable specifications/identification strategies that deliver a lower impact of monetary policy on house prices.

The forecast error variance decomposition presented in Table 1 provides some additional evidence on the quantitative impact of the monetary shocks. Monetary shocks explain about 13 percent of housing price movements (corresponding to the median IRFs) on impact of the shock, with a slow decline to about 8 percent at longer horizons. We interpret this number as a sizable fraction – given that housing prices should be affected by many other shocks such as income and mortgage market innovations. At the same time monetary shocks seem to account for a trivial amount of inflation volatility at short horizons and a more sizable 15 percent at longer horizons. The monetary shock also explains little of GDP volatility at any horizon. While the former result is consistent with that of Christiano, Eichenbaum and Evans (2005), the latter result is at odds with their conclusion that monetary shocks explain a large amount of GDP volatility at longer horizons. The difference is likely due to the fact that we study a period of relative stability often labeled the "great moderation" while the work of CEE includes the more volatile 1970s and ends in 1995.

Perhaps of greater interest than the shape and significance of the impulse response function for housing prices to the monetary shock is the quantitative importance of the recent policy choices on the housing price boom. To answer this question we construct a counterfactual where we create a new set of "data" for all variables in our VAR after setting the structural monetary shocks from QI-2001 on equal to zero. This experiment is in the same spirit of Uhlig (2001) who constructs a counterfactual to identify whether or not markets were surprised by the decline in interest rates. That is, he backs out the structural shocks to determine whether or not policy differed from the past in this period. Here we back out the monetary shocks to determine their effect on housing prices. The two left panels of Figure 6 show the actual (solid line) and the counterfactual (dash-and-dotted line) paths for the Fed Funds rate and the housing factor from this procedure. The two right panels of Figure 6 show the difference between actual and counterfactual for the Fed Funds rate and the housing factor (cumulated), respectively.

The difference between the actual and the counterfactual Fed Funds rate is not very large, but is not negligible either. Toward the end of 2004, the counterfactual Fed Funds rate is 80 basis points higher than the actual fed funds rate. Other multivariate models used in policy analysis that are estimated using a different sample and different variables, such as the Atlanta Fed BVAR for instance (see Zha 1998), produce a counterfactual Fed Funds rate that is not very different from the one shown here. Consistent with the results for the impulse response functions, the counterfactual housing factor is generally below the actual one. At the end of the sample the cumulated difference between the two is 1.2 percent. In absolute terms this is not a negligible number. Relatively to the overall growth in house prices in the recent period the number is quite small, as evidenced from the fact that in the bottom left panel the difference between the actual and the counterfactual is hardly visible.

The size of the state-level response to monetary policy shocks varies according to each state's exposure to the common cycle. States with larger exposure to the common factor will have a stronger response to the monetary shock than those with smaller factor loadings. Figure 7 shows the impact of policy shocks across the forty-eight states, relative to the overall impact of the national housing factor. Specifically, the gray bars in Figure 7 represent the component of the average growth rates in the 2001-2005 period that can be attributed to the national factor, namely the quantities $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}} \beta_n^0 f_t^0/20$ that were also shown in Figure 1 (except that in Figure 1 we include the sample mean μ_n for each state, while here we exclude it). The white bars quantify the impact of policy shocks. For each state they show the quantity $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}} \beta_n^0 \; (f_t^0 - \tilde{f}_t^0)/20$, where by $\tilde{f}_t^0$ we denote the counterfactual factor. In other words, the white bars show how smaller growth in house prices would have been in absence of policy shocks in state n. Figure 7 shows that the impact of monetary policy shocks has been small: always less than one percentage point in terms of average growth, consistently with the aggregate numbers reported in Figure 6. Again, these figures are an upper bound among all the specifications/identification schemes we considered.

The impulse response functions and the counterfactual seem to tell a different story about the effects of monetary policy. The IRF in Figure 5 shows a potentially large role for monetary policy shocks on housing prices while the counterfactual reveals a limited role for monetary shocks in driving the recent house price boom. The answer is twofold. First, while in absolute terms the impact of policy shocks on the house factor is sizable, relatively to magnitude of fluctuations in house prices observed in the last five years it is fairly small. Second, the evidence points to a limited role for monetary policy shocks in the post 2001 sample period. While there were more expansionary then contractionary shocks over this period, the cumulated impact of the shocks on interest rates and house prices is not large, as shown in Figure 6.

In summary, this section's findings are: i) the Fed, while following a slightly more expansionary policy than in the past did not deviate substantially from "business as usual" in the recent period; and ii) to the extent that the Fed did pursue an overly accommodative policy, the impact of this policy on house prices has been small relative to the overall housing price increase of the last five years. Given that in the popular press loose monetary policy is sometimes blamed for the housing bubble, our findings may be relevant to the current debate. Our result that the impact of monetary policy shocks on house prices was small in the recent period does not of course imply that the low interest rate environment experienced by the US economy is not responsible for the housing boom. Here, we only consider the component of the low interest rate that is attributable to policy shocks – that is, to the Fed deviating from its historical policy rule in an expansionary way. Had the Fed reacted differently to the environment – had it followed a different rule – the results might have been quite different. This is a much more difficult question that (because of the Lucas' critique) goes beyond the scope of this paper.

4.4 Alternative identification schemes

The identification scheme we use here is not without controversy. A more conventional approach to take is to use the recursive procedure first used in Sims (1980). This approach imposes a contemporaneous ordering of the shocks. For example, if GDP is ordered before the federal funds rate then the fed funds rate can respond to GDP shocks but not vice versa. We have implemented such a strategy for our dataset to assess the robustness of our identification approach. We order the house factor first, followed by GDP, inflation, mortgage rates, reserves and the federal funds rate last, consistently with the approach used in Christiano, Eichenbaum and Evans (2005). The results from the impulses response functions show that the house factor moves positively, but statistically insignificantly in response to a contractionary federal funds rate shock. In response to a contractionary shock reserves the house factor falls, but again the response is not statistically significant.[13] The response to the reserve shock is consistent with our intuition and previous results. The response to the funds rate shock is a bit puzzling, but in these IRFs inflation also does not fall in response contractionary policy shock and the responses of most variables in the VAR are not statistically significant. The avoidance of the so called price puzzle is one of the reasons we choose to use the sign restriction approach in the first place. Our identification scheme takes into account that a monetary shock simultaneously affects both the federal funds rate and the reserve market, hence we are more comfortable with that identification approach. The main point for us is that the more conventional identification approach will not lead to a large role for monetary policy shocks in driving house process.

Iacoviello (2005) also uses a recursive identification procedure on a four variable VAR of house price, interest rates, inflation and output. With the federal funds rate ordered first and the houses and output ordered last he finds that contractionary monetary shocks lead to a reduction in house prices. This is consistent with the results we report with the sign identification. The difference between his result and our recursive scheme is probably due to both sample period (Iacoviello start in 1974 and ends in 2003) and variables in the VAR. However, our sign restriction impulse response functions are quantitatively similar to his recursive estimation procedure in both sign and magnitude.

5 Conclusions

We use a dynamic factor model estimated via Bayesian methods to disentangle the relative importance of the common component in OFHEO house price movements from state- or region-specific shocks. Our sample consists of quarterly data from 1986 to 2005. We find that historically fluctuations in house prices have mainly been driven by the local (state- or region-specific) component. Indeed, growth rates in OFHEO house price index are less synchronized across states than are the growth rates in real per capita income, which are a measure of the business cycle at the state level. In the recent (2001-2005) period, however, "local bubbles" have been important in some states, but that overall the increase in house prices is a national phenomenon. We then use a VAR to investigate the extent to which

[13] Due to space constraints we do not report these impulse responses, but the results are available from the authors upon request.

expansionary monetary policy is responsible for the common component in house price movements. We find the impact of policy shocks on house prices to be small relative to the size of the recent housing price increase.

References

Bernanke, Ben S. and Jean Boivin. 2003. “Monetary policy in a data-rich environment.” *Journal of Monetary Economics* 50 (3), 525-546.

Bernanke, Ben S., Jean Boivin, and Piotr, Eliasz. 2005. “Measuring the Effects of Monetary Policy: A Factor-augmented Vector Autoregressive (FAVAR) Approach.” *Quarterly Journal of Economics*, 120 (1), 387-422.

Calhoun, Charles A.. 1996. “OFHEO House Price Indexes: HPI Technical Description.” Manuscript, Office of Federal Housing Enterprise Oversight, Washington, D.C. (*http://www.ofheo.gov/Media/Archive/house/hpi_tech.pdf*).

Canova, Fabio and Gianni De Nicoló. 2002. “Monetary Disturbances Matter for Business Cycle Fluctuations in the G-7.” *Journal of Monetary Economics*, **49**(6), 1131-1159.

Chib, Siddhartha and Edward Greenberg. 1994. “Bayes inferences in regression models with ARMA(p,q) errors.” *Journal of Econometrics* **64**, 183-206.

Chirinko, Robert S., Leo de Haan, and Elmer Sterken, 2004. “Asset Price Shocks, Real Expenditures, and Financial Structure: A Multi-Country Analysis.” *Working Paper*, Emory University.

Christiano, Lawrence J., Martin Eichenbaum, and Charles Evans, 2005. “Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy,” *Journal of Political Economy*, **113**, 1-45.

Cogley, Timothy and Thomas J. Sargent. 2002. “Evolving Post-World War II U.S. Inflation Dynamics.” *NBER macroeconomics annual*, Cambridge and London: MIT Press, 331-73.

Davis, Morris A. and Jonathan Heathcote. 2004. “The Price and Quantity of Residential Land in the United States” *Mimeo*, Georgetown University.

Del Negro, Marco and Christopher Otrok, 2005. “Monetary Policy and the House Price Boom Across U.S. States.” *Federal Reserve Bank of Atlanta WP* 2005-24.

Doz, Catherine, Domenico Giannone, and Lucrezia Reichlin, 2006. "A Quasi-Maximum Likelihood Approach For Large Approximate Dynamic Factor Models." *CEPR Discussion Paper* 5724.

Forni, Mario, and Lucrezia Reichlin. 2001. "Federal Policies and Local Economies: Europe and the U.S." *European Economic Review* 45, 109-134.

Fratantoni, Michael, and Scott Schuh. 2003. "Monetary Policy, Housing, and Heterogeneous Regional Markets." *Journal of Money, Credit, and Banking* 35 (4), 557-589.

Geweke, John F. 1977 "The Dynamic Factor Analysis of Economic Time Series," in D. Aigner and A. Goldberger, eds., *Latent Variables in Socioeconomic Models.* Amsterdam, North Holland, 365-383.

Glaeser, Edward L. and Joseph Gyourko. 2002. "The Impact of Zoning on Housing Affordability." *NBER Working Paper 8835*, Cambridge.

Iacoviello, Matteo. 2005. " House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle." *American Economic Review* 95 (3), 739-764.

Iacoviello, Matteo and Raoul Minetti, 2003, "The Credit Channel of Monetary Policy: Evidence from the housing market." *Working Paper*, Boston College.

Kim, Chang-Jin, and Charles R. Nelson. 1999. "State-space models with regime switching." *The MIT Press,* Cambridge, Massachusetts.

Kose, M. Ayhan, Christopher Otrok, and Charles H. Whiteman, 2005. "International Business Cycles: World, Region and Country Specific Factors." *American Economic Review* 93 (4), 1216-1239.

Kose, M. Ayhan, Christopher Otrok, and Charles H. Whiteman, 2003. "Understanding the Evolution of World Business Cycles" manuscript, University of Virginia

McCarthy, Jonathan and Richard W. Peach. 2004 "Are Home Prices the Next 'Bubble'?" *Federal Reserve Bank of New York Economic Policy Review*, December, 1-17.

Otrok, Christopher and Charles H. Whiteman, 1998. "Bayesian leading indicators: measuring and predicting economic conditions in Iowa." *International Economic Review* 39 (4), 997-1014.

Otrok, Christopher and Marco Terrones. 2004. "House Prices, Interest Rates, and Macroeconomic Fluctuations: International Evidence." *Manuscript.* University of Virginia.

Primiceri, Giorgio. 2005 "Time Varying Structural Vector Autoregressions and Monetary Policy." *Review of Economic Studies* 72, 821-852

Quah, Danny and Thomas J. Sargent. 1993. "A Dynamic Index Model for Large Cross Sections." in *NBER Studies in Business Cycles, vol. 28.* Chicago and London: University of Chicago Press, 285-306.

Sargent, Thomas J. and Christopher A. Sims. "Business Cycle Modeling Without Pretending to Have Too Much A Priori Economic Theory," in Christopher A. Sims et al., eds., *New Methods in Business Cycle Research.* Minneapolis: Federal Reserve Bank of Minneapolis, 1977, 45-108.

Sims, Christopher A., 1980, "Macroeconomics and Reality." *Econometrica* **48**(1), 1-48.

Stock, James H., and Mark H. Watson. 1989. "New indices of coincident and leading indicators." *NBER macroeconomics annual*, Cambridge and London: MIT Press, 351-393.

Uhlig, Harald. 1994. "What Macroeconomists Should Know About Unit Roots: A Bayesian Perspective." *Econometric Theory* 10, 645-671.

Uhlig, Harald, 2001, "Did the Fed Surprise the Markets in 2001? A case study for VARs with Sign Restrictions." *Manuscript.* Humboldt University.

Uhlig, Harald. 2005. "What are the Effects of Monetary Policy? Evidence from an Agnostic Identification Procedure." *Journal of Monetary Economics* 52 (2),381-419.

Waggoner, Daniel F. and Tao Zha. 2000. "Likelihood-Preserving Normalization in Multiple Equation Models." *Federal Reserve Bank of Atlanta WP* 2000-8.

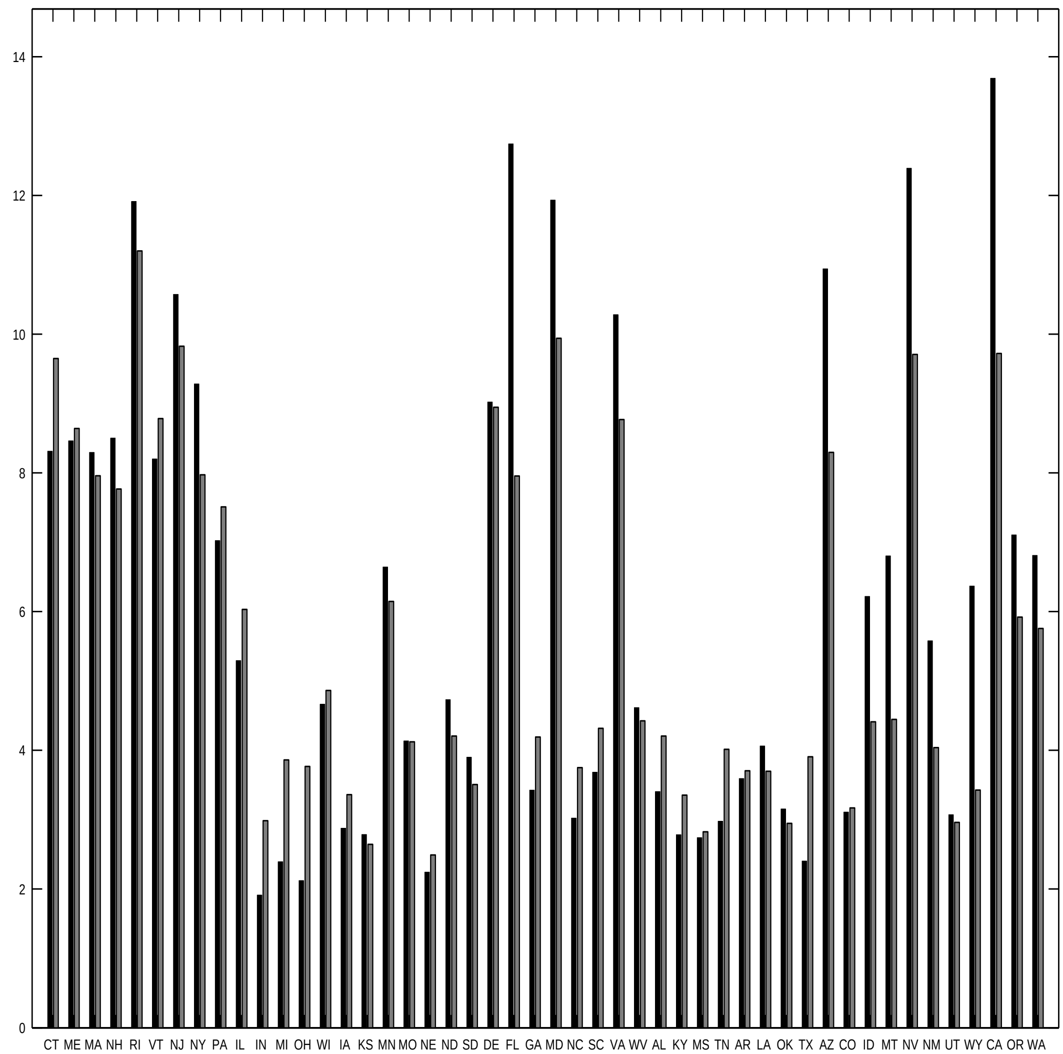
Zha, Tao. 1998. "A Dynamic Multivariate Model for Use in Formulating Policy." *Federal Reserve Bank of Atlanta Quarterly Review*, First Quarter, 16-29.

Table 1: Forecast Error Variance due to Monetary Shocks

	House	Reserves	GDP	Inflation	Fed Funds	Mortgage
horizon						
1	13.46	0.11	6.68	2.73	26.08	18.38
2	11.41	62.35	5.17	4.75	8.75	18.61
3	8.91	0.46	1.89	7.79	3.08	32.54
4	9.27	22.51	2.93	5.02	0.31	21.11
5	9.15	9.46	2.17	7.35	1.48	22.26
6	9.23	15.26	0.88	10.47	6.90	10.25
7	9.10	9.92	0.89	12.34	9.61	6.35
8	9.09	9.67	0.91	13.62	9.80	6.66
9	9.03	8.07	0.92	14.65	8.71	6.29
10	8.89	8.82	0.92	14.68	7.82	7.74

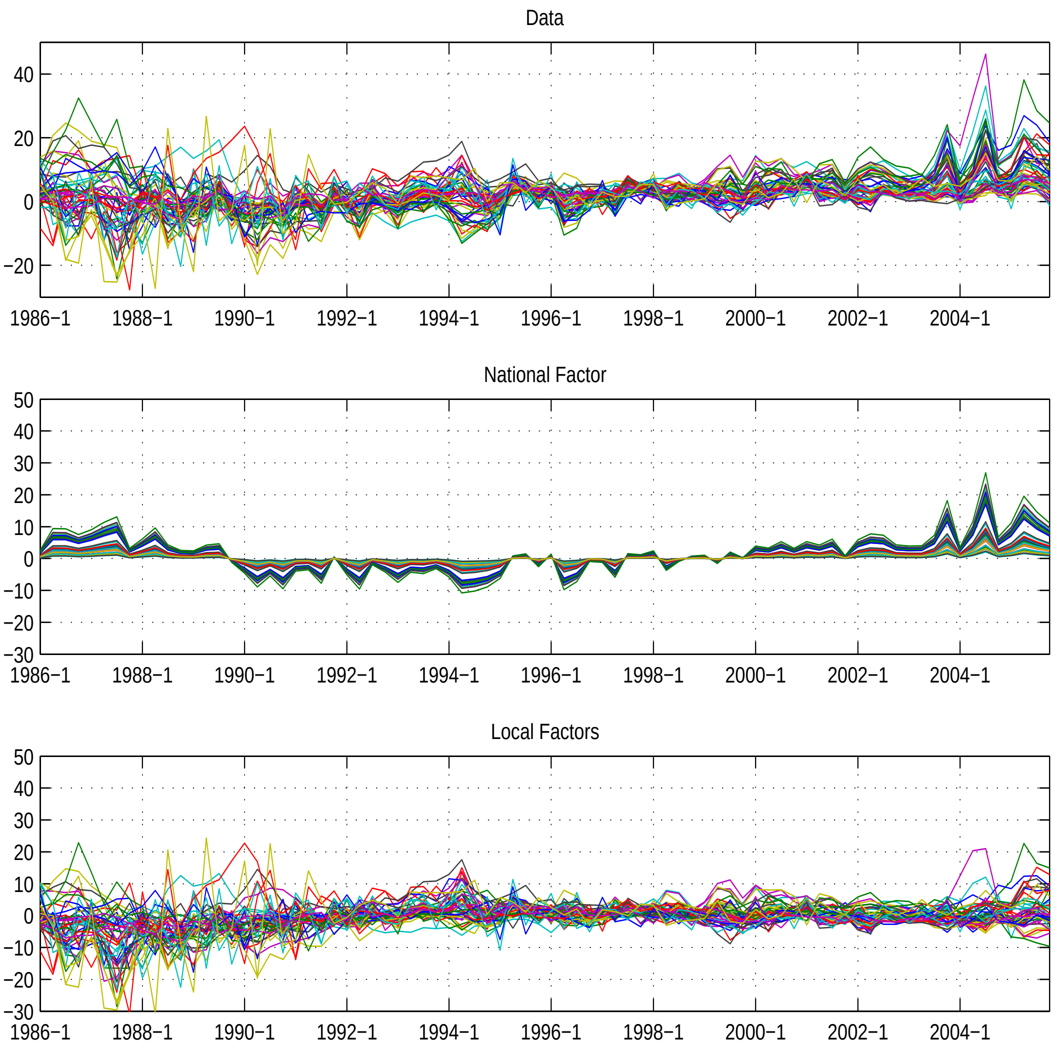
Notes: The figure displays for all of the variables included in the VAR the percent of the variance in the reduced form innovation at horizons 1 through 10 attributable to monetary policy shocks. The VAR is estimated with quarterly data from QI-2006 to QIV-2005.

Figure 1: Cross-state Heterogeneity in the Recent House Price Boom: The Role of the National Factor



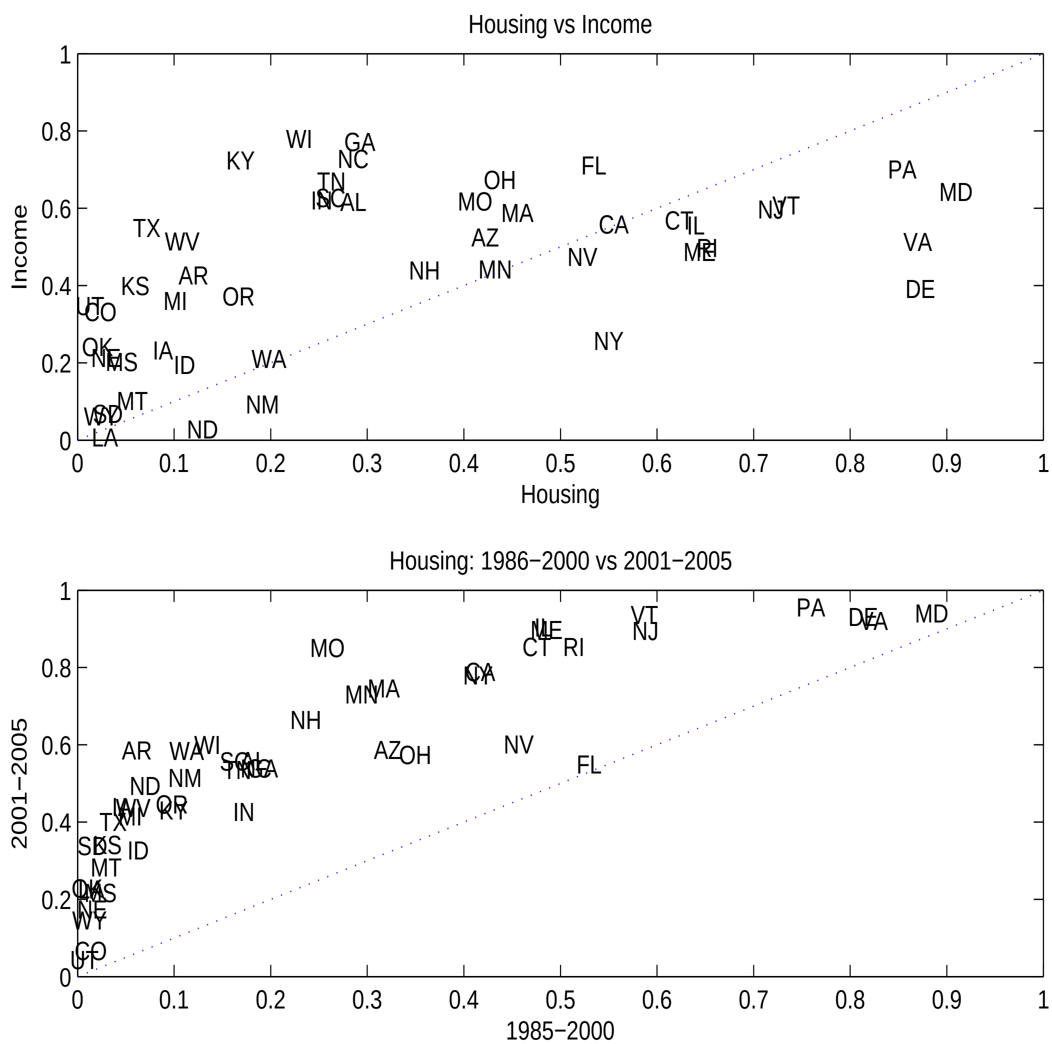
Notes: The figure shows for each of the forty-eight contiguous states the annualized average growth rates in real OFHEO house price index (dark gray bars) for the 2001-2005 period: $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}}(y_{n,t}-)/20$. The light gray bar represents the component of the average growth rates for the same period that can be attributed to the national factor: $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}}(\mu_n + \beta_n^0 f_t^0)/20$.

Figure 2: State-level Real House Price Growth: National and Local Factors



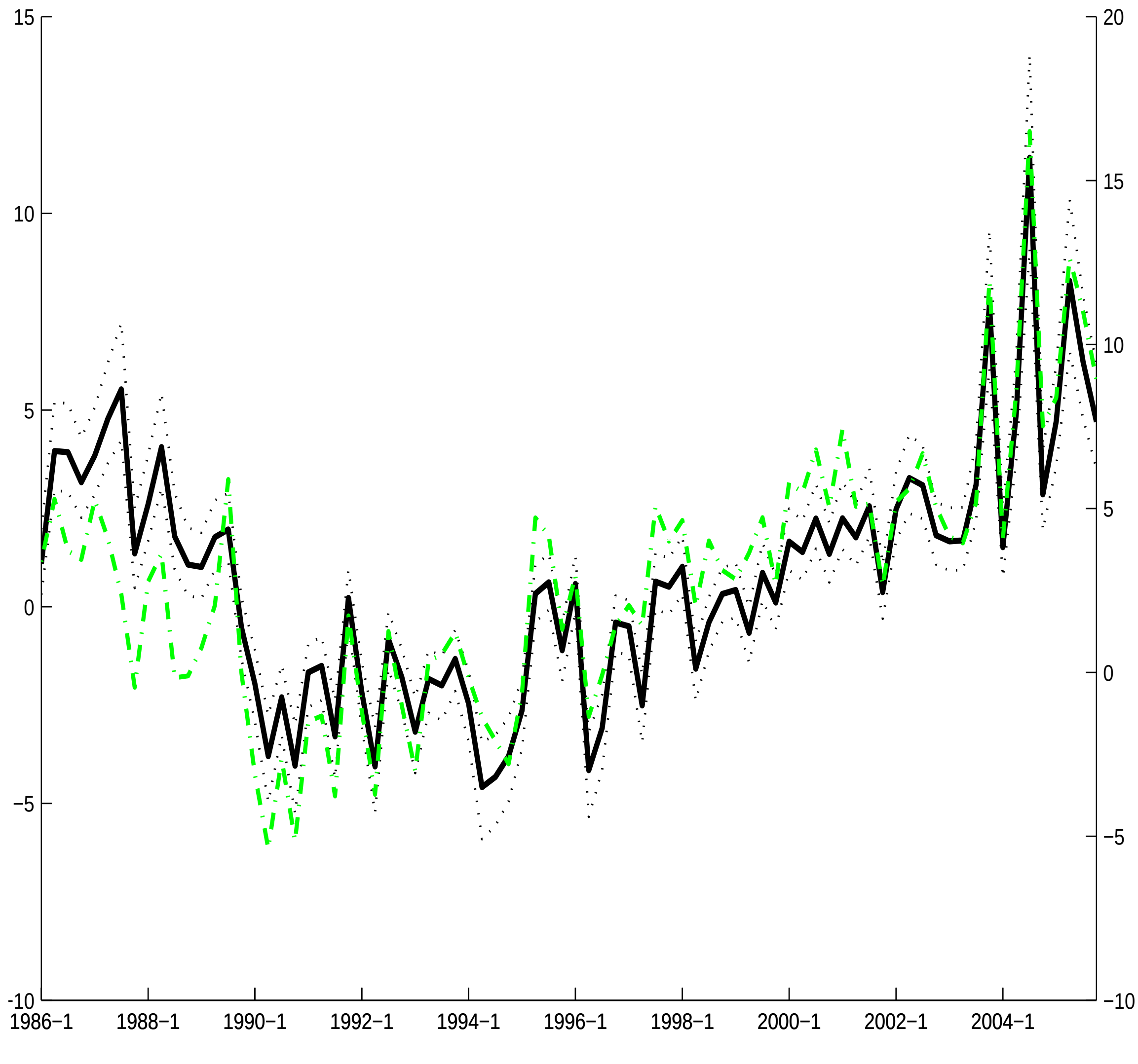
Notes: The top chart shows the growth rates in OFHEO house price indexes for the forty-eight contiguous U.S. states from QI-1986 to QIV-2005. The nominal house price indexes are deflated using the core PCE inflation. The middle chart (national factor) shows the impact of national shocks on state n, that is, the component $\beta_n^0 \ f_t^0$ of equation (1). The bottom chart (local factors) shows the joint impact of regional and state specific shocks, that is, the $\beta_n^r \ f_t^r \ + \ \epsilon_{n,t}$ component of equation (1). For all estimated quantities we show the median of the posterior distribution. See section 3 for a description of the data.

Figure 3: State-level Variance Decompositions



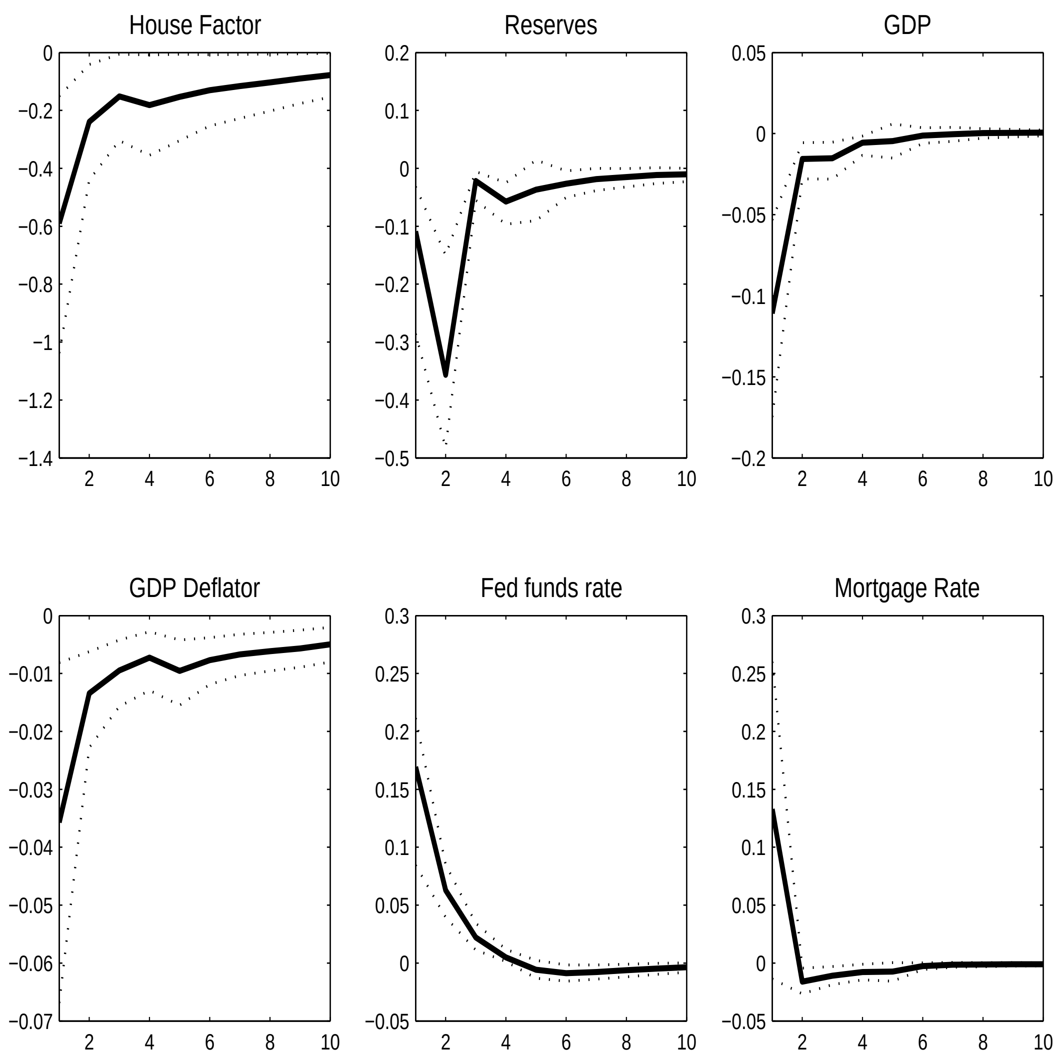
Notes: The top panel shows for each state the variance of fluctuations due to the national factor as a fraction of the variance of all components (referred to in the paper as $v_n(1,T)$, see equation (7)) for housing on the horizontal axis and for income on the vertical axis. The sample period is QI-1986 to QIV-2005. The panel shows the the median of the posterior distribution of $v_n(1,T)$. The bottom panel plots for each state the magnitude $v_n(t_0,t_1)$ for the house price data in the sub-sample 2001-2005 on the vertical axis, and 1986-2000 on the horizontal axis. Again, $v_n(.,.)$ represents the variance of house prices fluctuations in state n due to the national factor as a fraction of the variance of all components. The figure shows the median of the posterior distribution of $v_n(t_0,t_1)$ obtained from estimating the model over the entire sample period.

Figure 4: The National House Price Factor and the U.S. OFHEO House Price Index



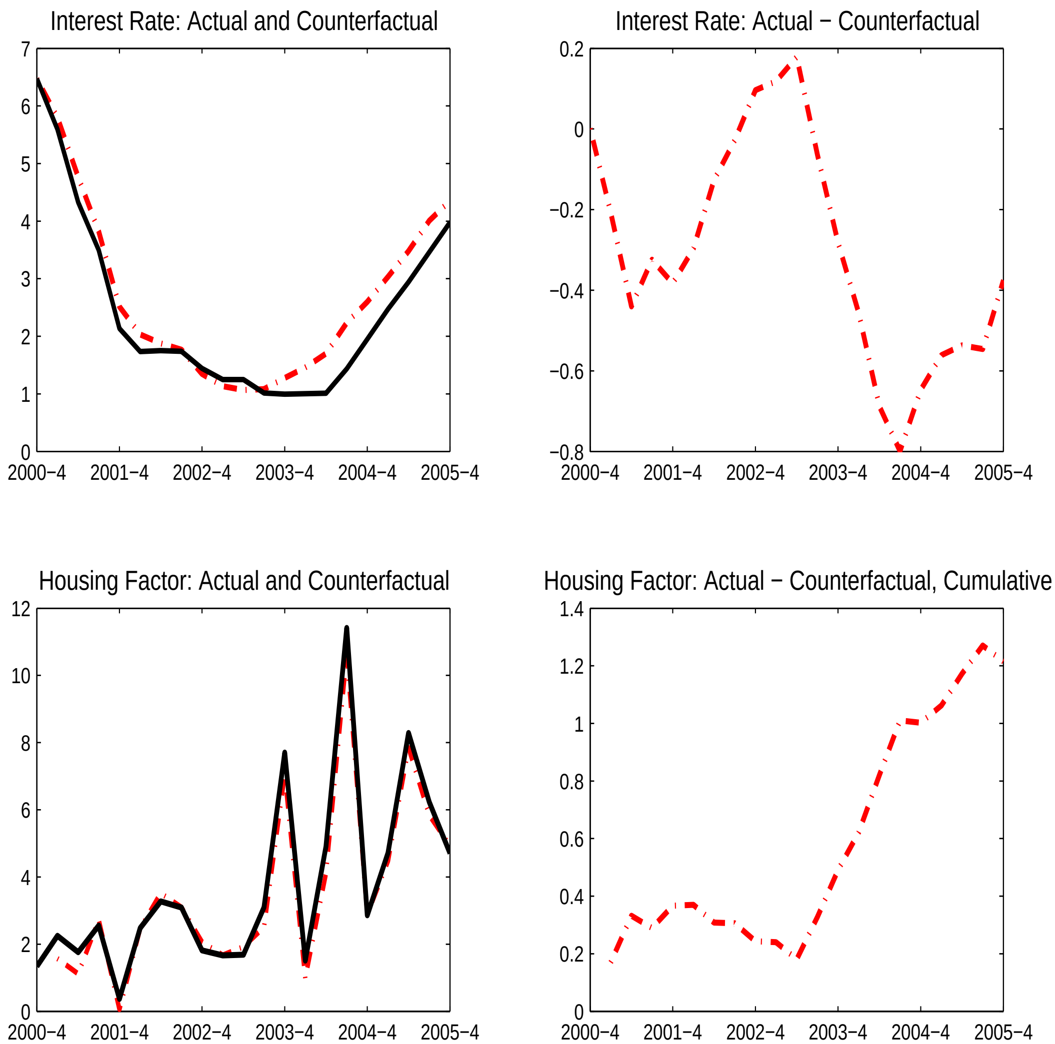
Notes: The figure plots the national housing factor f_t^0 (black, scale on the left axis), the ninety percent bands (dotted lines), as well as the OFHEO U.S. price index (gray, right axis). f_t^0 is the median posterior estimate of the factor from equation (1).

Figure 5: Housing and Monetary Policy: Impulse Responses



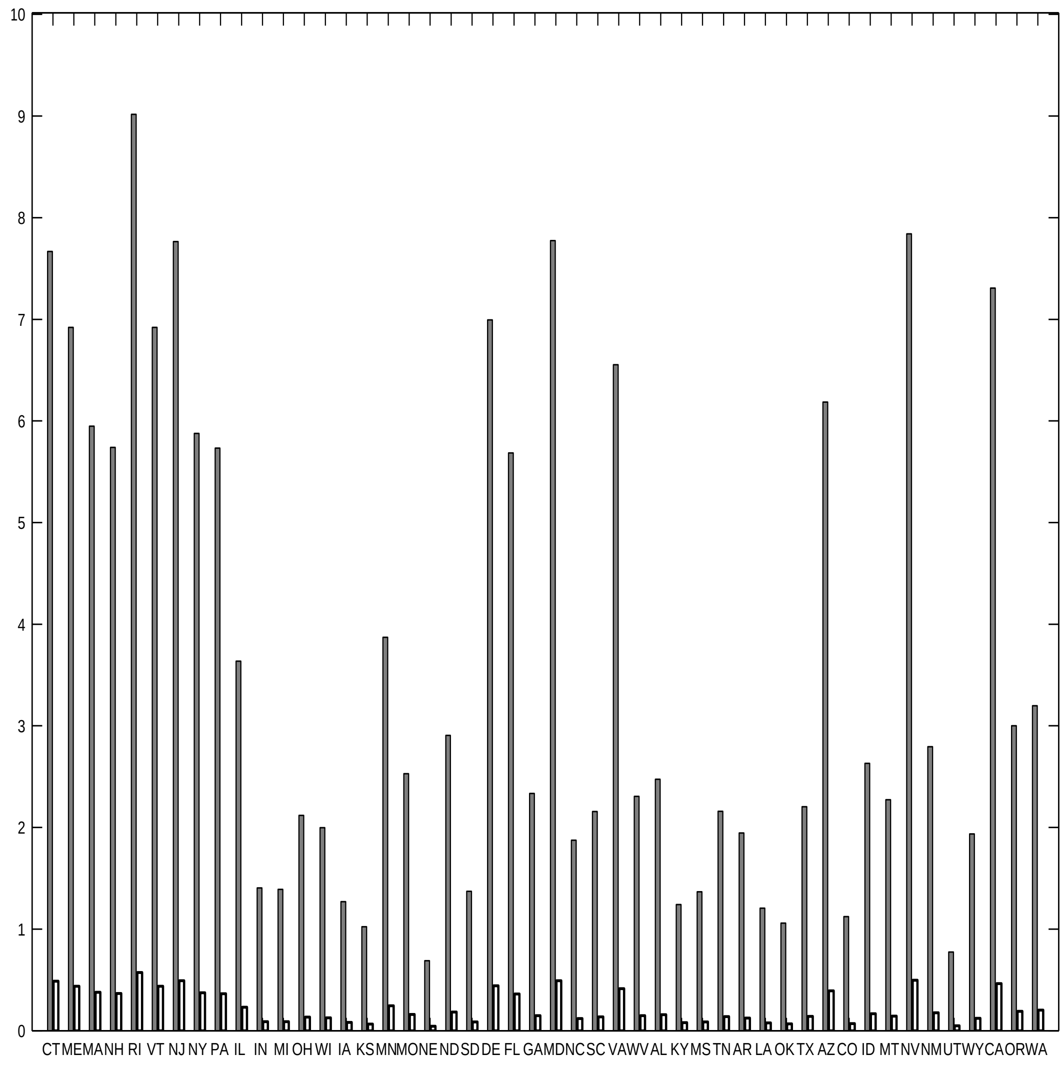
Notes: The figure displays the impulse response function of all of the variables to the monetary policy shock, along with the 68 percent posterior coverage intervals. The VAR is estimated with quarterly data from QI-2006 to QIV-2005.

Figure 6: Actual and Counterfactual Interest Rate and Housing Factor



Notes: The figure shows the actual (solid line) and the counterfactual (dash-and-dotted line) paths for the Fed Funds rate and the housing factor. The counterfactual paths are obtained by shutting down monetary policy shocks from QI-2001 to the end of the sample.

Figure 7: Cross-state Heterogeneity in the Recent House Price Boom: Assessing The Role of Monetary Policy Shocks



Notes: The gray bars in the figure represent the component of the average growth rates in the 2001-1004 period that can be attributed to the national factor, namely the quantities $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}} \beta_n^0 f_t^0/20$, for each of the forty-eight contiguous states. These are the same quantities also shown in Figure 1. The white bars quantify the impact of policy shocks. For each state the figure shows the quantity $\sum_{t=\text{QI-2001}}^{\text{QIV-2005}} \beta_n^0 \ (f_t^0 - \tilde{f}_t^0)/20$, where by $\tilde{f}_t^0$ we denote the counterfactual factor.