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Long-run Equilibrium and Short-Run Adjustment in U.S. Housing Markets

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This paper examines the long-run equilibrium between real house prices and macroeconomic fundamentals in U.S. housing markets, as well as the short-run adjustment of real house prices back to the equilibrium. Pooled mean-group and mean-group estimation techniques developed by Pesaran and Smith (1995) and Pesaran et al. (1999) are applied to a panel of the 51 U.S. states over the period of 1976Q3 to 2012Q4. Our results suggest a common long-run relationship over the sample period between real house prices and their economic fundamental determinants in the 51 U.S. states. However, the speed of adjustment of real house prices varies vastly across states, with a half-life estimate of 22 quarters on average, and the deviations of real house prices from the equilibrium range from –30% to 46% across states over time.

Keywords

House Price, Panel Unit Root, Panel Cointegration, Pooled Mean-group Rstimator, Mean-Group Estimator

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1. Introduction

The relationship between house prices and key macroeconomic variables is of great concern to policymakers and researchers, especially after the meltdown of the U.S. housing market which started around 2006 and the subsequent global financial crisis. In this context, there are two main streams of literature. One stream argues that there is no bubble in the U.S. housing market and that changes in house prices reflect movements in macro fundamentals, such as personal income, unemployment, mortgage rates, etc. For instance, Leung (2014) builds a dynamic stochastic general equilibrium model to justify that house prices and income are co-integrated. Courchane and Holmes (2014) find that U.S. house prices were closely aligned with economic fundamentals before 2008 when the mortgage markets crashed. Nneji et al. (2013) investigate whether intrinsic bubbles and rational speculative bubbles resulted in deviations in U.S. house prices from economic fundamentals during 1960–2011. Holly et al. (2010) study the determinants of real house prices in a panel of 49 U.S. states from 1976 to 2007 and conclude that the rising real house prices were in line with real income. Other influential studies include Malpezzi (1999), McCarthy and Peach (2004), Himmelberg et al. (2005), and Smith and Smith (2006). Empirical studies on the impact of macroeconomic variables on house prices with the use of international data are also rich (e.g., Apergis, 2003; Borowiecki, 2009; Deng et al., 2009; Kholodilin et al., 2010).

The other strand of the literature finds no evidence of a long-run relationship between house prices and macro-fundamentals, which implies that house prices are not in line with fundamentals, and thus, housing bubbles may exist. Stiglitz (1990) defines a house price bubble as a situation in which house price growth is not supported by changes in its fundamentals. For instance, Peláez (2012) suggests a disequilibrium existed between house prices and per capita income during 2003–2007 in the U.S. Arshanapalli and Nelson (2008) find evidence of a U.S. housing bubble from 2000 to 2007 with a cointegration test. Gallin (2006) uses both U.S. national-level data and a panel of 95 U.S. cities and concludes that house prices and income are not cointegrated. Meen (2002) and Shiller (2005) use aggregate U.S. data and find that changes in fundamentals do not explain the surge in U.S. house prices after 2000.

In the literature, the fundamental model of equilibrium house prices is one of the influential theories about the evolution of equilibrium house prices.[1] This model compares observed house prices with their fundamental values that are estimated based on the long-run relationship between house prices and macroeconomic fundamentals (Kaufmann and Mühleisen, 2003; Klyuev, 2008; Holly et al., 2010). *In this paper, we examine the long-run relationship between*

[1] An alternative theory includes the asset price approach, which compares observed price-rent ratios with time-varying discount factors that are determined by the user cost of owning a house (Campbell and Shiller, 1987; Wang, 2000; Mikhed and Zemčík, 2009a).

real house prices and macroeconomic fundamentals and the short-run adjustment of real house prices to the equilibrium in the 51 U.S. states over the period 1976Q3–2012Q4. [2] What distinguishes our work from previous empirical studies on the relationship between house prices and macroeconomic fundamentals are the following: (1) our work applies the pooled mean-group (PMG) and mean-group (MG) estimation techniques developed by Pesaran and Smith (1995) and Pesaran et al. (1999) to study both the long-run and the short-run behaviors of real house prices at the U.S. state level—and, in particular, the deviations of real house prices from their fundamentals and the speed of adjustment of real house prices to macroeconomic disturbances. The PMG and MG estimations are implemented in levels of data and allow for non-stationarity and cointegration across a panel of data well suited to capturing housing market characteristics, and (2) our study uses the most recent quarterly state-level data, which better capture the correction of real house prices across the 51 U.S. states after the recent collapse of the housing market in the U.S.[3] Our analysis aims to provide disaggregate perspectives of U.S. housing markets and show the common characteristics of such markets across the entire U.S.

The recent empirical literature has extensively investigated the relationship between house prices and economic fundamentals. The findings vary with the econometric models and the data used. For instance, Mikhed and Zemčík (2009a) confirm the existence of housing bubbles in 23 U.S. metropolitan statistical areas (MSAs) from the first half of 1978 to the second half of 2006 by using the panel unit root test by Pesaran (2007) and the panel cointegration test by Pedroni (1999, 2004) to account for cross-sectional dependence. Mikhed and Zemčík (2009b) use aggregate quarterly U.S. data for 1980Q2 to 2008Q2 and annual data on 22 U.S. MSAs from 1978 to 2007 to show that house prices do not align with fundamentals in sub-samples before 1996 and from 1997 to 2006. Clark and Coggin (2011) apply both unit root and cointegration tests to the U.S. nation-wide and divided into four regions over the period of 1975Q1 to 2005Q2, since these tests are technically immune to the cross-sectional dependence problem and explicitly allow for structural breaks. They find that U.S. house prices and fundamental economic variables are unit root variables that are not cointegrated. Vector error correction models (VECMs) have recently been applied to handle both the non-stationarity and endogeneity problems in the study of house price determinants. Such models also distinguish between long-run relationships and short-run adjustments. For instance, Wheaton et al. (2014) estimate VECMs separately for 68 U.S. MSAs by using quarterly data for house prices and residential construction permits from

[2] The United States consists of the 50 states and the District of Columbia, which are referred to as the 51 states hereafter.

[3] We do not consider MSA-level data in this study because MSAs correspond to labor market areas within which workers are willing to commute, and thus they cannot represent the entire U.S. housing market. For example, Pan and Wang (2013) find that a common positive long-run relationship among house prices, personal income, and labor force growth does exist in 286 U.S. MSAs but not in U.S. non-MSAs.

1980Q1 to 2012Q2. Panel error correction models, which are combinations of panel data and error correction models, have also been applied to study housing markets (e.g., Hendershott et al., 2002; Brounen and Jennen, 2009; Ott, 2014).

In this paper, to investigate the possible long-run relationship between real house prices and their macroeconomic fundamental variables, we first conduct three panel unit root tests—those by Im et al. (2003), Maddala and Wu (1999), and Pesaran (2007)—to verify the order of integration for each variable in level and first difference (quarter-to-quarter change). Then, we apply the panel cointegration tests of Westerlund (2007) to investigate the existence of a long-run relationship between real house prices and their economic fundamental variables. In contrast to previous studies, which have used the residual-based cointegration tests by Engle and Granger (1987), Phillips and Ouliaris (1990), or Pedroni (1999, 2004), our study uses the Westerlund tests which are based on structural (rather than residual) dynamics, have good size accuracy, and are more powerful than the residual-based tests. Moreover, the Westerlund tests accommodate individual specific short-run dynamics and allow for cross-sectional dependence. Three out of four tests reject the null hypothesis of no panel cointegration for the full sample from 1976Q3 to 2012Q4, thus suggesting the existence of a long-run relationship between real house prices and their fundamental values over the full sample period. This result provides empirical justification for applying the PMG and MG estimations to real house prices and economic fundamentals.

The PMG and MG estimators are two important methods for estimating non-stationary dynamic panels, while allowing for heterogeneous parameters across groups. Compared to the methods used in previous studies, the PMG and MG estimations are conducted in levels and take into account any non-stationarity and cointegration in the panel data. The PMG estimator imposes a homogeneity restriction on the long-run relationship between variables, while the MG estimator does not (Koetter and Poghosyan, 2010). These estimators have been previously used to study house price determinants in various countries. For example, Kholodilin et al. (2010) analyze house price determinants in an international sample of countries, Stepanyan et al. (2010) study selected countries from the former Soviet Union, and Koetter and Poghosyan (2010) focus on regional housing markets in Germany. In this paper, we employ the PMG and MG estimators to examine real house price determinants in the U.S. by using a panel of the 51 U.S. states over the period of 1976Q3 to 2012Q4. Our results show that the PMG estimator is preferable, which suggests a common long-run relationship among real house prices, real personal income per worker, population growth, unemployment rates, and the net cost of borrowing across the 51 U.S. states. However, there is substantial heterogeneity in the speed of the short-run adjustment of real house prices and the deviations of real house prices from their long-run equilibrium across states. In addition, the deviations of real house prices from fundamentals in the post-crisis period (2010Q4) are greater than those during the peak (2007Q1) which imply that the economic fundamentals deteriorated even more rapidly than real house prices in the post-crisis period. This is consistent with the finding of Stepanyan et al. (2010) for the former Soviet Union countries. Overall, our results can provide important empirical insight into modeling house price dynamics.

The remainder of this paper is organized as follows. Sections 2 and 3 describe the empirical methods and data, respectively. The empirical results are presented in Section 4. Finally, Section 5 concludes and sheds light on some policy implications.

2. Econometric Model and Techniques

2.1 Testing for Panel Unit Roots and Cointegration

Before investigating the possible long-run relationship between real house prices and their macroeconomic fundamental variables, we begin with three panel unit root tests from Im et al. (2003), Maddala and Wu (1999), and Pesaran (2007) to verify the order of integration for each variable in level and first difference. The null hypothesis for all three tests is that all panels contain unit roots. Specifically, the Im-Pesaran-Shin (IPS) unit root test is based on the individual augmented Dickey-Fuller (ADF) regression:

$$\Delta y_{it} = \alpha_i + \delta_i t + \lambda_i y_{i,t-1} + \sum_{j=0}^{p_i} \phi_{ij} \Delta y_{i,t-j} + \varepsilon_{it}, \quad (1)$$

where i and t indicate state and time, respectively, and y denotes real house prices. The IPS statistic is the average of the t-statistics (denoted as t_i) for λ_i's in the individual ADF regressions:

$$t_{IPS} = \frac{\sqrt{N}(\bar{t} - E[t_i|\rho_i=0])}{\sqrt{\mathrm{var}[t_i|\rho_i=0]}} \rightarrow N(0,1), \quad (2)$$

where $\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t_i$.

The Maddala and Wu (MW) test statistic is obtained by $P = -2\sum_{i=1}^{N} ln\, p_i$, and combines the p-values from the individual ADF tests. Both the IPS and MW tests assume no cross-sectional dependence in the panel data. However, house prices and macroeconomic fundamentals show strong cross-sectional dependence (Mikhed and Zemčík 2009a, 2009b; Holly et al., 2010), which should be taken into account in testing for unit roots and cointegration.

Pesaran (2007) proposes a panel unit root test robust to cross-sectional dependence, known as the CIPS test. It is based on the cross-section augmented Dickey-Fuller (CADF) regression:

$$\Delta y_{it} = \alpha_i + \delta_i t + \lambda_i y_{i,t-1} + \eta_i \bar{y}_{t-1} + \sum_{j=0}^{p_i} \phi_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta \bar{y}_{t-j} + \varepsilon_{it}, \quad (3)$$

where $\bar{y}_t$ is the cross-section mean of $\bar{y}_{it}$. The CIPS statistic is the cross-section average of $\tilde{t}_i$:

$$t_{CIPS} = \frac{1}{N} \sum_{i=1}^{N} \tilde{t}_i(N,T), \quad (4)$$

where $\tilde{t}_i$ is the t-statistic for λ_i in the individual CADF regression.

If the variables were found to be non-stationary based on the unit root tests, then we would have needed to further examine whether they were cointegrated. According to the economic theory, if real house price developments are in line with economic fundamentals, non-stationary real house prices should be cointegrated with other non-stationary economic variables with the same order of integration. We apply the panel cointegration tests of Westerlund (2007) to investigate the existence of a long-run relationship between real house prices and other key economic fundamental variables. The tests allow for a large degree of heterogeneity, both in the long-run cointegrating relationship and the short-run dynamics, and dependence within as well as across the cross-sectional units. The null hypothesis is that of no cointegration in the panel. In particular, the data-generating process for the error-correction tests is:

$$\Delta y_{it} = \delta_i' d_t + \alpha_i \left(y_{i,t-1} - \beta_i' x_{i,t-1} \right) + \sum_{j=1}^{p_i} \phi_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \phi_{ij} \Delta x_{i,t-j} + e_{it}, \quad (5)$$

where d contains the deterministic components and x represents economic fundamental variables. Westerlund (2007) proposes four tests based on the least squares estimate of α_i in Eq.5 and its t-ratio.[4]

The group-mean statistics are calculated as:

$$G_\tau = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{\alpha}_i}{SE(\hat{\alpha}_i)} \text{ and } G_\alpha = \frac{1}{N} \sum_{i=1}^{N} \frac{T\hat{\alpha}_i}{\hat{\alpha}_i(1)} T \quad (6)$$

where $SE(\hat{\alpha}_i)$ is a conventional standard error of $\hat{\alpha}_i$. The panel statistics are computed as

$$P_\tau = \frac{\hat{\alpha}}{SE(\hat{\alpha})} \text{ and } P_\alpha = T\hat{\alpha} \quad (7)$$

[4] Details of the test procedures are provided in Westerlund (2007).

2.2 PMG and MG Estimations

To further study the long-run equilibrium between real house prices and macroeconomic fundamentals in U.S. housing markets, as well as the short-run adjustment of real house prices, we apply the PMG and MG estimators to a panel of the 51 U.S. states over the period 1976Q3–2012Q4. The PMG and MG estimators have been proposed to estimate non-stationary dynamic panels in which the parameters are heterogeneous across groups. The main difference between the two estimators is that the PMG estimator imposes a homogeneity restriction on the long-run relationship between variables while the MG estimator does not. Such homogeneity restrictions imposed by the theory can be tested empirically by using the Hausman test.

The house price determinants frequently studied in the housing literature include real income per capita, population growth, unemployment rates, and real interest rates (e.g., Muellbauer and Murphy, 1997; Meen, 2002; Barker, 2005).[5] Holly et al. (2010) provide a theoretical model that justifies the existence of cointegration between real house prices and real income per capita, as well as a role for the real interest rate and demographic factors.[6] Compatible with the long-run theory and the cointegrating relationship among the variables of interest, we describe the long-run relationship between real house prices and their fundamentals in the following log-linear form[7]:

$$RHP_{it} = \beta_0 + \beta_{1i}RINC_{it} + \beta_{2i}POP_{it} + \beta_{3i}UR_{it} + \beta_{4i}RMORT_{it-1}$$

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[5] Other factors have also been considered in the literature, such as building costs (Shiller, 2005), ownership costs of housing (Himmelberg et al., 2005), and rent (Mikhed and Zemčík, 2009a, 2009b). However, data for these variables are not available for the 51 U.S. states at quarterly frequency, therefore, we do not include them in our study.

[6] In their theoretical model, the house price-income ratio, also known as the affordability index, is stationary. This implies that the log of the real house price index will be cointegrated with the log of real income per capita with the cointegrating vector given by $(1, -1)$, if the log of the real house price index is an integrated variable of order 1, i.e., I(1). In the long run, therefore, the elasticity of real house prices to real income is unity. In addition, they also consider the possible effect of population growth rates on the log of the real house price index at the state level. In aggregate time-series analysis, it is difficult to identify the effects of slowly moving variables such as population growth on real house prices. However, in the panel context, the cross-section dimension can be used to identify such effects. For a given level of real income per capita, real house prices are expected to be higher in states with a higher population growth rate.

[7] This is a semi-loglinear specification for real house prices and the fundamentals, which is commonly used in the empirical literature on the long-run relationship between house prices and the determinants (see, e.g., Terrones and Otrok, 2004; Ahearne et al., 2005; Almeida et al., 2006; Égert and Mihaljek, 2007; Iossifov et al., 2008, and Stepanyan et al., 2010). This specification reflects the fact that population growth is considered stationary in the steady state of the economy, and the level of economic development (approximated by the level of real income per capita), combined with other determinants, influence the level of real house prices in the long run.

$$+\mu_i + \eta D + \varepsilon_{it}, \tag{8}$$

where i and t indicate state and time, respectively; RHP is the (log) real house price index; and $RINC$ is the (log) real personal income per worker. Based on the theory in Holly et al. (2010), the elasticity of real house prices to real income is unity in the long run. In addition to real income per capita, other factors such as changes in demographics, unemployment rates, and the net cost of borrowing also play a role in the determination of real house prices at the state level. POP represents the rate of change in population, UR is the unemployment rate, and $RMORT$ is the net cost of borrowing defined by the real long-term mortgage rate net of real house price appreciation or depreciation as in Holly et al. (2010) and Kholodilin et al. (2010), which is included in Eq.8 with a lag to avoid simultaneity.[8] *A priori* we would expect that a rise in population growth and a fall in the unemployment rate would be associated with higher real house prices, while a rise in the net cost of borrowing would negatively influence real house prices. Finally, μ_i is the state-specific fixed effect and D represents the vector of dummy variables that capture the impact of policy interventions [9] and common shocks to the economy (e.g., the Interstate Banking and Branching Deregulation Index developed by Rice and Strahan (2010), and the recent financial crisis). One feature of the model in which we are interested is the extent to which real house prices are driven by fundamentals such as real income per capita, population growth, unemployment rate, and the net cost of borrowing. If the variables are integrated of order one (i.e. I(1)) and co-integrated, then the error term ε_{it} is stationary (i.e. I(0)) for all i . The autoregressive distributed lags (ARDLs)(p,q,q,q,q), dynamic panel representation of the long-run Eq.8 is:

$$RHP_{it} = \sum_{j=1}^{p} \lambda_{ij} RHP_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}^{1} RINC_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}^{2} POP_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}^{3} UR_{i,t-j} + \sum_{j=0}^{q+1} \delta_{ij}^{4} RMORT_{i,t-j} + \mu_i + \eta D + \varepsilon_{it} \tag{9}$$

The model specification in the error-correction form of Eq.9 is as follows:

$$\Delta RHP_{it} = \alpha_i (RHP_{it-1} - \beta_{0i} - \beta_{1i} RINC_{i1} - \beta_{2i} POP_{it} - \beta_{3i} UR_{it} - \beta_{4i} RMORT_{it-1}) + \sum_{j=1}^{p-1} \gamma_{ij} \Delta RHP_{it-j} + \sum_{j=0}^{q-1} \theta_{ij}^{1} \Delta RINC_{it-j} + \sum_{j=0}^{q-1} \theta_{ij}^{2} \Delta POP_{it-j} + \sum_{j=0}^{q-1} \theta_{ij}^{3} UR_{it-j} + \sum_{j=1}^{q} \theta_{ij}^{4} RMORT_{it-j} + \varepsilon_{it} \tag{10}$$

where $\alpha_i = -\left(1 - \sum_{J=1}^{p} \lambda_{ij}\right)$, $\beta_{0i} = \frac{\mu_i}{-\alpha_i}$, $\beta_{1i} = \frac{\sum_{j=0}^{q} \delta_{ij}^{1}}{-\alpha_i}$, $\beta_{2i} = \frac{\sum_{j=0}^{q} \delta_{ij}^{2}}{-\alpha_i}$,

[8] The appendix defines the variables used in more detail.

[9] There have been significant policy market interventions and bank deregulation during the sample period, for instance, the Community Reinvestment Act of 1977 (October 12, 1977), the Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994 (September 29, 1994), and the Financial Services Modernization Act of 1999 (November 12, 1999).

$\beta_{3i} = \frac{\sum_{j=0}^{q} \delta_{ij}^{3}}{-\alpha_i}$, and $\beta_{4i} = \frac{\sum_{j=1}^{q+1} \delta_{ij}^{4}}{-\alpha_i}$.

The error-correction term ($RHP_{it-1} - \hat{\beta}_{0i} - \hat{\beta}_{1i} RINC_{it} - \hat{\beta}_{2i} POP_{it} - \hat{\beta}_{3i} UR_{it} - \hat{\beta}_{4i} RMORT_{it-1}$) represents the temporary deviations of real house prices from their fundamental values at the state level. The homogeneity restriction imposed by the PMG estimator is on the coefficients of long-run real house price determinants β_1, β_2, β_3, and β_4, restricting all the long-run parameters to be the same across states. This restriction can be relaxed to restricting only the subset of the long-run parameters to be the same across states. The intercept β_{0i}, speed of the adjustment parameter α_i and short-run adjustment coefficients θ_{ij}^{1}, θ_{ij}^{2}, θ_{ij}^{3}, and θ_{ij}^{4} vary across states. We expect a negative speed of adjustment parameter α_i, which suggests that real house prices react to disequilibrium in the real estate market: real house prices decrease following positive deviations from the long-run equilibrium in the real estate market, while they increase following negative deviations from the long-run equilibrium.

3. Data

The data for house prices and macroeconomic variables at the U.S. state level cover the 50 states and the District of Columbia over the period 1976Q3–2012Q4. We obtain the quarterly house price all-transactions index (estimated by using sales prices and appraisal data) from the Federal Housing Finance Agency (FHFA). The index is based on transactions and appraisals, and then adjusted for appraisal bias.[10] Following previous literature, we use personal income per worker, population growth, and unemployment rates as the house price determinants to estimate the house price deviations. The U.S. state-level data on personal income are obtained from the U.S. Bureau of Economic Analysis (BEA). To obtain the real house price index and real personal income, the house price index and personal income are divided by the consumer price index (CPI), which is available for four census regions in quarterly frequency from the U.S. Bureau of Labor Statistics (BLS).[11] States in a particular census region share the same CPI. Inflation rates are calculated based on the CPI as well.[12] State population data are obtained from the U.S. Census Bureau.[13] Civil

[10] The FHFA house price index includes only homes with mortgages that conform to Freddie Mac and Fannie Mae guidelines. Jumbo loans over $417,000 are not included. This index is equally weighted regardless of the value of the house.
[11] The U.S. Census Bureau groups the 50 states and the District of Columbia into four census regions, namely, Northeast, Midwest, South, and West regions.
[12] We do Census X12 multiplicative seasonal adjustment for the CPI, and then calculate the annual rate of inflation based on the CPI.
[13] Raw population data is available in yearly frequency and we convert the annual data to quarterly frequency by using constant match in EViews.

labor force and unemployment rates at the state level are taken from the BLS.[14] Mortgage rates are available for four census regions in quarterly frequency from the Federal Reserve Bank of St. Louis (FRED). States in a particular census region share the same mortgage rates. Following Holly et al. (2010) and Kholodilin et al. (2010), we construct the net cost of borrowing as the real long-term mortgage rate net of real house price appreciation or depreciation. The long-term interest rate is adjusted by using the housing price index and not the CPI based on the considerations of a household, which makes the decision about the long-term investment of buying a housing asset. This rate compares the interest income from a bank deposit with capital gains from changes in housing prices. A detailed description of the data is provided in the Appendix.

Table 1 reports the summary statistics. We observe that the real house price index (1980=100) varied across the states over time, with a minimum of 52 (Hawaii, 1981Q4), a maximum of 345 (Massachusetts, 2005Q3), and an average of 135. The macroeconomic variables are more dispersed. Specifically, real personal income per worker varied from US$18,179 (Mississippi, 1980Q2) to US$60,802 (District of Columbia, 2010Q4), with an average of US$30,254. Unemployment rates ranged from 2.1% (New Hampshire, 1987Q1) to 18% (West Virginia, 1983Q1), with an average of 6%.

Table 1 Summary Statistics

State Variable	Mean	Sd	Min	Max
Real house price index	134.70	38.20	51.60	344.60
Real personal income per worker	30254.40	5682.00	18179.40	60801.90
Population	5,190,000	5,760,000	393,000	38,000,000
Unemployment rate (%)	6.08	2.13	2.10	18.03
Net cost of borrowing (%)	7.59	4.55	-111.93	61.01
Inflation rate (%)	3.88	3.38	-10.35	19.59
Log of real house price index	0.27	0.25	-0.66	1.24
Log of real personal income per worker	5.70	0.18	5.20	6.41
Growth rate of population (%)	1.02	1.10	-5.99	8.63
Speed of adjustment	-0.03	0.02	-0.09	0.00
House price deviations (%)	-0.09	31.21	-153.76	143.38

Note: 1. The sample contains 7,446 observations in the 51 U.S. states over the period 1976Q3–2012Q4.
2. Real house price index (1980=100). Personal income per worker is deflated by the consumer price index (base year is 1980).
3. We do Census X12 multiplicative seasonal adjustment for the consumer price index (CPI), then calculate annual rate of inflation based on the CPI.

[14] Civil labor force and unemployment rates are reported in monthly frequency and we convert monthly to quarterly frequency by using the average method in EViews.

4. Empirical Results

4.1 Panel Unit Root and Cointegration Test Results

Table 2 summarizes the panel unit root results based on the full sample from 1976Q3 to 2012Q4. To conduct the testing, we include time trends for the real house price index and real personal income per worker, since both variables exhibit a clear upward trend over time in the sample. The number of lags for each variable in each state is chosen automatically by using the Bayesian information criterion (BIC) with a maximum of four lags[15]. Boldface values denote sampling evidence in favor of unit roots. For the full sample, all three panel unit root tests suggest that the real house price index and real personal income per worker are integrated of order one, or I(1), and that the remaining four variables are stationary, or I(0), at the 5% significance level.

Table 2 Panel Unit Root Tests

Variable	**Level**			**First Difference**		
	IPS	MW	Pesaran	IPS	MW	Pesaran
Real house price index (log)	-0.462 (0.32)	113.119 (0.21)	-2.241 (0.85)	-41.488 (0)	656.214 (0)	-3.723 (0)
Real personal income per worker (log)	-0.086 (0.47)	118.776 (0.12)	-2.268 (0.79)	-79.185 (0)	1329.642 (0)	-5.081 (0)
Population growth	-7.199 (0)	214.361 (0)	-2.297 (0)	-80.715 (0)	1307.337 (0)	-4.988 (0)
Unemployment rate	-7.262 (0)	187.716 (0)	-2.010 (0.04)	-41.322 (0)	1022.729 (0)	-4.572 (0)
Net cost of borrowing	-43.502 (0)	643.433 (0)	-3.604 (0)	-1.00E+02 (0)	2875.171 (0)	-6.128 (0)
Inflation rate	-52.808 (0)	1361.37 (0)	-5.141 (0)	-1.10E+02 (0)	3187.687 (0)	-6.19 (0)

Note: The sample contains the 51 U.S. states over the period 1976Q3–2012Q4. *P*-values are reported in the parentheses. Boldface values denote sampling evidence in favor of unit roots. The null hypothesis is that of a unit root. The lags are chosen automatically by using the BIC with maximum four lags. Trend option is included in the testing for the real house price index and real personal income per worker in level. IPS represents the Im-Pesaran-Shin (2003) unit root test; MW denotes Maddala and Wu (1999) unit root test; and CIPS stands for Pesaran (2007) unit root test.

Given the panel unit root test results for the full sample, we proceed to conduct the panel cointegration tests of Westerlund (2007) on the two I(1) variables (i.e., real house price index and real personal income per worker). Table 3 reports the results. Three of the four statistics and the corresponding *p*-values suggest rejecting the null of no panel cointegration. Therefore, our results confirm the existence of a long-run relationship between real house prices and their

[15] Results are consistent when eight lags are used as the maximum.

fundamental values over the period of 1976Q3 to 2012Q4, which satisfies the assumptions in the following PMG and MG estimations.

Table 3 Panel Cointegration Tests of Westerlund (2007) on the Real House Price Index and Real Personal Income Per Worker

Statistic	Value	Critical Value	P-value
G_τ	-2.104	-2.593	0.01
G_α	-7.07	0.094	0.54
P_τ	-12.27	-1.943	0.03
P_α	-5.672	-2.319	0.01

Note: The sample contains the 51 U.S. states over the period 1976Q3–2012Q4. The null hypothesis is that of no cointegration in the panel. The lags are chosen automatically by using the BIC with maximum four lags.

4.2 Long-Run Equilibrium between Real House Prices and Macroeconomic Fundamentals

Table 4a reports the estimation results of the PMG and MG specifications of Eq.10. The upper panel reports the average long-run coefficient estimates in Eq.10 for the log of the real house price index. As a benchmark, Model I includes the log of real personal income per worker, population growth rates, unemployment rates, and the net cost of borrowing as the macroeconomic fundamentals for the log of the real house price index.[16] Notably, the Hausman tests with p-values greater than 0.05 suggest that the PMG estimator is preferable, since we fail to reject the null hypothesis that the difference between the PMG and MG models is not systematic. This suggests a common long-run relationship among real house prices, real personal income per worker, population growth, unemployment rates, and the net cost of borrowing across the 51 U.S. states. In the PMG estimations that restrict common long-run coefficients across states, we find a statistically significant positive long-run relationship between the log of the real house price index and the log of real personal income per worker, and the income elasticity of house prices is 0.417. The impact of population growth on real house prices is also significantly positive. In line with our expectations, unemployment rates and the net cost of borrowing have a significantly negative effect on real house prices in the long run. Overall, our results confirm that the long-run equilibrium real house prices increase with rising demand due to higher income and population growth, but lower unemployment rates and the net cost of borrowing.

[16] We also considered inflation rates as an explanatory variable in the PMG and MG estimations. However, we ran into convergence problems when we added inflation rates to the general specifications.

Table 4a PMG and MG Estimation Results for Real House Prices

Dependent variable:	Model I		Model II		Model III	
Real House Prices	PMG	MG	PMG	MG	PMG	MG
Long-run coefficient						
Real personal income	0.417** (0.174)	0.131 (1.282)	0.414*** (0.157)	3.628 (3.099)	0.346** (0.162)	-2.378 (2.613)
Population growth	0.204*** (0.021)	0.092 (0.201)	0.166*** (0.015)	0.042 (0.130)	0.158*** (0.015)	0.619* (0.319)
Unemployment rate	-0.112*** (0.009)	-0.192*** (0.052)	-0.097*** (0.007)	0.064 (0.175)	-0.109*** (0.008)	-0.118 (0.073)
Net cost of borrowing	-0.012*** (0.004)	0.014 (0.058)	-0.007** (0.003)	-0.008 (0.035)	-0.010*** (0.003)	-0.121 (0.070)
Short-run coefficient						
Speed of adjustment	-0.031*** (0.002)	-0.059*** (0.010)	-0.043*** (0.003)	-0.066*** (0.010)	-0.043*** (0.003)	-0.068*** (0.011)
Change in real house prices	-0.067* (0.037)	-0.142*** (0.036)	-0.081** (0.037)	-0.151*** (0.035)	-0.096*** (0.036)	-0.162*** (0.035)
Change in real personal income	0.218*** (0.073)	0.208*** (0.078)	0.208*** (0.074)	0.199** (0.078)	0.209*** (0.074)	0.198** (0.077)
Change in population growth	-0.005* (0.002)	-0.006*** (0.002)	-0.006** (0.002)	-0.007*** (0.002)	-0.006** (0.002)	-0.007*** (0.002)
Change in unemployment rate	0.000 (0.002)	0.001 (0.003)	0.000 (0.002)	0.001 (0.003)	0.000 (0.002)	0.001 (0.003)
Change in net cost of borrowing	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
Deregulation			0.001 (0.001)	0.001 (0.001)	0.000 (0.001)	0.001 (0.001)
Dummy08q3					0.000 (0.001)	0.001 (0.002)
Intercept	0.000 (0.001)	-0.004 (0.003)	0.000 (0.002)	-0.006 (0.004)	0.000 (0.002)	-0.005 (0.005)
Statistics						
Hausman test (p-value)	0.19		0.77		0.56	
Number of observations	7344		7344		7344	

Note: The sample contains the 51 U.S. states over the period 1976Q3–2012Q4. The number of observations is 7,395 since two lags are used in the estimations. PMG represents pooled mean-group estimation and MG denotes mean-group estimation. Both real house prices and real personal income (per worker) are in logarithms. Robust standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Diagnostic tests confirm the validity of our model specification. Specifically, the standard error of regression varies from 0.01 in North Carolina to 0.07 in Hawaii, with an average of 0.02. The formal statistical tests of heteroscedasticity reject the null hypothesis of equality of error variances at the 5% level in 28 states. At the 5% level, there is no evidence of residual serial correlation in the equations for 33 states. Ramsey's RESET tests for functional form show no evidence of misspecification in 29 states.[17]

Previous empirical studies indicate that bank deregulation led to inter and intra-state mergers and acquisitions, as well as a general broadening of the geographic scope of banking operations, which enabled banks to diversify deposit collection across locations, and lower the cost of funding. Rice and Strahan (2010) construct a time-varying index of interstate branching deregulation that captured differences in banking regulatory constraints between 1994 and 2005. [18] Favara and Imbs (2015) use the Rice–Strahan deregulation index to evaluate the consequences of deregulation on mortgage credit and house prices and find that U.S. branching deregulation between 1994 and 2005 affected the supply of mortgage loans, and via the effect on credit, the increase in house prices. Therefore, in Model II, we include the index of interstate branching deregulation constructed by Rice and Strahan (2010). This index captures the impacts of the Riegle–Neal Interstate Banking and Branching Efficiency Act of 1994 and the Financial Services Modernization Act of 1999.[19] It is shown that bank deregulation plays a positively insignificant role in the adjustment of real house prices. [20] The findings for the other macroeconomic variables remain unchanged.

The meltdown of the U.S. housing market around 2006 triggered the subsequent global financial crisis. In Model III, we add a time dummy variable (Dummy08q3) to control for the impact on the U.S. housing markets during the recent 2007–2008 financial crisis.[21] The time dummy is insignificant because its impact on real house prices is partially captured by other macroeconomic fundamentals.

Since economic fundamentals could have varying impacts on real house prices in the long run versus the short run, we examine alternative model specifications in Table 4b, which allow for different economic fundamentals in the long-run equilibrium versus the short-run adjustment of real house prices. Specifically, Model IV includes real personal income per worker and population growth, Model V considers real personal income per worker and the unemployment rate, and Model VI incorporates real personal income per worker, population growth, and the net cost of borrowing in the long run. The lags of real house prices and all four economic fundamentals in Eq.10 are

[17] If the sample is larger than the 30, one can ignore the normality issue if it exists, per the central limit theorem.
[18] The index ranges from 0 (most restricted) to 4 (least restricted).
[19] We considered the Community Reinvestment Act of 1977 in the estimations. It turns out to be insignificant.
[20] Bank deregulation remains insignificant when sub-samples 1994Q1–2015Q4 or 1994Q1–2012Q4 are used.
[21] The recent financial crisis was triggered by the Lehman Brothers who filed for bankruptcy on September 15, 2008.

considered in the short run. Consistent with the results in Table 4a, the Hausman tests suggest that the PMG estimator is preferred in all model specifications. All long-run coefficients are statistically significant with the expected signs.

Table 4b PMG and MG Estimation Results for Real House Prices

Dependent variable: Real House Prices	Model IV		Model V		Model VI	
	PMG	MG	PMG	MG	PMG	MG
Long-run coefficient						
Real personal income	0.921*** (0.207)	1.458 (1.281)	1.155*** (0.145)	0.710 (0.474)	0.979*** (0.202)	1.073 (1.105)
Population growth	0.394*** (0.031)	-0.105 (0.391)			0.374*** (0.029)	0.242** (0.111)
Unemployment rate			-0.164*** (0.011)	-0.132*** (0.046)		
Net cost of borrowing					-0.011*** (0.004)	0.043 (0.046)
Short-run coefficient						
Speed of adjustment	-0.024*** (0.002)	-0.049*** (0.008)	-0.032*** (0.003)	-0.057*** (0.008)	-0.024*** (0.002)	-0.049*** (0.008)
Change in real house prices	-0.058 (0.039)	-0.071** (0.037)	-0.047 (0.037)	-0.061* (0.035)	-0.051 (0.040)	-0.127*** (0.037)
Change in real personal income	0.232*** (0.069)	0.208*** (0.070)	0.182*** (0.068)	0.168** (0.070)	0.236*** (0.069)	0.216*** (0.073)
Change in population growth	-0.005** (0.002)	-0.007*** (0.002)	0.005* (0.003)	0.005 (0.003)	-0.005** (0.002)	-0.007*** (0.002)
Change in unemployment rate	0.000 (0.001)	-0.001 (0.001)	0.001 (0.002)	0.002 (0.002)	0.000 (0.001)	-0.001 (0.001)
Change in net cost of borrowing	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.001)
Intercept	0.002 (0.001)	-0.001 (0.003)	-0.001 (0.001)	-0.005** (0.002)	0.002 (0.001)	0.000 (0.003)
Statistics						
Hausman test (p-value)	0.13		0.45		0.21	
Number of observations	7344		7344		7344	

Note: The sample contains the 51 U.S. states over the period 1976Q3–2012Q4. The number of observations is 7,395 since two lags are used in the estimations. PMG represents pooled mean-group estimation and MG denotes mean-group estimation. Both real house prices and real personal income (per worker) are in logarithms. Robust standard errors are in parentheses. Robust standard errors are in parentheses. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

4.3 Short-Run Adjustments of Real House Prices to the Long-Run Equilibrium

The lower panels of Tables 4a and 4b report the short-run coefficient estimates of the PMG and MG specifications. In the short run, the lags of real house prices are included to allow for momentum in real house prices, following Case and Shiller (1989) and most recently, Favara and Imbs (2015). The coefficient is negatively significant, which implies that higher real house prices in the previous periods could lead to a subsequent reverse of real house prices to the equilibrium.

The speed of the adjustment of real house prices to the long-run equilibrium is measured by the coefficient α_i in Eq.10. All model specifications of the PMG estimations show a significantly negative speed of adjustment, which ranges from –0.024 to –0.043. This finding indicates that real house prices adjust to the long-run equilibrium in response to a shock: following positive deviations from the long-run equilibrium in the real estate market, real house prices decrease, and vice versa. Following the literature, the half-life of the adjustment is approximated by $-\ln(2)/\ln(1+\alpha_i)$, which indicates that the time necessary for a deviation from the long-run equilibrium is halved. For the benchmark Model I, the coefficient α_i of –0.031 suggests that roughly 3% of the real house price deviations in the previous quarter from the equilibrium are adjusted this quarter, and the half-life estimate is around 22 quarters or 5.5 years, larger than the half-life of 3.5 years obtained by Holly et al. (2010), who used annual data for U.S. states (excluding Alaska and Hawaii) from 1975 to 2003.[22]

Figure 1 is a plot of the speed of adjustment coefficients for the 51 U.S. states from the PMG estimation of the benchmark Model I. The reference line indicates the average speed of adjustment (–0.031) over the sample. We observe large variations in the speed of adjustment across the 51 U.S. states, which range from –0.085 for South Dakota to –0.0003 for Tennessee. There are 23 states with faster speeds of adjustment than the average.

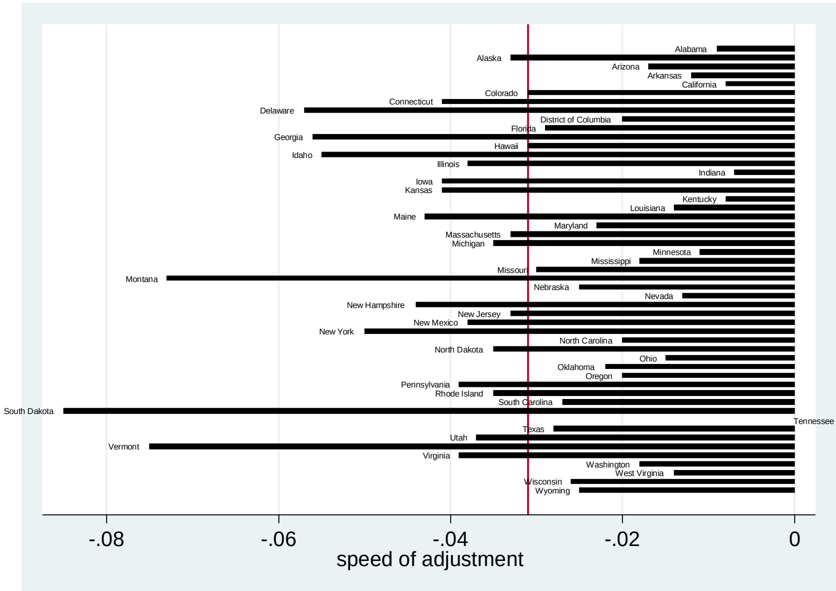
4.4 Real House Price Deviations from the Long-run Equilibrium

Based on the benchmark Model I, we investigate the magnitude of real house price deviations from their fundamental values in the U.S., calculated by the error correction term in Eq.10. Figure 2 presents the average deviation of real house prices from their fundamental values across the states over the sample period 1976Q3–2012Q4. We observe that real house prices positively deviated from their long-run equilibrium by more than 30% between 1982Q1 and 1983Q3, which coincided with the oil price shocks and economic recessions in the early 1980s. This suggests that the U.S. housing markets were overheated

[22] Koetter and Poghosyan (2010) obtain a half-life estimate of 6.79 years for the adjustment of house prices to the long-run equilibrium in Germany.

compared to the undesirable macroeconomic condition in the early1980s. Real house price deviations gradually reversed to the equilibrium after peaking in 1982Q4. A similar pattern occurred during the recent financial crisis. Real house price deviations substantially surged in 2007–2008 after the financial crisis triggered by the bankruptcy of the Lehman Brothers. The deviation peaked at 46% in 2010Q1 and then gradually declined.

Figure 1 Speed of Adjustment Coefficients from PMG Estimation



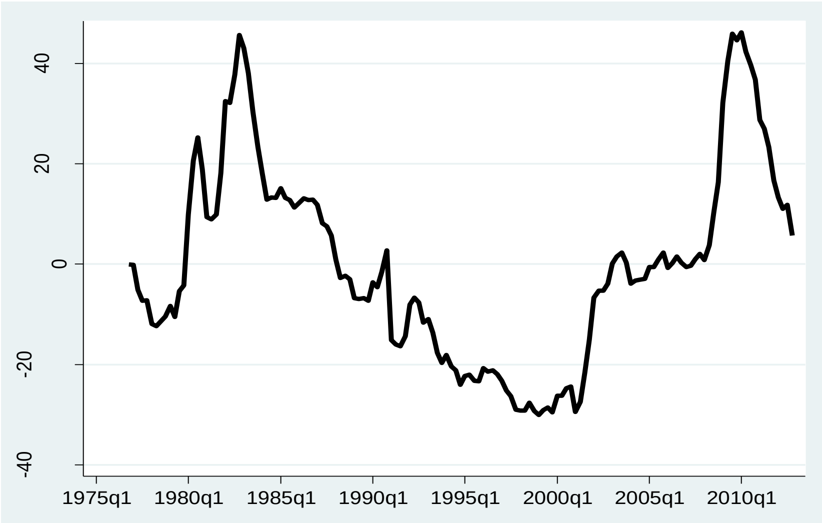
Note: Each bar shows the average of speed of adjustment coefficients over the sample period 1976Q3–2012Q4 in each state.

To further compare the deviations of real house prices from economic fundamentals before, during, and after the crisis, following the insight of Stepanyan et al. (2010), we examine three critical periods: pre-crisis (2005Q1), peak (2007Q1), and post-crisis (2010Q4).[23] Table 5 presents the means of the variables across the 51 states for the three periods. We observe larger positive real house price deviations (37%) during the post-crisis period than during the pre-crisis and peak periods. Real house prices continued to decline after the economy peaked (–1.43% of real house price changes). Real personal income per worker decreased by 0.16%, unemployment rates increased to unprecedented levels (8.58%), and the net cost of borrowing reached 4.98%, on

[23] The Lehman Brothers filed for bankruptcy on September 15, 2008, which marked the start of the recent global financial crisis. According to the business cycle reference dates from the NBER, the most recent recession started in December 2007 and ended in June 2009. We choose 2010Q4 as the post-crisis period. We have also tested other time periods after 2009Q2, with no significant difference in the conclusions.

average, across the 51 states during the post-crisis period. Thus, the economic fundamentals deteriorated even more rapidly than the decline in real house prices during the post-crisis period.

Figure 2 Average Real House Price Deviations from PMG Estimation



Note: The solid black line plots the average of real house price deviations across 51 U.S. states for each time period in the sample.

Table 5 Summary Statistics for the Pre-crisis, Peak, and Post-crisis Periods

51 States State variable	Pre-Crisis 2005Q1	Peak 2007Q1	Post Crisis 2010Q4
Percentage change of real house price index (%)	1.57	-0.43	-1.43
Percentage change of real personal income per worker (%)	-0.86	0.58	-0.16
Growth rate of population (%)	0.94	1.01	0.64
Unemployment rate (%)	4.99	4.22	8.58
Net cost of borrowing (%)	3.55	5.51	4.98
Inflation rate (%)**	2.58	4.55	3.61
House price deviations (%)	-0.61	-0.58	36.78

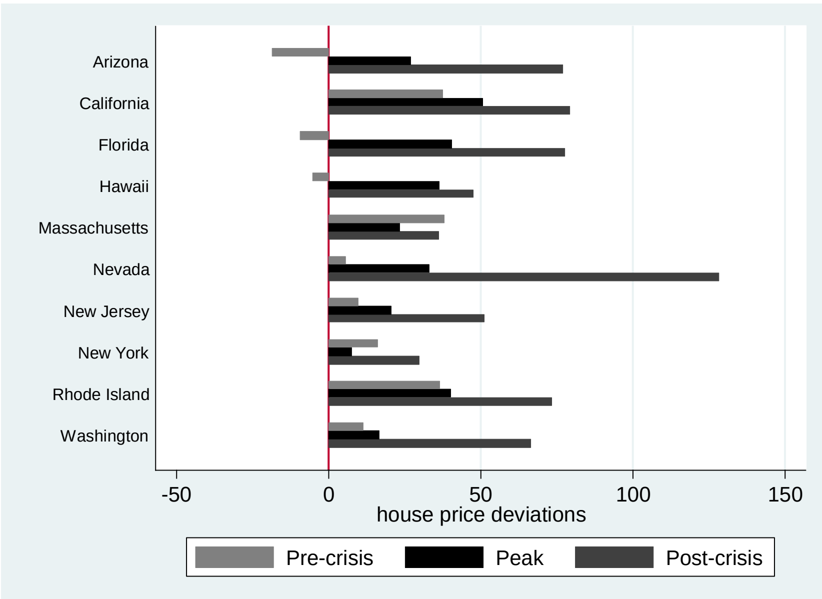
Note: The sample contains the 51 U.S. states.

Figure 3 illustrates real house price deviations in 10 selected states[24] (Arizona, California, Florida, Hawaii, Nevada, Massachusetts, Rhode Island, New Jersey,

[24] Due to limitations in space, we report on 10 selected states. Results from other states are available upon request.

New York, and Washington) during the three periods: pre-crisis (2005Q1), peak (2007Q1), and post-crisis (2010Q4). One common observation is that the real house price deviations during the post-crisis period are larger than those during the peak. The large positive real house price deviations in the post-crisis period imply that, despite the relatively low current levels, real house prices still have some room for downward adjustment. This finding is consistent with the results of Stepanyan et al. (2010), who analyze the house price determinants in 11 selected former Soviet Union countries.

Figure 3 Real House Price Deviations in 10 Selected States During the Pre-crisis, Peak, and Post-crisis Periods



Note: Each bar shows real house price deviations in a selected state for a specific period of time. Pre-crisis is 2005Q1; Peak represents 2007Q1; and Post-crisis is 2010Q4.

5. Conclusions

This paper applies the PMG and MG estimators to examine real house price determinants in the 51 U.S. states from 1976Q3 to 2012Q4. The empirical results show that real personal income per worker, population growth, unemployment rates, and the net cost of borrowing jointly contribute to real house price developments in the full sample. Our results confirm that the equilibrium real house prices increase with rising demand due to higher income and population growth, and lower unemployment rates and the net cost of borrowing, thus providing evidence of a house price adjustment to the long-run equilibrium. The short-run adjustment estimate indicates that about 3% on average of the real house price deviations in the previous period from the long-run equilibrium are adjusted during this period. We observe large variations in house price deviations from the long-run equilibrium and in the speed of short-run adjustment across the 51 U.S. states, which provide evidence of substantial heterogeneity of the U.S. housing markets.

Moreover, the deviations of real house prices from economic fundamentals in the post-crisis period are greater than those during the pre-crisis and peak periods, thus implying that economic fundamentals have deteriorated even more rapidly than real house prices in the post-crisis period—which could lead to a further decline in real house prices.

To summarize, our study provides new evidence of the relationship between real house prices and their economic determinants from the perspectives of both long-run equilibrium and short-run adjustment, and can shed some light on policy implications for the U.S. housing markets and the macro-economy. First, we confirm the existence of a common long-run relationship among real house prices, real personal income per worker, population growth, unemployment rates, and the net cost of borrowing across the 51 U.S. states. When housing markets are undesirable, policies that promote real personal income and employment or alleviate the net cost of borrowing should help long-run housing market recovery. Second, we show that the heterogeneity across the U.S. states in terms of deviations of real house prices from their fundamentals and the speed of adjustment should be taken into account when making policies. Third, we believe these results to be relevant for the theoretical debate between competing approaches to modeling house price dynamics. Finally, our results, which are based on the U.S. state-level data for the period of 1976Q3 to 2012Q4, suggest that a long time series is required to better capture the correction of real house prices after the recent housing market collapse in the U.S. Therefore, the use of a longer time series to fully capture the effects of the financial crisis on house price dynamics would be of great interest in future research.

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Appendix List of Variables and Their Descriptions

Variable	Source	Frequency*	Symbol
Nominal house price index	FHFA	Q	HP
Consumer price index	FRED	M	CPI
Real house price index	ln(HP/CPI)	Q	RHP
Personal income	BEA	Q	INC
Civil labor force	BLS	M	LF
Real personal income per worker	ln(INC/CPI/LF)	Q	RINC
Inflation rate (%) **	$\ln(CPI_t/CPI_{t-1})*100$	M	INF
Population	Census	A	POPUL
Population growth (%)	$\ln(POPUL_t/POPUL_{t-1})*100$	Q	POP
Unemployment rate (%)	BLS	M	UR
Nominal long-term mortgage rate (%)	FRED	Q	MORT
Net cost of borrowing (%)	MORT-ln(HP t /HP t-1)*100	Q	RMORT

Note: FHFA is Federal Housing Finance Agency; BEA is Bureau of Economic Analysis; BLS is Bureau of Labor Statistics; Census is U.S. Census Bureau; FRED is Federal Reserve Economic Data from Federal Reserve Bank of St. Louis. Labor force is from BLS in monthly frequency, 1976M1-2012M12. Both mortgage and inflation rates are at the census-region level. There are four census regions in the U.S. They consist of 50 states and the District of Columbia. Real house price index (1980=100). Personal income per worker is deflated by the consumer price index (base year is 1980).

* Frequency is for the raw data, M represents monthly; Q represents quarterly; and A represents annual. If the raw data is reported in monthly frequency, we convert monthly to quarterly frequency by using the average method in EViews. If the raw data is reported in yearly frequency, we convert the annual data to quarterly frequency by using constant match in EViews.

**We do Census X12 multiplicative seasonal adjustment for the consumer price index (CPI), then calculate annual rate of inflation based on the CPI.