The dynamic impact of monetary policy on regional housing prices in the United States

Fischer, Manfred M.; Huber, Florian; Pfarrhofer, Michael; Staufer-Steinnocher, Petra

Published in:
Real Estate Economics

DOI:
10.1111/1540-6229.12274

Published: 01/01/2021

Document Version
Publisher's PDF, also known as Version of record

Citation for published version (APA):
Fischer, M. M., Huber, F., Pfarrhofer, M., & Staufer-Steinnocher, P. (2021). The dynamic impact of monetary policy on regional housing prices in the United States. Real Estate Economics, 49(4), 1039 - 1068. https://doi.org/10.1111/1540-6229.12274
The Dynamic Impact of Monetary Policy on Regional Housing Prices in the United States

Manfred M. Fischer,* Florian Huber,** Michael Pfarrhofer** and Petra Staufer-Steinnocher*

This article uses a factor-augmented vector autoregressive model to examine the impact of monetary policy shocks on housing prices. To simultaneously estimate the model parameters and unobserved factors, we rely on Bayesian estimation and inference. Policy shocks are identified using high-frequency surprises around policy announcements as an external instrument. Impulse response functions reveal differences in regional housing price responses, which in some cases are substantial. The heterogeneity in policy responses is found to be significantly related to local regulatory environments and housing supply elasticities. Moreover, housing prices responses tend to be similar within states and adjacent regions in neighboring states.

Introduction

The housing market is one of the most important, but at the same time most volatile sectors of the economy, and hence of crucial concern for economic policy makers in general, and central banks in particular (Moulton and Wentland 2018). The notion of a national housing market disregards the fact that housing activities substantially vary across the United States. Moench and Ng (2011) emphasize that of the four regions defined by the United States Census Bureau, the West Region (including California, Nevada and Arizona) and the Northeast Region (including New York and Massachusetts) have, from a historical perspective, shown more active housing markets than the Midwest Region (including Illinois, Ohio and Minnesota) and the South

*Vienna University of Economics and Business, Welthandelsplatz 1, A-1020 Vienna or manfred.fischer@wu.ac.at, petra.staufer-steinnocher@wu.ac.at.
**Paris Lodron University of Salzburg, Moenchsberg 2A, A-5020 Salzburg or florian.huber@sbg.ac.at, michael.pfarrhofer@sbg.ac.at.

© 2019 The Authors. Real Estate Economics published by Wiley Periodicals, Inc. on behalf of American Real Estate and Urban Economics Association

This is an open access article under the terms of the Creative Commons Attribution License, which permits use, distribution and reproduction in any medium, provided the original work is properly cited.
Region (including Florida, Texas and North Carolina). Another factor motivating regional disaggregation of the housing market is the volatility of regional housing markets relative to macroeconomic fluctuations (Fratantoni and Schuh 2003).

The literature on the impact of monetary policy–related variables on housing is fairly limited, in particular at the regional level. Previous work generally relies on two competing approaches. The first uses structural models to analyze the relationship (see Iacoviello and Minetti 2003, Iacoviello and Neri 2010, Ungerer 2015, Bahadir and Gumus 2018). The major strength of this model-based approach is to provide a theoretically grounded answer to the question of interest. However, such models necessarily strongly impose a priori restrictions on crucial parameters. The second, evidence-based approach, focuses on empirics and relies less directly on economic theory. Microeconomic event studies, for example, provide answers using information on individual transactions to identify causal effects of monetary policy shocks in short time frames around monetary policy announcements (Moulton and Wentland 2018). Macroeconomists instead typically use vector autoregressive (VAR) models to measure the impact of monetary policy innovations and other macroeconomic shocks over longer time horizons, exploiting information contained in time series data. Examples include Fratantoni and Schuh (2003), Iacoviello (2005), Del Negro and Otrok (2007), Jarocinski and Smets (2008), Iacoviello and Minetti (2008), Vargas-Silva (2008a,b), Moench and Ng (2011) and Choudhry (2018). VAR models are dynamic models of time series that allow the data rather than the researcher, to specify the dynamic structure of the model, and provide a plausible assessment of macroeconomic variables to monetary policy shocks without the need of a fully specified structural model.

This article lies in the tradition of the second approach, and differs from previous work in terms of both its focus and methodology. Like Fratantoni and Schuh (2003) and Del Negro and Otrok (2007), we focus on regional differences in response of housing prices. The coarseness of quarterly state-level observations used in previous research, however, may conceal important variations that is key for researchers to identify cross-regional differences in policy responses. Hence, we use monthly observations on housing prices and provide a comprehensive coverage of the United States at the level of metro- and micropolitan statistical areas,¹ to appropriately identify a monetary policy shock and the associated regional reactions.

¹For the definition of metropolitan and micropolitan statistical areas, see section “Regions and Data” along with Appendix A.
Similar to Vargas-Silva (2008a) and Moench and Ng (2011), we rely on a factor-augmented vector autoregressive (FAVAR) model to identify the impact of a monetary policy shock on housing prices, but use a fully Bayesian FAVAR model, based on a set of macroeconomic and financial variables, to explore regional housing price responses to a national monetary policy shock. In particular, we apply Markov chain Monte Carlo (MCMC) methods to estimate the model parameters and unobserved factors simultaneously, in contrast to previous approaches. Bayesian inference is advantageous because it directly addresses uncertainty surrounding latent factors and model parameters. Policy shocks are identified using high-frequency surprises around policy announcements as external instrument, where policy surprises are measured within a tight window of 30 minutes around the announcements by the Federal Reserve (see Kuttner 2001, Gürkaynak, Sack and Swanson 2005, Gertler and Karadi 2015).

The effects of monetary policy on housing prices in the regions are analyzed using the FAVAR model estimated over the period 1997:04 to 2012:06. Impulse response functions from the estimated model reveal a rich picture about how an expansionary monetary policy shock affects regional housing prices. Differences are evident, and in some cases, substantial. Regions within California, Florida and Nevada are found to be the most sensitive to monetary policy changes, exhibiting effects two times as large as the average response across the country. By contrast, some regions, for example, within Mississippi, Tennessee, Oklahoma and North Carolina are found to be the least responsive, showing no significant impact or even slightly negative responses. By linking the results to the housing supply elasticity literature (Gyourko, Saiz and Summers 2008, Saiz 2010, Howard and Liebersohn 2018, Vinson 2018), this article provides evidence that the measured cumulative cross-regional differential responses can partly be explained by housing supply elasticities and local regulatory environments.

The remainder of the article is organized as follows. The next section presents the FAVAR model along with the Bayesian approach for estimation, and specifics about identification of monetary policy shocks. Section “Data and Model Implementation” describes the data and the sample of regions, and outlines the model specification. The results are presented in section “Econometric Results,” combined with a brief discussion about the question why housing prices in some regions are more sensitive to monetary policy shocks than others. The final section concludes.
Methodology

The FAVAR Model

The econometric approach we employ in this study is a FAVAR model, as introduced by Bernanke, Boivin and Eliasz (2005). In our implementation, we let $H_t$ denote an $R \times 1$ vector of housing prices at time $t$ ($t = 1, \ldots, T$) for $R$ regions. The model postulates that regional housing prices depend on a number of latent factors, monetary and macroeconomic national aggregates and region-specific shocks. This relationship, henceforth termed the measurement equation, can be written as

$$
\begin{bmatrix}
H_t \\
M_t
\end{bmatrix} =
\begin{bmatrix}
\Lambda^F & \Lambda^M \\
0_{K \times S} & I_K
\end{bmatrix}
\begin{bmatrix}
F_t \\
M_t
\end{bmatrix} +
\begin{bmatrix}
\epsilon_t \\
0_{K \times 1}
\end{bmatrix},
$$

(1)

where $F_t$ is an $S \times 1$ vector of latent (unobservable) factors, which capture comovement at the regional level. $M_t$ is a $K \times 1$ vector of economic and monetary national aggregates that are treated as observable factors, and $\epsilon_t$ is an $R \times 1$ vector of normally distributed zero mean disturbances with an $R \times R$ variance–covariance matrix $\Sigma_\epsilon = \text{diag}(\sigma_1^2, \ldots, \sigma_R^2)$. These disturbances arise from measurement errors and special features that are specific to individual regional time series. $\Lambda^F$ is an $R \times S$ matrix of factor loadings, while $\Lambda^M$ denotes a coefficient matrix of dimension $R \times K$. The number of latent factors is much smaller than the number of regions, that is, $S \ll R$. Note that the diagonal structure of $\Sigma_\epsilon$ implies that any comovement between the elements in $H_t$ and $M_t$ stems exclusively from the presence of the latent factors.

The evolution of the factors $y_t = (F'_t, M'_t)'$ is given by the state equation, governed by a VAR process of order $Q$,

$$y_t = Ax_t + u_t,
$$

(2)

with $x_t = (y'_{t-1}, \ldots, y'_{t-Q})'$ and the associated $(S + K) \times Q(S + K)$-dimensional coefficient matrix $A$. Moreover, $u_t$ is an $(S + K)$-dimensional vector of normally distributed shocks, with zero mean and variance–covariance matrix $\Sigma_u$.

The parameters $\Lambda^F$, $\Lambda^M$ and $A$ as well as the latent dynamic factors $F_t$ are unknown and have to be estimated. To econometrically identify the model, we follow Bernanke, Boivin and Eliasz (2005) and assume that the upper $S \times S$-dimensional submatrix of $\Lambda^F$ equals an identity matrix $I_S$ while the first $S$ rows of $\Lambda^M$ are set equal to zero. This identification strategy implies that the first $S$ elements in $H_t$ are effectively the factors plus noise.
A Bayesian Approach to Estimation

The model described above is highly parameterized, containing more parameters as can reasonably be estimated with the data at hand. In this study, we use a Bayesian estimation approach to incorporate knowledge about parameter values via prior distributions. It is convenient to stack the free elements of the factor loadings in an \( L \)-dimensional vector \( \lambda = \text{vec}(\Lambda^F, \Lambda^M) \) with \( L = R(S + K) \), and the VAR coefficients in a \( J \)-dimensional vector \( a = \text{vec}(A) \) with \( J = (S + K)^2 Q \).

Prior distributions for the state equation. For the VAR coefficients \( a_j \) \((j = 1, \ldots, J)\) we impose the normal-gamma shrinkage prior proposed in Griffin and Brown (2010, 2017), and applied in a VAR framework by Huber and Feldkircher (2019),

\[
a_j | \xi_a, \tau^2_{a_j} \sim \mathcal{N}(0, 2 \xi_a^{-1} \tau^2_{a_j}),
\]

that is controlled by gamma priors on \( \tau^2_{a_j} \) \((j = 1, \ldots, J)\) and \( \xi_a \),

\[
\xi_a \sim \mathcal{G}(d_0, d_1),
\]

\[
\tau^2_{a_j} \sim \mathcal{G}(\vartheta_a, \vartheta_a),
\]

with hyperparameters \( d_0, d_1 \) and \( \vartheta_a \), respectively. \( \xi_a \) operates as global shrinkage parameter, and \( \tau^2_{a_j} \) as local scaling parameter. This hierarchical prior shows two convenient features. First, \( \xi_a \) applies to all \( J \) elements in \( a \). Higher values of \( \xi_a \) yield stronger global shrinkage toward the origin whereas smaller values induce only little shrinkage. Second, the local scaling parameters \( \tau^2_{a_j} \) place sufficient prior mass of \( a_j \) away from zero in the presence of strong overall shrinkage involved by large values for \( \xi_a \), in cases where the likelihood suggests nonzero values.

The hyperparameter \( \vartheta_a \) in Equation (5) controls the excess kurtosis of the marginal prior,

\[
p(a_j | \xi_a) = \int p(a_j | \xi_a, \tau^2_{a_j}) d\tau^2_{a_j},
\]

obtained after integrating over the local scales. Lower values of \( \vartheta_a \) generally place increasing mass on zero, but at the same time lead to heavy tails, allowing for large deviations of \( a_j \) from zero, if necessary. The hyperparameters \( d_0 \) and \( d_1 \) in Equation (4) are usually set to rather small values to induce heavy overall shrinkage. See Griffin and Brown (2010) for more details.

For the variance–covariance matrix \( \Sigma_u \) we use an inverted Wishart prior,

\[
\Sigma_u \sim \mathcal{IW}(v, \Sigma),
\]

(7)
with $v$ denoting prior degrees of freedom, while $\Sigma$ is a prior scaling matrix of dimension $(S + K) \times (S + K)$.

Prior distributions for the observation equation. For the factor loadings $\lambda_\ell$ ($\ell = 1, \ldots, L$) we employ a normal-gamma prior similar to the one used for the VAR coefficients in $a$. The set-up follows Kastner (2018) with a single global shrinkage parameter $\xi_\lambda$ that applies to all free elements $\lambda_\ell$ in the factor loadings matrix. Specifically, we impose a hierarchical Gaussian prior on $\lambda_\ell$ that depends on gamma priors for $\tau_{\lambda\ell}^2$ ($\ell = 1, \ldots, L$) and $\xi_\lambda$,

$$
\lambda_\ell | \xi_\lambda, \tau_{\lambda\ell}^2 \sim \mathcal{N}(0, 2 \xi_\lambda^{-1} \tau_{\lambda\ell}^2),
$$

$$
\xi_\lambda \sim \mathcal{G}(c_0, c_1),
$$

$$
\tau_{\lambda\ell}^2 \sim \mathcal{G}(\theta_\lambda, \theta_\lambda).
$$

The hyperparameters $c_0$, $c_1$ and $\theta_\lambda$ control the tail behavior and overall degree of shrinkage of the prior. For the measurement error variances $\sigma_r^2$ ($r = 1, \ldots, R$), we rely on a sequence of independent inverted gamma priors,

$$
\sigma_r^2 \sim \mathcal{G}^{-1}(e_0, e_1),
$$

where the hyperparameters $e_0$ and $e_1$ are typically set to small values to reduce prior influence on $\sigma_r^2$.

Estimation of the model parameters and the latent factors is based on the MCMC algorithm described in Appendix B. More specifically, we use Gibbs sampling to simulate a chain consisting of 20,000 draws, where we discard the first 10,000 draws as burn-in. It is worth noting that the MCMC algorithm shows fast mixing and satisfactory convergence properties.

Identification of Monetary Policy Shocks

The standard approach to identify monetary policy shocks in a VAR framework involves imposing a set of zero restrictions via a Cholesky identification scheme. This approach relies on the assumption that macroeconomic quantities in the system react to changes in the monetary policy instrument with a time lag. Timing restrictions on the impact of the policy indicator may be reasonable for the interactions between the funds rate and macroeconomic variables, but becomes problematic if financial variables are present in addition. Policy shifts not only influence financial quantities, but may also respond to them, directly or indirectly (Gertler and Karadi 2015). To circumvent the problem of simultaneity, we follow Gertler and Karadi (2015) and use high-frequency surprises as external instrument to identify monetary policy shocks.
The high-frequency variant of the external instruments identification approach employed in this article is based on surprises in the prices of three-months-ahead futures contracts of the federal funds rate that reflect expectations on interest rate movements further into the future, measured within a 30 minutes time window surrounding announcements by the Federal Open Market Committee (FOMC), the governing council of the Federal Reserve (Kuttner 2001, Gürkaynak, Sack and Swanson 2005, Gertler and Karadi 2015). The tight time frame around these announcements is chosen to reduce the likelihood of other events affecting prices of the futures contracts.

Financial markets internalize the behavior of the Federal Reserve (Fed) by anticipating changes in the policy instrument based on predicted movements in key macroeconomic quantities. For instance, facing a weakening economic outlook, federal funds rates futures would decline in advance of the policy announcement by the Fed. Depending on the specific monetary policy action conducted by the central bank, futures markets may either correctly predict the enacted policies or react to unexpected changes in the policy rate precisely around official announcements. Gürkaynak, Sack and Swanson (2005) provide evidence that the adjustment of the prices of futures contracts happens almost instantaneously, in contrast to fully anticipated changes that do not cause observable reactions. A convenient by-product of this approach is that it also reflects Fed information shocks in the context of forward guidance.

For illustrative purposes, the evolution of the effective federal funds rate over the observation period 1997:04 to 2012:06 is shown in Figure 1 (upper panel) along with the corresponding policy surprises around announcements (lower panel). The dashed red line refers to the zero line, while the light blue shaded vertical bars represent the recessions dated by the Business Cycle Dating Committee of the National Bureau of Economic Research. Large monetary policy surprises tend to occur in recessionary economic episodes, evidenced by unexpected innovations for both the period between 2001 and 2002, as well as during the Great Recession. Notice that decreases in the federal funds rate not necessarily reflect expansionary shocks. In June 2001, for instance, markets expected the Fed to further decrease the target rate, while the rate was decreased only slightly, translating into a contractionary monetary policy shock. The contrary is observable in the first half of 1997 or June 2006. Here, the Fed left the target rate unchanged while markets expected further increases, resulting in expansionary monetary policy shocks.

To implement the approach, we follow Paul (2018) and use high-frequency surprises as a proxy for the monetary policy shock. This is achieved by
integrating the surprises into Equation (2) as an exogenous variable $z_t$, to yield

$$y_t = Ax_t + \zeta z_t + u_t. \quad (10)$$

Hereby, $\zeta$ is a $Q(S + K)$-dimensional vector of regression coefficients that collects the impulses of the shocks. Paul (2018) shows that under mild conditions, the contemporaneous relative impulse responses can be estimated consistently.\(^2\) Note that the impact response of $y_t$ to changes in $z_t$ is given by $\zeta$. Higher order responses are obtained recursively by exploiting the state space representation of the VAR model in Equation (2).

---

\(^2\)Relative impulse responses are obtained by normalizing the absolute impulse responses, that is, the change in $y_{t+h}$ to a change in $z_t$, by the contemporaneous response of some element in $y_t$. 

---

**Figure 1** The federal funds rate and exogenous monetary policy surprises. [Color figure can be viewed at wileyonlinelibrary.com]
Data and Model Implementation

Regions and Data

To explore regional differences in the impact of monetary policy on housing prices, we need to define our notion of regions. Throughout the article, we use $R = 417$ regions, a subsample of the 917 core-based statistical areas (CBSAs). These 417 regions include 263 metropolitan and 154 micropolitan statistical areas. They have been selected based on the availability of the data over time. For the list of regions in the sample, see Appendix A.

Our data set consists of a panel of monthly time series ranging from 1997:04 to 2012:06. The $R \times 1$ vector of housing prices $H_t$ is constructed using the Zillow Home Value Index. A key advantage of this index is to provide a comprehensive coverage of CBSAs across the country, in contrast to the Federal Housing Finance Agency (FHFA) Index and the Standard & Poor’s Case-Shiller Index. The Zillow Home Value Index does not use a repeat sales methodology, but statistical models along with information from sales assessments to generate valuations for all homes (single family houses, town houses, apartments, condos and properties that are typically associated with the residential market) in any given region. These valuations are aggregated to determine the Zillow Home Value Index, measured in U.S. dollars.

We include $K = 7$ variables in the vector of observable national aggregates $M_t$: three economic variables, namely, housing investment (measured in terms of housing starts), the industrial production index and the consumer price index. The one-year government bond rate serves as policy indicator of the Fed. The advantage of using this longer rate rather than the federal funds
rate is that it incorporates—as Gertler and Karadi (2015) argue—measures of forward guidance and hence remains a valid measure of the monetary policy stance also in situations when the federal funds rate is constrained by the zero lower bound.\footnote{To support this view, we estimated the model using the federal funds rate as policy indicator for a robustness check. The results—available upon request—suggest similar responses compared to the one-year government bond rate.}

The FAVAR model developed in this article extends a standard macroeconomic autoregressive model with a set of three credit-spreads: the 10-year treasury yield minus the federal funds rate, the prime mortgage spread calculated over 10-year government bond yields and the Gilchrist and Zakrajšek (2012) excess bond premium. The excess bond premium may roughly be seen as the component of the spread between an index of yields on corporate fixed income securities and a similar maturity government bond rate that is left after removing the component due to default risk (Gertler and Karadi 2015). Gilchrist and Zakrajšek (2012) show that this variable provides a convenient summary of additional information that may be relevant to economic activity.

The economic variables capture housing, price and output movements. The mortgage spread is relevant to the cost of housing finance, and the excess bond premium to the cost of long-term credit in the business sector, while the term spread measures expectations on short-term interest rates (Gertler and Karadi 2015). All observable national aggregates are taken from the FRED database (McCracken and Ng 2016), with the exception of the excess bond premium and the mortgage spread that have been obtained from the data set provided in Gertler and Karadi (2015). All data series are seasonally adjusted, if applicable, and transformed to be approximately stationary.

**Model Implementation**

For implementation of the FAVAR, we have to specify the lag order $Q$ of the VAR process and the number of latent factors, $S$. As is standard in the literature, we pick $Q = 2$ lags of the endogenous variables. To decide on the number of factors, we use the deviance information criterion (Spiegelhalter et al. 2002) where the full data likelihood is obtained by running the Kalman filter and integrating out the latent states. This procedure yields $S = 1$, a choice that is also consistent with traditional criteria, for instance, the Bayesian information criterion or the Kaiser criterion, for selecting the number of factors.

A brief word on hyperparameter selection for the prior setup is in order. We specify $\vartheta_a = \vartheta_\lambda = 0.1$, a choice that yields strong shrinkage but, at the same time, allows for a flexible model formulation.
time, leads to heavy tails in the underlying marginal prior. Recent literature
(see, for example, Huber and Feldkircher 2019) integrates out $\vartheta_a, \vartheta_\lambda$, and
finds that, for U.S. data, the posterior is centered on values between 0.10
and 0.15. The hyperparameters on the global shrinkage parameters are set
equal to $c_0 = c_1 = d_0 = d_1 = 0.01$, a choice that is consistent with heavy
shrinkage toward the origin representing a standard in the literature (Griffin
and Brown 2010). The prior on $\Sigma_u$ is specified to be weakly informative,
\textit{i.e.}, $\nu = S + K + 1$ and $\Sigma = 10^{-2}I_{S+K}$. Similarly for the inverted gamma
prior on $\sigma_r^2 (r = 1, \ldots, R)$ we set $e_0 = e_1 = 0.01$ to render the prior only
weakly influential.

**Econometric Results**

**The Dynamic Factor and Its Loadings**

We briefly consider the estimated latent factor and its loadings, with two aims
in mind: first, to provide a rough intuition on how the latent factor captures
covariation in regional house price variations, and second, to indicate the
relative importance of individual regions shaping the evolution of the common
factor. The posterior mean of the negative latent factor (in solid red) shown in
Figure 2 provides evidence that the common factor co-moves with the average
growth rate of housing prices (in solid blue, calculated using the arithmetic
mean of the individual regional housing prices) nearly perfectly. The figure
illustrates that during the 2001 recession, housing price declines have been
mild, while being substantial during the Great Recession, with large variations

![Figure 2](https://wileyonlinelibrary.com)
across space. It is worth noting that home prices fell the most during the late 2000s in regions with the largest declines in economic activity (Beraja et al. 2017).

While Figure 2 provides intuition on the shape of the latent housing factor, the question on how individual regions are linked to it still needs to be addressed. For this purpose, Figure 3 reports the posterior mean of the region-specific factor loadings in the form of a geographic map in which thinner lines denote the boundaries of the regions, while thicker lines signify U.S. state boundaries. Visualization is based on a classification scheme with equal-interval breaks. We see that the great majority of regions exhibit negative loadings, and only 23 regions show positive values. Eighty regions have zero loadings or loadings where the 16th and 84th credible sets (68% posterior coverage) of the respective posterior distributions include zero. The pattern of factor loadings, evidenced by the map, indicates that the latent factor is largely
driven by regions located in California, Arizona and Florida. Regions in the rest of the country, with loadings being either small in absolute terms or not significantly different from zero, tend to play only a minor role in shaping national housing prices.

*Impulse Responses of Macroeconomic Quantities*

Impulse response functions represent the standard way to summarize the dynamic impact of policy shocks. We first consider the dynamic evolution of the endogenous variables included in $M_t$ in response to a monetary policy shock to illustrate that the results of the model are consistent with established findings in the literature. An expansionary monetary policy shock is modeled by taking the one-year government bond rate as the relevant policy indicator, rather than the federal funds rate that is commonly used in the literature. Gertler and Karadi (2015) show that the one-year bond rate has a stronger impact on market interests than the funds rate does, based on the assertion that forward guidance is more adequately reflected in the longer maturity yield. Normalization is achieved by assuming that a monetary policy shock yields a five basis points decrease in the policy indicator.

The impulse response functions of all the endogenous variables to the monetary policy shock are presented in Figure 4. All plots include the median response (in blue) for 72 months after impact along with 68% posterior coverage intervals reflecting posterior uncertainty. An unanticipated decrease in the government bond rate by five basis points causes a significant increase in real activity, with industrial production, housing investment and consumer prices all increasing over the next months after the impact. From a quantitative standpoint, the effects of the monetary shock on industrial production and consumer prices are considerably larger than the impact on housing investment, although uncertainty surrounding the size of impacts is large and posterior coverage intervals include zero during the first months after impact. Housing investment shows a reaction similar in shape to real activity measured in terms of the industrial production index, suggesting a positive relationship between expansionary monetary policy and housing investment at the national level.

Turning to the responses of financial market indicators, it should be noted that the one-year government bond rate falls by five basis points on impact by construction, then increases significantly before it turns nonsignificant after about nine months. The term spread reacts adversely on impact, and we find significant deviations from zero that die out after about 16 months. This result points toward an imperfect pass-through of monetary policy on long-term rates, implying that long-term yields display a weaker decline as compared
to short-term rates. The prime mortgage spread does not show a significant effect on impact, while responses between 10 and 20 months ahead indicate a slightly negative overall reaction to expansionary monetary policy. Consistent with Gilchrist and Zakrjašek (2012), one implication of this finding is that movements in key short-term interest rates tend to impact credit markets, with mortgage spreads showing a tendency to decline. The responses of the excess bond premium almost perfectly mirror the reaction of the mortgage spread. The effects, however, are much larger from a quantitative point of view.

**Notes:** The solid blue line denotes the median response, the dashed red line the zero line, and the shaded bands (in light blue) the 68% posterior coverage interval. Results are based on 10,000 posterior draws. Sample period: 1997:04 to 2012:06. Vertical axis: percentage changes for indices and housing investment; otherwise percentage points. Front axis: months after impact.
To sum up, the results obtained by the impulse response analysis provide empirical support that monetary policy shocks, identified by using high-frequency surprises around policy announcements as external instrument, generate impulse responses of the endogenous variables that are consistent with economic theory and the findings of previous empirical studies.\(^6\)

**Impulse Responses of Housing Prices**

Figure 5 displays the impulse response function of the latent factor over 72 months after impact to an expansionary monetary policy shock. The latent factor reacts positively after the shock, but the posterior coverage interval includes zero for the time horizon considered. Nevertheless, sufficient posterior mass is shifted away from zero reflecting positive reactions. This is consistent with economic theory, suggesting decreases in the cost of financing a home purchase via expanding the availability of credit, thereby increasing the demand for housing. As a result, real housing prices tend to increase.

**Figure 5** Reaction of the negative latent factor, following a monetary policy shock. [Color figure can be viewed at wileyonlinelibrary.com]

Notes: The solid blue line denotes the median response, the dashed red line the zero line and the shaded bands (in light blue) the 68% posterior coverage interval. Results are based on 10,000 posterior draws. Vertical axis: percentage points. Front axis: months after impact.

\(^6\)To allay potential concerns of the policy rate reaching the zero lower bound, we conducted various robustness checks (results are available upon request). Using a shadow rate to capture unconventional monetary policy actions leaves the results qualitatively unchanged. The same holds true when using the federal funds rate rather than the one-year government bond rate as policy indicator.
Figure 6 ■ Long-run responses of regional housing prices to a monetary policy shock, i.e., cumulative 72 months responses. [Color figure can be viewed at wileyonlinelibrary.com]

Notes: Visualization is based on a classification scheme that generates breaks in standard deviation measures ($SD = 0.61$) above or below the mean of 0.71. The number of regions is allocated to the classes in squared brackets. The responses based on 10,000 posterior draws have been accumulated. Thinner lines denote the boundaries of the regions, while thicker lines represent U.S. state boundaries. Sample period: 1997:04 to 2012:06. For the list of regions, see Appendix A.

While for reasons of space, we do not report the housing price responses of all the 417 regions, we summarize the long-run regional house price responses (i.e., cumulative 72 months responses, expressed in percentage points) in the form of a geographic map with a classification scheme that generates class breaks in standard deviation measures ($SD = 0.61$) above and below the mean of 0.71 (see Figure 6). Again thinner lines denote the boundaries of the regions and thicker lines those of the U.S. states. Some few regions show no significant impact or even negative responses. In more than 91% of the regions, however, the cumulative response of housing prices is positive.

Monetary policy shocks affect regions asymmetrically. Differences in policy responses are evident, and in some cases, substantial. The largest

---

The results are robust to an alternative identification scheme based on sign-restrictions (see Appendix C). Concerns on the validity of the identification scheme using external instruments, that may come from the period where interest rates nearly reached zero, are thus alleviated.
response among regions (Las Vegas–Henderson–Paradise, Nevada) exceeds the smallest (Tahlequah, Oklahoma) by 3.2 percentage points. Regions within California, Florida and Nevada—commonly referred to as Sand States—are noticeably more responsive to monetary policy changes. The top 10 most responsive regions are dominated by six Californian regions: Riverside–San Bernadino–Ontario, Madera, Merced, Clearlake, Modesto and Bakersfield. The first two slots in the ranking, however, are occupied by Las Vegas–Henderson–Paradise and Fernley, a micropolitan region (both within Nevada). Port St. Lucie, Clewiston and Key West (all Florida) round out the top 10, and bring the coastal eastern regions into picture.

By contrast, certain regions of the country are much less sensitive to monetary policy shocks (±0.25 standard deviation from zero). These regions are not concentrated in only a few states or areas. Rather, they span 14 states and 24 metro- and micropolitan regions. Given the narrowness of our definition, this emphasizes the point that less responsive regions, in terms of reactions to monetary policy shocks, are spread throughout much of the country. Clarksville (Tennessee–Kentucky), Tulsa, Enid and Bartlesville (Oklahoma), as well as Hickory–Lenoir–Morgenton and Fayetteville (North Carolina), and Baton Rouge (Louisiana) are found to be the least responsive. Note that 5% of the regions do not show significant results, while 3.6% (including, e.g., Salt Lake City, Utah) exhibit negative responses.

Metropolitan regions like Chicago–Naperville–Elgin (Illinois–Indiana–Wisconsin), Boston–Cambridge–Newton (Massachusetts–New Hampshire), Portland–Vancouver–Willsborough (Oregon–Washington), Savannah (Georgia) and San Jose–Sunnyvale–Santa Clara (California) respond to monetary policy changes in ways that closely mirror the average dynamic response across the United States (±0.25 standard deviation).

Figure 6 reveals substantial heterogeneity in the magnitude of the dynamic responses, but also indicates that regional responses tend to be similar within states and adjacent neighboring states. This spatial autocorrelation phenomenon becomes particularly evident in the case of Californian regions and is most likely due to the importance of new house construction industries, along with the spatial influence the Californian housing market has on regions in neighboring states, especially Nevada and Arizona.

**Explanation for the Differential Housing Price Responses**

Housing price responses vary substantially over space, with size and modest sign differences among the regions, as evidenced by Figure 6. This raises the question why housing prices in some regions are more responsive to monetary

---

The Dynamic Impact of Monetary Policy on Regional Housing Prices
policy shocks than in others. To address this issue, we link our results to the housing supply elasticity literature (Gyourko, Saiz and Summers 2008, Saiz 2010, Howard and Liebersohn 2018), more specifically, to local land use regulation as captured by the Wharton Residential Land Use Regulatory Index (WRLURI), and a measure of housing supply elasticity developed by Howard and Liebersohn (2018).

The WRLURI created by Gyourko, Saiz and Summers (2008) is an index comprised of 11 subindices that summarize information on different aspects of the local regulatory environment. The index calculated for our regions shows that much heterogeneity in land use regulatory environments exists across the regions. The two Michigan metropolitan regions, Ann Arbor and Jackson, and the Michigan micropolitan region, Adrian, represent the most heavily regulated markets, with WRLURI scores at least 2.9 standard deviation above the national mean of −0.18. The next most heavily regulated regions, according to the index, are Seattle–Tacoma–Bellevue (Washington), San Diego–Carlsbad and San Francisco–Oakland–Hayward (both within California), being about one standard deviation above the mean. Dallas–Fort Worth–Arlington (Texas) is a typical housing market near the mean in terms of land use regulatory environments. Bartlesville (Oklahoma), Lewiston (Idaho–Washington), Toledo (Ohio) and Tahlequah (Oklahoma) are examples for the least regulated regions, having WRLURI scores that are at least one standard deviation below the mean. All these examples emphasize that local land use regulation is neither uniformly high nor uniform across the country.

Figure 7 presents the estimated local land use regulation in form of a geographic map with a classification scheme that generates class breaks in standard deviation (0.82) measures above and below the mean of −0.18 (left panel), while the comparison with the corresponding cumulative impulse responses of housing prices is shown in the right panel. The figure clearly suggests that there exists a positive relationship between the sensitivity of housing price reactions and land use regulation. Regions characterized by tight regulations also tend to feature strong reactions of local housing markets. This can be attributed to the positive relationship between regulatory measures and housing prices that has previously been identified in the literature (see, for instance, Ihlanfeldt 2007, Glaeser and Ward 2009). We conjecture that this relationship directly translates into increased responsiveness of housing prices, leading to stronger reactions to national monetary policy shocks.

8We calculated WRLURI scores for the regions by taking a population-weighted average for all counties within a region. Note that the WRLURI does not provide full coverage for the United States, and systematically undercovers micropolitan regions. Hence, scores are missing for 15% of the regions in our sample.
Figure 7: Estimated local land use regulation, captured by WRLURI, and comparison with the housing price responses. [Color figure can be viewed at wileyonlinelibrary.com]

Notes: The left-hand panel presents the estimated local land use regulation (visualized in form of a geographic map based on a classification scheme with equal-interval breaks around the mean of −0.18), while the right panel shows the correlation between cumulative impulse responses of housing prices and the corresponding WRLURI scores (the circles represent the regions, with their size indicating population density). Thinner lines denote the boundaries of the regions, while thicker lines represent U.S. state boundaries. The solid red line in the right-hand panel denotes the correlation. For the list of regions see Appendix A.

In the next step, we assess how housing supply elasticity is linked to housing price responses in Figure 8. We use Howard and Liebersohn’s housing supply elasticity measure for commuting zones to construct elasticities for the regions. The elasticity measure estimates the effect of a change in housing units on housing prices, projecting this relationship onto three measures associated with land availability: the WRLURI index, population density and the coastal status (Howard and Liebersohn 2018). Estimated housing supply elasticities for the regions reveal that San Francisco–Oakland–Hayward, San Diego–Carlsbad, Santa Rosa, Napa and Vallejo–Fairfield (all within California) belong to the top 10 most inelastic regions, with elasticities below 0.72. The three Michigan regions Ann Arbor, Adrian and Jackson along with Philadelphia–Camden–Wilmington (Pennsylvania–New Jersey–Delaware–Maryland) and Trenton (New Jersey) complete the top 10 list. Housing supply is estimated to be quite elastic (3.73, with a standard deviation of 2.78) for the average region, represented by Salt Lake City (Utah). By contrast, Las Vegas–Henderson–Paradise (Nevada) and Santa Fe

9To classify coastal regions, we used NOAA’s (National Oceanic and Atmospheric Administration) definition of coastal counties. Any region that contains a coastal county was coded as coastal.
Figure 8  ■ Estimated elasticities and comparison with the housing price responses. [Color figure can be viewed at wileyonlinelibrary.com]

Notes: The left-hand panel presents the estimated housing supply elasticities (visualized in form of a geographic map based on a classification scheme with equal-interval breaks around the mean of 3.73), while the right panel shows the correlation between cumulative impulse responses of housing prices and the corresponding elasticities (the circles represent the regions, with their size indicating population density). Thinner lines denote the boundaries of the regions, while thicker lines represent U.S. state boundaries. The solid red line in the right-hand panel denotes the correlation. For the list of regions see Appendix A.

(New Mexico) stand out as prominent examples with most elastic housing supply, with at least one standard deviation above the national mean.

On the right in Figure 8, one observes a negative relationship between housing supply elasticities and price responses. In our specific example, we find that expansionary monetary policy directly translates into cheaper credit, leading to upward movements in housing demand. This increase in housing demand in face of a rather steep supply curve for housing yields a strong price reaction. This finding corroborates and extends the results in Glaesera, Gyourko and Saiz (2008), who report a negative relationship between supply elasticities and movements in property prices, especially in the context of excessive increases in housing prices. These results indicate higher effectiveness of monetary policy to influence housing prices by the central bank in regions characterized by low levels of supply elasticities.

Closing Remarks

This article uses a Bayesian FAVAR model to examine the impact of monetary policy shocks on housing prices across the United States. Bayesian inference is advantageous because it directly addresses uncertainty surrounding latent factors and model parameters. Monetary policy shocks are identified making
use of high-frequency surprises around policy announcements as external instrument. Impulse response functions reveal that monetary policy shocks affect regions asymmetrically. There is substantial heterogeneity in the magnitude of the regional housing price responses. The largest response exceeds the smallest by 3.2 percentage points. Regions within California, Florida and Nevada are noticeably more responsive than others. By contrast, the least responsive regions are spread throughout much of the country.

This heterogeneity in responses may be due to varying sensitivity of housing to interest rates across space, and regional differences in housing markets such as different local regulatory environments and supply elasticities. The article links the results to the housing supply elasticity literature and provides evidence that the variation in housing responses across space can be explained partly by different supply elasticities and regulatory environments.

Finally, it is worth noting that our analysis is confined to a linear setting, implying the underlying transmission mechanism to be constant over time. This assumption simplifies the analysis, but may be overly simplistic in turbulent economic times such as the collapse of the housing market around the Great Recession. Hence, an extension of the linear setting to allow for nonlinearities—in the spirit of Huber and Fischer (2018)—might be a promising avenue for future research.

The authors thank the Vienna University of Economics and Business for research support and gratefully acknowledge funding by the Austrian National Bank, Jubilaeumsfond Grant No. 17650. Our specific gratitude goes to Greg Howard (University of Illinois) for providing commuting zone-based data for calculating the regional housing supply elasticities and WRLURI scores. This article is a substantially revised version of a paper circulated under the title “The dynamic impact of monetary policy on regional housing prices in the US: Evidence based on factor-augmented vector autoregressions,” Working Paper Series in Regional Science No. 2018-1, WU Vienna University of Economics and Business, Vienna.

References

Arias J., J. Rubio-Ramirez and D. Waggoner. 2014. Inference Based on SVAR Identified with Sign and Zero Restrictions: Theory and Applications. Working Paper Series No. 1, Federal Reserve Bank of Atlanta.

Bahadir B. and I. Gumus. 2018. Transmission of Household and Business Credit Shocks in Emerging Markets: The Role of Real Estate. Real Estate Economics. First published online, https://doi.org/10.1111/1540-6229.12273

Beraja M., A. Fuster, E. Hurst and J. Vavra. 2017. Regional Heterogeneity and Monetary Policy. Staff Report No. 731, Federal Reserve Bank of New York.
Bernanke B.S., J. Boivin and P. Eliasz. 2005. Measuring the Effects of Monetary Policy: A Factor-Augmented Vector Autoregressive (FAVAR) Approach. The Quarterly Journal of Economics 120(1): 387–422.

Carter C.K. and R. Kohn. 1994. On Gibbs Sampling for State Space Models. Biometrika 81(3): 541–553.

Choudhry T. 2018. Economic Policy Uncertainty and House Prices: Evidence from Geographical Regions of England and Wales. Real Estate Economics. First published online, https://doi.org/10.1111/1540-6229.12266

Dedola L. and S. Neri. 2007. What Does a Technology Shock Do? A VAR Analysis with Model-Based Sign Restrictions. Journal of Monetary Economics 54(2): 512–549.

Del Negro M. and C. Otrok. 2007. 99 Luftballons: Monetary Policy and the House Price Boom across U.S. States. Journal of Monetary Economics 54(7): 1962–1985.

Fratantoni M. and S. Schuh. 2003. Monetary Policy, Housing, and Heterogeneous Regional Markets. Journal of Money, Credit, and Banking 34(4): 557–589.

Frühwirth-Schnatter S. 1994. Data Augmentation and Dynamic Linear Models. Journal of Time Series Analysis 15(2): 183–202.

George E.I., D. Sun and S. Ni. 2008. Bayesian Stochastic Search for VAR Model Restrictions. Journal of Econometrics 142(1): 553–580.

Gertler M. and P. Karadi. 2015. Monetary Policy Surprises, Credit Costs, and Economic Activity. American Economic Journal: Macroeconomics 7(1): 44–76.

Gilchrist S. and E. Zakrjšek. 2012. Credit Spreads and Business Cycle Fluctuations. The American Economic Review 102(4): 1692–1720.

Glaeser E.L. and B.A. Ward. 2009. The Causes and Consequences of Land Use Regulation: Evidence from Greater Boston. Journal of Urban Economics 65(3): 265–278.

———, J. Gyourko and A. Saiz. 2008. Housing Supply and Housing Bubbles. Journal of Urban Economics 64(2): 198–217.

Griffin J.E. and P.J. Brown. 2010. Inference with Normal-Gamma Prior Distributions in Regression Problems. Bayesian Analysis 5(1): 171–188.

———. 2017. Hierarchical Shrinkage Priors for Regression Models. Bayesian Analysis 12(1): 135–159.

Gürkaynak R.S., B. Sack and E. Swanson. 2005. The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models. The American Economic Review 95(1): 425–436.

Gyourko J., A. Saiz and A. Summers. 2008. A New Measure of the Local Regulatory Environment for Housing Markets: The Wharton Residential Land Use Regulatory Index. Urban Studies 45(3): 693–729.

Howard G. and J. Liebersohn. 2018. The Geography Channel of House Price Appreciation. Fisher College of Business Working Paper No. 2018(17).

Huber F. and M. Feldkircher. 2019. Adaptive Shrinkage in Bayesian Vector Autoregressive Models. Journal of Business & Economic Statistics 37(1): 27–39.

——— and M.M. Fischer. 2018. A Markov Switching Factor-Augmented VAR Model for Analyzing US Business Cycles and Monetary Policy. Oxford Bulletin of Economics and Statistics 80(3): 575–604.

Iacoviello M. 2005. House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle. The American Economic Review 95(3): 739–764.

——— and R. Minetti. 2003. Financial Liberalization and the Sensitivity of House Prices to Monetary Policy: Theory and Evidence. The Manchester School 71(1): 20–34.
______ and R. Minetti. 2008. The Credit Channel of Monetary Policy: Evidence from the Housing Market. *Journal of Macroeconomics* 30(1): 69–96.

______ and S. Neri. 2010. Housing Market Spillovers: Evidence from an Estimated DSGE Model. *American Economic Journal: Macroeconomics* 2(2): 125–164.

Ihlanfeldt K.R. 2007. The Effect of Land Use Regulation on Housing and Land Prices. *Journal of Urban Economics* 61(3): 420–435.

Jarocinski M. and F. Smets. 2008. House Prices and the Stance of Monetary Policy. ECB Working Paper 891, European Central Bank, Frankfurt.

Kastner G. 2018. Sparse Bayesian Time-Varying Covariance Estimation in Many Dimensions. *Journal of Econometrics*. First published online, https://doi.org/10.1016/j.jeconom.2018.11.007

Kuttner K.N. 2001. Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market. *Journal of Monetary Economics* 47(3): 523–544.

McCracken M.W. and S. Ng. 2016. FRED-MD: A Monthly Database for Macroeconomic Research. *Journal of Business & Economic Statistics* 34(4): 574–589.

Moench E. and S. Ng. 2011. A Hierarchical Factor Analysis of U.S. Housing Market Dynamics. *The Econometrics Journal* 14(1), C1–C24.

Moulton J.G. and S.A. Wentland. 2018. Monetary Policy and the Housing Market: Evidence from National Microdata. Paper presented at the 2018-SF Fed-UCLA Conference on Housing Financial Markets and Monetary Policy.

Paul P. 2018. The Time-Varying Effect of Monetary Policy on Asset Prices. Working Paper 2017-09, Federal Reserve Bank of San Francisco.

Rubio-Ramirez J.F., D.F. Waggoner and T. Zha. 2010. Structural Vector Autoregressions: Theory of Identification and Algorithms for Inference. *The Review of Economic Studies* 77(2): 665–696.

Saiz A. 2010. The Geographic Determinants of Housing Supply. *The Quarterly Journal of Economics* 125(3): 1253–1296.

Spiegelhalter D.J., N.G. Best, B.P. Carlin and A. van der Linde. 2002. Bayesian Measures of Model Complexity and Fit. *Journal of the Royal Statistical Society: Series B* 64(4): 583–639.

Uhlig H. 2005. What Are the Effects of Monetary Policy on Output? Results from an Agnostic Identification Procedure. *Journal of Monetary Economics* 52(2): 381–419.

Ungerer C. 2015. Monetary Policy, Hot Housing Markets and Leverage. Finance and Economics Discussion Series 2015-048, Divisions of Research & Statistics and Monetary Affairs, Federal Reserve Board, Washington, D.C.

Vargas-Silva C. 2008a. The Effect of Monetary Policy on Housing: A Factor-Augmented Vector Autoregression (FAVAR) Approach. *Applied Economics Letters* 15(10): 749–752.

______. 2008b. Monetary Policy and the US Housing Market: A VAR Analysis Imposing Sign Restrictions. *Journal of Macroeconomics* 30(3): 977–990.

Vinson P. 2018. House Prices and Consumption in the United States. *Real Estate Economics*. First published online, https://doi.org/10.1111/1540-6229.12271

**Appendix A: Regions Used in the Study**

Regions in this study are defined as CBSAs that—by definition of the United States Office of Management and Budget—are based on the concept of a core area of at least 10,000 population, plus adjacent counties having at least 25% of employed residents of the county who work in the core area. CBSAs
### Table A1  ■ The list of metropolitan statistical areas used.

| State (Census Bureau Region) | Region |
|------------------------------|--------|
| Alabama (South)              | Birmingham–Hoover, Daphne–Fairhope–Foley, Mobile, Montgomery, Tuscaloosa |
| Arizona (West)               | Flagstaff, Lake Havasu City–Kingman, Phoenix–Mesa–Scottsdale, Prescott, Sierra Vista–Douglas, Tucson, Yuma |
| Arkansas (South)             | Fayetteville–Springdale–Rogers*, Fort Smith*, Hot Springs, Jonesboro, Little Rock–North Little Rock–Conway |
| California (West)            | Bakersfield, Chico, El Centro, Fresno, Hanford–Corcoran, Los Angeles–Long Beach–Anaheim, Madera, Merced, Modesto, Napa, Oxnard–Thousand Oaks–Ventura, Redding, Riverside–San Bernardino–Ontario, Sacramento–Roseville–Arden–Arcade, Salinas, San Diego–Carlsbad, San Francisco–Oakland–Hayward, San Jose–Sunnyvale–Santa Clara, San Luis Obispo–Paso Robles–Arroyo Grande, Santa Cruz–Watsonville, Santa Maria–Santa Barbara, Santa Rosa, Stockton–Lodi, Vallejo–Fairfield, Visalia–Porterville, Yuba City |
| Colorado (West)              | Boulder, Colorado Springs, Denver–Aurora–Lakewood, Fort Collins, Grand Junction, Greeley, Pueblo |
| Connecticut (Northeast)      | Bridgeport–Stamford–Norwalk, Hartford–West Hartford–East Hartford, New Haven–Milford, Norwich–New London |
| Delaware (South)             | Dover |
| District of Columbia (South) | Washington–Arlington–Alexandria* |
| Florida (South)              | Cape Coral–Fort Myers, Crestview–Fort Walton Beach–Destin, Deltona–Daytona Beach–Ormond Beach, Gainesville, Homosassa Springs, Jacksonville, Lakeland–Winter Haven, Miami–Fort Lauderdale–West Palm Beach, Naples–Immokalee–Marco Island, North Port–Sarasota–Bradenton, Ocala, Orlando–Kissimmee–Sanford, Palm Bay–Melbourne–Titusville, Panama City, Pensacola–Ferry Pass–Brent, Port St. Lucie, Punta Gorda, Sebastian–Vero Beach, Sebring, Tallahassee, Tampa–St. Petersburg–Clearwater, The Villages |
| Georgia (South)              | Albany, Athens–Clarke County, Atlanta–Sandy Springs–Roswell, Augusta–Richmond County*, Columbus*, Dalton, Gainesville, Hinesville, Macon, Savannah, Valdosta, Warner Robins |
| Hawaii (West)                | Kahului–Wailuku–Lahaina, Urban Honolulu |
| Idaho (West)                 | Boise City, Idaho Falls, Lewiston* |
| State (Census Bureau Region) | Region |
|-----------------------------|--------|
| Illinois (Midwest)          | Bloomington, Chicago–Naperville–Elgin*, Davenport–Moline–Rock Island*, Kankakee, Springfield |
| Indiana (Midwest)           | Bloomington, Elkhart–Goshen, Evansville*, Fort Wayne, Lafayette–West Lafayette, Muncie, South Bend–Mishawaka*, Terre Haute |
| Iowa (Midwest)              | Des Moines–West Des Moines |
| Kansas (Midwest)            | Lawrence |
| Kentucky (South)            | Lexington-Fayette, Louisville–Jefferson County* |
| Louisiana (South)           | Alexandria, Baton Rouge, Houma–Thibodaux, Lafayette, Lake Charles |
| Maryland (South)            | Baltimore–Columbia–Towson, California–Lexington Park, Cumberland*, Hagerstown–Martinsburg*, Salisbury* |
| Massachusetts (Northeast)   | Barnstable Town, Boston–Cambridge–Newton*, Pittsfield, Springfield, Worcester* |
| Michigan (Midwest)          | Ann Arbor, Battle Creek, Bay City, Grand Rapids–Wyoming, Jackson, Lansing–East Lansing, Midland, Monroe, Muskegon, Saginaw |
| Minnesota (Midwest)         | Mankato–North Mankato, Minneapolis-St. Paul–Bloomington*, Rochester |
| Mississippi (South)         | Hattiesburg, Jackson |
| Missouri (Midwest)          | Columbia, Joplin, Springfield, St. Louis* |
| Nebraska (Midwest)          | Grand Island, Lincoln, Omaha–Council Bluffs* |
| Nevada (West)               | Las Vegas–Henderson–Paradise, Reno |
| New Hampshire (Northeast)   | Manchester–Nashua |
| New Jersey (Northeast)      | Ocean City, Trenton, Vineland-Bridgeton |
| New Mexico (West)           | Albuquerque, Las Cruces, Santa Fe |
| New York (Northeast)        | Albany–Schenectady–Troy, Binghamton, Elmira, Glens Falls, Ithaca, Kingston, New York–Newark–Jersey City*, Rochester, Syracuse, Watertown–Fort Drum |
| North Carolina (South)      | Asheville, Burlington, Charlotte-Concord–Gastonia*, Durham–Chapel Hill, Fayetteville, Greensboro–High Point, Hickory–Lenoir–Morganton, Raleigh, Rocky Mount, Wilmington, Winston–Salem |
| North Dakota (Midwest)      | Fargo* |
| Ohio (Midwest)              | Akron, Canton–Massillon, Cincinnati*, Cleveland–Elyria, Columbus, Dayton, Lima, Springfield, Toledo, Youngstown–Warren–Boardman* |
| Oklahoma (South)            | Oklahoma City, Tulsa |
Table A1  ■  Continued.

| State (Census Bureau Region) | Region |
|------------------------------|--------|
| Oregon (West)                | Albany, Bend–Redmond, Corvallis, Eugene, Grants Pass, Medford, Portland–Vancouver–Hillsboro*, Salem |
| Pennsylvania (Northeast)     | Allentown–Bethlehem–Easton*, Altoona, Erie, Harrisburg–Carlisle, Lancaster, Philadelphia–Camden–Wilmington*, Pittsburgh, Reading, Scranton–Wilkes–Barre–Hazleton, State College, York–Hanover |
| Rhode Island (Northeast)     | Providence–Warwick* |
| South Carolina (South)       | Columbia, Florence, Greenville–Anderson–Mauldin, Hilton Head Island–Bluffton–Beaufort, Myrtle Beach–Conway–North Myrtle Beach*, Spartanburg |
| Tennessee (South)            | Chattanooga*, Clarksville*, Cleveland, Jackson, Johnson City, Kingsport–Bristol–Bristol*, Knoxville, Nashville–Davidson–Murfreesboro–Franklin |
| Texas (South)                | Amarillo, Brownsville–Harlingen, College Station–Bryan, Dallas–Fort Worth–Arlington, El Paso, Killeen–Temple, Laredo, Midland, Texarkana* |
| Utah (West)                  | Ogden–Clearfield, Provo–Orem, Salt Lake City, St. George |
| Virginia (South)             | Charlottesville, Harrisonburg, Richmond, Roanoke, Staunton–Waynesboro, Virginia Beach–Norfolk–Newport News*, Winchester* |
| Washington (West)            | Bellingham, Kennewick–Richland, Longview, Olympia–Tumwater, Seattle–Tacoma–Bellevue, Spokane–Spokane Valley, Walla Walla, Yakima |
| West Virginia (South)        | Charleston |
| Wisconsin (Midwest)          | Appleton, Eau Claire, Fond du Lac, Janesville–Beloit, La Crosse–Onalaska*, Madison, Oshkosh–Neenah, Racine |

Note: Asterisks indicate that the metropolitan area lies mainly in the indicated state, but parts of it cross state borders.

may be categorized as being either metropolitan or micropolitan. The 917 CBSAs include 381 metropolitan statistical areas, which have an urban core population of at least 50,000, and 536 micropolitan statistical areas, which have an urban core population of at least 10,000 but less than 50,000. In this study, we use 263 metropolitan and 154 micropolitan statistical areas, due to limited availability of data. These 417 regions represent contiguous states (excluding Maine, Montana, South Dakota, Vermont and Wyoming) plus the District of Columbia and Hawaii.
Table A2  ■ The list of micropolitan statistical areas.

| State (Census Bureau Region) | Region |
|-----------------------------|--------|
| Arizona (West)              | Nogales, Payson, Safford |
| Arkansas (South)            | Batesville, Harrison, Paragould, Russellville, Searcy |
| California (West)           | Clearlake, Eureka–Arcata–Fortuna, Red Bluff, Susanville, Truckee–Grass Valley |
| Colorado (West)             | Durango, Glenwood Springs, Montrose, Sterling |
| Connecticut (Northeast)     | Torrington |
| Florida (South)             | Clewiston, Key West, Lake City, Okeechobee, Palatka |
| Georgia (South)             | Bainbridge, Calhoun, Cedartown, Dublin, Jesup, Moultrie, St. Marys, Thomaston, Tifton, Vidalia, Waycross |
| Hawaii (West)               | Hilo |
| Idaho (West)                | Burley |
| Illinois (Midwest)          | Effingham, Jacksonville |
| Indiana (Midwest)           | Angola, Auburn, Bedford, Connersville, Crawfordsville, Decatur, Frankfort, Greensburg, Huntington, Jasper, Kendallville, Logansport, Madison, Marion, New Castle, North Vernon, Peru, Plymouth, Richmond, Seymour, Vincennes, Wabash, Warsaw, Washington |
| Kansas (Midwest)            | Garden City |
| Kentucky (South)            | Danville, Murray |
| Louisiana (South)           | Opelousas |
| Maryland (South)            | Cambridge, Easton |
| Massachusetts (Northeast)   | Greenfield Town, Vineyard Haven |
| Michigan (Midwest)          | Adrian, Hillsdale, Holland, Ionia, Ludington, Owosso |
| Minnesota (Midwest)         | Owatonna, Willmar, Winona |
| Mississippi (South)         | Cleveland, Columbus, Corinth, Grenada, Laurel, Oxford, Picayune, Tupelo, Vicksburg |
| Missouri (Midwest)          | Mexico |
| Nebraska (Midwest)          | North Platte |
| Nevada (West)               | Elko, Fernley, Gardnerville Ranchos |
| New Hampshire (Northeast)   | Concord, Keene, Laconia |
| New York (Northeast)        | Amsterdam, Batavia, Corning, Cortland, Gloversville, Hudson, Olean, Oneonta, Plattsburgh, Seneca Falls |
Table A2  ■  Continued.

| State (Census Bureau Region) | Region |
|-----------------------------|--------|
| North Carolina (South)      | Albemarle, Morehead City, Sanford, Wilson |
| Ohio (Midwest)              | Ashtabula, Coshocton, Defiance, Findlay, Jackson, New Philadelphia–Dover, Portsmouth, Sandusky, Urbana, Wooster |
| Oklahoma (South)             | Ardmore, Bartlesville, Durant, Enid, Marion, McAlester, Tahlequah |
| Oregon (West)                | Coos Bay, Hermiston–Pendleton, Klamath Falls, Ontario*, Roseburg, The Dalles |
| Pennsylvania (Northeast)     | Indiana, Lock Haven, Oil City, Pottsville |
| South Carolina (South)       | Orangeburg |
| Tennessee (South)            | Cookeville, Lawrenceburg, Lewisburg, Martin, Paris, Sevierville, Shelbyville, Tullahoma–Manchester |
| Virginia (South)             | Danville, Martinsville |
| Washington (West)            | Oak Harbor, Port Angeles, Shelton |
| Wisconsin (Midwest)          | Baraboo, Marinette*, Whitewater–Elkhorn |

Note: Asterisks indicate that the micropolitan area lies mainly in the indicated state, but parts of it cross state borders.

Appendix B: The MCMC Algorithm

We estimate the model by running an MCMC algorithm. The full conditional posterior distributions are available in closed form implying that we can apply Gibbs sampling to obtain draws from the joint posterior distribution. More specifically, our MCMC algorithm involves the following steps:

(i) Simulate the VAR coefficients $a_j$ ($j = 1, \ldots, J$) conditional on the factors and remaining model parameters from a multivariate Gaussian distribution that takes a standard form (see, for instance, George, Sun and Ni 2008, for further information).

(ii) Simulate the latent factors $F_t$ by using forward filtering backward sampling (Carter and Kohn 1994, Frühwirth-Schnatter 1994).

(iii) The error variance–covariance matrix $\Sigma_u$ is simulated from an inverted Wishart posterior distribution with degrees of freedom equal to $\nu = v + T$ and scaling matrix equal to $P = \sum_{t=1}^{T} (y_t - Ax_t)'(y_t - Ax_t) + \Sigma$. 
(iv) Simulate the factor loadings $\lambda_\ell$ ($\ell = 1, \ldots, L$) from Gaussian posteriors (conditioned on the remaining parameters and the latent factors) by running a sequence of $(R - S)$ unrelated regression models.

(v) The measurement error variances $\sigma^2_r$ for $r = S + 1, \ldots, R$ are simulated independently from an inverse Gamma distribution $\sigma^2_r | \Xi \sim G^{-1}(\alpha_r, \beta_r)$ with $\alpha_r = \frac{1}{2} T + e_0$ and $\beta_r = \frac{1}{2} \sum_{t=1}^T (H_{rt} - \Lambda^F_r F_t - \Lambda^M_r M_t)^2 + e_1$. The notation $\Lambda^F_r$ indicates that the $r$th row of the matrix concerned is selected, and $\Xi$ stands for conditioning on the remaining parameters and the data.

(vi) Simulate $\tau^2_{a_j}$ ($j = 1, \ldots, J$) from a generalized inverted Gaussian distributed posterior distribution with

$$\tau^2_{a_j} | \Xi \sim GIG\left(\theta_a - \frac{1}{2}, a_j^2, \theta_a \xi_a\right). \quad (B1)$$

(vii) Draw $\xi_a$ from a Gamma distributed posterior given by

$$\xi_a | \Xi \sim G\left(c_0 + \theta_a J, c_1 + \frac{1}{2} \theta_a \sum_{\ell=1}^L \tau^2_{a\ell}\right). \quad (B2)$$

(viii) Simulate the posterior of $\tau^2_{\lambda_\ell}$ ($\ell = 1, \ldots, L$) from a generalized inverted Gaussian distribution,

$$\tau^2_{\lambda_\ell} | \Xi \sim GIG\left(\theta_\lambda - \frac{1}{2}, \lambda^2_\ell, \theta_\lambda \xi_\lambda\right). \quad (B3)$$

(ix) Finally, the global shrinkage parameter $\xi_\lambda$ associated with the prior on the factor loadings is simulated from a Gamma distribution,

$$\xi_\lambda | \Xi \sim G\left(d_0 + \theta_\lambda L, d_1 + \frac{1}{2} \theta_\lambda \sum_{\ell=1}^L \tau^2_{\lambda_\ell}\right). \quad (B4)$$

Steps described above are iterated for 20,000 cycles, where we discard the first 10,000 draws as burn-in.

Appendix C: Robustness Check—Comparing with an Identification Scheme Imposing Sign Restrictions

To assess the sensitivity of our results with respect to identification of the monetary policy shock, we use an alternative strategy based on contemporaneous sign restrictions (see Uhlig 2005, Dedola and Neri 2007). Technical implementation is achieved by adopting the algorithm proposed in Arias, Rubio-Ramirez and Waggoner (2014) that collapses to the procedure outlined in Rubio-Ramirez, Waggoner and Zha (2010) in the absence of zero restrictions. For each iteration of the MCMC algorithm, we draw a rotation matrix
and assess whether the following set of sign restrictions is satisfied. Consistent with economic common sense, output (measured in terms of the industrial production index), housing investment (measured in terms of housing starts) and consumer prices (measured in terms of the consumer price index) are bound to increase on impact. Moreover, we assume that the term-spread also widens on impact. Finally, consistent with the normalization adopted when using an external instrument, we assume that the one-year yield declines. If this is the case, we keep the rotation matrix and store the associated structural coefficients, while if the sign restrictions are not met, we reject the draw and repeat the procedure.

The results are displayed in form of a geographic map with a classification scheme that generates class breaks in standard deviation measures above and below the mean, see Figure C1. A comparison with Figure 6 provides evidence of the robustness of our results.

**Figure C1** Robustness check: Cumulative responses of regional housing prices to a monetary policy shock identified using sign restrictions. [Color figure can be viewed at wileyonlinelibrary.com]

Notes: Visualization is based on a classification scheme that generates breaks in standard deviation measures. The number of regions is allocated to the classes in squared brackets. The responses based on 10,000 posterior draws have been accumulated. Thinner lines denote the boundaries of the regions, while thicker lines represent U.S. state boundaries. Sample period: 1997:04–2012:06. For the list of regions see Appendix A.