House Prices and Economic Growth

Abstract

Using data for all 379 metropolitan statistic areas (MSAs) in the U.S. from the first quarter of 1983 to the fourth quarter of 2005, we empirically study the effects of house prices on Gross Metropolitan Product (GMP) in two econometric frameworks. We have three main findings. First, using panel regressions with a multifactor error structure, which control for unobserved macro economic factors and spatially correlated error terms, we find that a 10% increase in house prices is associated with a 0.5% increase in GMP in the same quarter. This effect survives after controlling for mortgage refinancing originations and new constructions. Second, using a panel VAR model, we find GMP Granger causes house prices but house prices do not Granger cause GMP. Third, after controlling for the Granger causality from GMP to house prices, we find that the component of house prices that is orthogonal to lagged GMP is still significant in explaining concurrent GMP. The results seem to suggest that house prices have significant but short lived effects on economic growth.

JEL classification: E23, E24, R11

Key words: housing market, common correlated effects estimators, CD test, Granger causality

House Prices and Economic Growth

I. Introduction

There seems to be a consensus among economists and policy makers that house prices have been playing an important role in fueling the growth of the economy. For instance, many believe that the strong housing market during the crash of the stock market in 2001 may have helped save the US economy from a more serious recession,[1] and recent cooling of the housing market generates concerns regarding a possible slow down of the economy.[2] Indeed, many economic theories indicate that house price changes should have real effects on the economy. For instance, Friedman's permanent income hypothesis suggests that people would change their consumption if house price changes affect their estimates of their permanent wealth. Recently, economists propose more mechanisms through which house price changes may affect the economy. For example, Aoki, Proudman and Vlieghe (2004), Lustig and Nieuwerburg (2004), and Ortalo-Magné and Rady (2004) argue that house price increases may help relax borrowing constraints and thus increase consumption (the collateral effect).[3] Furthermore, Li and Yao (2005) argue that the effect of house price changes may not be uniform across individual households. These theories shall apply to both closed and open economies, including regional economies.

Empirical evidence of the consumption effect of house prices is generally consistent with theories. Early evidence includes the positive effect of house price

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[1] On September 26th, 2005 Alan Greenspan in a speech to the American Bankers Association re-iterated the important role of house prices in the current economy and suggested that housing may have fueled consumption over the past few years.

[2] On December 12th 2006, Federal Reserve Open Market Committee stated that "Economic growth has slowed over the course of the year, partly reflecting a substantial cooling of the housing market."

[3] On the other hand, it is worth noting that a theory by Bajari, Benkardb, and Krainer (2005) suggests no aggregate change in welfare due to price increases in the existing housing stock.

increases on consumption found by Bhatia (1987). Engelhardt (1994), Engelhardt (1996), and Sheiner (1995), among others, find that house price changes may affect young renters' as well as homeowners' saving behavior. Recently, a fast growing literature shows that the wealth effect of housing is not only statistically significant, but also probably larger than the wealth effect from the stock market (see Benjamin, Chinloy and Jud (2004), Case, Quigley and Shiller (2005), Kishor (2006), and Lettau and Ludvigson (2003), Bostic, Gabriel and Painter (2006), Carroll, Otsuka and Slacalek (2006), Slacalek (2006), among others). Furthermore, Campbell and Cocco (2005) show that the aggregate effect on consumption varies across different age groups and there is evidence supporting both the wealth effect and the collateral effect. Up to now, the only evidence we know of that shows no aggregate effect of house price changes on the economy is Phang (2004).

While the literature on the wealth effect of house prices provides important insights into the economic implications of housing markets, the relation between house price and economic growth (not just consumption) is rarely analyzed despite its important policy implications. The wealth effect of house prices may only partially contribute to the impact of house prices on economic growth, and there are likely other mechanisms through which house prices directly affect economic production. First, increasing house prices may indicate excess demand or strengthening of the housing market, which in turn may lead to more construction and more residential investment. Moreover, increasing house prices are likely associated with increasing land prices, which may affect corporations' business decisions as well as investment decisions (e.g., see Chaney, Sraer and Thesmar (2007) for the corporate wealth effect). Both changes in residential and commercial investment/constructions may directly affect economic production. Second, increasing house prices are often associated with increasing trading volume, which would be associated with more service provided by real estate agencies and mortgage lenders.[4] Third, decreasing house prices may increase the default probability for mortgages, which in turn may disturb the financial market and thus negatively affect economic growth. Since house price changes seem to be able to affect the economy not only through the wealth effect, it is important to analyze the direct effects of house prices on economic growth, which is what this paper focuses on.

Using a unusually rich data set that covers *all* 379 (as defined in 2006) Metropolitan Statistic Areas (MSAs) in the U.S. from the first quarter of 1983 to the fourth quarter of 2005, we first study the effects of house price changes on Gross Metropolitan Product (GMP) growth using fixed effect contemporaneous panel data regressions with a multifactor error structure. The multifactor error structure, which is proposed by Pesaran (2006), controls for unknown variables that might correlate with dependent variables, and the corresponding estimators are called Common Correlated Effects (CCE) estimators. It is important to control for unknown variables for omitting them would bias the estimate results (see, e.g. Poterba (2000) in the context of stock market wealth effect which seems also applicable on housing effects and Campbell and Cocco (2005) for more discussions). Controlling for unknown variables distinguishes this part of our empirical analysis from most empirical work in the literature.

[4] It is worth noting that trading volume in the housing market may affect economic growth via other mechanisms as well. For example, changes in trading volume may affect the expectation regarding the trend of house prices, and thus affect the behavior of home buyers, sellers, and developers. We will leave this topic for future research.

The results suggest that MSAs that have more rapid house-price growth tend to have more rapid economic growth, after controlling for other explanatory variables for economic growth. Specifically, a 10% increase in house prices would be associated with a 0.5% increase in GMP (based on the CCE estimators). The results are statistically significant and seem robust to regression specifications. The results also appear to indicate that house prices affect the economic growth not only through the wealth effect or new constructions. After controlling for the amount of mortgage refinancing, which might help capture consumption, and single family permits, which help capture new constructions, house price changes still significantly correlate with GMP growth.

Second, we analyze the dynamic interactions between house price changes and GMP growth using panel Vector Autoregressions (VAR). We test the Granger Causality between GMP growth and house price changes in both directions. The Granger Causality helps us capture "causality" between these two variables in the lead-lag sense. We find that house prices do not Granger cause GMP, while GMP Granger causes house prices.

Finally, we analyze if the contemporaneous relation between house prices and GMP survives after we control the Granger causality from GMP to house prices. We estimate the component of house prices that is orthogonal to lagged GMP using the residuals in regressions of house prices on lagged GMP. This orthogonal component of house prices is statistically significant and positive in explaining GMP in all specifications of the contemporaneous regressions. Therefore, the explanatory power of house prices in the contemporaneous setting does not seem to be caused by the Granger causality from GMP to house prices. Overall, the results seem to indicate that house prices have significant but short lived effects on economic growth. To our knowledge, this paper seems to be one of if not the first to measure the statistical and economic significance of the direct effects of house prices on economic growth.

This rest of this paper is organized as follows. Section 2 describes the data. Section 3 discusses the empirical analysis. Section 4 concludes.

II. Data

The main part of our analysis utilizes a large panel data set that comprises five MSA level quarterly variables from 1983:1 to 2005:4 for each of the 379 MSAs in the U.S. (as defined in 2006).[5] The five MSA level variables are per capita GMP, the single-family house price index, average household income, population, and the unemployment rate. While our analysis focuses on the relation between per capita GMP and house prices, we include average household income, population, and the unemployment rate as important MSA level control variables. Theories, such as Ortalo-Magné and Rady (2004), suggest that income shocks might affect not only the economy but also the housing market. Changes in population help capture migration, which can affect the per capita output in a MSA, for migration may affect not only the population level but also the productivity of the labor force of the MSA. Migration may also affect the dynamics of house prices (see, e.g. Gabriel, Mattey and Wascher (1999) for direct evidence). Changes in the unemployment rate may capture changes in the magnitude of frictions in the labor market or transitions of the economy, which relates to relocation of labor force and affects both the economy and the housing market.

[5] The panel data set is unbalanced. For 212 MSAs, the OFHEO house price index begins in a time period that is later than 1983:1. However, the problem of missing data does not seem to be severe: 314 MSAs have their OFHEO house price indices starting before 1988:1. Figure 1 plots the histogram of MSAs with different starting points of the OFHEO house price indices.

In additional to MSA level variables, the data set also includes two national level time series in the same sample period, which are the national average 30-year fixed-rate mortgage rate, and the SP500 index. Moreover, in our analysis regarding whether house prices affect GMP after controlling for the wealth effect and new constructions, we utilize two more MSA level variables: mortgage refinancing originations and single family house permits. The sample period is 1987:1 to 2005:4 for mortgage refinancing origination (with missing values for the last two quarters in 1987), and 1983:1 to 2005:4 for single family house permits.

Our MSA level data have important advantages. In additional to widely understood benefits of panel data (e.g. allowing for the control of individual heterogeneity and having more power due to a large number of observations), the large cross section in our panel of MSAs allows us to benefit from the novel approach by Pesaran (2006), so that we are able to control for unobserved macro economic variables. Moreover, all MSAs are in the United Sates, and thus are homogenous in the sense that they are subject to the similar if not identical monetary policy, political environment, legal context, tax codes, and financial market conditions. As a result, MSA-level data seem superior to international data such as those used by Case, Quigley and Shiller (2001), because the economic effects of house prices may differ with a different economic, legal and tax environment, as pointed by Slacalek (2006), and such an environment is difficult to control in international data. Finally, the sample period in this paper covers both economic expansions and recessions so it does not seem to be biased.

However, our analysis at the MSA level also has caveats, which are mainly regarding how to interpret the results. It would be reasonable to assume that house price changes in a MSA would affect not only local economy but also other MSAs, for MSA economies are integrated. As a result, there would be cross sectional dependence in idiosyncratic errors, which, fortunately, is accommodated by the approach of Pesaran (2006). Nonetheless, readers should be cautious that our results should be interpreted as the effect of house prices on the growth of an *open* economy.

The data sources are as follows. Moody's economy.com provides quarterly GMP series for MSAs. The estimation of the GMP for a MSA in each quarter comprise of two steps. The first step estimates the "weighted" productivity for each NAICS Supersector industry (e.g., Manufacturing, Education & Health Services, etc) in the MSA by multiplying the U.S. level productivity for this industry with the ratio of industry employment to total employment in the MSA (Moody's economy.com seems to obtain industry employment data from BLS). The second step estimates the GMP with the product of the sum of "weighted" productivity with the total MSA employment. This two step estimation process is essentially equivalent to summing up the estimated products of all NAICS Supersector industries in the MSA.

The estimated GMP might contain measurement errors. The problem is to what extent this would bias our analysis. A main econometric insight regarding measurement errors is that measurement errors in explanatory variables would cause the attenuation bias (the coefficient estimators are biased toward 0), while measurement errors in dependant variables, which is our case, are not a problem unless the measurement errors correlate with explanatory variables. It seems plausible that, *holding constant everything else*, MSAs with higher productivity may have higher average household income, which in turn may be associated with higher house prices. Therefore, the measurement error, which equals estimated GMP minus true GMP, may be negatively correlated house prices. This will bias the estimator of house price coefficients downward. In our contemporaneous panel regressions, the coefficient estimators are all significantly positive; therefore, the *true* coefficients would be even more significant. In this sense, the results can be even stronger than they appear.

The sources for other quarterly variables are as follows. Moody's economy.com estimates quarterly population series using census data and migration flows among MSAs (data sources are Census Bureau and IRS respectively). It also compiles unemployment rates using BLS data, estimates average household income using BEA data, compiles total single family house permits using BOC data, and estimates mortgage refinancing originations. The Office of Federal Housing Enterprise Oversight (OFHEO) provides transaction-based quarterly home price indices.[6] The HSH associates provide the average mortgage rates, and the SP500 index is from CRSP.

We process the data by first converting nominal terms into real terms, and then calculating the first order differences of log values of the variables. We use CPI to adjust for inflation and obtain real terms for per capita GMP, the house price index, average household income, mortgage refinancing origination, and the SP500 index. We choose to work on log differences instead of the original variables or their logs (level) because all OFHEO house price indices are set to be 100 in 1995:1, and thus house price levels are not comparable across MSAs.

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[6] OFEHO house price indices are estimated with the repeat sale regression. Such indices control for time invariant attributes of houses that enter into the sample at least twice, and thus appear to be superior to median or mean sale prices. On the other hand, repeat sale regressions have some shortcomings and this is an active research area. In our analysis, we appreciate the large cross section and the long sample period of the OFEHO indices despite possible weakness of them.

Figures 2 to 4 illustrate the temporal behavior of the variables. Figure 2 plots the time series of OFHEO house price indices in quarterly frequency. Figure 3 plots the time series of the SP500 index (normalized to be one in 1983:1) and the average 30-year fixed-rate mortgage rate (percentage points). Figure 4 plots the 25%, 50%, and 75% percentiles of across MSA distributions of per capita GMP (in thousand dollars), average household income (in dollars), population (thousand people), unemployment rate (percentage points), mortgage refinancing originations (million dollars), and single family house permits over time.

Since our analysis uses log differences instead of levels, we report some important statistics on the log differences of the seven variables in table 1.[7] It is interesting to notice that the correlation between house prices and GMP is significantly positive, which is consistent with a positive effect of house prices on economic growth. However, both GMP and house prices significantly relate to many of other variables, such as household income, population, and single family house permits, which may indicate the existence of simultaneity. Furthermore, GMP significantly relates to all proposed control variables, which justifies the importance of the control variables in measuring the impact of house price changes on GMP growth rates.

III. Empirical Analysis

III.1 Cross-section Dependence Tests

We first analyze how house price changes may affect the economic growth of MSAs use the following multifactor error structure panel model:

[7] When we calculate statistics for MSA level variables, or correlations between a MSA level variable and a national level variable, we also provide t-statistics from testing null hypotheses that means of the across-MSA distributions of the statistics are zero.

$$gmp_{i,t} = \alpha_i + \beta hp_{i,t} + \rho' x_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T, \quad (1)$$

where α_i is a MSA specific intercept term that captures all time-constant economic variables that might differ across MSAs, $gmp_{i,t}$ and $hp_{i,t}$ are respectively the log difference of GMP and the log difference of the house price index from quarter t to $t+1$ for MSA i, $x_{i,t}$ is a vector of MSA level variables that may help determine $gmp_{i,t}$, including log differences of population, average household income, and the unemployment rate; and $u_{i,t}$ is the error term. We assume that the error term captures the common unobserved factors as well as possible spatial effects:

$$u_{i,t} = \gamma_i' f_t + \varepsilon_{i,t}, \quad (2)$$

Where f_t is a vector of unobserved common factors, which includes macro economic variables that are the same across all MSAs in a given time period but vary across time, such as interest rates, performance of the stock market, etc. $\varepsilon_{i,t}$ is a idiosyncratic error, which is assumed to be distributed independently of $x_{i,t}$ and f_t and across MSAs.

The model in equations (1) and (2) is reasonably general and flexible. First, the MSA specific intercept term (MSA fixed effect) captures unobserved heterogeneity in $gmp_{i,t}$ that remains constant across time. Second, $x_{i,t}$ controls for local variables that change both across MSAs and time periods that may affect $gmp_{i,t}$. Third, the model controls for macro economic variables, despite their being possibly unobserved, that affect the economic growth in all MSAs. Finally, the model is flexible enough to allow each MSA to respond to the common factors differently. As a result, factors that affect some but not all MSAs are also controlled since the MSAs that are not affected by these factors simply have zero coefficients.

To show the importance of including the multifactor error structure in the model, we conduct the CD (Cross-section Dependence) test of Pesaran (2004), which tests the existence of cross section dependence of error terms. Denote by T_i the set of dates over which time series observations on $gmp_{i,t}$ and $x_{i,t}$ are available, and by $\#T_i$ the number of the elements in the set. We first compute the OLS residuals based on full set of time series of observations for each MSA, using three variations of equation (1) respectively. The OLS regressions are run for each MSA separately, so they are not seemingly unrelated regressions. For each variation/specification, we calculate the CD statistic as follows. Denote the residuals by $e_{i,t}$ for $t \in T_i$, we compute the pair-wise correlations of $e_{i,t}$ and $e_{j,t}$ using the common set of observations in $T_i \cap T_j$, and have

$$\hat{\rho}_{i,t} = \frac{\sum_{t \in T_i \cap T_j} \left(e_{i,t} - \overline{e}_i\right)\left(e_{j,t} - \overline{e}_j\right)}{\left[\sum_{t \in T_i \cap T_j} \left(e_{i,t} - \overline{e}_i\right)^2\right]^{1/2} \left[\sum_{t \in T_i \cap T_j} \left(e_{j,t} - \overline{e}_j\right)^2\right]^{1/2}}, \quad (3)$$

where

$$\overline{e}_i = \frac{\sum_{t \in T_i \cap T_j} e_{i,t}}{\#\left(T_i \cap T_j\right)}. \quad (4)$$

The CD statistic is given by

$$CD = \sqrt{\frac{2}{N\left(N-1\right)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \sqrt{\#\left(T_i \cap T_j\right)} \hat{\rho}_{i,j} \right). \quad (5)$$

Under the null hypothesis that $\gamma_i' = 0$, therefore $\text{cov}\left(u_{i,t}, u_{j,t}\right) = 0$, $CD \sim N\left(0,1\right)$.

The three specifications we use to conduct the CD tests are:

$$gmp_{i,t} = \alpha_i + \beta_i hp_{i,t} + u_{i,t}\ , \qquad (6)$$

$$gmp_{i,t} = \alpha_i + \beta_i^1 hp_{i,t} + \beta_i^2 po_{i,t} + \beta_i^3 hi_{i,t} + \beta_i^4 ur_{i,t} + u_{i,t}\ , \qquad (7)$$

and

$$gmp_{i,t} = \alpha_i + \beta_i^1 hp_{i,t} + \beta_i^2 po_{i,t} + \beta_i^3 hi_{i,t} + \beta_i^4 ur_{i,t} + \beta_i^5 mr_t + \beta_i^6 sp_t + u_{i,t}\ . \qquad (8)$$

In above equations, $po_{i,t}$, $hi_{i,t}$, $ur_{i,t}$, mr_t and sp_t are log differences of population, average household income, the unemployment rate, the 30-year fixed rate mortgage rate, and the SP500 index respectively. Apparently, (6) does not have any control variables; (7) includes MSA level control variables; and (8) includes both MSA level variables and two national level variables. Table 2 reports the CD statistics for the three specifications, which are 673.80, 387.22, and 382.22 respectively. This provides strong evidence of the existence of cross section dependence of error terms. Furthermore, the tests show that the 30-year mortgage rate and the S&P500 index are not sufficient in capturing all factors that affect MSAs. As a result, it is important to use the multifactor error structure of equations (1) and (2) to control for unobserved factors that affect all MSAs.

It is worth noting that equations (6) to (8) allow different MSA have different parameters. This is governed by the procedure of the CD test, not because we try to analyze the heterogeneity among MSAs. While there is likely very interesting and important heterogeneity among MSAs, this paper focuses on the relation between house price changes and economic growth in a *typical* U.S. MSA. Therefore, in our analysis except the CD test, we always let all MSAs have the same parameters. The analysis of the heterogeneity will be future research.

III.2 Contemporaneous Panel Regressions

After the test, we estimate four different specifications of the model of equations (1) and (2), which helps us check the robustness of the results. The first specification is

$$gmp_{i,t} = \alpha + \beta hp_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T \quad (9)$$

which does not include the fixed effect (MSA specific intercept terms) or any local control variables. The second specification is

$$gmp_{i,t} = \alpha + \beta hp_{i,t} + \rho' x_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T \quad (10)$$

which includes MSA level control variables but does not include the fixed effect. The third specification is

$$gmp_{i,t} = \alpha_i + \beta hp_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T \quad (11)$$

which includes the fixed effect but not MSA level control variables. The fourth specification is

$$gmp_{i,t} = \alpha_i + \beta hp_{i,t} + \rho' x_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T \quad (12)$$

which includes both the fixed effect and MSA level control variables.

We provide two different estimators for each of the four specifications. One is the simple OLS estimator, and the other is the Common Correlated Effects estimators (CCE estimators) proposed by Pesaran (2006), which controls for unobserved macro economic factors. This estimator is constructed using regressions augmented with cross-sectional averages of all dependent and independent variables. Pesaran (2006) proves that the CCE estimators can use cross-sectional averages of all dependant and independent variables to capture the unobserved variables, for the unobserved variables converge to a linear combination of the cross-sectional averages. To illustrate, following Pesaran (2006), consider a simple but generic model

$$y_{i,t} = \alpha_i + \beta_i' x_{i,t} + u_{i,t}, i = 1,2,\dots,N; t = 1,2,\dots,T \ , \tag{13}$$

where $u_{i,t} = \Upsilon_i' f_t + \varepsilon_{i,t}$ and f_t is a vector of unobserved variables that help determine $y_{i,t}$. Suppose the $k \times 1$ vector $x_{i,t}$ in (1) is generated as

$$x_{i,t} = a_i + \Gamma_i' f_t + v_{i,t} \, , \tag{14}$$

where a_i is a vector of individual effects, Γ_i' is a factor loading matrix with fixed components, $v_{i,t}$ are the specific components of $x_{i,t}$ distributed independently of the common effects and across i, but assumed to follow general covariance stationary processes. Pesaran (2006) shows that

$$f_t - \left(CC'\right) C \left(\bar{z}_t - \bar{d}\right) \xrightarrow{p} 0, \text{ as } N \to \infty \, , \tag{15}$$

where $z_t = \begin{pmatrix} y_{i,t} \\ x_{i,t} \end{pmatrix}$, $d_t = \begin{pmatrix} 1 & \beta_i' \\ 0 & I_k \end{pmatrix} \begin{pmatrix} \alpha_i \\ a_i \end{pmatrix}$, and $C_i = \begin{pmatrix} \Upsilon_i & \Gamma_i \end{pmatrix} \begin{pmatrix} 1 & 0 \\ \beta_i & I_k \end{pmatrix}$. Equation (15) indicates that the cross-sectional averages of all dependent and independent variables span the same space spanned by the unobserved variables, which is the basic insight that lies behind the CCE estimators.

Table 3 reports the estimation results.[8] First, the coefficient of house prices is significant at 1% level across all four specifications and for both the OLS and CCE estimators. Second, CCE estimators are significantly lower than OLS estimators. For example, in the fourth specification, which controls for both the fixed effect and macro factors, the CCE estimator is 0.048, while the OLS estimator is 0.121. This seems to indicate that the OLS estimators might be upward biased, possibly due to omitted variables that correlate with both GMP and house prices. Economically, the results in

[8] It is worth noting that R2 is a spurious statistic for two stage regressions though table 3 reports them.

table 3 seem to suggest that house price changes strongly correlate with GMP growth. A 10% increase in the house price would be associated with about 0.5% (CCE estimator in specification 4) to 1.2% (OLS estimator in specification 4) of GMP growth in the same quarter.

To investigate whether house prices affect GMP through mechanisms other than the wealth effect and changes in new constructions, we reproduce regressions in table 3 with more control variables (in log differences) added. We hope to include consumption to control for the wealth effect, but consumption data are not available at the MSA level. A proxy for consumption that we can find is mortgage refinancing originations. If the collateral effect theories are correct, mortgage refinancing originations would correlate with consumption. To capture possible lag between refinancing and consumption, the regressions include current and one quarter lag of mortgage refinancing originations. Further lags are not significant in the regressions. We use quarterly single family home permits issued in each MSA to help capture new constructions for it is plausible that these two variables are highly correlated. In fact, based on the Bureau of Census, the lag between permit issuance and completion of new houses is about six months. To accommodate the variation of the lag, we use the current and three quarters of lags of permits to help capture new constructions.

Table 4 reports the regression results for the fourth specification in table 3, with the new control variables added. Other specifications provide similar results, which we do not report. Table 4 shows that, first, single family house permits (current and lags) are all significant at 1% level. This indicates that new constructions significantly affect GMP. Second, mortgage refinancing origination has a negative sign and is not always significant. Among other explanations such as measurement errors, mortgage refinancing originations may not positively correlate with consumption and thus might be a poor proxy for consumption. Finally, table 4 shows that, even after adding more control variables, coefficients for house prices are still statistically significant at 1% level. In addition, the magnitude of the coefficients in table 4 and table 3 is similar. This seems to indicate that house prices have direct and strong effects on GMP after controlling the effects of new constructions.

To check if the results in table 3 are robust across time, we split the sample period into roughly two equal sized time periods before 1994 period and after 1994 period and reproduce the regressions in table 5. Both the OLS and the CCE estimators are statistically significant at 1% level in both sub periods, while there seem to be some variations in magnitude. While this paper does not focus on the time variation of the economic effects of house prices or the causes of such variation, it is worth noting that the result in table 5 appears to be consistent with some recent research that documents structural changes in the mortgage and REIT market. For example, Bennett, Peach and Peristiani (1998) estimate a hazard model of loan survival for 1980s and 1990s, and find that the homeowners' propensity to refinance is higher in 1990s, likely because the mortgage market becomes more competitive and efficient in 90s. In the REIT market, Clayton and Mackinnon (2003) find significant changes in the determination of the REIT returns from 1980s to 1990s. All these related structural changes might help explain the changing economic effects of house prices, and we leave this interesting topic for future research.

III.3 Panel VAR Analysis

Another conventional framework to analyze causality between economic variables is Vector Autoregressions (VAR). In this framework, economists use Granger Causality to capture the "causality" relation, though it is well known that Granger Causality is essentially a lead-lag relation. Our analysis is based on the following bivariate VAR model.

$$\begin{aligned} gmp_{i,t} &= \alpha_i^{gmp} + \alpha_t^{gmp} + \sum_{s=1}^{4} \gamma_s gmp_{i,t-s} + \sum_{s=1}^{4} \beta_s hp_{i,t-s} + \beta^2 po_{i,t} + \beta^3 hi_{i,t} + \beta^4 ur_{i,t} + u_{i,t} \\ hp_{i,t} &= \alpha_i^{hp} + \alpha_t^{hp} + \sum_{s=1}^{4} \lambda_s gmp_{i,t-s} + \sum_{s=1}^{4} \rho_s hp_{i,t-s} + \rho^2 po_{i,t} + \rho^3 hi_{i,t} + \rho^4 ur_{i,t} + v_{i,t} \end{aligned} \quad (16)$$

Note that symbols in this equation are the same with the symbols in equations (9) to (12). In this VAR model, $gmp_{i,t}$ and $hp_{i,t}$ are endogenous variables. The lag order is chosen by running a preliminary VAR for each MSA separately, and about 95% of the MSAs have optimal (based on AIC) lag order that is equal to or shorter than 4[9]. This regression includes two way fixed effects: MSA specific intercept terms and time specific intercept terms. We use them to control for two kinds of unobserved heterogeneity – the effects of variables that differ across MSAs but remain constant across time and variables that vary across time but affect all MSAs in each time period. We are not using the multifactor error structure setting here, for theories have not yet validated this method in dynamic settings.

To obtain consistent estimators, we first subtract cross sectional averages from each variable to eliminate the time dummy, and then we take first order difference on both sides of each equation in (16). After this differencing, MSA dummies all disappear,

[9] Granger Causality tests have similar results with different lag orders such as 3 and 5.

and variables are differences of differences. It is well known that the differenced model can not be directly estimated, for the correlation between independent variables and the error term is not zero (see, e.g. Nickell (1981)). To overcome this problem, we use the well established instrumental variable approach of Holtz-Eakin, Newey and Rosen (1988), for this approach is appropriate for panels with large N and short T, which is our case, and this approach is valid even if there are unit roots and nonstationary variables (as a result of large N). Therefore, we do not conduct any unit root or cointegration tests. Since there are two endogenous variables in the equation, we use $t-2$ to $t-6$ lagged endogenous variables as instrumental variables. According to Holtz-Eakin, Newey and Rosen (1988), the coefficients of lagged endogenous variables are identified, for we have eight such coefficients for endogenous variables but ten instrumental variables.

Table 6 reports the regression results for the first equation in (16). None of the four lagged of house price changes is significant, while lagged GMP growth rates and the MSA level control variables are significant. This seems to suggest that lagged house price changes do not affect economic growth. To formally test this notion, we construct a F test (Granger Causality test) with the null hypothesis being all coefficients of the four lagged house price changes are equal to zero. The test statistic and the corresponding p value are also reported in Table 6, which does not reject the null hypothesis.

Table 7 reports the regression results for the second equation in (16). It is interesting to notice that all lagged house price changes and GMP growth are significant. Moreover, the coefficients for one and two quarter lags of GMP growth are negative, which appears to indicate that the effects of GMP growth on house prices are not necessarily positive in the short run. Another result is that the Granger Causality test significantly rejects the null hypothesis that lagged GMP does not affect house price changes.

III.4 Panel Regressions with Granger Causality Controlled

The significant coefficients of house price changes in the contemporaneous regressions reported earlier sharply contrast with the insignificant coefficients in the dynamic setting. It is important to investigate if the explanatory power of house prices is merely caused by the Granger causality from GMP to house prices. We use the following two step approach to shed light on this question. First, we estimate the component of house prices that is orthogonal to lagged GMP, using residuals from a regression of house price changes on lagged GMP growth in the past four quarters. We run this regression separately for each MSA, to accommodate possible heterogeneity. Second, we run the panel regressions in table 3 using the orthogonal component of house prices. If the explanatory power of house prices is not caused by the Granger causality from GMP to house prices, we should expect that the orthogonal component of house prices is significant.

Table 8 reports the results, which indicate that the orthogonal component of house prices is still statistically significant at 1% level in all specifications and for both the OLS and the CCE estimators. Moreover, the magnitude of the orthogonal component of house prices is similar with the magnitude in table 3. In short, table 8 provides strong evidence that the explanatory power of house prices is not caused by the Granger causality from GMP to house prices.

IV. Conclusions

Applying recent advances in panel econometrics on a rare data set that covers all 379 MSAs in the U.S. from 1983:1 to 2005:4, this paper controls for unobserved macro economic factors and provides unbiased estimation of the effects of local house price changes on MSA economic growth. Based on the CCE estimators, we find that a 10% increase in house prices is associated with about 0.5% increase in GMP in the same period. This effect is statistically significant at 1% level for different specifications of the model, and the effect survives after controlling for mortgage refinancing originations and new constructions.

Moreover, we find house price changes do not Granger cause GMP growth, but GMP growth Granger causes house prices changes, in a two way fixed effect panel VAR model with GMP and house prices being endogenous variables. We then control for the Granger causality from GMP to house prices, and find that the component of house prices that is orthogonal to lagged GMP is still significant in explaining concurrent GMP. This appears to indicate that the correlation between GMP and house prices in contemporaneous regressions is not caused by the Granger causality from GMP to house prices.

References:

Aoki, Kosuke, James Proudman, and Gertjan Vlieghe, 2004, House Prices, Consumption, and Monetary Policy: a Financial Accelerator Approach, *Journal of Financial Intermediation* 13, 414-435.

Benjamin, John D., Peter Chinloy, and G. Donald Jud, 2004, Real Estate Versus Financial Wealth in Consumption, *Journal of Real Estate Finance and Economics* 29, 341-354.

Bennett, Paul, Richard W. Peach, and Stavros Peristiani, 1998, Structural Change in the Mortgage Market and the Propensity to Refinance, *FRB of New York Staff Report* 45.

Bhatia, Kul B., 1987, Real Estate Assets and Consumer Spending, *Quarterly Journal of Economics* 102, 437-444.

Bostic, Raphael, Stuart Gabriel, and Gary Painter, 2006, Housing Wealth, Financial Wealth, and Consumption: New Evidence from Micro Data, *University of Southern California Working Paper*.

Campbell, John Y., and Joao F. Cocco, 2005, How Do House Prices Affect Consumption? Evidence from Micro Data, *Journal of Monetary Economics* Forthcoming.

Carroll, Christopher D., Misuzu Otsuka, and Jirka Slacalek, 2006, How Large is the Housing Wealth Effect? A New Approach, *NBER Working Paper No. W12746*.

Case, Karl E., John M. Quigley, and Robert J. Shiller, 2001, Comparing Wealth Effects: The Stock Market versus the Housing Market, *Cowles Foundation Discussion Paper No. 1335*.

Case, Karl E., John M. Quigley, and Robert J. Shiller, 2005, Comparing Wealth Effects: The Stock Market versus the Housing Market, *The B.E. Journal of Macroeconomics* 5.

Chaney, Thomas, David Sraer, and David Thesmar, 2007, The Corporate Wealth Effect: From Real Estate Shocks to Corporate Investment, *Available at SSRN: http://ssrn.com/abstract=965762*.

Clayton, Jim, and Greg Mackinnon, 2003, The Relative Importance of Stock, Bond, and Real Estate Factors in Explaining REIT Returns, *Journal of Real Estate Finance and Economics* 27, 39-60.

Engelhardt, Gary V., 1994, House Prices and the Decision to Save for Down Payments, *Journal of Urban Economics* 36, 209-237.

Engelhardt, Gary V., 1996, House Prices and Home Owner Saving Behavior, *Regional Science and Urban Economics* 26, 313-336.

Gabriel, Stuart A., Joe P. Mattey, and William L. Wascher, 1999, House Price Differentials and Dynamics: Evidence from the Los Angeles and San Francisco Metropolitan Areas, *Economic Review* Federal Reserve Bank of San Francisco.

Holtz-Eakin, Douglas, Whitney Newey, and Harvey S. Rosen, 1988, Estimating Vector Autoregressions with Panel Data, *Econometrica* 56, 1371-1395.

Kishor, N. Kundan, 2006, Does Consumption Respond More to Housing Wealth than to Financial Market Wealth? If So, Why?, *Journal of Real Estate Finance and Economics* Forthcoming.

Lettau, Martin, and Sydney Ludvigson, 2003, Understanding Trend and Cycle in Asset Values: Reevaluating the Wealth Effect on Consumption, *American Economic Review* forthcoming.

Li, Wenli, and Rui Yao, 2005, The Life-Cycle Effects of Hous Price Changes, *Journal of Money, Credit and Banking* Forthcoming.

Lustig, Hanno, and Stijn Van Nieuwerburg, 2004, Housing Collateral and Consumption Insurance Across US Regions, *University of Chicago Working Paper*.

Nickell, Stephen, 1981, Biases in Dynamic Models with Fixed Effects, *Econometrica* 49, 1417-1426.

Ortalo-Magné, François, and Sven Rady, 2004, Housing Transactions and Macroeconomic Fluctuations: A Case Study of England and Wales, *Journal of Housing Economics* 13, 287-303.

Pesaran, M. Hashem, 2004, General Diagnostic Tests for Cross Section Dependence in Panels, *mimeo, University of Cambridge*.

Pesaran, M. Hashem, 2006, Estimation and Inference in Large Heterogeneous Panels with Multifactor Error Structure, *Econometrica* 74, 967-1012.

Phang, Sock-Yong, 2004, House prices and aggregate consumption: do they move together? Evidence from Singapore, *Journal of Housing Economics* 13, 101-119.

Poterba, James M., 2000, Stock Market Wealth and Consumption, *Journal of Economic Perspectives* 14, 99-118.

Sheiner, Louise, 1995, Housing Prices and the Savings of Renters, *Journal of Urban Economics* 38, 94-125.

Slacalek, Jirka, 2006, What Drives Personal Consumption? The Role of Housing and Financial Wealth, *German Institute for Economic Research Working Paper*.

Table 1 Data Summary

Panel A summarizes the respective mean, median, standard deviation for the time series of log differences of per capita GMP (GMP), home price index (HP), population (PO), average household income (HI), unemployment rate (UR), mortgage refinancing origination (MT), single family house permits (PT), 30-year fixed rate mortgage rate (MR), and the SP500 index (SP). Panel B reports their 1 to 4-quarter autocorrelations. Panel C reports correlations among the variables. For statistics involving MSA level variables (GMP, HP, PO, HI, UR, MT and PT), the reported numbers are across MSA averages. For these variables, we also report the t-statistics with the null hypotheses being that the distributions of the variables have zero means. * denotes significance at the 5% level and ** at the 1% level.

	GMP	HP	PO	HI	UR	MT	PT	MR	SP
Panel A. Mean, median, and standard deviation									
Mean	**0.447%	**0.461%	**0.281%	**0.305%	**-0.092%	**1.232%	0.823%	-0.785%	1.472%
Median	**0.452%	**0.470%	**0.282%	**0.321%	**-0.476%	**2.916%	0.200%	-1.241%	1.563%
Sta. Dev.	**1.264%	**2.131%	**0.194%	**1.288%	**7.382%	**38.958%	16.605%	5.768%	6.831%
Panel B. Autocorrelation									
1 quarter	**0.292	*-0.043	**0.955	**-0.058	**0.230	**0.203	**-0.262	-0.093	0.002
2 quarter	**0.247	**0.182	**0.837	**0.123	**0.118	**0.050	*-0.013	-0.055	0.064
3 quarter	**0.082	**0.220	**0.678	**-0.121	**0.067	**0.092	**0.061	0.127	0.146
4 quarter	**0.066	**0.204	**0.532	**0.205	*-0.019	**-0.267	**-0.082	-0.187	0.057
Panel C. Correlation									
GMP	1	**0.211	**-0.060	**0.418	**-0.229	**-0.094	**0.054	0.007	**0.135
HP		1	**0.102	**0.168	**-0.029	**0.081	**0.051	**-0.056	**0.074
PO			1	**-0.049	0.016	**0.019	**0.027	0.007	**0.021
HI				1	**-0.112	**0.020	**0.057	**0.079	**0.181
UR					1	**0.167	**-0.031	**-0.069	**-0.109
MT						1	**0.157	**-0.101	**0.078
PT							1	*0.011	-0.006
MR								1	-0.147
SP									1

Table 2 Cross Section Dependence Tests

This table reports the cross section dependence tests for three different specifications of a linear model in which the log difference of GMP is regressed on the log difference of the house price index for the corresponding metropolitan area.

Specification 1: $gmp_{i,t} = \alpha_i + \beta_i hp_{i,t} + u_{i,t}$
CD test value: 673.80
Specification 2: $gmp_{i,t} = \alpha_i + \beta_i^1 hp_{i,t} + \beta_i^2 po_{i,t} + \beta_i^3 hi_{i,t} + \beta_i^4 ur_{i,t} + u_{i,t}$
CD test value = 387.22
Specification 3: $gmp_{i,t} = \alpha_i + \beta_i^1 hp_{i,t} + \beta_i^2 po_{i,t} + \beta_i^3 hi_{i,t} + \beta_i^4 ur_{i,t} + \beta_i^5 mr_t + \beta_i^6 mr2_t + \beta_i^7 sp_t + \beta_i^9 sp2_t + u_{i,t}$
CD test value = 382.22

Table 3 House Price Changes and Economic Growth

This table reports the estimation results for four different specifications of the following linear model:

$$gmp_{i,t} = \alpha_i + \beta hp_{i,t} + \rho' x_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T$$

$$x'_{i,t} = \left(po_{i,t}, hi_{i,t}, ur_{i,t} \right), u_{i,t} = \gamma'_i f_t + \varepsilon_{i,t},$$

where $gmp_{i,t}$ and $hp_{i,t}$ are the log differences of the GMP and the house price index from quarter t to $t+1$ for MSA i; $x_{i,t}$ is a vector of MSA level variables, including log differences of population, average household income, and unemployment rate; f_t is a vector of unknown common effects; and $\varepsilon_{i,t}$ are the idiosyncratic errors. The first specification does not include the MSA fixed effect or control variables; the second specification includes control variables but not the MSA fixed effect; the third specification includes the MSA fixed effect but not control variables; and the fourth specification includes both the MSA fixed effect and control variables. OLS stands for simple OLS estimators and CCE estimators are Common Correlated Effects estimators of Pesaran (2006). * denotes significance at the 5% level and ** at the 1% level.

Estimator	HP	PO	HI	UR	Fixed effect	R2
			Specification I			
OLS	**0.110				No	0.04
	[0.004]					
CCE	**0.042				No	0.26
	[0.004]					
			Specification II			
OLS	**0.115	**-0.210	**0.263	**-0.029	No	0.20
	[0.004]	[0.056]	[0.008]	[0.001]		
CCE	**0.039	-0.141	**0.145	**-0.027	No	0.35
	[0.005]	[0.072]	[0.009]	[0.001]		
			Specification III			
OLS	**0.110				Yes	0.04
	[0.04]					
CCE	**0.039				Yes	0.27
	[0.004]					
			Specification IV			
OLS	**0.121	**-0.405	**0.261	**-0.030	Yes	0.21
	[0.004]	[0.081]	[0.008]	[0.001]		
CCE	**0.048	*-0.332	**0.145	**-0.027	Yes	0.36
	[0.005]	[0.131]	[0.009]	[0.001]		

Table 4 House Price Changes and Economic Growth

This table reproduces the estimation results for the fourth specification in table 3 with six extra control variables added to the regressions. The extra control variables are the current and one quarter lag of the log difference in mortgage origination (refinance), the current and one to three quarter lags of the log difference in single family house permits. * denotes significance at the 5% level and ** at the 1% level.

Estimators	OLS	CCE
HP	**0.123	**0.047
	[0.004]	[0.005]
PO	**-0.443	**-0.349
	[0.073]	[0.129]
HI	**0.261	**0.141
	[0.009]	[0.009]
UR	*-0.026	**-0.026
	[0.001]	[0.001]
MT	**-0.004	-0.000
	[0.000]	[0.000]
MT lag1	**0.002	**-0.000
	[0.000]	[0.000]
PT	**0.003	**0.003
	[0.000]	[0.000]
PT lag1	**0.004	**0.003
	[0.001]	[0.000]
PT lag2	** 0.003	**0.002
	[0.001]	[0.000]
PT lag3	**0.002	**0.001
	[0.001]	[0.000]
R2	0.23	0.36

Table 5 House Price Changes and Economic Growth: Sub-sample Analysis

This table reproduces the estimation results for the fourth specification in table 3 for two sub periods: the period from 1983 to 1994 and the period from 1995 to 2005. * denotes significance at the 5% level and ** at the 1% level.

Estimator	HP	PO	HI	UR	Fixed effect	R2
			1983 to 1994			
OLS	**0.080	0.034	**0.238	**-0.039	Yes	0.19
	[0.007]	[0.064]	[0.010]	[0.002]		
CCE	**0.060	*0.155	**0.163	**-0.030	Yes	0.29
	[0.007]	[0.060]	[0.013]	[0.002]		
			1995 to 2005			
OLS	**0.139	**-0.463	**0.261	**-0.025	Yes	0.20
	[0.005]	[0.044]	[0.006]	[0.001]		
CCE	**0.023	**-0.501	**0.126	**-0.024	Yes	0.40
	[0.005]	[0.038]	[0.006]	[0.001]		

Table 6 Do House Price Changes Granger Cause Economic Growth

This table reports the estimation of the following dynamic model

$$gmp_{i,t} = \alpha_i + \alpha_t + \sum_{s=1}^{4} \gamma_s gmp_{i,t-s} + \sum_{s=1}^{4} \beta_s hp_{i,t-s} + \beta^2 po_{i,t} + \beta^3 hi_{i,t} + \beta^4 ur_{i,t} + u_{i,t} .$$

Time period fixed effects are eliminated by subtracting cross sectional means on both sides. The MSA fixed effects are eliminated by taking differences on both sides. The model is then estimated with instrumental variable regression to overcome the correlation between independent variables and the error term. * denotes significance at the 5% level and ** at the 1% level. The Granger Causality Test is essentially a F test with the null hypothesis being that all coefficients of the four lagged house price changes (log differences) are all 0.

Variable	Estimator	Standard dev.	T value
HP lag1	-0.004	0.006	-0.69
HP lag2	-0.011	0.009	-1.25
HP lag3	-0.011	0.009	-1.26
HP lag4	-0.009	0.006	-1.51
GMP lag1	**0.90	0.020	14.22
GMP lag2	**0.116	0.013	9.19
GMP lag3	-0.006	0.010	-0.59
GMP lag4	**-0.178	0.009	-18.52
PO	**-0.949	0.273	-3.48
HI	**0.038	0.006	6.85
UR	**-0.021	0.001	-18.39
Granger Causality Test: Does HP Granger cause GMP			
F statistic		1.624	
P value		0.16	

Table 7 Does Economic Growth Granger Cause House Price Changes

This table reports the estimation of the following dynamic model

$$hp_{i,t} = \alpha_i + \alpha_t + \sum_{s=1}^{4} \gamma_s gmp_{i,t-s} + \sum_{s=1}^{4} \beta_s hp_{i,t-s} + \beta^2 po_{i,t} + \beta^3 hi_{i,t} + \beta^4 ur_{i,t} + u_{i,t} .$$

Time period fixed effects are eliminated by subtracting cross sectional means on both sides. The MSA fixed effects are eliminated by taking differences on both sides. The model is then estimated with instrumental variable regression to overcome the correlation between independent variables and the error term. * denotes significance at the 5% level and ** at the 1% level. The Granger Causality Test is essentially a F test with the null hypothesis being that all coefficients of the four lagged GMP growth changes (log differences) are all 0.

Variable	Estimator	Standard dev.	T value
HP lag1	**-1.047	0.006	-164.29
HP lag2	**-0.772	0.009	-88.36
HP lag3	**-0.433	0.009	-50.30
HP lag4	**-0.144	0.006	-24.10
GMP lag1	**-0.125	0.020	-6.22
GMP lag2	*-0.027	0.013	-2.07
GMP lag3	*0.025	0.010	2.41
GMP lag4	**0.050	0.009	5.55
PO	**0.305	0.056	5.41
HI	0.007	0.006	1.12
UR	0.001	0.001	0.89
Granger Causality Test: Does GMP Granger cause HP			
F statistic		26.392	
P value		0.00	

Table 8 Orthogonalized House Price Changes and Economic Growth

This table reports the estimation results for four different specifications of the following linear model:

$$gmp_{i,t} = \alpha_i + \beta hp_{i,t} + \rho' x_{i,t} + u_{i,t}, i = 1, 2, \ldots, N; t = 1, 2, \ldots, T$$

$$x'_{i,t} = \left(po_{i,t}, hi_{i,t}, ur_{i,t} \right), u_{i,t} = \gamma'_i f_t + \varepsilon_{i,t},$$

where $gmp_{i,t}$ is the log difference of the GMP; $hp_{i,t}$ is the component of the log difference of the house price index that is orthogonal to lagged $gmp_{i,t}$ in past four quarters; $x_{i,t}$ is a vector of MSA level variables, including log differences of population, average household income, and unemployment rate; f_t is a vector of unknown common effects; and $\varepsilon_{i,t}$ are the idiosyncratic errors. The first specification does not include the MSA fixed effect or control variables; the second specification includes control variables but not the MSA fixed effect; the third specification includes the MSA fixed effect but not control variables; and the fourth specification includes both the MSA fixed effect and control variables. OLS stands for simple OLS estimators and CCE estimators are Common Correlated Effects estimators of Pesaran (2006). * denotes significance at the 5% level and ** at the 1% level.

Estimator	HP	PO	HI	UR	Fixed effect	R2
			Specification I			
OLS	**0.160				No	0.01
	[0.014]					
CCE	**0.071				No	0.26
	[0.008]					
			Specification II			
OLS	**0.080	**-0.160	**0.291	**-0.029	No	0.17
	[0.012]	[0.060]	[0.009]	[0.001]		
CCE	**0.035	-0.119	**0.147	**-0.027	No	0.35
	[0.012]	[0.075]	[0.009]	[0.001]		
			Specification III			
OLS	**0.172				Yes	0.03
	[0.017]					
CCE	**0.054				Yes	0.26
	[0.017]					
			Specification IV			
OLS	**0.101	**-0.336	**0.290	**-0.029	Yes	0.17
	[0.014]	[0.095]	[0.009]	[0.001]		
CCE	**0.052	*-0.286	**0.147	**-0.027	Yes	0.35
	[0.014]	[0.141]	[0.009]	[0.001]		

Figure 1 Starting Quarters for MSA OFHEO House Price Indices

This figure plots the histogram of MSAs with different starting quarters for OFEHO house price indices in the sample period, which is from 1983:1 to 2005:4.

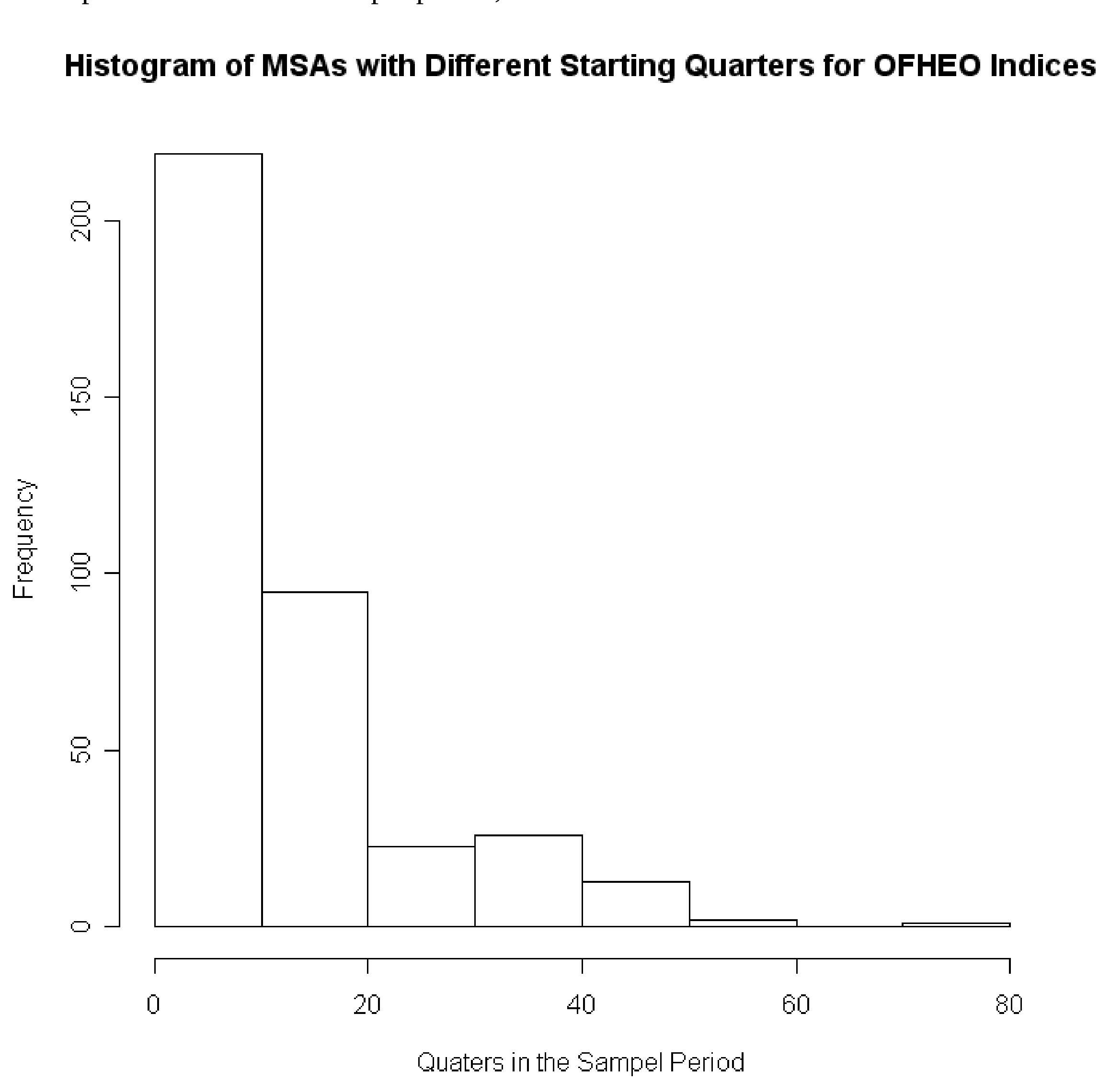


Figure 2 Home Price Indices
This figure plots OFHEO quarterly home price indices for all 379 MSAs in the U.S. from 1983:1 to 2005:4.

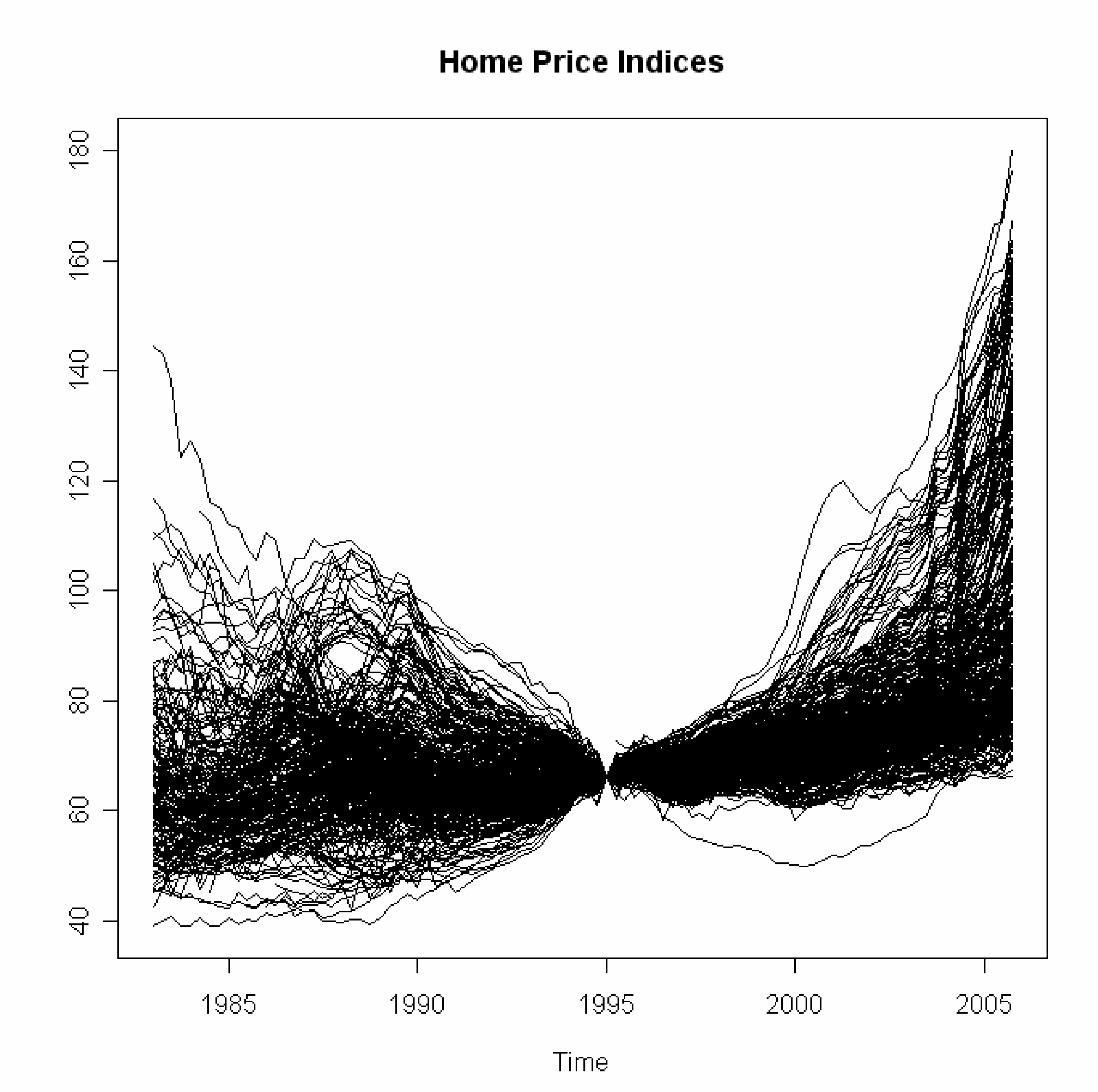


Figure 3 SP500 Index and 30-year Fixed-rate Mortgage Rate from 1983:1 to 2005:4

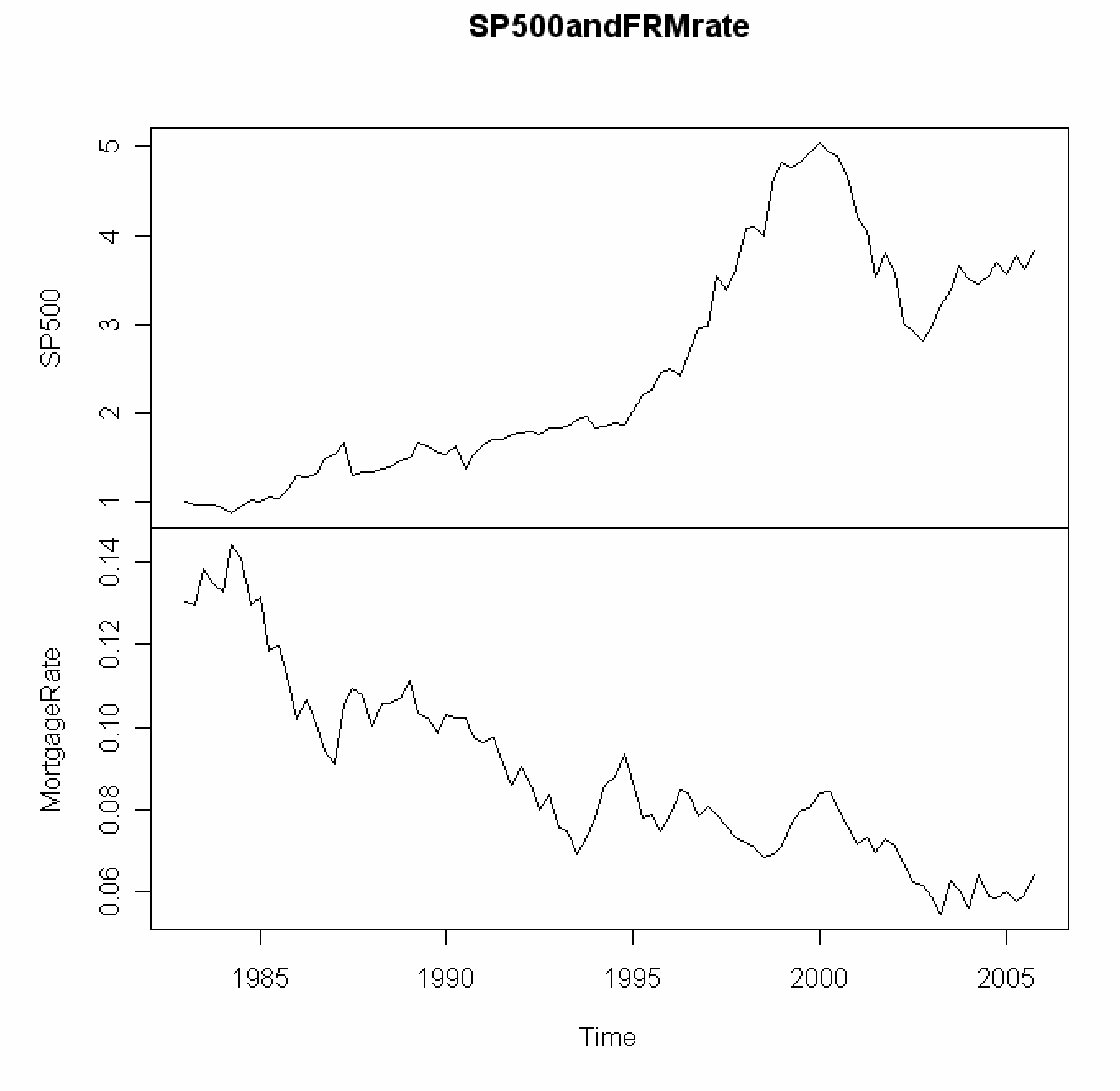
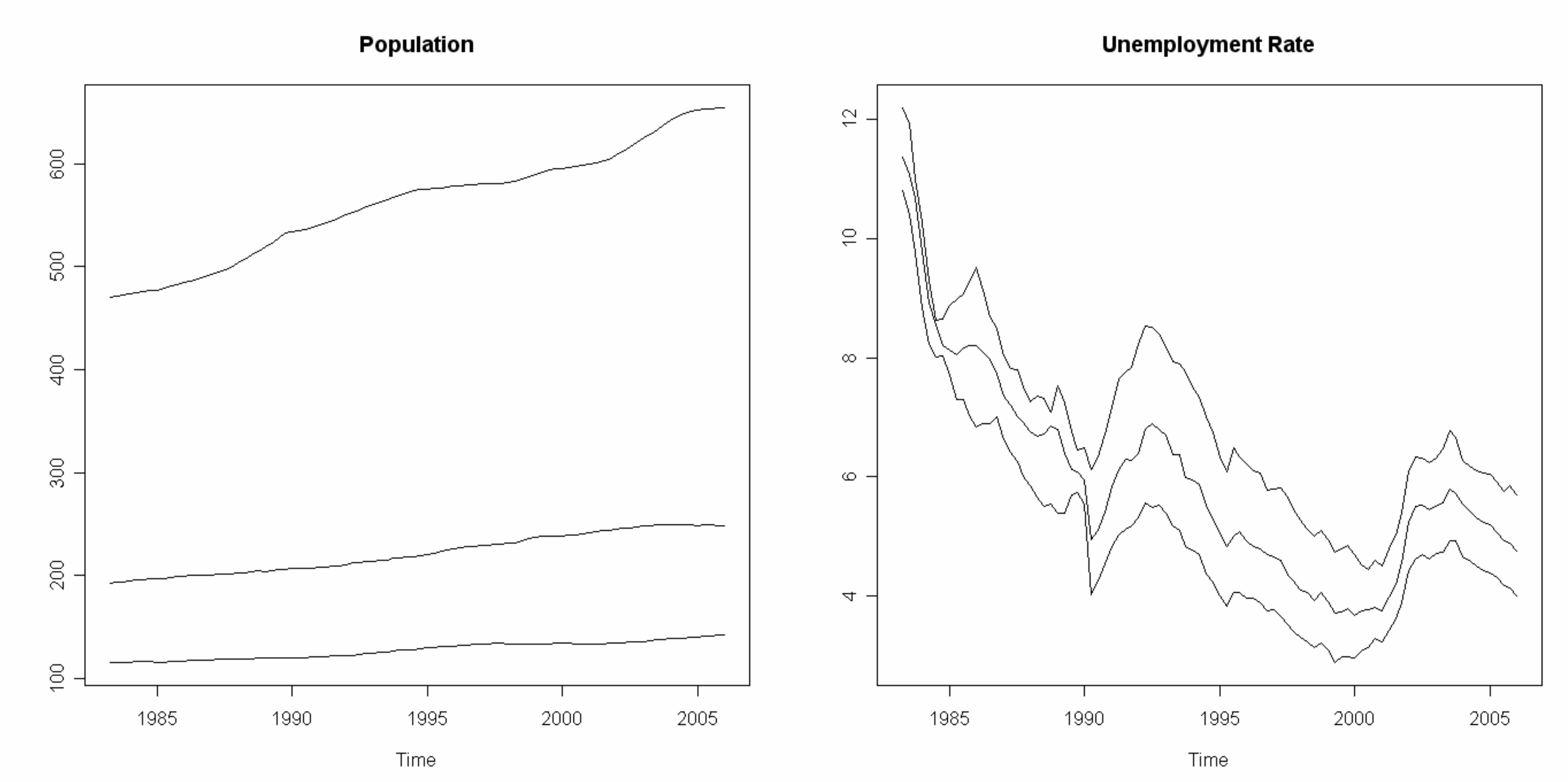
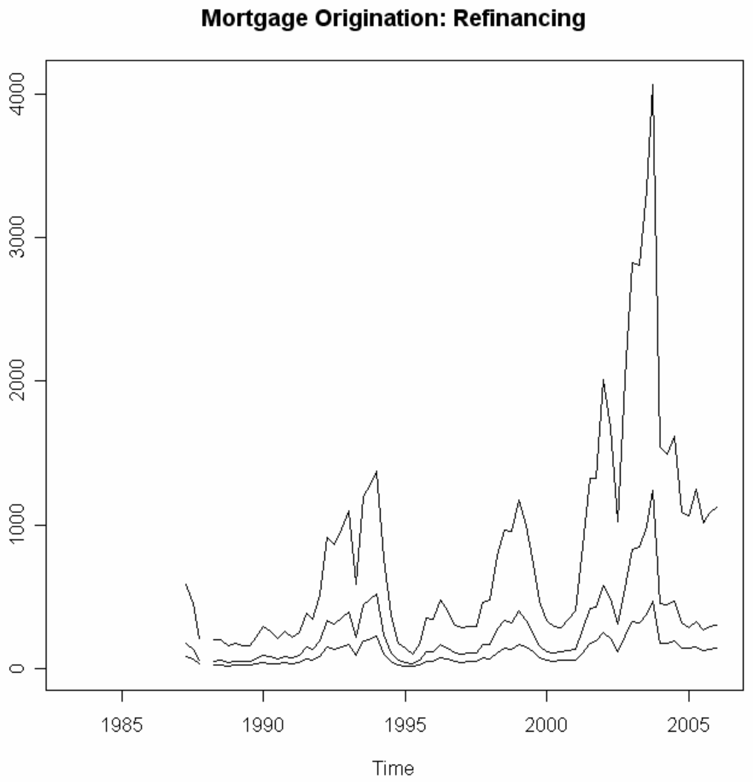
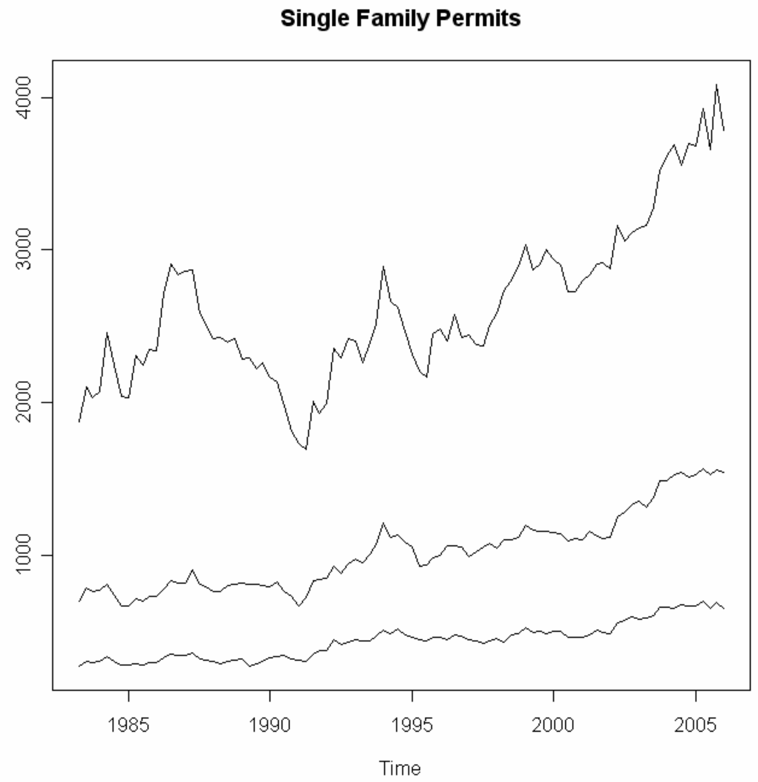


Figure 4

This figure visualizes the time series of 25%, 50%, and 75% percentiles of the per capita GMP (in thousand dollars), average household income (in dollars), population (in 1000 people), unemployment rate (in percentage points), mortgage originations (refinancing and in million dollars), and permits for single family homes, across all 379 MSAs in the U.S. from 1983:1 to 2005:4.


Population
600
500
400
300
200
100
1985
1990
1995
2000
2005
Time
Unemployment Rate
12
10
8
6
4
1985
1990
1995
2000
2005
Time


Mortgage Origination: Refinancing
0
1000
2000
3000
4000
1985
1990
1995
2000
2005
Time

Single Family Permits
1000
2000
3000
4000
1985
1990
1995
2000
2005
Time