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LESSONS FROM REGIONAL DIVERGENCES IN A COMMON MONETARY POLICY SETTING

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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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2 European Central Bank, Kaiserstrasse 29, 60311 Frankfurt am Main, Germany. Tel: (+49) (0)69 1344 8446; e-mail: Isabel.Vansteenkiste@ecb.int

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Internet
http://www.ecb.int

Fax
+49 69 1344 6000

Telex
411 144 ecb d

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Abstract

In this paper, we seek to quantify the importance of state-level housing price spillovers and interest rate shocks to house price developments in the United States. The econometric approach involves an application of the recently developed global VAR (*GVAR*) as presented in Dées, DiMauro, Pesaran, and Smith (2005) and Pesaran, Schuermann, and Weiner (2004) to the 31 biggest US states over the period 1986-2005. Such an approach allows not only for the empirical derivation of the impact of common shocks (such as interest rate shocks) on US house price developments, but also for an analysis of the importance of interstate housing price spillovers. Beyond real house prices and real income per capita, each state-specific vector error correction model also includes nation-wide variables – measured as a weighted average of other states –. These individual state models are then linked in a consistent and cohesive manner.

Impact elasticities indicate strong interregional linkages for both real house prices and real income per capita. An analysis of generalised impulse responses indicates that the importance of housing price spillovers is state dependent, with shocks occurring in states with relatively lower land supply elasticities having much stronger spillover effects that those in the other states. As regards real interest rates, the impact appears to be relatively small with an increase of 100 basis points in the real 10-year government bond yield resulting in a long run fall in house prices of between 0.5 and 2.5%. This would suggest, in line with DelNegro and Otrok (2005) that the decline in long-term interest rates is not the primary factor that has driven the recent surge in house prices in the United States.

J.E.L. classification: C32, E44, R10, R31
Keywords: housing, monetary policy, global VAR (GVAR).

Non-technical summary

In recent years, home prices have risen rapidly in the US, both in nominal and real terms. In more detail, US nominal house values rose by almost 40% over the decade of the 1990s. In many metropolitan areas, house values doubled again during 2000-2005 and as a result, at the end of 2005, nominal and real house prices stood respectively about 31% and 23% above their national levels in mid-2003. This development at the national level however masks considerable heterogeneity across cities, states and regions. Indeed, while price increases in some areas have remained relatively constrained, they have been particularly rapid in the US coastal areas. Such a divergence in developments across areas triggers two important policy questions. First, given the increase in state-level dispersion, how important has a common factor — i.e. the decline in long term interest rates — been in driving US house price developments and second, how does a house price shock in one state spill over across other states in the US.

To address both questions, we make use of the rather novel Global Vector Autoregression (GVAR) approach. Such a model may be particularly suitable for our purpose since it explicitly allows for interdependencies that exist between state-level and nationwide factors. This enables us to analyse both the effects of common shocks (i.e. interest rates) on state-level developments as well as the importance of spillovers from state-specific housing price shocks.

We apply the GVAR approach to the 31 biggest states in the United States over the period 1986 until 2005. The results from our analysis would suggest that historically:

- House price spillovers are present at the US state level. However, their magnitude only becomes important when house price shocks occur in certain states (namely those with low land supply elasticity). The largest response was found when imposing a 10% shock on Californian house prices. This would result in a long run responses in the range of 0.6% to 4.6% in the other states. Such an outcome would clearly provide an upper bound for the type of spillovers we might expect from house price shocks in the euro area. Indeed, given that labour mobility is much lower in the euro area (especially across borders) and financial integration is less advanced, spillover effects in the euro area will likely be small.
- As regards real interest rates, the impact on the US housing market appears to be relatively small with an increase of 100 basis points in the real 10-year government bond yield resulting in a long run fall in house prices of between 0.5 and 2.5%. This would suggest that while the role of the decline in long-term interest rates in recent years is non-negligible, at the same time, it has not been the main driver behind recent US house price developments.

1 Introduction

In recent years, home prices have risen rapidly in the US, both in nominal and real terms. In more detail, at the end of 2005, nominal and real house prices stood respectively about 31% and 23% above their levels in mid-2003. This development at the national level, however, has masked considerable heterogeneity across cities, states and regions, consistent with the importance of local conditions in determining housing market activity and prices. In examining the geographic distribution of housing across US cities, property price inflation in areas with lower land supply elasticity – such as coastal areas – has been much more elevated than that of other areas. Such a development has translated into a general flattening in the distribution of house price increases across the 387 Metropolitan Statistical Areas (MSAs) covered by the Office of Federal Housing Enterprise Oversight (OFHEO) in the last years, along with the development of considerable positive skew (see Chart 1a).[1] Such an increase in housing price dispersion across US cities contrasts with the general declining trend in income dispersion over the last decades. As a result, as is shown in Chart 1b the kernel density of US average housing price inflation for the period 2000-2005 indicates a large positive skew which is not present in the income per capita data over the same period.

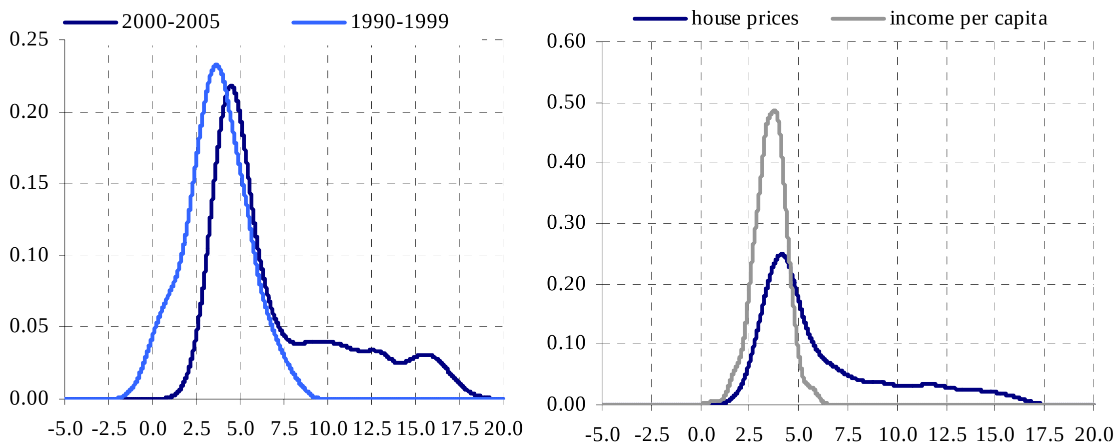

Figure 1: Kernel density of US average housing price inflation and income in 387 MSA

From the perspective of the current debate, these developments raise two important questions. First, given the increasing importance of regional dispersion, how important has a common factor – i.e. the decline in long term interest rates – been in driving housing price developments and second, how does a house price shock in one state spill over across other states in the United States. The answer to both questions has important policy implications. As regards the first question, our alternative approach allows us to verify the finding by DelNegro and Otrok (2005) that the role of the declining long term interest rates has been limited in the current house price rally. Moreover it

[1]The data used to compute the kernel densities in Chart 1 is the average year-on-year growth rates for house prices and income per capita for the 387 metropolitan statistical areas over the sample periods considered (namely 1990-1999 and 2000-2005).

allows us to assess the divergences in house price responses across states to an interest rate shock. As regards the second question, if indeed stata-level housing price spillovers are important, then a housing price correction in one state may in fact have important implications for the overall macroeconomic outlook.

Taking the last two arguments together, if indeed state-level divergences would be important, it would suggest that this may matter for the efficacy of monetary policy (see in this context Carlino and DeFina (1999)). For example, the aggregate macroeconomic response to a state-level housing shock will depend on the state where the shock originated. Similarly, the response to a monetary tightening will depend on issues such as whether the most rapidly expanding regions are the most interest sensitive. More generally, the aggregate effects of monetary policy depend on the distribution of regional sensitivities to monetary policy and on the initial distribution of regional economic conditions at the time of monetary tightening. Both distributions vary over time, so small changes in the configuration of heterogeneity can produce economically significant changes in aggregate responses. However, while regional divergences will matter for monetary policy, obviously, at the same time, such a result would not imply that monetary policy should try to control or target the conditions of particular geographic regions. It would only indicate that regional economic conditions may have a significant influence on the aggregate and hence warrant monitoring.

To address the issues raised above, we make use of a Global Vector Autoregression (GVAR) modelling approach in the spirit of Dées, DiMauro, Pesaran, and Smith (2005) (henceforth DdPS) and Pesaran, Schuermann, and Weiner (2004) (henceforth PSW).[2] Such a model may be particularly suitable for our purpose since it explicitly allows for interdependencies that exist between state-level and nationwide factors. This enables us to analyse both the effects of common shocks (i.e. interest rates) on state-level developments as well as the importance of spillovers from state-specific housing price shocks. Our results suggest that historically, housing price spillovers are present across states but their magnitude becomes only important when housing price shocks occur in certain states (in particular those with lower land supply elasticity). As regards interest rates, while their role is non-negligible, at the same time, the magnitude suggests that they have not been the main driver behind the recent house price developments.

While there exist an established literature 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 (see DelNegro and Otrok (2005)). Iacoviello and Minetti (2003) 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 US. Chirinko, de Haan, and Sterken (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. As regards the study of state-level or regional divergences in the United States, the literature is more extensive. Important studies in this context are those by Blanchard and Katz (1992), Quah (1995) and Carlino and DeFina (1999). Overall, studies in this literature indicate the importance of taking into account regional heterogeneity and suggest that considering the response of

[2]In this sense, whilst the 'GVAR' nomenclature is retained, the term 'global' applies to US economy as a whole –and not other countries/ regions– in contrast to the DdPS global model application.

the various regions may be of importance to avoid the issue of aggregation bias. For instance Carlino and DeFina (1999) provide evidence on heterogeneous sensitivity of regions to monetary policy. In their study, the authors find that after 2 years, the dynamic response of real income to a monetary policy tightening varies widely across US states. In some states, such as Michigan, Indiana, and Arizona, income declines on average by about 2% whereas in other states, such as Texas, Wyoming, and New York, income declines only by about 0.5%. Both estimates deviate significantly from the national average of 1.2%.

Looking at the literature that combines the two strands above, perhaps the two closest studies ours are Fratantoni and Schuh (2003) and DelNegro and Otrok (2005). Both study the effects of monetary policy on regions in the US. Fratantoni and Schuh (2003) do this over the period 1986-1998. They find that the response of housing investment appreciation to monetary policy varies by region. DelNegro and Otrok (2005) use a factor model to extract the common cycle in house price fluctuations. Their focus is on the role of monetary policy in the latest housing boom. Overall, the authors attribute only a limited role to interest rate developments in explaining the recent boom in house prices. Our paper however differs from previous studies in the literature both in terms of methodology and focus. In terms of methodology, we make use of the rather novel GVAR approach to address the link between state-level, national and interstate factors. In terms of focus, in line with DelNegro and Otrok (2005) we consider the role of interest rates on housing price developments. However we also go beyond that and consider the importance of housing price spillovers across states to assess the impact of a state-level shock on national developments.

The remainder of the paper is as follows. Section 2 describes the model approach, section 3 describes the data, section 4 discusses the model test and estimation results, section 5 presents the dynamic properties of the model and section 6 concludes.

2 The empirical model

To determine the role of nationwide factors and spillovers on state-level house price developments we estimate a GVAR model, as in DdPS and PSW. In the model, individual state-specific vector error-correcting models are estimated in which the state-specific variables are related to corresponding state-specific weighted averages of the other states' variables plus deterministic variables, such as time trends, and nationwide (weakly) exogenous variables, such as the real interest rate.[3] These individual state models are then linked in a consistent and cohesive manner.

More specifically, in line with DdPS, we assume we have $N+1$ states, indexed by $i = 0, 1, 2, ..., N$. For each state, we assume that state-specific variables x (namely real house prices p and real income per capita y) are related to corresponding state-specific weighted averages of the other states' variables x^* (here comprising p^* and y^*) plus deterministic variables, such as a time trend (t), and a nationwide (weakly) exogenous

[3] Modelling the relationship between house prices and fundamentals such as income as an error correction specification is common practice in the housing literature (see for instance Malpezzi (1999), Capozza, Hendershott, Mack, and Mayer (2002) and Meen (2002)). More recently Gallin (2003) has however questioned the underlying assumption of these traditional models, namely that house prices and income are linked by a stable long-run relationship. However, as is shown in Holly, Pesaran, and Yamagata (2006) and in our section 4 in our case this assumption appears to be reasonable.

variable, namely the real interest rate (rir). For simplicity, we confine our exposition here to a first-order dynamic specification as in PSW. In this case we can relate the $k_i \times 1$ state-specific variables, $x_{it} = (p_{it}, y_{it})$, to $x^*_{it} = (p^*_{it}, y^*_{it})$ and $d_t = (t, rir_t)$ and write:

$$x_{it} = a_{io} + a_{i1}t + \Phi_i x_{i,t-1} + \Lambda_{i0} x^*_{i,t} + \Lambda_{i1} x^*_{i,t-1} + \psi_{i0} d_t + \psi_{i1} d_{t-1} + \varepsilon_{it}$$

where Φ_i is a $k_i \times k_i$ matrix of lagged coefficients, Λ_{i0} and Λ_{i1} are $k_i \times k^*_i$ matrices of coefficients associated with the foreign-specific variables, ψ_{i0} and ψ_{i1} are $k_i \times s$ matrices of coefficients associated with the common country-wide variables and ε_{it} is a $k_i \times 1$ vector of idiosyncratic state-specific shocks. We assume in this model that the idiosyncratic shocks, ε_{it}, are serially uncorrelated with mean 0 and a nonsingular covariance matrix, $\Sigma_{ii} = (\sigma_{ii,ls})$ where $\sigma_{ii,ls} = cov(\varepsilon_{ilt}, \varepsilon_{ist})$, or written more compactly, $\varepsilon_{it} \sim iid(0, \Sigma_{ii})$. The assumption that the state-specific variance-covariance matrices are time invariant can be relaxed, but for the analysis of quarterly observations, this time invariant assumption may not be overly restrictive. This state-specific model can now be consistently estimated separately, treating d_t and x^*_{it} as weakly exogenous $I(1)$ with respect to the parameters of this model.

The weak exogeneity assumption in the context of cointegrating models implies no long run feedbacks from x_{it} to x^*_{it}, without necessarily ruling out lagged short run feedbacks between the two sets of variables. In this case x_{it} is said to be l*ong run forcing* x^*_{it}, and implies that the error correction terms of the individual country VECMs do not enter in the marginal model of x^*_{it} (see DdPS). The weak exogeneity of these variables can then be tested in the context of each of the state-specific models (see section 4). Once the individual state models are estimated all the endogenous variables need to be solved simultaneously.

All state-specific models together with the relations linking the (weakly) exogenous variables of the state-specific models to the variables in the rest of the model, provide a complete system. However, due to data limitations for even moderate values of N, a full system estimation of the model may not be feasible. To avoid this difficulty, we follow PSW and estimate the parameters of the state-specific models separately, treating the foreign-specific variables as weakly exogenous on the grounds that most states are small relative to the size of the overall US economy (see section 4 for further details on this).

Overall, the nationwide model, associated with the state-specific models can now be given by:

$$Gx_t = a_o + a_1 t + Hx_{i,t-1} + \psi_0 d_t + \psi_1 d_{t-1} + \varepsilon_t$$

where a_o, a_1, ψ_0, ψ_1, G, H, and ε_t can be defined as: ($j = 0$ or 1)

$$a_j = \begin{pmatrix} a_{0j} \\ a_{1j} \\ \dots \\ a_{Nj} \end{pmatrix}, \varepsilon_t = \begin{pmatrix} \varepsilon_{0t} \\ \varepsilon_{1t} \\ \dots \\ \varepsilon_{Nt} \end{pmatrix}, \psi_j = \begin{pmatrix} \psi_{0j} \\ \psi_{1j} \\ \dots \\ \psi_{Nj} \end{pmatrix}, G = \begin{pmatrix} A_0 W_0 \\ A_1 W_1 \\ \dots \\ A_N W_N \end{pmatrix}, H = \begin{pmatrix} B_0 W_0 \\ B_1 W_1 \\ \dots \\ B_N W_N \end{pmatrix}$$

whereby W_i is a $(k_i \times k^*_i) \times k$ matrix of fixed constants defined in terms of the state-specific weights. W_i can be viewed as the link matrix that allows the state–specific models to be written in terms of the global variable vector x_t.

In general, such a GVAR model allows for interactions among the different economies through three separate but interrelated channels:

- Contemporaneous dependence of x_{it} on x^*_{it} and on its lagged values.
- Dependence of the state-specific variables on common exogenous variables, such as the real interest rate.
- Nonzero contemporaneous dependence of shocks in state i on the shocks in state j, measured via the cross-country covariances, Σ_{ij}

In what follows, we will first consider the data used for the analysis and then proceed with the various steps in estimating the model. First and foremost this implies analysing the integrating properties of the variables, determining the cointegrating rank and testing for weak exogeneity of the exogenous and starred variables. All these steps were implemented in *PcGive*. After the individual state VECM models have been estimated, they are then linked together to create the nationwide model from which the generalised impulse responses can be derived. This step was implemented in *Matlab*.

3 The data

For the purpose of our study we require three data series: *(1)* real house prices (p), *(2)* real income per capita data (y) and *(3)* real interest rates (rir). We choose to collect the data at a US state level, even though more disaggregate data series are available (e.g. MSAs) for our purpose. State-level analysis should however be disaggregate enough to establish the main conclusion, while keeping the computational burden relatively contained.[4]

As a proxy for nominal house price developments, the Housing Price Index (*HPI*) which is published by the Office of Federal Housing Enterprise Oversight (*OFHEO*) is used. The index is a repeat-sales price index for existing homes. 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. The index does however not control for changes to the house through improvements or neglect; that is, while it does hold some characteristics constant, it is not a true quality-adjusted price. In addition, the repeat-sales sample excludes homes with jumbo, FHA or VA mortgages (which captures changes in the value of single-family homes, see Calhoen (1996)). Although the housing price data has been criticized for its construction, it seems to remain however the best data available to the public at the state (or more disaggregated) level. The HPI data are nominal. The series are deflated by using the personal consumption expenditure deflator less food and energy (core PCE).[5] The HPI data are available from 1975, but in the estimation for this paper, in line with DelNegro and Otrok (2005), only data beginning in the first quarter of 1986 is used. There are various reasons to opt for the shorter sample. First, as illustrated in Figure 2, and as noted by DelNegro and Otrok (2005), state-level HPI data are rather noisy for a

[4] One of the problems related to using more disaggregate data (e.g. MSAs) is that selecting those areas among the 387 available which do not violate the weak exogeneity assumption may be rather tedious (see section 4 for more details on the testing for weak exogeneity of the data).

[5] The deflators used in this paper are national deflators as state-level PCE or CPI deflators are not available. Moreover, the CPI which exists for some MSAs suggests that using a national deflator is not a bad proxy as inflation differentials between the MSAs has been rather limited over the time period we consider.

number of states before the mid-eighties, with sharp appreciations immediately followed by sharp depreciations. The noise abates considerably for most states after the mid-eighties. Further, large structural changes in the credit market, such as the end of the regulation Q[6] coupled with the shift towards nationwide funding of housing activity based on mortgage securitisation following the *S&L* crisis provide another (related) reason for leaving the first part of the sample out of the analysis.[7] Finally, the shorter sample gives a period with one monetary policy regime. The sample ends in the last quarter of 2005. In summary, we have 19 years of data for the 48 contiguous US states. The real per capita personal income data for the same states are computed by deflating the nominal per capita income data from the Bureau of Economic Analysis using PCE inflation. However, since the weak exogeneity assumption for the model doesn't hold when including all 48 contiguous US states, we decided to eliminate some, in particular the smaller states which seemed to cause the problem. As a result, all states which contain less than 1% of the total US population were dropped, leaving us with 31 US states which, in terms of population, however still cover 90% of the total US population (see Annex A for the states included in our analysis).[8] Finally, for the interest rate, we use the real 10-year government bond yield, whereby the interest rate was deflated using PCE inflation.[9]

One important input in the calculation of the GVAR model is to construct the state-specific starred variables. To compute them state specific weights are required. Here, we rely on distance weights.[10] These distance weights were calculated in two steps. First the average distance between one state and the others was calculated. To obtain the distance between states, the average of all the distances between the Metropolitan Statistical Areas (MSAs) was calculated. Next, once the distances for all states were obtained, the inverse of these numbers were taken. After, these numbers were then rescaled so as to add up to one for each state (see PSW). The resulting 31×31 matrix of the weights is presented in Table A1 to A3 in Appendix B.

The shares of each region are displayed in the columns. The matrix plays a key role in linking up the models of the different states together and shows the degree to which one state depends on the remaining states. For example, not surprisingly, the weights show that New Jersey and Pennsylvania have a larger weight on the New York state than for instance Arizona has. Based on this matrix of weights, starred variables can be calculated for both the real house prices and income per capita data. To illustrate

[6] Regulation Q was a Federal Reserve Board regulation that limits the interest rate that banks can pay on savings deposits and as such restricts mortgage finance. The regulation ended in March 1986.

[7] In particular the shift to a more market-based financial structure has led to a reduction in the volatility of mortgage lending, whereby the availability of funds is no longer limited by conditions at local depository institutions or the strength of regional economies, and monetary policy no longer has a direct effect on the supply of mortgages through high-powered money expanding bank balance sheets. As a result, the correlation between growth of real deposits and real mortgage flows dropped from 0.72 between 1960 and 1984 to 0.06 between 1985 and 2004 (see Schnure (2005)). Consequently, the cyclicality of mortgage flows has reduced (see also Peek and Wilcox (2006)).

[8] The states dropped are AK, AR, DC, DE, HI, ID, KS, ME, MT, ND, NE, NH, NM, NV, RI, SD, UT, VT, WV and WY.

[9] We also estimated the model using the real Federal Funds Rate. However, this did not change our results substantially. For this reason, we only report the result with the real 10-year government bond yield.

[10] In the GVAR model of DdPS trade weights were used to compute each country's starred variables. In our case alternatives could have been using the squared distance or the number of nearest neighbours.

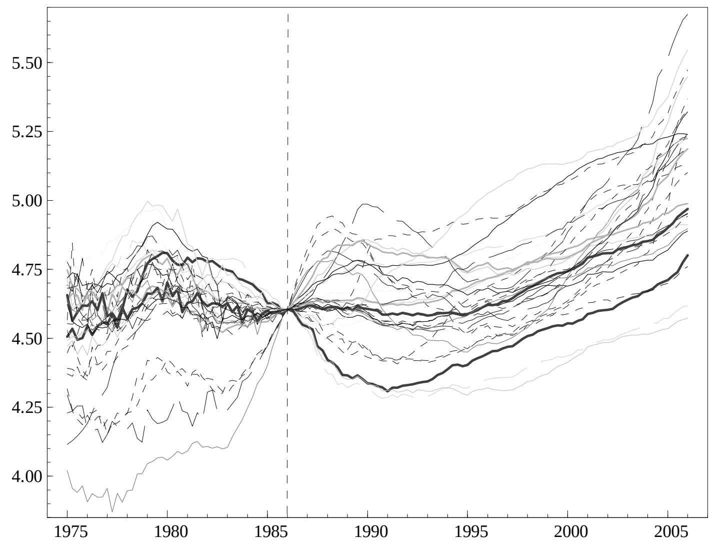

Figure 2: The log of real house prices in the 31 largest US states (1986Q1=100).

the results, the starred price variables and their state-specific counterparts are shown in Appendix B.

4 Model testing and estimation results

4.1 Integration properties of the series

The first step to take before estimating the individual state-specific cointegrating VAR models is to verify whether the variables included are approximately integrated of order one. To ascertain the order of integration of the variables, we present Augmented Dickey-Fuller (ADF) statistics for the log-levels, first and second differences of the starred and state-specific variables. To ensure comparability, all these statistics are computed over the sample 1986Q3-2005Q4, starting with an underlying AR process of order 5, with a linear trend in case of the levels (except for the interest rate data) and an intercept term (except for the level series). The order of the ADF test statistics reported in Table 1 are selected according to the Akaike Information Criterion.

Generally speaking, the results of these tests show that series are $I(1)$ across all states with only a few exceptions. In more detail, house prices in Louisiana and Maryland and personal income in Florida, Louisiana and Maryland may be $I(2)$. On the basis of this we use level series to determine the long-run cointegrating vectors of the model and first differenced series when calculating the VECM models. For the real interest rate data, which is the same for all states, the ADF test statistic for the level series equals -2.8 and for the first differenced series -4.9 suggesting this data is also $I(1)$.

Table 1: ADF Unit Root Test Statistics (based on AIC order selection)

	p	dp	ddp	y	dy	ddy	p^*	dp^*	ddp^*	y^*	dy^*	ddy^*
AL	-0.0	-4.4	-9.8	-1.7	-3.9	-5.9	-0.4	-3.5	-6.1	-2.3	-3.5	-4.9
AZ	0.4	-4.4	-7.7	-1.7	-4.4	-3.6	-0.8	-3.5	-6.1	-2.4	-3.1	-4.6
CA	-2.3	-3.3	-5.4	-2.5	-3.1	-4.3	-0.6	-3.6	-6.2	-2.3	-3.0	-4.4
CO	-2.0	-3.2	-6.5	-1.9	-1.6	-4.6	-0.6	-3.5	-5.9	-2.4	-3.3	-4.7
CT	-2.5	-3.7	-5.8	-2.5	-3.0	-6.4	-2.1	-2.5	-6.0	-2.6	-3.2	-5.3
FL	2.8	-3.7	-8.6	-2.4	-2.7	-4.1	-0.6	-3.6	-5.9	-2.3	-3.5	-5.0
GA	0.4	-3.8	-10.2	-2.0	-3.0	-4.4	-0.4	-3.6	-5.6	-2.3	-3.5	-5.0
IA	-0.8	-3.2	-9.9	-2.1	-5.5	-7.3	-1.1	-3.7	-5.4	-2.4	-3.5	-5.1
IL	-1.4	-3.1	-7.5	-2.3	-4.3	-6.5	-1.2	-3.7	-5.2	-2.4	-3.8	-5.3
IN	-1.7	-3.7	-7.4	-2.5	-4.9	-6.7	-1.2	-3.7	-5.3	-2.4	-3.6	-5.3
KY	-1.0	-3.0	-6.6	-1.8	-3.8	-6.0	-0.9	-3.8	-5.6	-2.4	-3.7	-5.2
LA	-1.0	-1.9	-6.2	-2.5	-2.9	-5.4	-0.4	-3.5	-6.3	-2.3	-3.4	-4.9
MA	-3.0	-3.1	-5.4	-2.7	-4.6	-5.6	-2.0	-3.7	-6.0	-2.6	-3.2	-5.4
MD	-1.1	-1.5	-6.0	-2.5	-2.3	-5.0	-1.5	-3.6	-5.9	-2.5	-3.3	-5.3
MI	-1.9	-3.1	-6.9	-3.3	-4.3	-7.2	-1.0	-3.8	-5.5	-2.5	-3.6	-5.3
MN	-0.4	-3.8	-8.2	-2.4	-4.7	-6.2	-1.0	-3.6	-5.2	-2.4	-3.6	-5.2
MO	-1.8	-3.3	-6.2	-2.3	-3.8	-5.5	-0.7	-3.6	-5.5	-2.3	-3.6	-5.1
MS	0.1	-5.6	-10.2	-1.8	-4.4	-6.8	-0.5	-3.5	-5.5	-2.3	-3.4	-4.9
NC	-0.3	-3.1	-8.7	-2.2	-3.9	-5.1	-0.6	-3.5	-5.8	-2.4	-3.3	-5.1
NJ	-2.9	-3.3	-5.6	-2.4	-3.3	-5.8	-1.7	-3.6	-6.3	-2.6	-3.2	-5.3
NY	-1.2	-3.4	-9.2	-2.5	-3.8	-7.4	-1.7	-3.8	-6.0	-2.5	-3.3	-5.2
OH	-1.8	-3.2	-7.2	-2.6	-3.8	-7.3	-1.1	-3.7	-5.6	-2.5	-3.6	-5.3
OK	-1.4	-3.3	-6.5	-2.3	-3.5	-6.1	-0.9	-3.5	-5.6	-2.3	-3.4	-4.8
OR	-1.4	-3.6	-6.8	-2.3	-3.6	-4.8	-0.8	-3.3	-6.3	-2.5	-2.9	-4.4
PA	-1.3	-3.7	-8.7	-2.7	-3.1	-7.1	-1.5	-3.6	-5.8	-2.5	-3.2	-5.2
SC	0.7	-4.0	-7.6	-1.9	-3.6	-5.6	-0.6	-3.6	-5.7	-2.4	-3.4	-4.9
TN	-0.4	-4.1	-9.9	-2.1	-4.2	-5.5	-0.7	-3.7	-5.5	-2.3	-3.5	-5.1
TX	-1.5	-4.0	-8.0	-2.0	-3.3	-3.5	-0.6	-3.6	-5.8	-2.3	-3.3	-4.8
VA	0.1	-3.9	-6.7	-2.2	-3.5	-5.3	-1.1	-3.7	-6.1	-2.5	-3.4	-5.1
WA	-2.1	-3.5	-6.8	-3.0	-4.1	-4.3	-0.7	-3.5	-6.3	-2.4	-3.1	-4.5
WI	-0.6	-4.1	-6.9	-2.6	-3.7	-6.5	-1.0	-3.6	-5.8	-2.4	-3.8	-5.3

Note: the ADF statistics are based on univariate AR(p) models with p≤5. The statistics for the level, first difference, and second differences of the variables are all computed on the basis of the same sample period, namely 1986Q3-2005Q4. The ADF statistics for all the level variables are based on regressions including a linear trend. The 95% critical value of the ADF statistics for regressions with trend is -3.47, and for regressions without trend -2.90.

4.2 State-specific models

In view of the above results, we use the level series for both our endogenous and our weakly exogenous variables in the state-specific models. The next step of the analysis is to estimate state-specific cointegrating VAR models and identify the rank of their cointegrating space. We rely for this on the trace test statistics for each of the 31 states as set out in Pesaran, Shin, and Smith (2000).[11] The results of this are reported in Table 2. The statistics are computed using a VAR specification with restricted trend coefficients. The order of the VAR is determined by the AIC criterion and is also reported in the table. Based on the trace test, we find for all regions one cointegrating relation. A prima facie, this result contrasts with Gallin (2003) who finds no cointegration using a panel of 95 MSAs. However, in this analysis the other regions' house prices and income levels do not feed into each separate equation. Moreover, more recently Holly, Pesaran, and Yamagata (2006) also find that real house prices and real income per capita at the US state level are cointegrated by estimating a model which allows for unobserved common factors that could potentially be correlated with the observed regressors.

Table 2: Cointegration rank statistics and lag structure (based on AIC)

H_o	$r \leq 0$	$r \leq 1$	$r \leq 2$	lags	H_o	$r \leq 0$	$r \leq 1$	$r \leq 2$	lags
AL	124.87**	75.86**	5.29	2	MO	115.34**	77.50**	6.51	2
AZ	152.31**	99.76**	10.04	2	MS	169.05**	108.64**	8.52	2
CA	150.34**	106.08**	10.23	2	NC	143.85**	71.31**	5.41	1
CO	149.34**	103.28**	6.47	2	NJ	159.74**	88.42**	8.37	1
CT	148.77**	83.95**	2.14	2	NY	117.38**	74.25**	3.59	2
FL	144.23**	86.98**	7.89	2	OH	128.78**	79.24**	9.13	1
GA	120.92**	78.57**	3.74	2	OK	205.56**	116.79**	8.46	2
IA	145.43**	96.07**	9.39	2	OR	153.26**	98.86**	12.08	3
IL	151.8**	93.48**	11.35	2	PA	155.87**	108.99*	11.51	3
IN	153.87**	88.95**	10.87	2	SC	106.84**	67.89*	6.96	2
KY	120.37**	84.08**	7.16	2	TN	122.61**	77.34**	3.46	2
LA	170.41**	119.60**	9.36	3	TX	225.25**	118.96**	5.04	1
MA	125.56**	80.77**	4.23	2	VA	181.55**	87.50**	10.23	1
MD	145.47**	93.14**	7.30	2	WA	116.53**	74.77**	3.01	2
MI	107.9**	71.41**	3.16	2	WI	136.45**	88.09**	7.15	2
MN	147.77**	70.95**	3.44	1					

Note: the model contains unrestricted intercepts and restricted trend coefficients with I(1) endogenous variables p and y and I(1) exogenous variables p^*, y^* and rir.

4.3 Testing weak exogeneity of the state-specific foreign variables

One of the key assumptions underlying our estimation approach is the weak exogeneity of the state-specific foreign variables. To test this assumption, we can run first-difference

[11] We could also consider the maximum eigenvalue test statistics. However, since it has been shown, using Monte Carlo experiments, that the maximum eigenvalue test is generally less robust to departures from normal errors than the trace test, we prefer to use the latter in our analysis.

regressions of the foreign variables and test the significance of the sector-specific error-correction terms in the regressions. For instance, to test that foreign house prices (p^*) in the California (CA) model are weakly exogenous, we need to test the joint hypothesis that:

$$\delta_{CA} = 0$$

in the regression

$$\Delta p^*_{CA,t} = a_{CA} + \delta_{CA} ECM_{CA,t-1} + \phi'_{CA} \Delta z_{CA,t-1} + \phi_{CA,0} \Delta rir_{t-1} + \varsigma_{CA,t}$$

where ECM is the estimated error-correction term associated with the cointegrating relation found in the CA model, $\Delta z_{CA,t-1} = (\Delta p'_{CA,t-1}, \Delta p'^*_{CA,t-1}, \Delta y'_{CA,t-1}, \Delta y'_{CA,t-1})$. The F-statistics for testing the weak exogeneity of all the state-specific foreign variables and the real interest rate variable are summarised in Table 3. In none of the cases the weak exogeneity tests turned out to be statistically significant.

Table 3: F-Statistics for Testing the Weak Exogeneity of the State-specific foreign variables and the real interest rate

	p^*	y^*	rir		p^*	y^*	rir		p^*	y^*	rir
AL	2.39	0.96	0.00	LA	2.11	2.29	1.77	OK	0.04	0.48	0.14
	[0.13]	[0.17]	[0.98]		[0.15]	[0.14]	[0.19]		[0.84]	[0.49]	[0.71]
AZ	0.99	0.93	3.06	MA	3.54	0.54	0.11	OR	3.29	3.37	1.62
	[0.32]	[0.34]	[0.09]		[0.06]	[0.46]	[0.74]		[0.07]	[0.07]	[0.21]
CA	2.08	0.13	0.09	MD	2.97	2.85	0.12	PA	2.46	0.15	0.67
	[0.15]	[0.72]	[0.77]		[0.09]	[0.106]	[0.73]		[0.12]	[0.70]	[0.42]
CO	0.60	3.96	0.06	MI	3.83	1.30	1.47	SC	0.01	0.21	0.84
	[0.44]	[0.05]	[0.82]		[0.05]	[0.26]	[0.23]		[0.91]	[0.65]	[0.36]
CT	0.02	1.12	0.10	MN	0.71	2.44	0.87	TN	2.26	1.26	0.16
	[0.88]	[0.29]	[0.75]		[0.40]	[0.12]	[0.35]		[0.14]	[0.27]	[0.69]
FL	1.07	1.97	2.40	MO	1.05	0.71	1.05	TX	1.87	1.39	0.37
	[0.30]	[0.17]	[0.13]		[0.31]	[0.40]	[0.31]		[0.18]	[0.24]	[0.55]
GA	2.88	0.91	0.04	MS	0.25	1.51	3.02	VA	0.20	1.32	0.88
	[0.09]	[0.34]	[0.84]		[0.62]	[0.22]	[0.09]		[0.65]	[0.25]	[0.35]
IA	0.16	1.01	0.06	NC	0.25	0.05	2.35	WA	1.32	0.15	2.89
	[0.69]	[0.32]	[0.80]		[0.62]	[0.83]	[0.13]		[0.25]	[0.70]	[0.09]
IL	0.89	0.33	2.44	NJ	0.09	0.62	0.61	WI	1.30	0.02	2.26
	[0.35]	[0.57]	[0.12]		[0.77]	[0.43]	[0.44]		[0.26]	[0.88]	[0.14]
IN	0.26	2.79	1.30	NY	1.28	1.49	0.91				
	[0.61]	[0.109]	[0.26]		[0.26]	[0.23]	[0.34]				
KY	0.11	3.96	0.00	OH	3.60	0.51	3.42				
	[0.74]	[0.05]	[0.99]		[0.06]	[0.48]	[0.07]				

Note: the model contains unrestricted intercepts and restricted trend coefficients with I(1) endogenous variables hpi and dy and I(1) exogenous variables p^*, y^* and rir. The figures is square brackets are the estimated t-values of the tests.

4.4 Contemporaneous effects of starred variables on their sector specific counterparts

Table 4 presents the contemporaneous effects of the starred variables on their state-level counterparts for robust t-ratios, computed using White's heteroscedasticity-consistent variance estimator. These values can be interpreted as impact elasticities between domestic and starred variables. Most of them are significant and have a positive sign. They are particularly informative as regards the linkages across states. Focusing on the New York state, for instance, we can see that a 1% change in the other states' weighted house prices in a given quarter leads to a statistically significant increase of 0.11% in the house prices in New York in the same quarter. Similar elasticities are obtained across the different states, though the magnitude of the effects tends to be slightly stronger.

We can also observe high elasticities between real income per capita, implying a strong comovement across states in the income level. Moreover, in all cases, again, this elasticity is significant and in most cases it is higher than the elasticities we found for house price developments.

Table 4: Contemporaneous effects of foreign variables on their domestic counterparts in state-specific models

		p^*	y^*		p^*	y^*		p^*	y^*
AL	coef	0.409	0.549	LA	0.750	0.613	OK	0.827	0.681
	t-ratio	[11.93]	[17.30]		[5.60]	[8.44]		[7.99]	[10.98]
AZ	coef	0.237	0.682	MA	0.674	0.842	OR	0.580	0.692
	t-ratio	[13.74]	[8.95]		[20.76]	[14.74]		[8.36]	[9.44]
CA	coef	0.612	0.826	MD	0.161	0.392	PA	0.698	0.432
	t-ratio	[8.99]	[12.30]		[14.74]	[17.62]		[20.20]	[22.16]
CO	coef	0.454	0.948	MI	0.820	0.928	SC	0.705	0.730
	t-ratio	[7.38]	[11.38]		[8.94]	[8.28]		[11.12]	[12.89]
CT	coef	0.500	0.601	MN	0.378	0.864	TN	0.366	0.681
	t-ratio	[26.67]	[15.47]		[11.57]	[15.25]		[11.39]	[16.43]
FL	coef	0.430	0.703	MO	0.130	0.397	TX	0.815	0.876
	t-ratio	[11.10]	[12.32]		[11.46]	[16.24]		[11.58]	[18.44]
GA	coef	0.503	0.422	MS	0.328	0.447	VA	0.961	0.979
	t-ratio	[17.78]	[14.93]		[4.79]	[9.94]		[15.21]	[12.08]
IA	coef	0.325	0.559	NC	0.906	0.568	WA	0.947	0.980
	t-ratio	[8.16]	[8.41]		[11.84]	[15.65]		[6.88]	[10.00]
IL	coef	0.246	0.387	NJ	0.944	0.680	WI	0.239	0.365
	t-ratio	[14.95]	[16.45]		[32.07]	[17.12]		[8.38]	[17.27]
IN	coef	0.997	0.611	NY	0.105	0.360			
	t-ratio	[12.31]	[20.78]		[15.78]	[10.55]			
KY	coef	0.670	0.836	OH	0.376	0.981			
	t-ratio	[8.12]	[15.65]		[13.55]	[17.09]			

Note: The figures in parenthesis are the t-statistics.

5 Dynamic properties of the integrated model

Due to the simultaneous nature of the state-specific models, a more satisfactory approach to the analysis of dynamics and interdependencies (both on impact and over time) among the various factors would be via impulse response functions computed from the solution to the GVAR model. In our analysis we rely on the computation of generalised impulse response functions (GIRFs) as advanced by Koop, Pesaran, and Potter (1996) for non-linear models and developed further in Persaran and Shin (1998) for vector error-correcting models.[12] In the absence of strong a priori beliefs on the ordering of the variables and/or sectors in the GVAR model, the GIRFs provide useful information with respect to changes in real income per capita, house prices or interest rates. Although the approach is silent as to the specific structural factors behind the changes, the GIRFs can be quite informative about the dynamics of the transmission of shocks. The GIRFs identify the shocks as interept shifts in the various equations using a historical variance-covariance matrix of the errors.

Impulse responses are presented for 24 quarters following the imposition of a shock. Charts 7 to 10 display the bootstrap estimates of the GIRFs.[13] Moreover, we present the associated 90% confidence bands for the maximum and minimum response.

From the GIRFs and the GVAR's eigenvalues, we can conclude that the model is stable.

We investigate in this paper the implication of four different shocks:

- A 10% shock to Californian house prices
- A 10% shock to New York house prices
- A 10% shock to Texan house prices
- A 100 bps increase in the real 10-year government bond yield

We choose these three states not only because of the ranging geographical location but also for their differences in the housing market. In fact, California at present is generally seen to have an overvalued housing market and supply is lacking.[14] Similarly, New York has seen rapidly rising house prices in recent years.[15] This contrasts however

[12]The GIRF is an alternative to the Orthongonalised Impulse Responses (OIR) of Sims (1980). The OIR approach requires the impulse responses to be computed with respect to a set of orthogonalised shocks, whilst the GIR approach considers shocks to individual errors and integrates out the effects of the other shocks using the observed distribution of all the shocks without any orthogonalisation. Unlike the OIR, the GIRF is invariant to the ordering of the variables and the countries in the GVAR model, which is clearly an important consideration given various possible alternative orderings. Even if a suitable ordering of the variables in a given state model can be arrived at from economic theory or general *a priori* reasoning, it is not clear how to order states in the application of the OIR to the GVAR model.

[13]The computations are carried out using a sieve bootstrap procedure as reported in Dées, DiMauro, Pesaran, and Smith (2005).

[14]In this context see for instance Peterson (1996) who finds that California has the lowest housing affordability currently in the US whilst Himmelberg, Mayer, and Sinai (2005) find that southern California is one of the two areas in the US that in the fourth quarter of 2005 appeared relatively expensive, based on imputed rent-to-income ratios.

[15]In a recent study by Holly, Pesaran, and Yamagata (2006) it is shown that although no house price bubble is present at the national level, there may be a bubble present in Califoria, New York and Massachusetts.

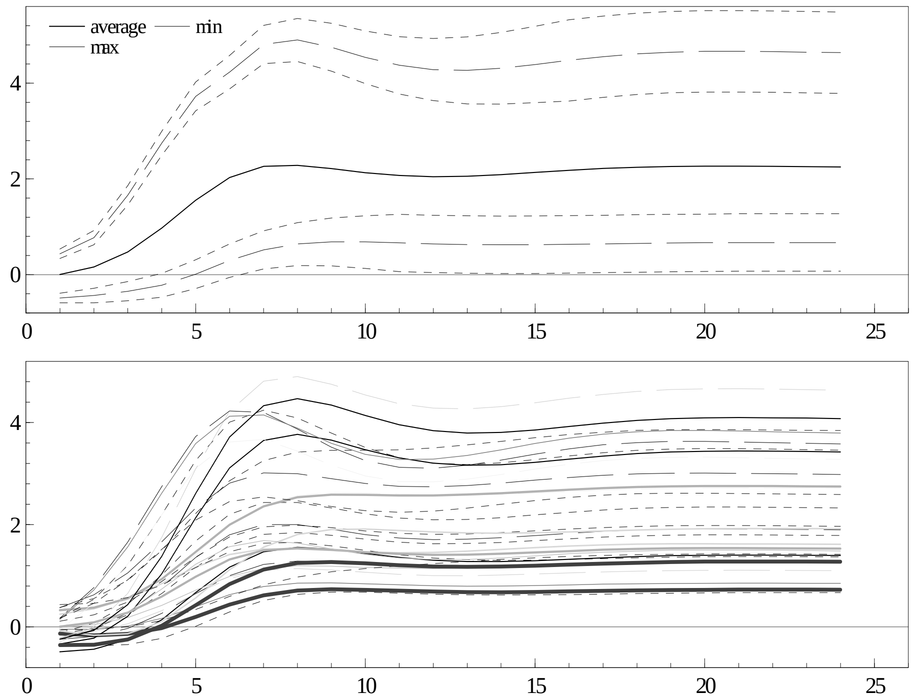

Figure 3: Impulse response to a 10% shock in house prices in California

with Texan house prices which by various measures do not appear to be overvalued. In part this may be related to the fact that in this region land supply is relatively abundant. For that reason, we would also *a priori* expect that a shock to house prices in Texas would spillover much less than one in California or New York.

Figure 3 displays the impacts of shocks to California house prices on house prices in the rest of the US. Moreover the top panel of the chart also includes the 90% confidence bands for the maximum and minimum response. On impact, a 10% increase in Californian house prices in the long run causes prices in all housing markets to increase as well, but by smaller amounts: between 0.6% and 4.6% in the long run. Moreover, all responses are statistically significant. The effect peaks after around 8 quarters and then to come down marginally. The strongest responses are found in Oregon, followed by Washington, Arizona and Texas whereas the states with the weakest responses are Connecticut and Massachusetts (see Table 5 for details). This result suggests that in fact closer states tend to be more affected by the shock than states which are located further away. However, at the same time, distance seems not to be the sole factor. Indeed, when considering the distance weights that were calculated as an input into the GVAR model, the ordering for California in the impulse responses deviates somewhat from that.

When looking at the impact of a 10% shock to house prices in New York we tend to find the similar result in that nearby states show the strongest impact. In this case we find it to be New Jersey, where house prices after 24 quarters are about 2.4% higher.

Table 5: Impulse response results after 24 quarters

10% shock to CA p		10% shock to NY p		10% shock to TX p		100 bps shock to rir	
CT	0.669	AZ	0.346	WA	0.176	NJ	-2.288
MA	0.729	CA	0.379	MA	0.192	CT	-2.071
MD	0.853	MS	0.443	WI	0.224	AZ	-1.983
IN	1.100	CO	0.571	CA	0.289	TX	-1.967
MI	1.258	OK	0.654	MI	0.331	CO	-1.884
TN	1.274	TN	0.661	MD	0.335	MA	-1.834
OH	1.291	GA	0.672	CT	0.339	NY	-1.820
NC	1.370	WA	0.709	NC	0.360	OK	-1.767
AL	1.400	AL	0.722	NY	0.368	FL	-1.652
NJ	1.423	OR	0.744	OR	0.375	VA	-1.559
VA	1.489	IA	0.772	NJ	0.392	MD	-1.545
SC	1.534	MN	0.798	SC	0.404	MN	-1.399
KY	1.617	WI	0.844	KY	0.426	CA	-1.272
GA	1.789	MO	0.928	VA	0.470	IL	-1.260
MS	1.900	KY	0.986	AZ	0.500	PA	-1.240
PA	1.922	IL	1.025	PA	0.505	MS	-1.233
NY	1.971	TX	1.052	IL	0.518	LA	-1.193
FL	2.331	SC	1.210	IN	0.527	MO	-1.190
WI	2.590	LA	1.352	GA	0.613	WI	-1.124
IL	2.711	IN	1.413	MN	0.682	SC	-1.092
LA	2.749	MI	1.434	TN	0.713	IA	-1.074
IA	2.987	FL	1.557	OH	0.723	OR	-1.050
OK	3.280	VA	1.706	FL	0.786	AL	-1.039
MN	3.425	NC	1.772	IA	0.864	WA	-1.028
MO	3.460	OH	1.791	AL	0.900	GA	-0.998
CO	3.584	MD	1.865	MO	0.910	TN	-0.991
TX	3.796	MA	1.977	CO	0.943	MI	-0.985
AZ	3.844	PA	2.007	MS	1.012	IN	-0.924
WA	4.078	CT	2.113	LA	1.072	KY	-0.829
OR	4.637	NJ	2.406	OK	1.219	OH	-0.781
CA	10.000	NY	10.000	TX	10.000	NC	-0.757

Note: The results are for each impulse response sorted according to the magnitude

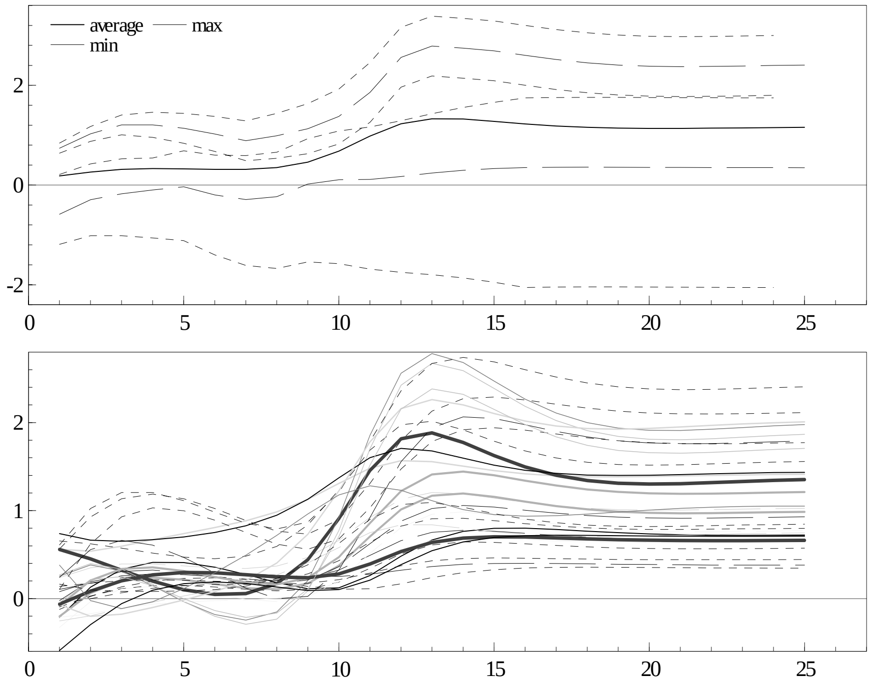

Figure 4: Impulse response to a 10% shock in house prices in New York

These responses are also statistically significant. The lowest responses are found in Arizona and California. These responses are however stastistically not significant. On average, the long run response to a 10% shock to New York house prices is around 1.15%, so half the response we found in the case of California. Moreover, the transmission of a shock to New York house prices to the other regions is also slower than is the case for California, and peaks after around 13 quarters (see Figure 4). The faster and stronger transmission of a house price shock in California to other states may reflect the strong economic linkages California has with the other US states, even those that are geographically very distant. This may reflect the importance of the growth in aerospace and information technology which orginates largely from California and has also been an important factor stimulating growth in other US states (see for this argument also Holly, Pesaran, and Yamagata (2006)).

Finally, a shock to Texan house prices has the lowest impact of the three state-level housing shocks we consider (see Figure 5). In fact, after 24 quarters a 10% shock to Texan house prices still only results in an average response of around 0.51% in the other states. The strongest impact is found in this case to be in Oklahoma and Louisiana while the weakest is recorded in Washington and Massachusetts. However none of the responses are statistically significant. This confirms our a priori belief than in fact a shock to Texan house prices results relatively swiftly in an increase in housing supply, which would keep spillover effects to other regions limited.

A last shock we consider is a 100 basis points increase in the real 10-year government bond yield. The results, as shown in Figure 6 would suggest that such an increase

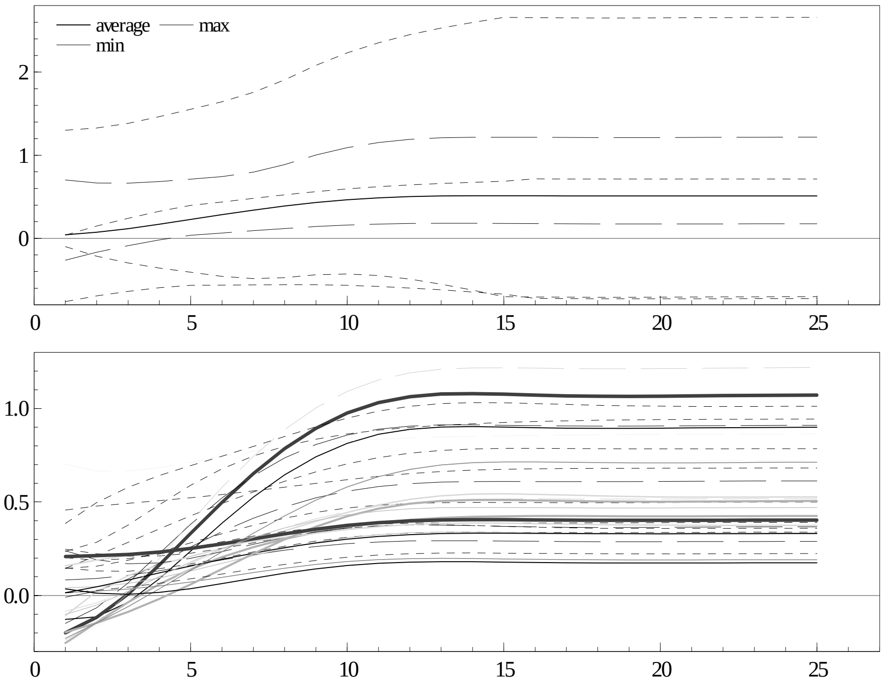

Figure 5: Impulse response to a 10% shock in house prices in Texas

would have only a limited, but statistically significant, impact on house prices across the US, resulting in lower house prices by between 0.7% and 2.3% in the longer run. This finding would suggest that monetary policy has not played a major role in the recent US house price rally. This is further evidenced when we consider in Table 6 the contribution of the decline in long-term interest rates to house price inflation since 2000. On average 36% of the change in house prices since 2000Q1 appears to be explained by the movements in interest rate. In some states, however, the role of real interest rates has been much smaller such as California and Florida. On the other hand, in Texas house price developments appear to be explained by up to 90% by the movement in the real long term interest rate.

Finally, when considering the extent and nature of state-level heterogeneity, we find it to be generally consistent with Carlino and DeFina (1999) and Fratantoni and Schuh (2003). In the first study, the responses of income in nine Census regions to monetary tightening are reported. In the second, the response of income, housing investment and housing appreciation is reported for a range of MSAs. In fact for the latter the variation in housing appreciation is larger, however, this may be because the regions considered are more disaggregated. Overall, this state-level heterogeneity we find in response to a monetary policy tightening could occur for various reasons. However, given that our model is of an empirical nature, it remains silent as regards the structural reasons for our findings. However, various studies have already looked into this issue and offer plentiful options. In our case, the following may be the most plausible: *(1)* different industrial composition with some industries more interest rate

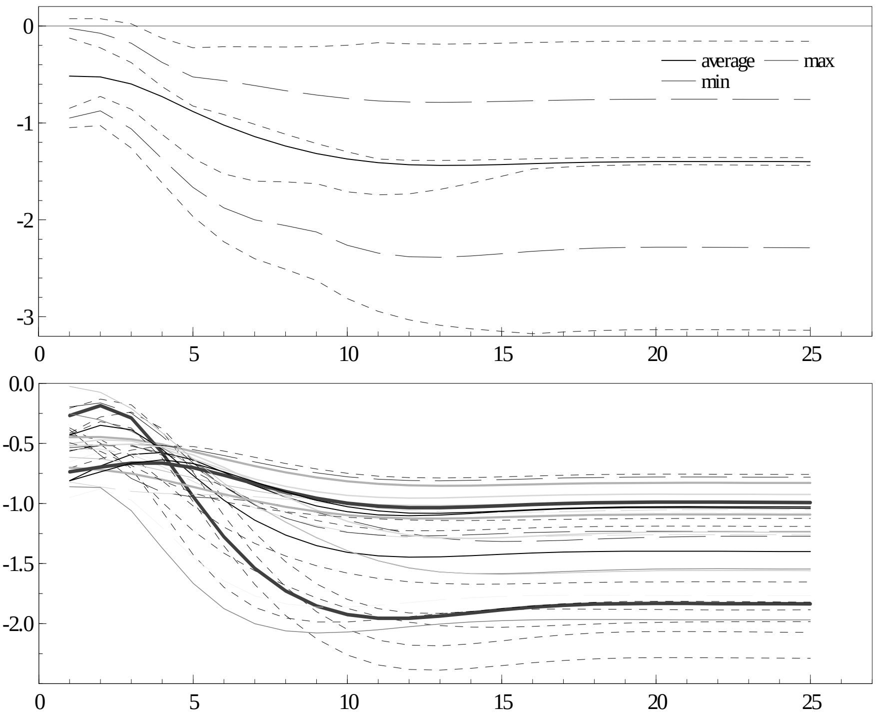

Figure 6: Impulse response to a 100 bps increase in the real 10-year government bond yield

sensitive than others (Hayo and Uhlenbrock (2000)), *(2)* differences in demographic composition (Mankiw and Weil (1989)), *(3)* variation in labour market conditions and human capital stocks (Johnes and Hyclak (1999)), *(4)* variation in federal, state, and tax policies (Poterba (1984), Poterba (1991)), *(5)* state-level differences in government land regulation and public goods provisions (Malpezzi (1996), Mayer and Somerville (1997)) and *(6)* financial conditions such as leverage and loan-to-value ratios (Lamot and Stein (1999)).

6 Concluding Remarks

In this paper we presented a quarterly nationwide US model in which state level real house prices and income per capita are related to state-specific foreign variables and US real interest rates by means of vector error correction models at the state level. We estimated the model for the 31 biggest states in the US over the period 1986-2005. Our approach advances research in this area in terms of methodology and focus. In terms of methodology, we make use of the rather novel GVAR approach, akin to Dées, DiMauro, Pesaran, and Smith (2005) and Pesaran, Schuermann, and Weiner (2004), to address the link between state-level and national and interstate factors. In terms of focus, in line with DelNegro and Otrok (2005) we consider the role of monetary policy on housing price developments. However we also go beyond that and consider the importance of spillovers across regions to assess the impact of a state-level housing price shock to

Table 6: Contribution of interest rate developments to house price inflation between 2000Q1 and 20005Q4

AL	38.45	FL	16.70	KY	37.43	MN	25.81	NY	24.57	SC	35.75
AZ	23.79	GA	33.49	LA	35.71	MO	35.38	OH	43.72	TN	41.67
CA	11.93	IA	49.03	MA	27.19	MS	53.35	OK	73.33	TX	90.77
CO	64.19	IL	29.72	MD	17.27	NC	29.14	OR	18.93	VA	19.91
CT	31.69	IN	60.14	MI	47.89	NJ	27.84	PA	23.30	WA	18.91
										WI	31.91

national developments. The results suggest that the importance of spillovers is state dependent. The largest response was found when imposing a 10% shock to Californian house prices. This would result in long run responses in the range of 0.6% to 4.6% in the other states. Such an outcome would clearly provide also an upper bound for the type of spillovers we might expect from housing shocks in the euro area. Indeed, given that labour mobility is much lower in the euro area (especially across borders) and financial integration is less advanced, spillover effects in the euro area will likely be small.

As regards real interest rates, the impact on the US housing market appears to be relatively small with an increase of 100 basis points in the real 10-year government bond yield resulting in a long run fall in house prices of between 0.5 and 2.5%. This would suggest, in line with DelNegro and Otrok (2005) that interest rate movements have not been an important factor in explaining the recent house price rally in the United States.

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Appendices

A States covered

AL	ALABAMA
AZ	ARIZONA
CA	CALIFORNIA
CO	COLORADO
CT	CONNECTICUT
FL	FLORIDA
GA	GEORGIA
IA	IOWA
IL	ILLINOIS
IN	INDIANA
KY	KENTUCKY
LA	LOUISIANA
MA	MASSACHUSETTS
MD	MARYLAND
MI	MICHIGAN
MN	MINNESOTA
MO	MISSOURI
MS	MISSISSIPPI
NC	NORTH CAROLINA
NJ	NEW JERSEY
NY	NEW YORK
OH	OHIO
OK	OKLAHOMA
OR	OREGON
PA	PENNSYLVANIA
SC	SOUTH CAROLINA
TN	TENNESSEE
TX	TEXAS
VA	VIRGINIA
WA	WASHINGTON
WI	WISCONSIN

B Aggregation weights and house price developments

TABLE A1. Distance weights

	AL	AZ	CA	CO	CT	FL	GA	IA	IL	IN	KY
AL	0	0.0159	0.0137	0.0150	0.0097	0.0445	0.0714	0.0183	0.0209	0.0221	0.0343
AZ	0.0089	0	0.0523	0.0314	0.0043	0.0109	0.0081	0.0107	0.0085	0.0074	0.0074
CA	0.0068	0.0463	0	0.0204	0.0037	0.0087	0.0064	0.0085	0.0068	0.0059	0.0059
CO	0.0116	0.0430	0.0315	0	0.0055	0.0128	0.0103	0.0186	0.0131	0.0108	0.0104
CT	0.0135	0.0105	0.0103	0.0098	0	0.0186	0.0158	0.0124	0.0137	0.0156	0.0148
FL	0.0292	0.0128	0.0114	0.0109	0.0088	0	0.0393	0.0113	0.0119	0.0126	0.0158
GA	0.0696	0.0141	0.0125	0.0130	0.0111	0.0583	0	0.0158	0.0183	0.0206	0.0307
IA	0.0185	0.0193	0.0173	0.0245	0.0091	0.0174	0.0164	0	0.0534	0.0301	0.0233
IL	0.0236	0.0170	0.0153	0.0192	0.0111	0.0205	0.0212	0.0594	0	0.0733	0.0404
IN	0.0261	0.0155	0.0141	0.0165	0.0132	0.0227	0.0250	0.0351	0.0768	0	0.0620
KY	0.0403	0.0155	0.0139	0.0159	0.0125	0.0283	0.0370	0.0270	0.0421	0.0617	0
LA	0.0382	0.0198	0.0159	0.0180	0.0073	0.0307	0.0259	0.0171	0.0167	0.0155	0.0193
MA	0.0123	0.0102	0.0100	0.0095	0.1068	0.0173	0.0143	0.0117	0.0126	0.0141	0.0133
MD	0.0186	0.0117	0.0111	0.0111	0.0352	0.0239	0.0229	0.0151	0.0179	0.0222	0.0222
MI	0.0178	0.0143	0.0135	0.0152	0.0150	0.0184	0.0180	0.0304	0.0409	0.0464	0.0271
MN	0.0136	0.0180	0.0172	0.0226	0.0086	0.0143	0.0127	0.0479	0.0275	0.0203	0.0160
MO	0.0275	0.0198	0.0168	0.0231	0.0091	0.0216	0.0219	0.0514	0.0505	0.0334	0.0333
MS	0.0709	0.0181	0.0150	0.0172	0.0083	0.0351	0.0363	0.0192	0.0202	0.0195	0.0266
NC	0.0267	0.0123	0.0114	0.0114	0.0179	0.0360	0.0402	0.0147	0.0175	0.0213	0.0264
NJ	0.0154	0.0110	0.0106	0.0103	0.0765	0.0209	0.0185	0.0133	0.0152	0.0178	0.0173
NY	0.0145	0.0113	0.0109	0.0108	0.0548	0.0186	0.0166	0.0147	0.0166	0.0193	0.0173
OH	0.0229	0.0138	0.0128	0.0140	0.0180	0.0229	0.0248	0.0239	0.0349	0.0586	0.0416
OK	0.0206	0.0274	0.0205	0.0333	0.0068	0.0190	0.0164	0.0256	0.0199	0.0161	0.0170
OR	0.0063	0.0242	0.0437	0.0177	0.0038	0.0079	0.0060	0.0086	0.0067	0.0059	0.0057
PA	0.0179	0.0119	0.0114	0.0115	0.0352	0.0219	0.0211	0.0163	0.0197	0.0247	0.0229
SC	0.0380	0.0130	0.0118	0.0120	0.0137	0.0481	0.0779	0.0151	0.0179	0.0212	0.0294
TN	0.0670	0.0158	0.0138	0.0157	0.0114	0.0336	0.0516	0.0231	0.0306	0.0361	0.0861
TX	0.0189	0.0285	0.0201	0.0246	0.0059	0.0197	0.0153	0.0159	0.0136	0.0119	0.0132
VA	0.0223	0.0121	0.0114	0.0115	0.0243	0.0280	0.0291	0.0156	0.0189	0.0240	0.0264
WA	0.0063	0.0211	0.0332	0.0169	0.0039	0.0079	0.0060	0.0088	0.0069	0.0060	0.0057
WI	0.0167	0.0161	0.0151	0.0184	0.0112	0.0169	0.0160	0.0517	0.0502	0.0362	0.0232

Note: Distance weights are based on own computations and are derived as the inverse of the average distance between the metropolitan statistical areas of the various states, corrected for rescaling so that the column sums to one.

TABLE A2. Distance weights

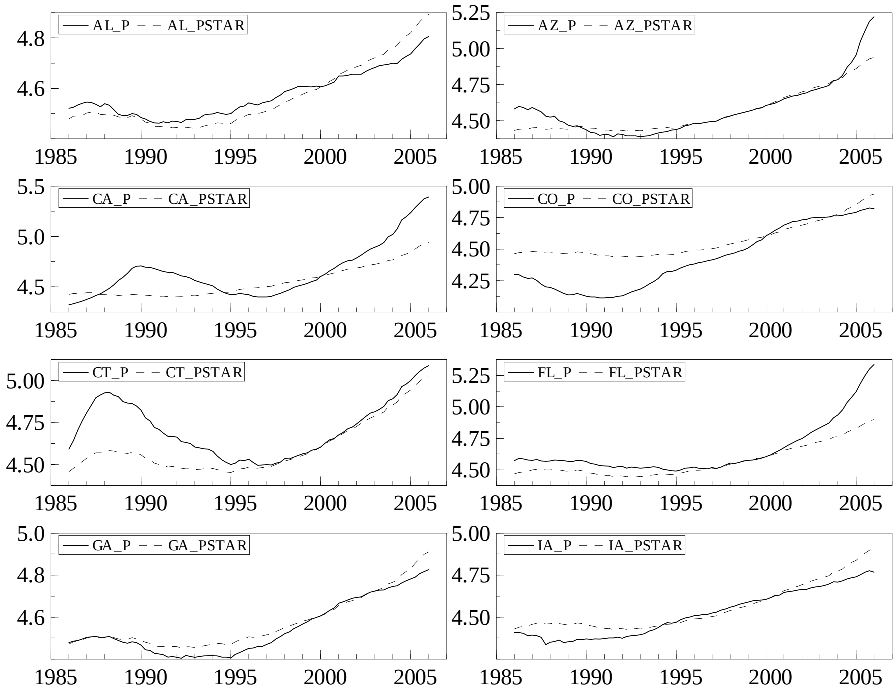
	LA	MA	MD	MI	MN	MO	MS	NC	NJ	NY
AL	0.0457	0.0084	0.0118	0.0174	0.0157	0.0264	0.0740	0.0247	0.0110	0.0121
AZ	0.0133	0.0039	0.0042	0.0078	0.0117	0.0107	0.0106	0.0064	0.0044	0.0053
CA	0.0095	0.0034	0.0035	0.0066	0.0099	0.0080	0.0078	0.0052	0.0038	0.0045
CO	0.0166	0.0050	0.0054	0.0114	0.0201	0.0171	0.0138	0.0081	0.0057	0.0069
CT	0.0121	0.1009	0.0310	0.0202	0.0138	0.0121	0.0121	0.0229	0.0759	0.0632
FL	0.0241	0.0077	0.0099	0.0118	0.0108	0.0136	0.0241	0.0219	0.0098	0.0101
GA	0.0302	0.0095	0.0142	0.0171	0.0143	0.0205	0.0369	0.0362	0.0129	0.0134
IA	0.0207	0.0081	0.0097	0.0301	0.0560	0.0500	0.0203	0.0137	0.0097	0.0124
IL	0.0225	0.0097	0.0128	0.0449	0.0357	0.0547	0.0238	0.0182	0.0122	0.0156
IN	0.0219	0.0113	0.0166	0.0534	0.0277	0.0379	0.0240	0.0233	0.0150	0.0190
KY	0.0272	0.0107	0.0166	0.0311	0.0216	0.0376	0.0327	0.0287	0.0145	0.0169
LA	0	0.0064	0.0080	0.0134	0.0147	0.0247	0.0818	0.0145	0.0079	0.0090
MA	0.0113	0	0.0233	0.0186	0.0132	0.0112	0.0112	0.0197	0.0442	0.0489
MD	0.0152	0.0251	0	0.0255	0.0156	0.0155	0.0158	0.0429	0.0633	0.0453
MI	0.0165	0.0130	0.0166	0	0.0319	0.0245	0.0169	0.0190	0.0162	0.0234
MN	0.0153	0.0078	0.0086	0.0270	0	0.0241	0.0145	0.0114	0.0089	0.0118
MO	0.0308	0.0080	0.0102	0.0248	0.0290	0	0.0323	0.0160	0.0099	0.0120
MS	0.0938	0.0073	0.0096	0.0158	0.0160	0.0297	0	0.0181	0.0092	0.0104
NC	0.0188	0.0145	0.0295	0.0200	0.0142	0.0166	0.0204	0	0.0231	0.0209
NJ	0.0133	0.0422	0.0561	0.0221	0.0144	0.0133	0.0135	0.0298	0	0.0614
NY	0.0129	0.0401	0.0345	0.0275	0.0163	0.0139	0.0131	0.0232	0.0528	0
OH	0.0186	0.0149	0.0262	0.0553	0.0225	0.0242	0.0200	0.0297	0.0215	0.0279
OK	0.0358	0.0061	0.0072	0.0148	0.0207	0.0351	0.0284	0.0116	0.0072	0.0086
OR	0.0083	0.0035	0.0035	0.0068	0.0108	0.0077	0.0070	0.0051	0.0038	0.0047
PA	0.0150	0.0255	0.0806	0.0310	0.0171	0.0162	0.0156	0.0336	0.0531	0.0624
SC	0.0229	0.0115	0.0191	0.0184	0.0141	0.0182	0.0259	0.0736	0.0165	0.0163
TN	0.0333	0.0097	0.0146	0.0236	0.0189	0.0345	0.0446	0.0288	0.0131	0.0147
TX	0.0425	0.0053	0.0062	0.0112	0.0144	0.0202	0.0269	0.0103	0.0063	0.0073
VA	0.0170	0.0187	0.0656	0.0244	0.0155	0.0168	0.0181	0.0754	0.0352	0.0300
WA	0.0081	0.0036	0.0036	0.0071	0.0113	0.0077	0.0069	0.0051	0.0039	0.0048
WI	0.0170	0.0100	0.0118	0.0601	0.0569	0.0294	0.0170	0.0151	0.0119	0.0161

Note: Distance weights are based on own computations and are derived as the inverse of the average distance between the metropolitan statistical areas of the various states, corrected for rescaling so that the column sums to one.

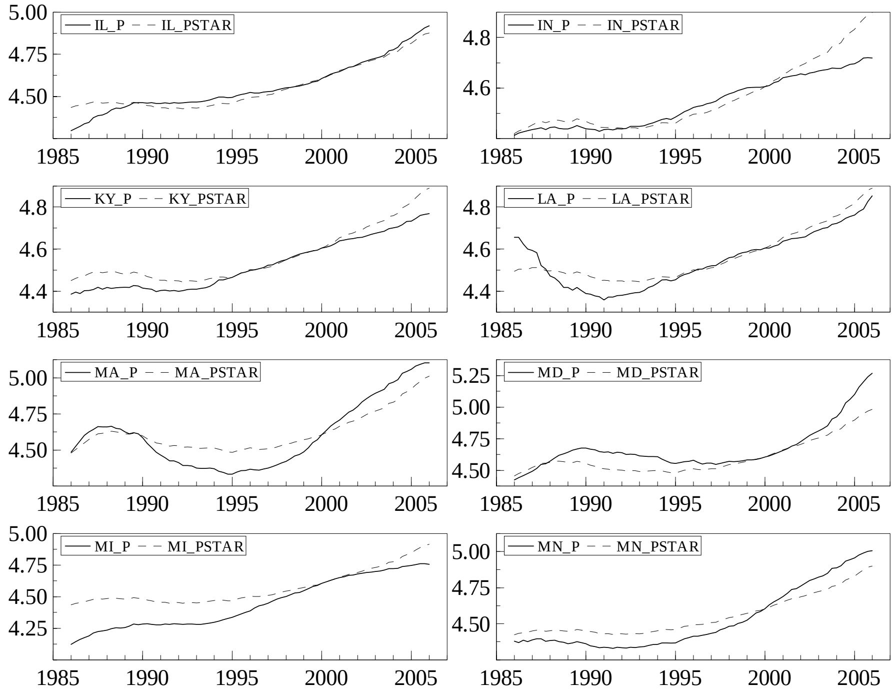
TABLE A3. Distance weights

	OH	OK	OR	PA	SC	TN	TX	VA	WA	WI
AL	0.0193	0.0233	0.0120	0.0132	0.0376	0.0595	0.0270	0.0168	0.0123	0.0170
AZ	0.0065	0.0174	0.0260	0.0049	0.0072	0.0079	0.0229	0.0051	0.0232	0.0092
CA	0.0054	0.0115	0.0415	0.0042	0.0058	0.0061	0.0143	0.0043	0.0323	0.0076
CO	0.0090	0.0290	0.0260	0.0065	0.0092	0.0107	0.0271	0.0066	0.0255	0.0144
CT	0.0210	0.0106	0.0100	0.0359	0.0188	0.0140	0.0117	0.0252	0.0105	0.0157
FL	0.0127	0.0141	0.0099	0.0106	0.0312	0.0196	0.0185	0.0138	0.0101	0.0112
GA	0.0203	0.0181	0.0111	0.0152	0.0751	0.0446	0.0213	0.0213	0.0114	0.0158
IA	0.0204	0.0294	0.0166	0.0122	0.0151	0.0209	0.0231	0.0118	0.0174	0.0532
IL	0.0331	0.0254	0.0145	0.0163	0.0199	0.0307	0.0220	0.0160	0.0152	0.0574
IN	0.0582	0.0215	0.0133	0.0215	0.0247	0.0379	0.0202	0.0213	0.0139	0.0433
KY	0.0412	0.0226	0.0127	0.0198	0.0342	0.0899	0.0222	0.0233	0.0132	0.0277
LA	0.0130	0.0338	0.0131	0.0092	0.0189	0.0247	0.0507	0.0107	0.0132	0.0144
MA	0.0184	0.0101	0.0098	0.0276	0.0166	0.0127	0.0112	0.0206	0.0104	0.0148
MD	0.0347	0.0128	0.0106	0.0935	0.0298	0.0204	0.0140	0.0776	0.0111	0.0188
MI	0.0477	0.0171	0.0133	0.0234	0.0187	0.0215	0.0165	0.0187	0.0142	0.0626
MN	0.0164	0.0203	0.0178	0.0109	0.0121	0.0145	0.0179	0.0101	0.0192	0.0500
MO	0.0212	0.0413	0.0152	0.0124	0.0188	0.0319	0.0300	0.0131	0.0157	0.0311
MS	0.0161	0.0308	0.0128	0.0110	0.0245	0.0380	0.0368	0.0130	0.0130	0.0165
NC	0.0271	0.0141	0.0105	0.0268	0.0788	0.0276	0.0160	0.0613	0.0109	0.0165
NJ	0.0253	0.0115	0.0102	0.0547	0.0228	0.0162	0.0126	0.0369	0.0107	0.0168
NY	0.0282	0.0117	0.0107	0.0552	0.0194	0.0157	0.0126	0.0271	0.0114	0.0197
OH	0	0.0170	0.0122	0.0372	0.0276	0.0302	0.0171	0.0321	0.0128	0.0327
OK	0.0126	0	0.0170	0.0086	0.0137	0.0185	0.0590	0.0091	0.0170	0.0179
OR	0.0054	0.0101	0	0.0042	0.0056	0.0058	0.0116	0.0042	0.1261	0.0080
PA	0.0425	0.0132	0.0110	0	0.0260	0.0203	0.0141	0.0491	0.0115	0.0214
SC	0.0235	0.0157	0.0107	0.0193	0	0.0354	0.0181	0.0319	0.0111	0.0160
TN	0.0287	0.0236	0.0125	0.0168	0.0394	0	0.0243	0.0210	0.0128	0.0222
TX	0.0101	0.0466	0.0155	0.0072	0.0125	0.0151	0	0.0079	0.0152	0.0128
VA	0.0360	0.0138	0.0107	0.0481	0.0420	0.0249	0.0151	0	0.0112	0.0186
WA	0.0055	0.0098	0.1229	0.0043	0.0056	0.0058	0.0111	0.0043	0	0.0083
WI	0.0271	0.0199	0.0150	0.0155	0.0156	0.0195	0.0180	0.0138	0.0160	0

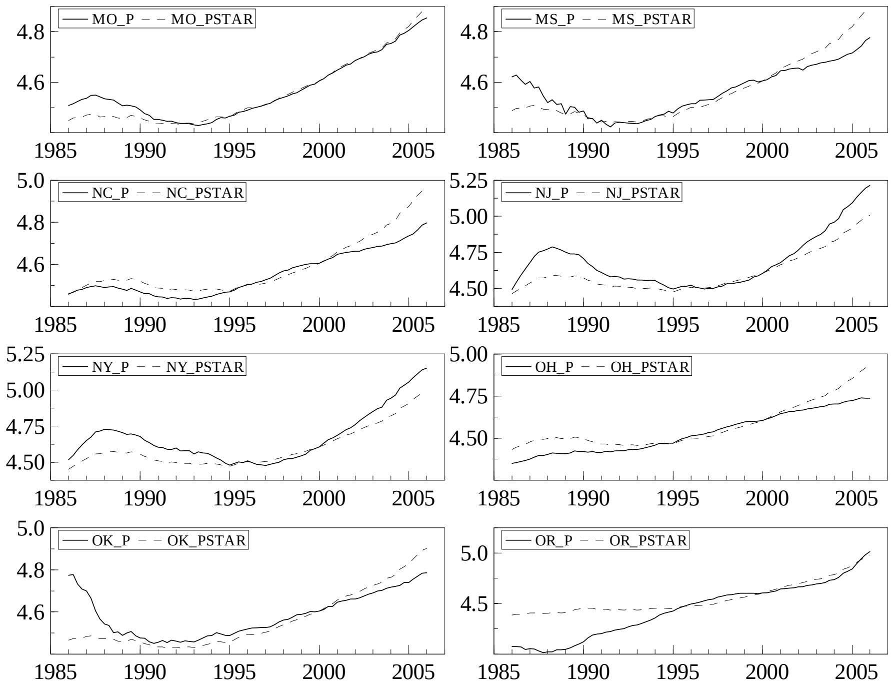
Note: Distance weights are based on own computations and are derived as the inverse of the average distance between the metropolitan statistical areas of the various states, corrected for rescaling so that the column sums to one.



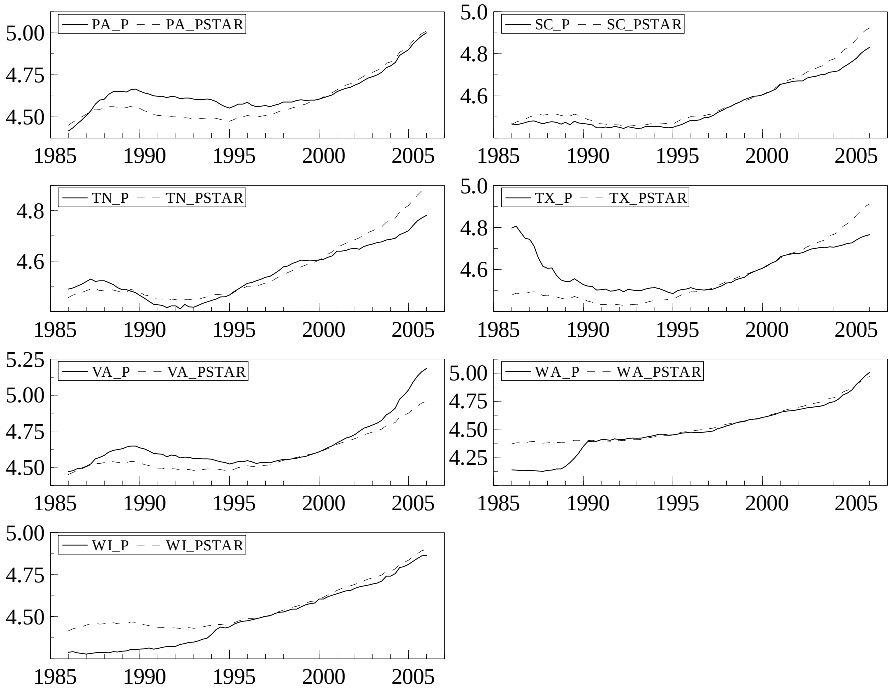
House price developments in state and state-specific starred prices



House price developments in state and state-specific starred prices



House price developments in state and state-specific starred prices



House price developments in state and state-specific starred prices

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