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Is Housing the Business Cycle? Evidence from U.S. Cities*

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Abstract

We analyze the relationship between housing and the business cycle in a set of 51 U.S. cities. Most surprisingly, we find that declines in house prices are often not followed by declines in employment. We also find that national permits are a better leading indicator for a city's employment than a city's own permits. [JEL: C32, E32, R11]

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1 Introduction

In a recent paper, Leamer (2007) argues that real estate markets are grossly understudied by macroeconomists interested in understanding business cycles. He asserts several stylized facts about the behavior of the national housing market over the business cycle including: 1) residential investment leads the business cycle and a fall in residential investment is a reliable harbinger of a recession, and 2) volumes, rather than house prices, are what matter for business cycles.

This paper furthers our understanding of the relationship between housing and the business cycle by exploiting the cross-sectional variation in both business cycles and housing markets across Metropolitan Statistical Areas (MSAs). Because housing is fundamentally a non-tradable good, the housing market is a city-level phenomenon; in the short to medium run, consumers may find it difficult to substitute between housing across cities. Glaeser and Gyourko (2007), for example, document that time dummies explain only about a quarter of the variation in city-level house price changes, suggesting that most of the variation in house prices comes from city-specific factors. Del Negro and Otrok (2007) similarly find that state and regional factors, rather than national factors, drive the majority of the movement in house prices.

In addition, a growing literature has documented substantial heterogeneity in both the timing and magnitude of business cycles at different levels of disaggregation. For example, Owyang, Piger, and Wall (2005) use state-level data to analyze what determines growth rates during recession and expansion phases; Owyang, Piger, Wall, and Wheeler (2008) apply a similar framework to U.S. cities. These studies find that the timing (and, potentially, the number) of recessionary experiences vary across regions.[1]

MSA-level variation enables us to identify empirical regularities in the relationship be-

[1] Carlino and DeFina (1998) and Fratantoni and Schuh (2003) show that the effects of monetary policy differ substantially across regions. The majority of this literature attributes these differences to industry composition and the makeup of the regional banking sector. See also Carlino and Sill (2001) and Crone (2005). Housing, which may be thought of as a key component in the propagation of monetary policy, has been largely neglected due to the highly aggregated (i.e., states or larger) geographic unit of analysis.

tween housing and the business cycle more robustly than with national data alone. While large real declines in house prices at the national level have been rare during the post-war period, several cities in our sample experienced large and sustained declines in house prices over the last 25 years. Understanding how house prices affect employment in these cities may help us to understand whether the national recession that began in December 2007 was due to factors that simultaneously lowered house prices and employment or whether declines in house prices themselves played a key role in driving economic activity.

We first seek to determine if the stylized facts Leamer identifies at the national level are also true for our sample of MSAs. Because residential investment is not available at the MSA level, we proxy for residential investment using permit data. We ask whether housing variables are robust leading indicators of employment after we control for national-level factors. Smets (2007) points out that the appearance of residential investment as a leading indicator may be due to cycles in interest rates. Since housing is known to be sensitive to changes in interest rates (e.g., Hamilton, 2008), interest rate shocks may drive both residential investment and employment.

We also assess the relationship between housing prices and employment over the business cycle. Earlier work has not reached a consensus on the relationship between the business cycle and house prices: Iacoviello (2005) finds that a decline in aggregate housing prices leads to a decline in gross domestic product (GDP). Davis and Heathcote (2005) find that the contemporaneous correlation between national-level HP-filtered house prices and output is 65 percent over the 1971-2001 period. In contrast, Kan, Kwong, and Leung (2004) examine annual city-level data and find that the contemporaneous correlation between house prices and output growth is less than 15 percent.

Consistent with Leamer's findings, we find that house prices are poor leading indicators, but that increases in a city's permits do not always raise employment in that city. Instead, national permits are a more consistent leading indicator for employment at the city level. At the national level, we find that permit shocks raise employment after we control for financial factors and oil prices. We also explore the possibility that the housing market influences employment over the business cycle in a nonlinear fashion. In particular, we use a Markov-switching model with time-varying transition probabilities (TVTPs) to ascertain whether housing variables are significant in influencing the probabilities that cities move between expansion and recession regimes. The findings from our Markov-switching model do not support the notion that housing variables influence the probability of being in a recession.

The remainder of the paper proceeds as follows: Section 2 describes our data and documents the correlations between housing indicators and employment at the national and MSA level. Section 3 presents the results from the VAR used to identify whether, and which, housing variables are good leading indicators for MSA-level employment. Section 4 provides a preliminary comparison of local business cycles and the MSA-level housing variables. Section 5 presents the results of our TVTP Markov-switching model formalizing the notion that housing drives cyclical fluctuations. Section 6 concludes.

2 A Preliminary Look at Local Housing Cycles

2.1 Data

Our goal is to examine the relationship between housing and economic activity for a cross-section of cities. Unfortunately, residential investment – a key housing variable – is unavailable at the MSA level. We can, however, proxy for residential investment using permits, either in units or values. There is a substantial amount of high-frequency variation in our permit series, with all series exhibiting strong low-frequency variation. This requires us to somehow filter the data before establishing anything about the relationships among the series. We isolate the business cycle relationships by filtering the data in the frequency domain with the optimal band-pass filter suggested by Corbae, Ouliaris, and Phillips (2002) and Corbae and Ouliaris (2006).[2]

[2]This approach has the advantage of not forcing us to take a stand on the relationship between the variables at low frequencies. The filter extracts the component of each series associated with cycles of 6 to 32 quarters. We also looked at linearly detrending the series; the results suggested an even weaker

The U.S. Census Bureau (2009) summarizes the relationship between permits and residential investment at the aggregate level as follows: "Current surveys indicate that construction is undertaken for all but a very small percentage of housing units authorized by building permits. A major portion typically get under way during the month of permit issuance and most of the remainder begin within the three following months." Indeed, the contemporaneous correlation between band-pass filtered residential investment and band-pass filtered permit values at the national level is 98 percent.

In addition to seasonally-adjusted permits (both units and values), we use house prices as a secondary measure of the housing market. Our nominal house price series are the Freddie Mac Conventional Mortgage House Price Indices (CMHPIs). Real house and permit values are constructed for the analysis. Our measure of economic activity is MSA-level seasonally-adjusted non-farm employment.

Data availability dictates our choice of cities (51 in total) and sample period. Table 1 lists the cities in our sample, the deflator used to convert nominal variables into real variables, and the sample period for which all MSA-level series are available.[3] Although permit data begin for most cities in our sample in 1982Q1, we chose to use 1983Q1 as the starting point for our analysis to avoid starting our analysis in the middle of a national recession. We also include a few cities for which permit data do not become available until 1984Q1. Our data end in 2008Q4.

Finally, we include several national-level variables that may affect both housing variables and employment. To control for the stance of monetary policy, we include the federal funds rate. We also include national-level core consumer price index (CPI) inflation, the 30-year conventional mortgage rate, and the spread between 3-month commercial paper and 90-day Treasury bills as a measure of credit market turmoil.

relationship between housing variables and employment than when we band-pass filtered the series.

[3] Our sample excludes a few large and economically important cities which lack either permit or employment data for the full sample: permit data for New York begin in 1988 while employment data for Boston and Washington, DC, start in 1990.

2.2 Correlations

One of Leamer's findings is that residential investment – though a small component of GDP – is a large factor in economic fluctuations around turning points. A similar relationship appears to exist between our national housing variables and national employment (Figure 1). The major peaks and troughs in employment correspond fairly closely with NBER recession dates; exceptions are that the 1991 and 2001 troughs are delayed in the employment data because of the jobless recoveries. For house prices, the cyclical component is dominated by the period from 2004-2008. As a result, in computing a simple correlation over the whole sample period, house prices appear to lead the business cycle. However, as the top panel of Figure 1 illustrates, there does not appear to be a strong relationship between the two variables for the initial portion of the sample, even at turning points. As the bottom two panels show, permits – measured in units or values – appear to be a more consistent leading indicator of the national cycle with the highest correlation occurring at a lead of 5 quarters (Table 2).

Housing's performance as a leading indicator is weaker when examined at the city level, where the average correlation between MSA house prices and MSA employment is roughly half that of national house prices and national employment. Figures 2 and 3 demonstrate the cross-sectional diversity in the data by plotting both MSA-level housing series along with employment for a representative group of cities. Table 2 shows the average correlations between the housing series and employment at the MSA level. Altogether, house prices are procyclical in about 80 percent of our MSAs. However, the pattern differs substantially across MSAs. In 22 of the 51 MSAs, house prices lead employment, while 9 MSAs actually lag. Some MSAs (Augusta, Durham, New Orleans, Pittsburgh, and Wichita) even exhibit counter-cyclical house price behavior. As with the national data, the weak cyclical relationship between house prices and employment appears to be driven by the 2004-2008 period for many cities. Excluding the 2004-2008 period reduces the number of procyclical cities to 33, less than two-thirds.[4]

Permits, measured both in values and units, exhibit a more consistent pattern, leading the cycle at the national level and in 80 percent of our MSAs. The timing of the relationship among our MSAs is similar to the timing of the relationship at the national level. The similarity between the results for units and values suggests that a simple change in the composition of housing cannot explain the leading indicator property of residential investment. Rather, it appears to be the case that the cyclical pattern in housing is driven by changes along the extensive margin rather than the intensive margin. Finally, the correlation between lags of MSA permits and MSA employment is higher than that of lags of national permits and MSA employment.

The heterogeneity in the within-MSA correlations suggests a national series – a weighted average of a distribution of idiosyncratically moving series – may be a poor proxy for city-level effects. Table 2 verifies this notion, summarizing the contemporaneous correlations between MSA-level employment, house prices, and permits with their national counterparts. While the average of the correlations of MSA employment with national-level employment is 69 percent, they range from a low of −23 percent (New Orleans) to a high of 95 percent (Chicago). Not surprisingly, the lowest concordance with national-level employment occurs in cities heavily involved in energy production – e.g., Bakersfield (44 percent), Baton Rouge (44 percent), Oklahoma City (53 percent), and Tulsa (38 percent). With an average correlation of over 60 percent, MSA-level house prices exhibit substantial concordance with national house prices but, again, the individual MSA-national correlations vary substantially.

Much of the synchronicity between MSA and national house prices is driven by the 2004-2008 period; prior to that period, city-level house prices exhibited substantially less concordance, with several cities experiencing long declines in real house prices at the same

[4] The average correlation between U.S. house prices and MSA employment is higher than the average of those between MSA house prices and MSA employment. We considered whether changes in national house prices drive MSA employment. We found that national house prices have a countercyclical relationship with employment in more than a third of our MSAs and conclude that neither national nor MSA house prices are good leading indicators for MSA employment.

time. The correlation over the 1983-2003 period is 54 percent while the correlation over the 2004-2008 period is 77 percent. MSA-level permits exhibit similar concordance with their national counterpart: Over the full sample, the average correlation of city-level permits with national-level permits is 57 percent when permits are measured in units and 68 percent when permits are measured in values.

3 Is There a Linear Relationship Between Housing and Employment?

Correlations for the band-pass-filtered data are suggestive of the relationship between employment and the housing market. Because there may be causal relationships between both employment and the housing market, we use a VAR to look at the relationship between the variables, which allows us to identify the effect of housing market innovations on the business cycle. Our VAR approach both exploits the cross-sectional variation in the data and allows us to include national-level variables. We continue to use ideal band-passed data in our VAR analysis to ensure that we accurately extract the cyclical components of the data. By controlling for the national-level variables most likely to affect housing and employment, we are able to identify the effect of shocks to our housing variables on employment.

Because we have a limited number of observations (104) with which to estimate the VAR and many parameters to estimate for a VAR with four lags, we use principal components to reduce the dimensionality of the national data.[5] For each city i, we estimate the model

$$X_{t,i} = A_{0,i} + \sum_{k=1}^{p} A_{k,i} X_{t-k,i} + C\varepsilon_t, \tag{1}$$

where

$$X_{t,i} = \left[F_{t1}, F_{t2}, MSA'_{t,i}\right]',$$

[5] Bernanke, Boivin, and Eliasz (2005) used this data reduction technique to analyze the effects of monetary policy shocks without making restrictive assumptions about which data the Federal Reserve can and cannot observe. Boivin, Giannoni, and Mihov (2009) also used this technique to study the effect of price stickiness using industry-level data while controlling for a host of national factors.

C is lower triangular, $\Sigma = CC'$, F_{t1} and F_{t2} are two national-level financial factors, $MSA_{t,i}$ is a vector of MSA-level variables, and p is set to 4.[6] We control for the type of national-level financial variables that Smets (2007) suspects drive the leading indicator property of housing at the national level by constructing F_{t1} and F_{t2} from $Nat_t = [\pi_t, FFR_t, MR_t, Spread_t]'$: π_t is national-level core CPI inflation, FFR_t is the federal funds rate, MR_t is the conventional mortgage rate, and $Spread_t$ is the spread between the 3-month commercial paper rate and the 3-month T-bill yield.

3.1 Do MSA Housing Variables Affect MSA Employment?

Results from the preceding section suggested a relationship, albeit varying across cities, between MSA employment and a city's housing market. To assess this relationship formally, we estimate (1) with

$$MSA_{t,i} = [PermVal_{t,i}, Emp_{t,i}]' \tag{2}$$

to ascertain how much of the strength of the relationship is due to national financial factors affecting permits. Table 3 contains these results.[7] We define a city or the United States as reacting primarily positively if the cumulative impulse response function after 19 quarters is positive.

At the national level, employment rises immediately after the positive permit shock, but experiences a significant and sustained decline after four years. At the MSA level, however, including financial factors changes the predictive power of permits for many cities. Figure 4 depicts the diversity across select cities of the employment response to a permit shock. In

[6] Because the cities are small relative to the U.S. economy, we assume that the financial factors do not respond contemporaneously to MSA-level shocks. On the other hand, MSA-level variables respond to shocks to the financial factors. Since our interest is not in identifying shocks to our national-level financial factors, the ordering of these factors does not matter.

[7] We considered a specification in which we included MSA house prices in the VAR as a control variable; we ordered house prices ahead of permit values since house prices have been shown to respond slowly to changes in macroeconomic conditions (see, for example, Mankiw and Weil, 1989; and Poterba, 1991). The reactions to permit shocks when we included house prices in the VAR were qualitatively similar to that of our benchmark case; a few impulse responses became insignificant when we included house prices but the signs and magnitudes of the reactions were usually the same. We also explored using house prices as the measure of housing activity. We found that employment responded insignificantly to house price shocks both at the national level and in almost all of our MSAs.

our eight selected cities, Atlanta, Baltimore, Orlando, and Tampa have similar responses to a permit shock with the peak response in employment occurring about four quarters after the shock. Employment in Tulsa also responds significantly positively although the peak response occurs five quarters after the shock. Employment in Chicago and San Francisco, however, responds negatively to a permit shock. Altogether, employment responds positively to a permit shock in slightly less than half the MSAs in our sample (25). Similar to the reaction at the national level, many (17) of these cities subsequently experience a significant decline in employment starting several quarters after the shock. In 9 MSAs, employment falls immediately after a permit shock.[8]

We view this as weak evidence that permits lead the cycle. The lack of a significant response in some MSAs, the level at which we believe we can accurately identify a permit shock, suggests that part of the pattern Leamer finds at the national level is due to financial factors. Since national permits are simply an aggregation of individual cities' permits, national permits fall when there is a shock that affects a number of cities. For the permit shock to be pervasive (and perhaps simultaneous) enough to yield a response in multiple cities, the driver may, in fact, be a national – perhaps financial – shock. However, at the MSA level, permits themselves continue to have some explanatory power although the response is negative in many cases.[9]

3.2 Do National Housing Variables Affect MSA Employment?

The preceding discussion assumes that local housing conditions affect the local business cycle. However, national housing conditions may affect local economic conditions (perhaps with

[8] These results are, for the most part, sensitive to the exclusion of the national financial variables. For a few cities (Baton Rouge, Lexington, Milwaukee, Minneapolis, San Diego, and Virginia Beach), excluding the financial variables reverses the sign of the impulse response. For others (Cleveland, Columbus, Greensboro, and Stockton), it renders the impulse response insignificant.

[9] Our identification strategy at the city level assumes that national-level financial factors cannot simultaneously respond to permits or employment. Since this is a much less plausible assumption at the national level, we also consider an alternative ordering of the VAR in which the financial factors are ordered after economic activity indicators to test the robustness of the effect of permit shocks on employment at the national level. The results at the national level are very similar to that of our benchmark model.

varying timings and degrees) even though no national housing market exists. Suppose, for example, shocks to aggregate wealth through an increase in house prices raised the demand for autos, thereby affecting Detroit's local cycle.[10] We, therefore, explore the possibility that it is not a city's own housing market that influences its business cycle but rather the general tenor of the U.S. housing market as a whole. We estimate (1) with

$$MSA_{t,i} = \left[PermVal_{t,US}, Emp_{t,US}, Emp_{t,i}\right]. \tag{3}$$

Because house prices appear to have very little explanatory power for employment, we exclude them to reduce the dimensionality of the VAR. However, we include national-level employment as a control to capture other linkages between MSA and national business cycles not captured by our financial factors or permits.

Table 4 summarizes our results from estimating (3) and Figure 5 shows the impulse responses and their associated 80 percent coverage intervals for a representative group of cities.[11] Here, we see much more consistency in the response of employment to a national permit shock. Indeed, in about 1/4 of the cities, employment does not respond significantly positively to an increase in permits at the national level; most of these cities are dependent on energy production. Bakersfield, Tulsa, and Oklahoma City are all oil drilling centers while Baton Rouge is a major refining center. Pittsburgh is heavily dependent on steel production, which is a very energy-intensive industry. Although Austin's industry is not strongly linked to oil itself, its geographic proximity to Oklahoma City, Dallas, and Houston likely makes it much more dependent on oil than the other cities in our sample.

To understand why these cities do not respond positively to a national-level permit shock, Figure 6 plots the band-pass filtered real spot price of West Texas Intermediate crude and U.S. real permit values. The correlation between the two series is -48 percent, with the real price of oil often falling and permits rising simultaneously. Based on this, we test to

[10] We thank the referees for bringing these external effects to our attention.

[11] These results are robust to a reordering of the variables, $X_{t,i} = \left[PermVal_{t,US}, Emp_{t,US}, F_{t,1}, F_{t,2}, Emp_{t,i}\right]$.

see whether the relationship from Table 4 is spurious by reestimating (1) including Oil_t, the band-pass filtered real spot price of West Texas Intermediate.[12] While the strength of the relationship between national permits and MSA employment falls after we control for oil prices, there remains more consistency in the relationship between national permits and MSA employment than between MSA permits and MSA employment. Further, the relationship between national permits and national employment remains robust.[13] We conclude that employment at the MSA level is influenced, at least in part, by the national housing market.

4 A Preliminary Look at Recessions and Housing

Much research (e.g., Burns and Mitchell, 1946; Hamilton, 1989) suggests that the business cycle is an inherently nonlinear phenomenon with asymmetric dynamics. As periods of economic decline are rapid and sharp, with expansions being slow-paced and steadier, studying only the linear relationship among variables may fail to capture important business cycle relationships. Suppose we believe that national housing variables drive business cycle turning points. That is, declines in permits do not proportionally lower employment but instead raise the probability of a recession. We might, in turn, expect local housing markets to affect the probability that cities experience recessions. Recent studies – e.g., Owyang, Piger, Wall, and Wheeler (2008) – have highlighted the heterogeneity in business cycle turning points across regions. Using a similar framework, we can construct sets of city-level turning points with which we can compare MSA-level housing variables.[14]

Suppose employment growth in city i follows the process

$$\Delta Emp_{i,t} = \alpha_{s_{i,t}} + \varepsilon_{i,t}, \qquad \varepsilon_{i,t} \sim N\left(0, \sigma_i\right), \tag{4}$$

[12] We exclude national employment in this specification to ensure that we are not artificially biasing our results against permits as an explanatory variable. We order oil first in this specification as the price of oil is determined in world markets, making it exogenous.

[13] For brevity, we do not report these results here. The complete set of results is available upon request.

[14] Because Hurricane Katrina so heavily dominates the employment dynamics of New Orleans, we drop New Orleans in our analysis of MSA-level recessions.

where the average growth rate, $\alpha_{s_{i,t}}$, depends on the regime according to

$$\alpha_{s_{i,t}} = \alpha_{0,i} + \alpha_{1,i} s_{i,t}, \tag{5}$$

normalized such that $\alpha_{1,i} < 0$. The unobserved regime, $s_{i,t} \in \{0, 1\}$ follows a first-order Markov-switching process with transition probabilities:

$$\begin{aligned} \Pr(s_{i,t} = 0 | s_{i,t-1} = 0) &= p_{it} \\ \Pr(s_{i,t} = 1 | s_{i,t-1} = 1) &= q_{it}, \end{aligned} \tag{6}$$

where $s_{i,t} = 0$ indicates the city is in expansion and $s_{i,t} = 1$ indicates the city is in recession. We estimate the model using the Gibbs sampler with standard, fairly diffuse priors.

Table 5 summarizes the recession dates in each city and displays the vast heterogeneity in the timing of the turning points across cities.[15] While our national turning points are similar to those defined by the NBER, troughs tend to be dated later with employment. The cities, on the other hand, experience several recessions that do not correspond with national-level recessions. In particular, Akron, Ann Arbor, Austin, Baton Rouge, Minneapolis-St. Paul, Oklahoma City, Salt Lake City, Tulsa, Boulder, and Wichita experience a recession during the mid-1980s when the nation as a whole was in a prolonged expansion. While the majority of cities experience recessions around the time of the aggregate, the starting and ending dates typically differ. A handful of cities (e.g., Austin, Oklahoma City, Tulsa, and Boulder) never experience the national recession of the early 1990s. Several cities (e.g., Buffalo, Cincinnati, Columbus, Indianapolis, St. Louis, San Diego, and Virginia Beach) do not recover from the 2001 recession. Oklahoma City and Tulsa continued to expand through the last period in

[15] In Table 5, we identify a recession as any period in which the posterior recession probability exceeds 0.7 and the recession lasts at least three quarters. Varying the cutoff yields similar results as few probabilities stray from either 0 or 1. An exception is Durham where we do not see evidence of switching in employment and thus exclude it from our analysis. Because MSA employment is more volatile than national employment, we omit from Table 5 periods in which $s_{i,t} = 1$ for only one or two quarters. We also count it as a single recession if a city experiences one quarter of expansion sandwiched between two recessionary periods.

our sample, 2008Q4.

Figures 7 compares filtered national and MSA permits (in values) and the recession probabilities for a few representative cities: Atlanta, Austin, Baltimore, Chicago, Orlando, Tulsa, San Francisco, and Tampa. Consistent with Leamer's results at the national level, a fall in MSA permits does tend to precede a local recession, even for the 2001 recession, which was associated more with the internet bubble than housing. However, the timing of this relationship is not consistent across cities. The business cycle peak can be simultaneous with a decline in housing (Austin 2001; Baltimore 1990) or after a long decline (Chicago 2001; Atlanta 1991 and 2001). Even for the same city (say, Orlando), the business cycle peak can follow the downturn in housing by a few quarters (1990) or a few years (2000). In most cities, permits usually begin recovering during the recession, often not far from the start of the recession, which explains the negative correlation between leads of permits and employment in Table 2. To use Leamer's terminology, the most noticeable false positive and false negative occur for Tulsa. Tulsa experiences a trough in permits in 1985, but it does not enter recession until a full year later after permits have been trending upward.

Similar conclusions can be drawn from examining the relationship between national permits and MSA turning points in Figure 7. National permits increase steadily prior to and during recessions in both Austin and Tulsa during the mid-1980s. Austin does not experience even a slight increase in its recession probability despite the decline in national permits in the early 1990s. Similarly, Tulsa remains in robust expansion in 2008Q4 despite the very sharp decline in national permits that began more than a year earlier. These findings may reflect the fact that many energy-producing regions have business cycles which are disjoint from the national cycle (see Hamilton and Owyang, 2009).

Figure 8 shows similar plots for MSA and national house prices and local recessions. We see little consistency in the relationship between MSA house prices and recessions. House prices are close to flat in Orlando and Tampa around both the 1990 and 2001 recessions. Large declines in house prices in Austin and Tulsa that begin around 1993, declines that are much larger than the fall in house prices either city experiences in 2007 or 2008, do not cause a recession in either city. In fact, both Austin and Tulsa experience peaks in house prices in the midst of recessions in the mid-1980s and early 2000s. The relationship between national house prices and MSA recessions is not much stronger, but it is worth noting that the rise and fall of house prices in the 2004-2008 period heavily dominate the house price series shown in Figure 8.

5 Does Housing Drive Recessions?

In this section, we formalize the local housing market's influence on the probability that the local economy is in a recession phase. By looking at MSA-level data rather than focusing on national data alone, we have substantially more recessions with various turning points. To address the issue, we modify the Markov-switching model from the previous section to include TVTPs along the lines suggested by Goldfeld and Quandt (1973). This relaxes the assumption that the probabilities governing the transition between expansion and recession are constant. Instead, we will assume that these transition probabilities are functions of national- or MSA-level housing variables.

Suppose that a city's employment growth continues to be characterized by (4) and (5); however, the unobserved regime, $s_{i,t} = s_{i,t}(z_{i,t}) \in \{0,1\}$, is a homogeneous Markov process with transition probabilities:

$$\begin{aligned} \Pr(s_{i,t} = 0 | s_{i,t-1} = 0, z_{i,t}) &= p_i(z_{i,t}) \\ \Pr(s_{i,t} = 1 | s_{i,t-1} = 1, z_{i,t}) &= q_i(z_{i,t}), \end{aligned} \tag{7}$$

which are now determined by a vector of variables, $z_{i,t}$. We allow the lagged six-quarter-moving-average of our first financial factor, house price growth, and permit growth to influence the transition probabilities. Using the six-quarter-moving-average has the advantage of both moderating the high-frequency variation in permits and allowing more distant lags to affect the probability of being in a recession in a parsimonious specification.

The transition probabilities can be defined in terms of a set of latent variables

$$s^*_{i,t} = \beta_{0,i} + \beta_{z,i} z_{i,t} + \beta_{s,i} s_{i,t-1} + u_{i,t}, \qquad u_{i,t} \sim N(0,1), \tag{8}$$

defined such that

$$\Pr(s_{i,t} = 1) = \Pr\left(s^*_{i,t} > 0\right).$$

Equation (8) produces the desired interpretation. The transition probability depends jointly on the past regime and the vector of $z_{i,t}$. Letting $\phi(\bullet)$ denote the standard normal probability density function,

$$\frac{\partial \Pr(s_{i,t} = 0 | s_{i,t-1} = 0)}{\partial z_{i,t}} = -\beta_{z,i} \phi\left(-\beta_{0\,i} - \beta_{z,i} z_{i,t}\right)$$

and

$$\frac{\partial \Pr(s_{i,t} = 1 | s_{i,t-1} = 1)}{\partial z_{i,t}} = \beta_{z,i} \phi\left(-\beta_{0\,i} - \beta_{z,i} z_{i,t} - \beta_{s,i}\right),$$

such that a negative value for $\beta_{z,i}$ indicates that an increase in $z_{i,t}$ raises the probability of staying in an expansion and lowers the probability of staying in a recession. The model can be estimated via a straightforward application of the Gibbs sampler detailed in the appendix (see also Filardo and Gordon, 1998).

Table 6 shows the 68 percent coverage intervals for each of our candidate leading indicator variables on the transition probabilities in our TVTP Markov-switching models. The first two columns show the significance of the influence of the moving average of permit values on $p(\mathbf{z}_{i,t})$ and $q(\mathbf{z}_{i,t})$ when $\mathbf{z}_{i,t} = [1, \Delta PermMA_{i,t-1}]$. The next two columns show the effect of lagged house price growth on employment by setting $\mathbf{z}_{i,t} = [1, \Delta \Pr iceMA_{i,t-1}]$. The final columns show the 68 percent coverage intervals when $\mathbf{z}_{i,t} = [1, Fin1MA_{i,t-1}]$, where $Fin1$ is the first principal component constructed from the financial variables described in the previous section. Neither of our housing variables significantly affect the regime transition probabilities in any of our cities or at the national level. The parameter estimates for the regime means and the influence of the current regime on transition probabilities are highly significant and have the expected signs, but are not shown in the interest of space. However, there is no evidence that financial factors influence the transition probabilities either.

6 Conclusions

We assessed the robustness of the relationship between housing and employment over the business cycle in a set of 51 U.S. cities. In most of the cities we considered, we found permits to be a good leading indicator of employment. Perhaps because of the disparate business cycle experiences of the cities, however, this relationship failed to hold across the board. We also find that national permits are a better leading indicator for MSA employment than a city's own permits. Housing prices, on the other hand, are not a good leading indicator for employment at either the national or the MSA level.

These results stand in some contrast to the previous literature linking housing to business cycles. At the national level, housing appears to be an important driver of cyclical fluctuations. If this causality truly held, we would expect the link to be preserved at the MSA level, as housing spillovers into business cycle indicators (e.g., employment) should remain localized. This may be especially important to verify at the local level in light of findings by Del Negro and Otrok (2007) and others that housing shocks are primarily local phenomena. Our results, however, appear at odds with the finding that there exists a direct channel (whether it be wealth effects or housing capital effects) from local housing markets into local employment.

Our results should not, however, be seen as evidence that there is no relationship between housing prices and employment over the business cycle. Clearly, theoretical models such as Davis and Heathcote (2005) illustrate that there is a relationship. However the lack of consistency in the relationship at the city-level suggests that the relationship may be more complicated than simple causal stories wherein a rise in house prices raises wealth, leads households to consume more, and then leads to an economic expansion. Furthermore, the extent to which we find frequency-dependence in the relationships suggests that increases in house prices may have a different impact in the short term than in the long term. Finally, our results do not address the question of whether or not housing should play a role in the formulation of monetary policy: We examine only empirical linkages between housing markets and recessions and leave the question of modeling optimal monetary policy in the presence of housing for future research.

While housing played a pivotal role in the recession beginning in December 2007, our findings suggest caution in assuming that future recessions will fit the current pattern. The diversity we find across cities in the relationship between housing and employment – in many cities, a rise in permits and especially house prices appears to be associated with a subsequent *decline* in employment – is further evidence that, as Temin (1998) finds, business cycles are tremendously diverse.

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Technical Appendix

6.1 Estimating the VAR

We estimate the model using the Gibbs sampler, a Bayesian technique in which the full joint posterior is obtained from iterative draws from each parameter's conditional distribution (see Gelfand and Smith, 1990; Casella and George, 1992; Carter and Kohn, 1994). We assume that the set of parameters consisting of the VAR coefficients and the variance-covariance matrix are characterized by normal-inverse Wishart priors – that is,

$$A, \Sigma \sim NW(\alpha, S, \nu, \delta),$$

where $\{\alpha, S, \nu, \delta\}$ are hyperparameters that govern the shape of the prior. We set $\alpha = \mathbf{0}_N$, $S = \mathbf{I}_N$, $\nu = 0$, and $\delta = \mathbf{I}_n$, where n is the number of variables in the VAR and N is the total number of VAR coefficients (in this case, $N = pn^2 + n$). Sampling from the posterior is a straightforward implementation of Chib and Greenberg's (1996) SUR sampler where the posteriors are formed from $5,000$ draws after discarding the first $5,000$ draws.[16]

6.2 Estimating the TVTP Markov-Switching Model

The sampler is broken into four blocks to estimate the joint posterior for the entire parameter vector

$$\Theta_i = \left[\sigma_i, \alpha_{0,i}, \alpha_{1,i}, \{s_{i,t}\}_{t=1}^{T}, \beta_{0,i}, \beta_{z,i}, \beta_{s,i}\right]'.$$

[16] For each draw of the model parameters, we can compute the impulse responses to a shock to any system variable. To determine whether the impulse responses are significantly different from zero, we examine the 10th and 90th percentiles of their posterior distributions.

A single iteration of the Gibbs sampler begins with a draw from the posterior distribution of σ_i conditional on values for $\Theta_{-\sigma_i}^{(j-1)}$. The prior for σ_i^2 is inverse-gamma, i.e.,

$$\sigma_i^{-2} \sim \mathbf{\Gamma}\left(1,1\right).$$

The posterior is then

$$\sigma_i^{-2} | \left\{s_{i,t}\right\}_{t=1}^{T}, \alpha_{0,i}, \alpha_{1,i} \sim \mathbf{\Gamma}\left(\frac{1+T}{2}, \frac{2T}{\left(1+T\right)\left(1+\tilde{y}_i'\tilde{y}_i\right)}\right),$$

where $\tilde{y}_i$ is a $T \times 1$ vector with row t, $\tilde{y}_{i,t} = y_{i,t} - \alpha_{0,i}^{(j-1)} - \alpha_{1,i}^{(j-1)} s_{i,t}^{(j-1)}$, and $y_{i,t} = \Delta Emp_{i,t}$.

We then draw $\alpha_{0,i}^{(j)}$, $\alpha_{1,i}^{(j)}$ conditional on $\sigma_i^{(j)}$ and $\Theta_{-\sigma_i}^{(j-1)}$. If the prior for $\boldsymbol{\alpha}_i = \left[\alpha_{0,i}, \alpha_{1,i}\right]'$ is

$$\boldsymbol{\alpha}_i \sim \mathbf{1}_{\alpha_{1i}<0} N\left(\hat{\boldsymbol{\alpha}}_i, \sigma_i I_2\right),$$

where $\mathbf{1}_{\alpha_{1i}<0}$ is an indicator function that identifies $s_{i,t} = 1$ as the recession state, $\hat{\boldsymbol{\alpha}}_i = \left[1.5 * \bar{y}_i, -2 * \bar{y}_i\right]$, and $\bar{y}_i = \frac{1}{T}\sum_{t=1}^{T} y_{i,t}$, then the posterior from which we draw $\boldsymbol{\alpha}_i^{(j)} = \left[\alpha_{0,i}^{(j)}, \alpha_{1,i}^{(j)}\right]'$ is

$$\boldsymbol{\alpha}_i | \left\{s_{i,t}\right\}_{t=1}^{T}, \sigma_i \sim \mathbf{1}_{\alpha_{1i}<0} N\left(\bar{\boldsymbol{\alpha}}_i, \bar{A}_i \sigma_i\right),$$

where $\bar{A}_i = \left(\sigma_i I_2 + \ddot{Y}_i'\ddot{Y}_i\right)^{-1}$, $\bar{\boldsymbol{\alpha}}_i = \bar{A}_i\left(\frac{1}{\sigma_i} I_2 \hat{\boldsymbol{\alpha}}_i + \ddot{Y}_i' y_i\right)$, $\ddot{Y}_i$ is a $T \times 2$ matrix with row t $\left[1, s_{i,t}\right]$, and $\ddot{y}_i$ is a $T \times 1$ vector with row t as $y_{t,i}$.

We next draw $\left\{s_{i,t}^{(j)}\right\}_{t=2}^{T}$ recursively from a Bernoulli distribution using

$$P\left(s_{i,t}|\ddot{y}_i\right) \propto P\left(s_{i,t}|s_{i,t-1}, z_{i,t}\right) P\left(s_{i,t+1}|s_{i,t}, z_{i,t+1}\right) f\left(y_{i,t}|s_{i,t}\right).$$

Given $\pi^{(j-1)}$ and $\boldsymbol{\beta}_i = \left[\beta_{0,i}, \beta_{z,i}, \beta_{s,i}\right]$, we apply Bayes' rule to obtain $\pi^{(j)}$, the conditional posterior probability that $s_{i,1} = 0$, as

$$\Pr\left(\pi^{(j)}|s_{i,2} = 0\right) = \frac{p\left(z_{i,2}\right)\pi^{(j-1)}}{p\left(z_{i,2}\right)\pi^{(j-1)} + \left(1 - q\left(z_{i,2}\right)\right)\left(1 - \pi^{(j-1)}\right)}$$

and

$$\Pr\left(\pi^{(j)}|s_{i,2}=1\right)=\frac{\left(1-p\left(z_{i,2}\right)\right)\pi^{(j-1)}}{\left(1-p\left(z_{i,2}\right)\right)\pi^{(j-1)}+q\left(z_{i,2}\right)\left(1-\pi^{(j-1)}\right)}.$$

We then use a data augmentation technique to draw a set of latent variables $\left\{s_{i,t}^{*(j)}\right\}_{t=0}^{T}$ from truncated normal distributions (see Tanner and Wong, 1987; and Albert and Chib, 1993). If $s_{i,t}^{(j)}=0$, $s_{i,t}^{*(j)}$ is drawn from the truncated normal distribution $N\left(-\beta_{0,i}-\beta_{z,i}z_{i,t},1\right)$, where the truncation is at 0 from the right since $s_{i,t}^{*(j)}$ is negative by definition if $s_{i,t}^{(j)}=0$. If $s_{i,t}^{(j)}=1$, $s_{i,t}^{*(j)}$ is drawn from the truncated normal distribution $N\left(-\beta_{0,i}-\beta_{z,i}z_{i,t}-\beta_{s,i},1\right)$, where the truncation is at 0 from the left.

We use $\left\{s_{i,t}^{*(j)}\right\}_{t=0}^{T}$ to generate draws for $\boldsymbol{\beta}_i$. The prior for $\boldsymbol{\beta}_i$ is

$$\boldsymbol{\beta}_i\sim N\left(0,\begin{bmatrix}1&0&0\\0&1000&0\\0&0&1\end{bmatrix}\right).$$

The prior over $\beta_{z,i}$ is relatively diffuse as we have little information regarding the likely scale of the coefficient; however, the results were qualitatively very similar for substantially more diffuse and less diffuse priors over $\beta_{z,i}$. It can be shown (see, for example, Filardo and Gordon, 1998) that the posterior distribution of $\boldsymbol{\beta}_i$ has the form

$$\boldsymbol{\beta}_i\sim N\left(\bar{\boldsymbol{\beta}}_i,\bar{\mathbf{A}}_{\beta,i}\right),$$

where $\bar{\mathbf{A}}_{\beta,i}=\left(\mathbf{I}_3+\mathbf{W}_i'\mathbf{W}_i\right)^{-1}$, $\bar{\boldsymbol{\beta}}_i=\bar{\mathbf{A}}_{\beta,i}\mathbf{W}_i'\mathbf{s}_i^*$, $\mathbf{s}_i^*=[s_{i,0},s_{i,1},...,s_{i,T}]'$, and $\mathbf{W}_i$ is a $T\times 3$ matrix with rows $[1,\ z_{i,t},\ s_{i,t}]$. We repeat this process $J=10,000$ times for each city and discard the first $5,000$ draws as burn-in iterations. See Filardo and Gordon (1998) for additional details on the Gibbs sampling approach to estimating TVTP Markov-switching models.

Table 1: Cities

MSA Name	Deflator Used	Sample Start Date
USA	**USA**	**1983Q1**
Akron OH	Cleveland-Akron, OH	1983Q1
Albuquerque NM	West Urban	1983Q1
Ann Arbor MI	Detroit-Ann Arbor-Flint, MI	1983Q1
Atlanta-Sdy Sprgs-Marietta GA	Atlanta, GA	1983Q1
Augusta-Richmond County GA-SC	South Urban	1983Q1
Austin-Round Rock TX	South Urban	1983Q1
Bakersfield CA	West Urban	1983Q1
Baltimore-Towson MD	South Urban	1983Q1
Baton Rouge LA	South Urban	1983Q1
Buffalo-Niagra Falls NY	Northeast Urban	1983Q1
Chicgo-Naprvlle-Joliet IL-IN-WI	Chicago-Gary-Kenosha, IL-IN-WI	1983Q1
Cincinnati-Middletown OH-KY-IN	Cincinnati-Hamilton, OH-KY-IN	1983Q1
Cleveland-Elyria-Mentor OH	Cleveland-Akron, OH	1983Q1
Columbia SC	South Urban	1983Q1
Columbus OH	Mid-West Urban	1983Q1
Durham NC	South Urban	1983Q1
Greensboro-High Point NC	South Urban	1983Q1
Greenville SC	South Urban	1983Q1
Indianapolis IN	Mid-West Urban	1983Q1
Jacksonville FL	South Urban	1983Q1
Las Vegas-Paradise NV	West Urban	1983Q1
Lexington-Fayette KY	South Urban	1983Q1
Louisville KY-IN	South Urban	1983Q1
Memphis TN-MS-AR	South Urban	1983Q1
Milwaukee-Waukesha-West Allis WI	Milwaukee-Racine, WI	1983Q1
Minnplis-St. Paul-Bloomngtn MN-WI	Minneapolis-St. Paul, MN-WI	1983Q1
New Orleans-Metairie-Kenner LA	South Urban	1983Q1
Oklahoma City OK	South Urban	1983Q1
Orlando-Kissimmee FL	South Urban	1983Q1
Pensacola-Ferry Pass-Brent FL	South Urban	1983Q1
Phoenix-Mesa-Scottsdale AZ	West Size Class A Urban	1983Q1
Pittsburgh PA	Pittsburgh PA	1983Q1
St. Louis MO-IL	St. Louis MO-IL	1983Q1
Salt Lake City UT	West Urban	1983Q1
San Diego-Carlsbad-San Marcos CA	San Diego, CA	1983Q1
Sarasota-Bradenton-Venice FL	South Urban	1983Q1
Stockton CA	West Urban	1983Q1
Tucson AZ	West Urban	1983Q1
Tulsa OK	South Urban	1983Q1
Boulder CO	Denver-Boulder-Greeley, CO	1984Q1
Charleston-North Charleston SC	South Urban	1984Q1
Charlotte-Gastonia-Concrd NC-SC	South Urban	1984Q1
Nashvlle-Davdsn-Murfreesboro TN	South Urban	1984Q1
Palm Bay-Melbourne-Titusvll FL	South Urban	1984Q1
Portlnd-Vancouver-Beavrtn OR-WA	Portland-Salem, OR-WA	1984Q1
Richmond VA	South Urban	1984Q1
San Francisco-Oakland-Fremont CA	San Francisco-Oakland-San Jose, CA	1984Q1
Santa Rosa-Petaluma CA	West Urban	1984Q1
Tampa-St. Petrsbrg-Clearwater FL	South Urban	1984Q1
VA Bch-Norfolk-Nwprt News VA-NC	South Urban	1984Q1
Wichita KS	Mid-West Urban	1984Q1

Table 2: Average correlation between series x at quarter t and series y at t+j

							j							
x	y	-6	-5	-4	-3	-2	-1	0	1	2	3	4	5	6
Bandpass Filtered Series:														
US Res. Inv.	US Permits (V)	30%	43%	56%	70%	83%	94%	98%	96%	86%	72%	57%	43%	29%
US Emp	MSA Emp							69%						
US House Prices	MSA House Prices							62%						
US Permits (U)	MSA Permits (U)							57%						
US Permits (V)	MSA Permits (V)							68%						
US House Prices	US Emp	-33%	-26%	-16%	-5%	8%	20%	31%	41%	47%	50%	50%	47%	42%
MSA House Prices	MSA Emp	5%	8%	12%	15%	18%	22%	24%	26%	26%	25%	22%	19%	15%
US House Prices	MSA Emp	-23%	-18%	-12%	-5%	3%	12%	20%	27%	33%	36%	38%	37%	34%
US Permits (U)	US Emp	-57%	-53%	-46%	-37%	-26%	-14%	-1%	11%	23%	35%	44%	49%	48%
MSA Permits (U)	MSA Emp	-31%	-27%	-22%	-15%	-8%	0%	8%	17%	25%	32%	36%	37%	36%
US Permits (U)	MSA Emp	-38%	-36%	-31%	-25%	-17%	-8%	1%	10%	20%	29%	35%	39%	39%
US Permits (V)	US Emp	-60%	-56%	-49%	-40%	-29%	-15%	-1%	14%	28%	41%	51%	55%	55%
MSA Permits (V)	MSA Emp	-36%	-33%	-28%	-21%	-13%	-3%	7%	18%	27%	35%	41%	43%	42%
US Permits (V)	MSA Emp	-39%	-36%	-32%	-26%	-18%	-8%	2%	12%	22%	31%	38%	41%	42%

Note: For MSA level correlations, the table reports the average correlation across MSAs. The sample is 1983Q1 - 2008Q4 for most cities; see Table 1 for the set of cities for which the sample is 1984Q1 - 2008Q4. U denotes permits in units; V denotes real permit values. House prices are real house prices.

Table 3: Responses to Orthogonalized MSA Permit Shock

	Responds Primarily Positively	10% and 90% of Posterior Both Positive	10% and 90% of Posterior Both Negative		Responds Primarily Positively	10% and 90% of Posterior Both Positive	10% and 90% of Posterior Both Negative
US	**No**	**0-8**	**11-19**	**US Alt Ordering**	**No**	**0-8**	**11-19**
Akron	No	Never	Never	Boulder	Yes	0-5	Never
Albuquerque	Yes	3-8	Never	Charleston	Yes	0-6	13-16
Ann Arbor	No	Never	5-7	Charlotte	No	0-5	13-16
Atlanta	No	0-8	13-19	Nashville	Yes	0-12	16-19
Augusta	No	Never	Never	Palm Bay	No	0-7	11-15
Austin	No	0-1	Never	Portland OR	No	0-6	11-17
Bakersfield	No	Never	0-2	Richmond	No	0-6	13-17
Baltimore	No	0-5	Never	San Francisco	Yes	15-17	1-6
Baton Rouge	No	Never	0-6	Santa Rosa	No	Never	Never
Buffalo	No	18-19	0-2	Tampa	No	0-8	13-18
Chicago	No	Never	2-6,19	Virginia Beach	No	Never	0-6, 10-13
Cincinnati	Yes	0-4	Never	Wichita	No	0-4	7-12
Cleveland	No	Never	7-16				
Columbia	No	0-8	Never				
Columbus	No	Never	Never				
Durham	Yes	14-15	0-6				
Greensboro	No	Never	11				
Greenville SC	No	Never	Never				
Indianapolis	No	0-4	Never				
Jacksonville	No	0-7	15-19				
Las Vegas	No	0-4	11-16				
Lexington	No	6-7	0-3,11,16-18				
Louisville	No	Never	Never				
Memphis	No	0-6	14-18				
Milwaukee	Yes	Never	0-11				
Minneapolis	Yes	8-14	0-6,19				
New Orleans	Yes	0-3,10-14	5-8,18-19				
Oklahoma City	No	Never	Never				
Orlando	No	0-4	10-18				
Pensacola	Yes	0-5	Never				
Phoenix	No	0-7	10-19				
Pittsburgh	No	16-19	2-10				
Salt Lake City	No	Never	4-8				
San Diego	No	19	0-2,7-16				
Sarasota	No	0-4	Never				
St. Louis	No	0-5	14-17				
Stockton	No	Never	Never				
Tucson	Yes	0-9	18-19				
Tulsa	No	2-7	10-14				

Notes: 1) Variables in VAR: 4 lags of national financial factors, permit values of MSA i, and employment of MSA i with that ordering for the Choleski Decomposition except for US Alt Ordering where the financial factors are ordered last. 2) "Primarily Positively" is defined as Yes if the cumulative median IRF is positive after 19 quarters. 3) Series are filtered using Corbae, Ouliaris, and Phillips (2002) ideal bandpass filter. Filter extracts components associated with cycles of 6-32 quarters.

Table 4: Responses to Orthogonalized US Permit Shock

	Responds Primarily Positively	10% and 90% of Posterior Both Positive	10% and 90% of Posterior Both Negative
US	**No**	**0-8**	**12-18**
Akron	No	Never	14-19
Albuquerque	No	0-4	Never
Ann Arbor	No	Never	Never
Atlanta	No	0-7	12-16
Augusta	No	Never	Never
Austin	No	Never	3-14
Bakersfield	No	Never	Never
Baltimore	No	0-5	13-16
Baton Rouge	No	Never	0-1
Buffalo	No	0-2	Never
Chicago	Yes	0-8	12-15
Cincinnati	No	0-8	13-19
Cleveland	No	0-7	12-18
Columbia	Yes	0-9	Never
Columbus	No	0-5	13-15
Durham	No	Never	13-15
Greensboro	No	0-4	13-19
Greenville SC	No	0-3	5-8
Indianapolis	No	0-7	13-16
Jacksonville	No	0-8	14-18
Las Vegas	No	0-9	13-16
Lexington	Yes	0-5	Never
Louisville	No	0-5	13-17
Memphis	Yes	0-7	13-17
Milwaukee	No	0-4	12-15
Minneapolis	No	0-10	14-18
New Orleans	No	12-15	0-6
Oklahoma City	No	Never	5-15
Orlando	No	0-8	13-17,19
Pensacola	Yes	0-7	16
Phoenix	Yes	0-10	12-19
Pittsburgh	No	Never	0-14
Salt Lake City	No	0-5	12-14
San Diego	Yes	0-9	13-16
Sarasota	Yes	0-7	Never
St. Louis	No	0-6	12-18
Stockton	No	8	13-16
Tucson	Yes	0-9	Never
Tulsa	No	19	9-12

	Responds Primarily Positively	10% and 90% of Posterior Both Positive	10% and 90% of Posterior Both Negative
US Alt Ordering	**No**	**0-8**	**14-19**
Boulder	Yes	0-4	Never
Charleston	Yes	0-5	Never
Charlotte	Yes	0-3	Never
Nashville	Yes	0-10	13-19
Palm Bay	Yes	0-10	Never
Portland OR	No	0-10	15-16
Richmond	No	0-9	13-18
San Francisco	Yes	0-4	Never
Santa Rosa	Yes	0-3	Never
Tampa	No	0-10	12-19
Virginia Beach	No	1-7	13
Wichita	Yes	0-5	Never

Notes: 1) Variables in VAR: 4 lags of national financial factors, national permit values, national employment, and employment of MSA i with that ordering for the Choleski Decomposition. 2) "Primarily Positively" is defined as Yes if the cumulative median IRF is positive after 19 quarters. 3) Series are filtered using Corbae, Ouliaris, and Phillips (2002) ideal bandpass filter. Filter extracts components associated with cycles of 6-32 quarters.

Table 5: MSA Recessions

				Recessions			
City	R0	R1	R2	R3	R4	R5	R6
USA		90Q3-92Q2	00Q4-04Q1	07Q3-			
Akron	-83Q3	85Q2-86Q3	90Q3-93Q3	95Q4-98Q4	99Q4-03Q1	03Q4-04Q3	05Q3-
Albuquerque		87Q1-87Q3	90Q4-92Q1	95Q2-96Q4	97Q4-99Q3	01Q1-05Q2	06Q3-
Ann Arbor		86Q2-89Q4	90Q3-96Q1	96Q4-00Q4	01Q3-		
Atlanta		90Q4-92Q1	00Q4-04Q2	07Q2-			
Augusta	-83Q3	91Q1-98Q1	99Q4-				
Austin		86Q1-89Q3	01Q1-04Q1	08Q2-			
Bakersfield	-83Q2	86Q2-87Q1	88Q4-89Q3	91Q3-96Q3	02Q1-04Q1	07Q2-	
Baltimore	-83Q2	90Q3-93Q2	95Q3-96Q1	01Q2-03Q3	07Q4-		
Baton Rouge	-83Q3	85Q2-87Q3	00Q3-05Q1	08Q1-			
Buffalo	-83Q3	90Q3-93Q4	95Q1-98Q3	99Q4-			
Chicago		90Q4-92Q2	01Q2-04Q1	07Q3-			
Cincinnati	-83Q2	91Q1-93Q2	00Q1-				
Cleveland		90Q4-92Q2	00Q2-05Q3	06Q2-			
Columbia SC		91Q1-93Q1	97Q4-98Q3	99Q3-03Q3	07Q4-		
Columbus	-83Q2	90Q4-92Q1	92Q4-93Q2	95Q3-97Q1	00Q2-		
Greensboro		01Q2-02Q1	03Q2-03Q4	08Q2-			
Greenville SC		01Q3-02Q2	03Q1-03Q3	08Q4-			
Indianapolis		89Q4-91Q4	96Q1-97Q3	00Q3-			
Jacksonville		90Q4-92Q2	00Q4-03Q3	07Q1-			
Las Vegas		91Q1-92Q1	01Q3-02Q4	06Q4-			
Lexington		90Q1-91Q4	00Q2-04Q3	07Q2-			
Louisville		91Q1-91Q3	00Q3-03Q4	07Q4-			
Memphis	-83Q2	90Q1-92Q3	98Q4-05Q1	06Q1-			
Milwaukee	-83Q2	90Q4-91Q3	00Q1-04Q2	07Q2-			
Minneapolis-St. Paul	-83Q2	85Q2-86Q2	89Q4-92Q1	00Q4-05Q1	06Q1-		
Oklahoma City	-83Q2	85Q1-87Q2	02Q4-03Q3				
Orlando		90Q4-91Q4	00Q4-03Q2	07Q2-			
Pensacola		90Q3-91Q2	98Q3-00Q4	01Q4-02Q3	07Q2-		
Phoenix		86Q4-87Q2	88Q2-93Q1	99Q4-04Q1	06Q3-		
Pittsburgh	-85Q2	91Q1-92Q2	93Q1-93Q3	95Q3-96Q1	01Q3-06Q3	07Q4-	
St. Louis		89Q2-92Q4	98Q2-99Q1	00Q1-			
Salt Lake City	-83Q2	86Q3-87Q4	01Q2-04Q2	07Q4-			
San Diego	-83Q2	90Q3-95Q1	01Q2-				
Sarasota		91Q1-91Q4	97Q2-98Q3	00Q4-01Q3	02Q4-03Q2	06Q4-	
Stockton	-83Q3	89Q1-89Q3	90Q4-96Q2	01Q3-02Q2	03Q1-03Q3	05Q3-	
Tucson	-83Q2	85Q4-93Q3	94Q3-97Q2	98Q1-99Q1	00Q2-		
Tulsa	-83Q3	86Q1-87Q2	01Q4-03Q3				
Boulder		84Q4-86Q4	01Q2-04Q1	08Q3-			
Charleston		90Q4-92Q2	93Q1-96Q2	00Q3-02Q1	03Q1-03Q4	06Q1-06Q3	07Q4-
Charlotte		90Q3-92Q1	00Q4-04Q1	08Q1-			
Nashville		89Q1-92Q1	00Q3-03Q4	06Q3-07Q2	08Q1-		
Palm Bay		86Q2-87Q1	90Q3-92Q2	93Q2-96Q2	97Q4-99Q4	01Q1-03Q2	06Q1-
Portland OR		98Q3-99Q1	01Q2-04Q1	08Q2-			
Richmond		89Q4-92Q2	01Q1-03Q2	07Q2-			
San Francisco		91Q2-92Q4	01Q2-04Q1	08Q2-			
Santa Rosa		91Q2-92Q4	93Q3-94Q1	01Q3-05Q2	06Q4-		
Tampa		90Q4-92Q2	01Q1-03Q2	06Q4-			
Virginia Beach		89Q1-93Q4	94Q4-96Q2	98Q4-99Q3	00Q2-		
Wichita		85Q1-85Q3	90Q2-94Q1	94Q4-95Q3	98Q4-05Q4	07Q4-	

Table 6: 68% Coverage Interals for Coefficients from TVTP Markov-Switching Models

	Permit Values		House Prices		Fin Factor 1	
USA	-0.042	0.035	-0.141	0.129	-0.036	0.037
Akron	-0.026	0.022	-0.179	0.152	-0.038	0.038
Albuquerque	-0.021	0.019	-0.113	0.112	-0.037	0.041
Ann Arbor	-0.014	0.011	-0.116	0.094	-0.037	0.041
Atlanta	-0.032	0.023	-0.174	0.153	-0.037	0.039
Augusta	-0.026	0.023	-0.183	0.173	-0.038	0.041
Austin	-0.022	0.014	-0.076	0.068	-0.039	0.036
Bakersfield	-0.023	0.021	-0.055	0.049	-0.038	0.037
Baltimore	-0.040	0.034	-0.091	0.075	-0.038	0.038
Baton Rouge	-0.017	0.014	-0.108	0.109	-0.037	0.037
Buffalo	-0.023	0.019	-0.159	0.137	-0.037	0.040
Chicago	-0.022	0.018	-0.132	0.114	-0.038	0.037
Cincinnati	-0.029	0.025	-0.242	0.199	-0.037	0.042
Cleveland	-0.033	0.025	-0.198	0.169	-0.037	0.040
Columbia SC	-0.028	0.026	-0.254	0.246	-0.039	0.038
Columbus	-0.032	0.025	-0.246	0.186	-0.036	0.044
Durham	-0.021	0.021	-0.178	0.153	-0.038	0.039
Greensboro	-0.027	0.024	-0.218	0.185	-0.038	0.038
Greenville SC	-0.034	0.032	-0.307	0.274	-0.038	0.037
Indianapolis	-0.034	0.025	-0.285	0.216	-0.034	0.041
Jacksonville	-0.025	0.021	-0.096	0.085	-0.039	0.037
Las Vegas	-0.022	0.016	-0.061	0.053	-0.039	0.037
Lexington	-0.029	0.023	-0.216	0.209	-0.036	0.039
Louisville	-0.033	0.026	-0.254	0.207	-0.037	0.042
Memphis	-0.029	0.021	-0.168	0.152	-0.036	0.041
Milwaukee	-0.029	0.023	-0.178	0.157	-0.038	0.037
Minneapolis-St. Paul	-0.029	0.025	-0.116	0.118	-0.038	0.038
New Orleans	-0.018	0.017	-0.104	0.110	-0.038	0.037
Oklahoma City	-0.021	0.017	-0.102	0.099	-0.044	0.036
Orlando	-0.029	0.021	-0.070	0.065	-0.037	0.038
Pensacola	-0.023	0.022	-0.087	0.075	-0.038	0.039
Phoenix	-0.028	0.023	-0.067	0.057	-0.039	0.037
Pittsburgh	-0.039	0.035	-0.242	0.234	-0.038	0.037
St. Louis	-0.032	0.027	-0.178	0.157	-0.039	0.039
Salt Lake City	-0.025	0.022	-0.085	0.067	-0.039	0.037
San Diego	-0.022	0.017	-0.056	0.053	-0.036	0.045
Sarasota	-0.018	0.014	-0.062	0.054	-0.038	0.037
Stockton	-0.015	0.014	-0.047	0.041	-0.036	0.038
Tucson	-0.022	0.019	-0.086	0.077	-0.039	0.037
Tulsa	-0.021	0.020	-0.129	0.132	-0.041	0.037
Boulder	-0.023	0.024	-0.117	0.116	-0.043	0.042
Charleston	-0.027	0.028	-0.118	0.129	-0.044	0.041
Charlotte	-0.035	0.036	-0.234	0.244	-0.044	0.042
Nashville	-0.027	0.029	-0.148	0.159	-0.043	0.042
Palm Bay	-0.023	0.022	-0.058	0.059	-0.045	0.043
Portland OR	-0.025	0.025	-0.109	0.123	-0.043	0.041
Richmond	-0.033	0.033	-0.139	0.146	-0.044	0.041
San Francisco	-0.024	0.024	-0.064	0.063	-0.043	0.042
Santa Rosa	-0.022	0.022	-0.057	0.059	-0.044	0.042
Tampa	-0.022	0.022	-0.075	0.077	-0.043	0.040
Virginia Beach	-0.030	0.031	-0.089	0.092	-0.043	0.041
Wichita	-0.023	0.023	-0.200	0.201	-0.043	0.042

Notes: 1) Columns show the 16th and 84th percentile of the posterior distribution of the coefficient on the variable indicated from the bivariate TVTP Markov-Switching Model. 2) Independent variable is 6-quarter moving average.

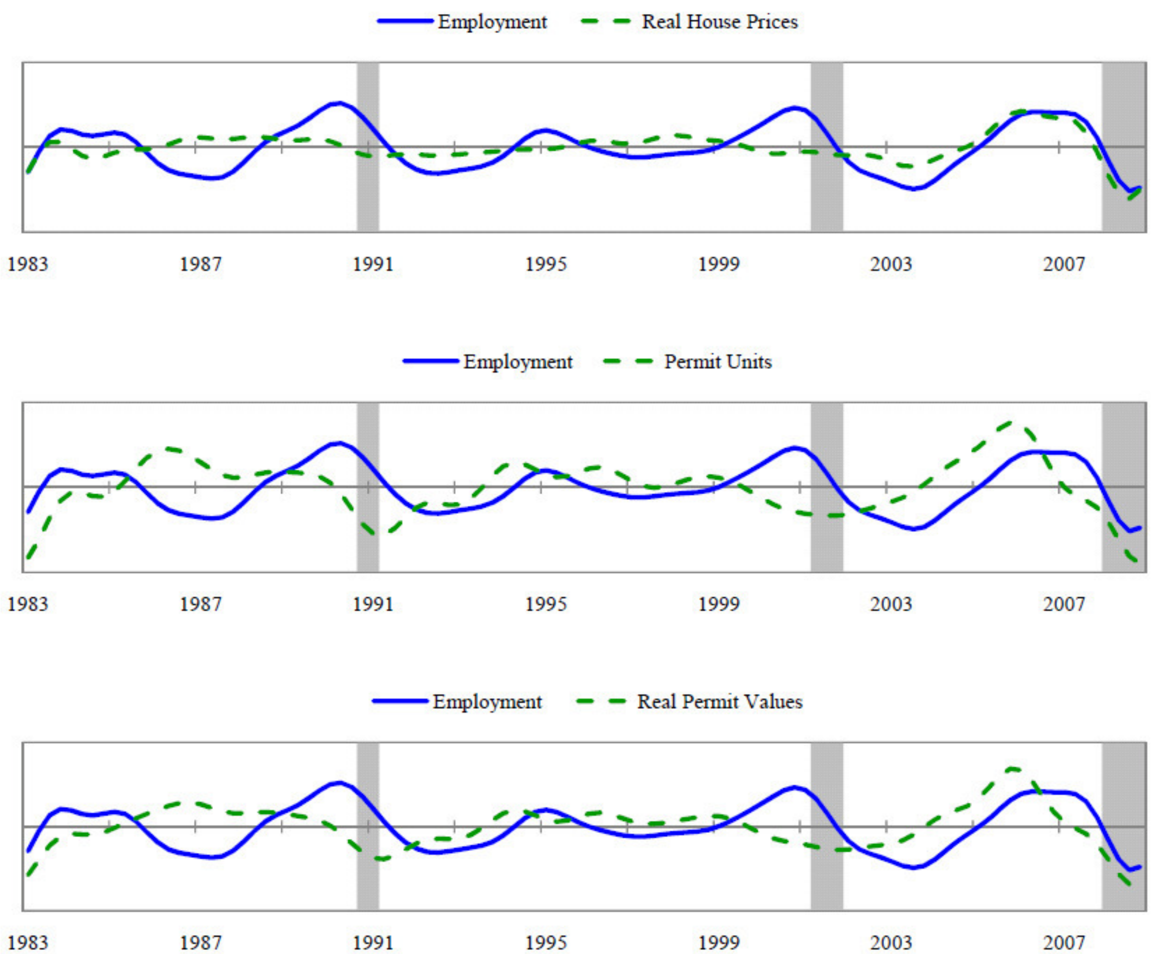


Figure 1: Band-Passed U.S. Employment and Band-Passed U.S. Housing Variables
Shaded regions denote NBER Recessions

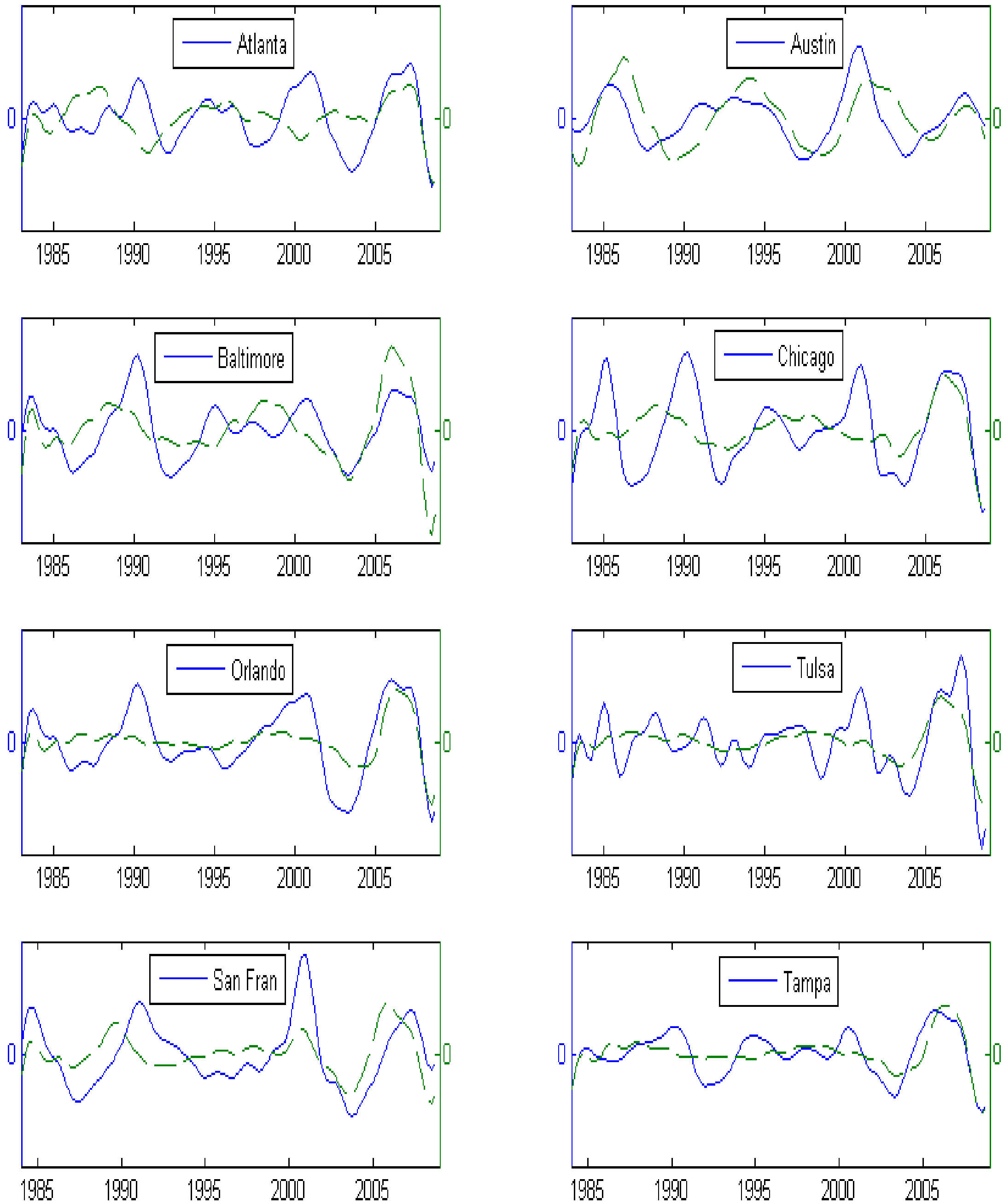


Figure 2: Band-Passed Employment (Solid Line) and Band-Passed Real House Prices (Dashed Line) for Selected Cities

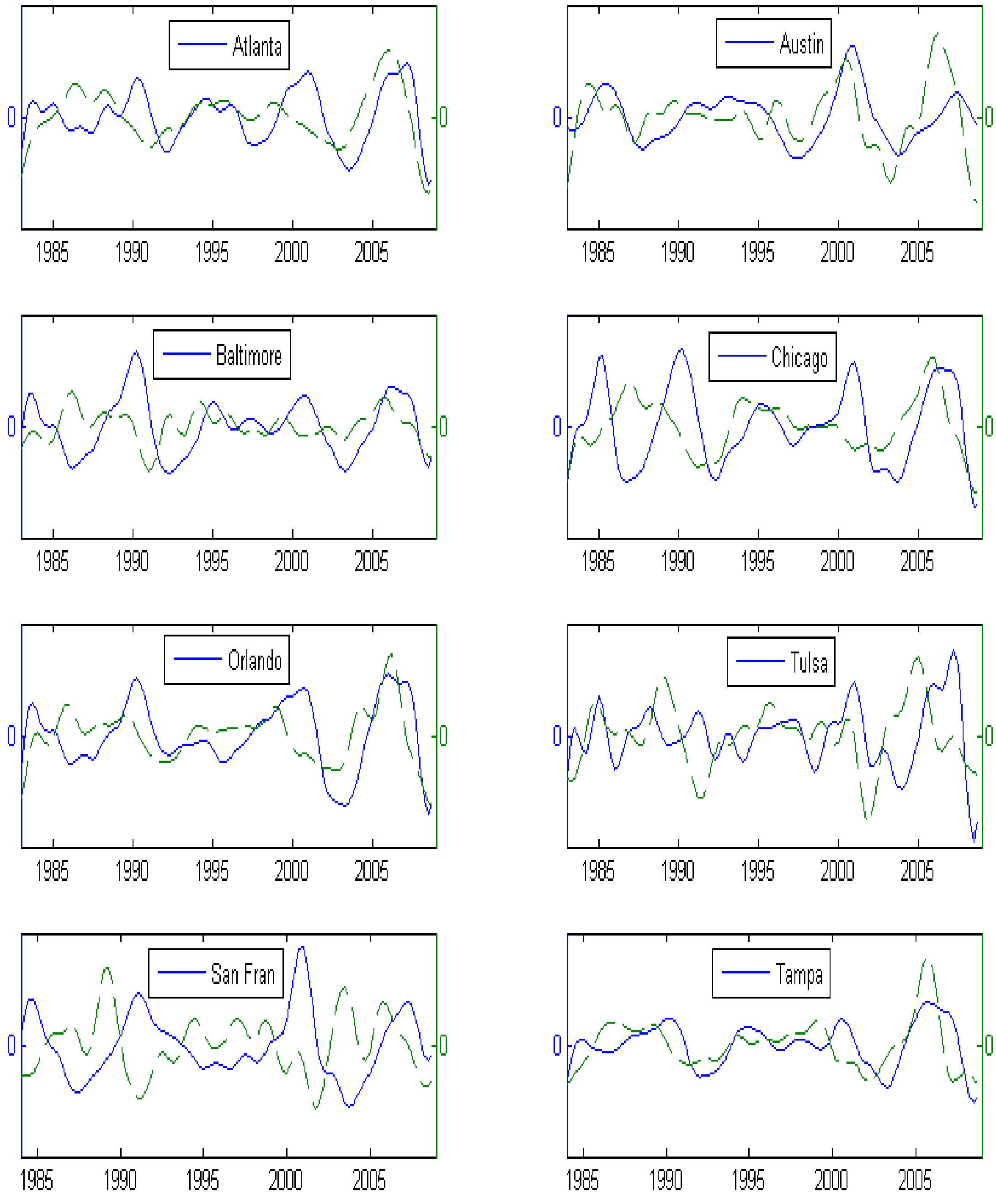


Figure 3: Band-Passed Employment (Solid Line) and Real Permit Values (Dashed Line) for Selected Cities

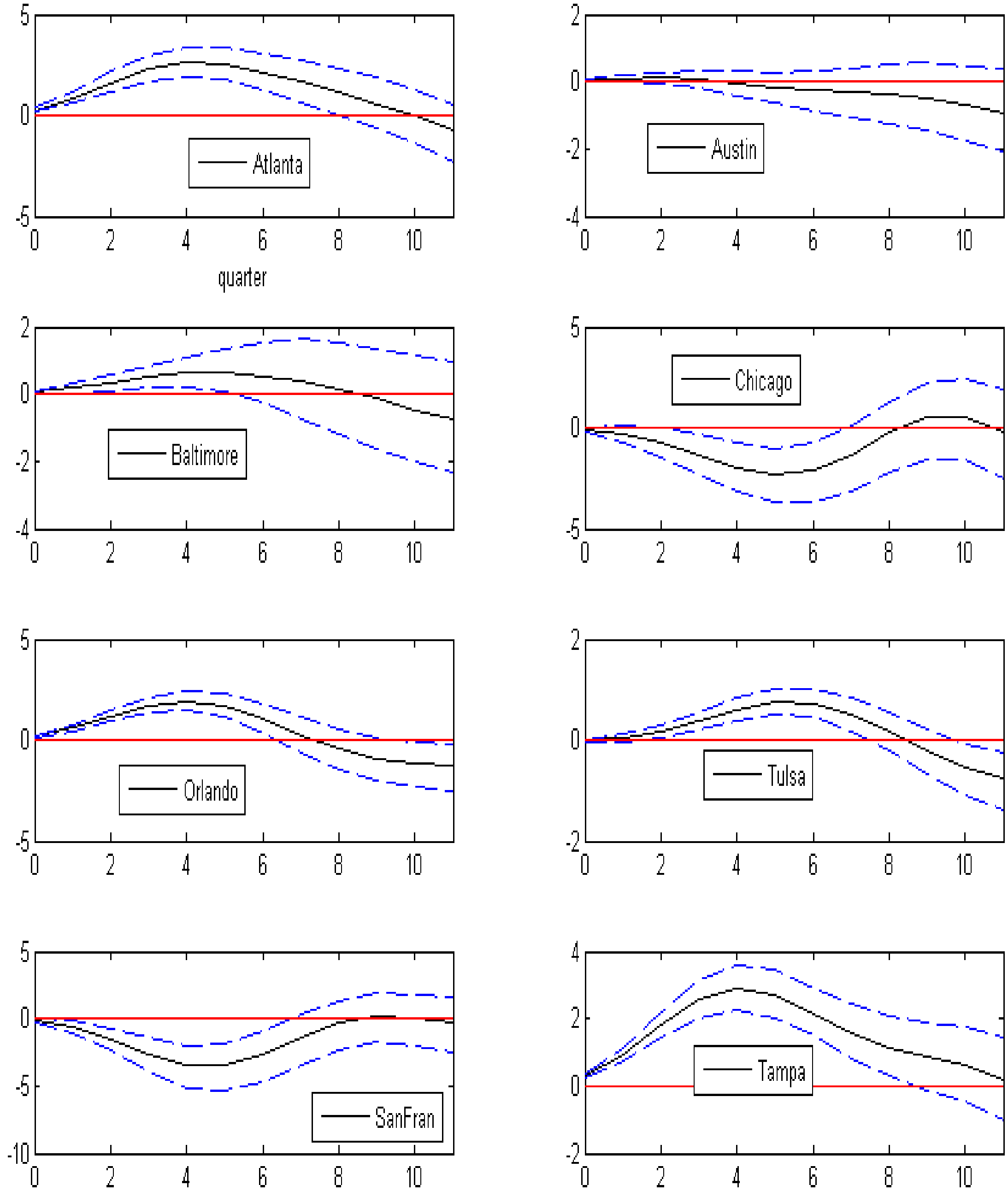


Figure 4: Impulse Responses of MSA Employment to an MSA Permit Shock, Selected Cities
Dashed lines denote 10th and 90th percentiles of the posterior distribution

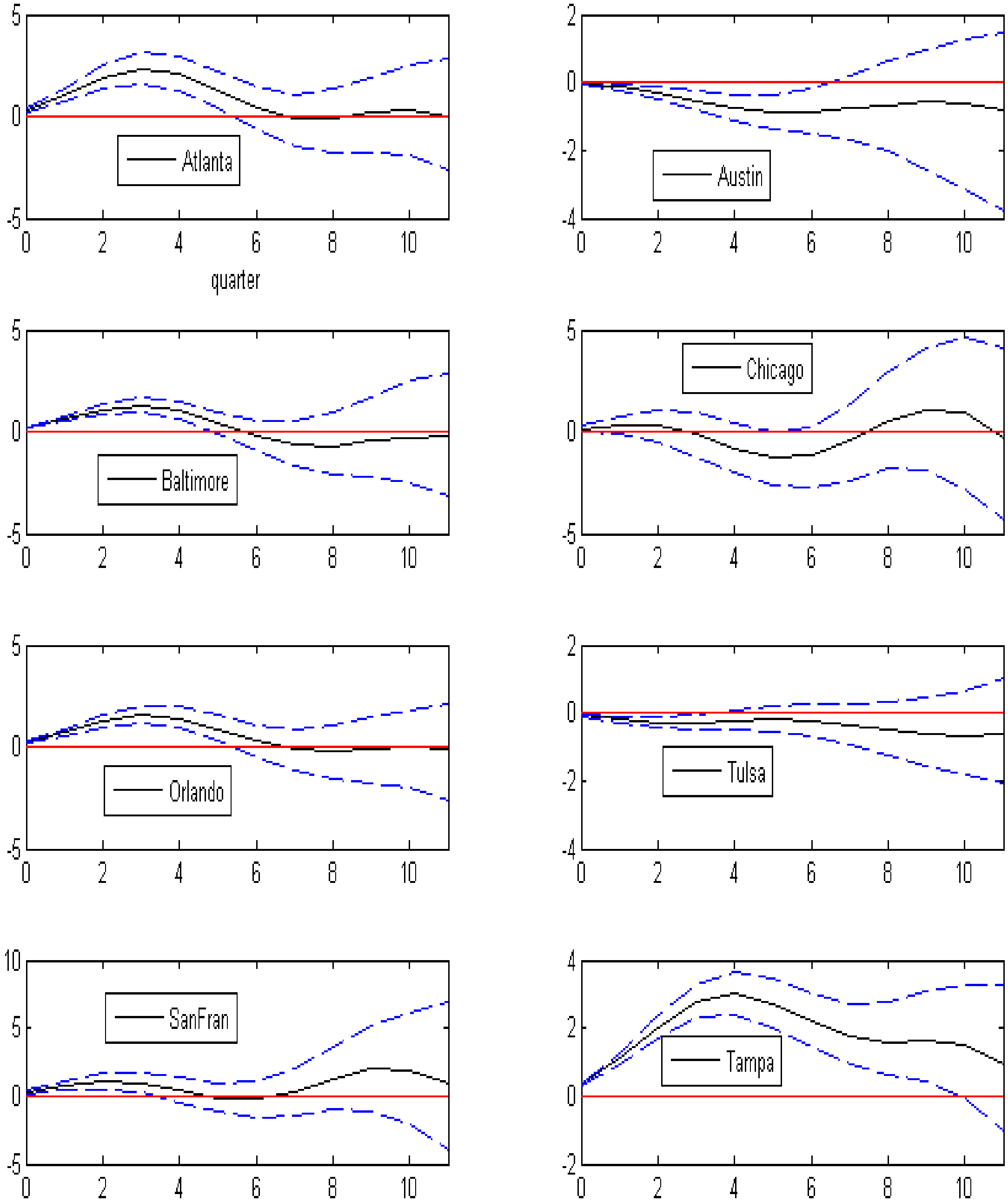


Figure 5: Impulse Responses of MSA Employment to U.S. Permit Shock, Selected Cities
Dashed lines denote 10th and 90th percentiles of the posterior distribution

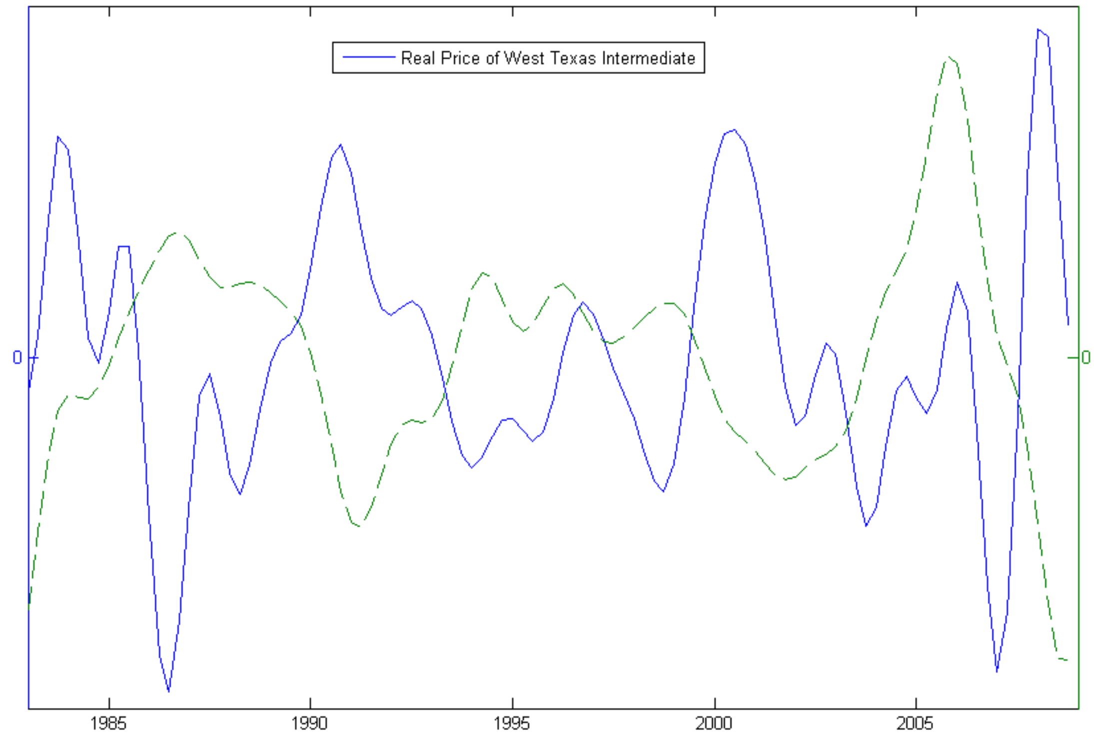
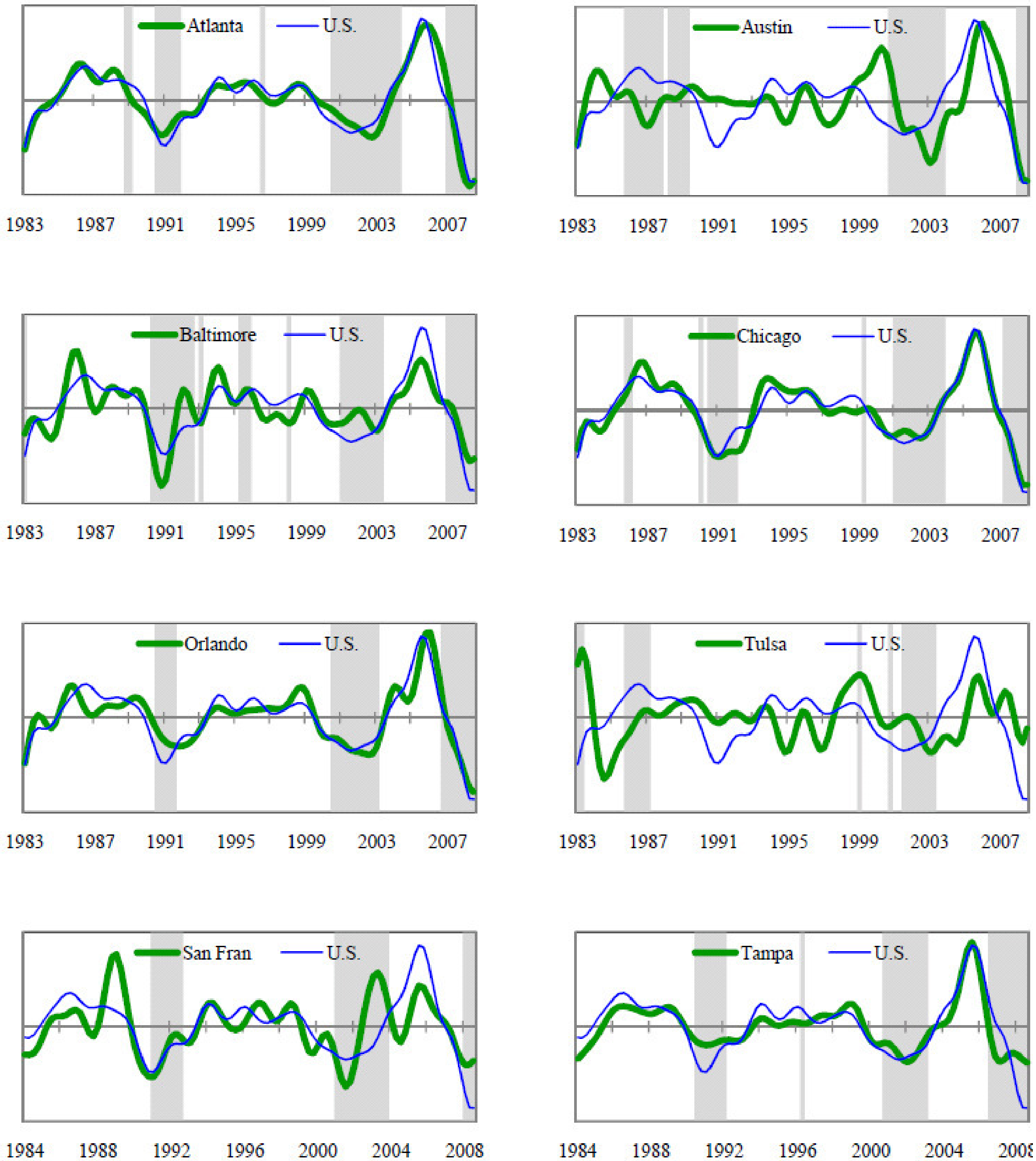


Figure 6: Bandpassed Real Oil Price vs. Bandpassed Real U.S. Permits



Recession Dates for Selected Cities and Band-pass Filtered MSA and U.S. Real Permit Values

Shaded regions indicate that the probability the city is in a recession is greater than 60%. MSA and U.S. permit values are on different scales.

Recession Dates for Selected Cities and Band-pass Filtered MSA and U.S. Real House Prices

Shaded regions indicate that the probability the city is in a recession is greater than 60%. MSA and U.S. real house prices are on different scales.