The Economic Impact of Anticipated House Price Changes - Evidence from Home Sales

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If realized house prices have the wealth effect and the collateral effect on the economy, anticipated house price changes should have similar economic effects. This paper empirically analyzes the effects of single family home sales, which are shown to be able to predict house prices in the literature, on economic production, using 372 Metropolitan Statistic Areas in the U.S. from the first quarter of 1981 to the second quarter of 2008 in a panel Vector Error Correction Model. Changes in home sales are found to Granger cause the growth rate of per capita Gross Metropolitan Product, and the dynamic effects are visualized with impulse response functions. Supporting evidence for the economic impact of home sales is also found in contemporaneous regressions.

House price changes appear to have significant effects on the economy. Friedman's permanent income hypothesis implies that home owners would increase (decrease) their consumption if home values, and thus their expected life time wealth, increase (decrease). Recent research (Aoki, Proudman and Vlieghe (2004), Lustig and Nieuwerburgh (2010), and Ortalo-Magné and Rady (2004), among others) suggests that increases in home values may help relax home owners' borrowing constraints and thus increase their consumption. The literature generally provides supporting evidence for both the wealth effect and the collateral effect. A strong correlation between house prices and consumption is found at both the aggregate level (see, e.g., Benjamin, Chinloy and Jud (2004), Case, Quigley and Shiller (2005)) and the household level (see, e.g., Bostic, Gabriel and Painter (2009), Campbell and Cocco (2007), and Haurin and Rosenthal (2006)). Miller, Peng and Sklarz (2010) find a strong relationship between house prices and economic production using panel data of all Metropolitan Statistic Areas (MSA) in the U.S., and that the collateral effect is about three times stronger than the wealth effect.

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This paper hypothesizes that *anticipated* house price changes, which affect *future* house collateral values and expected life time wealth, have similar economic effects as *realized* house price changes. This paper uses changes in the volume of home sales as a proxy for anticipated house price changes, and tests if home sales affect economic production. Economists have long realized the power of home sales in capturing housing demand (see, e.g., Berkovec and Goodman (1996)) and thus predicting future house prices. Stein (1995), Genesove and Mayer (1997, 2001), , Chan (2001), Engelhardt (2003); Ortalo-Magné and Rady (2004), among others, suggest that sellers may be slow to reduce their reservation prices in busting housing markets.[1] As a result, decreasing housing demand is captured more by decreasing home sales than by decreasing house prices. Moreover, Wheaton (1990) and Novy-Marx (2009) suggest that sellers would gradually adapt to new market conditions and adjust their reservation prices. Consequently, home sales gradually return to a more "normal" level as house prices begin to reflect changed market conditions. The predictive power of home sales on house prices is substantiated by Clayton, Miller and Peng (2010) and Wheaton and Lee (2009) using Granger causality tests.

This paper tests the effects of home sales on economic production in a Vector Error Correction Model (VECM), using a panel data set on 372 out of 379 Metropolitan Statistical Areas (MSAs) in the U.S. from the first quarter of 1981 to the second quarter of 2008. The VECM consists of three equations, which respectively include the log differences of per capita Gross Metropolitan Production (GMP), house price indexes from the Office of Federal Housing Enterprise Oversight (OFHEO), and house transaction turnover as endogenous variables. The estimation results and formal Granger causality tests provide strong evidence for the predictive power of home sales on

[1] For example, falling home values reduce homeowners' home equity. Therefore, to ensure that the sale proceeds would be sufficient to repay mortgages and provide down payments on next homes, home sellers ask for prices that might be higher than fair market values, which increases the time on the market and reduces the trading volume.

house prices and the impact of home sales on economic production. A robustness check based on the S&P Case-Shiller house price indexes of 19 MSAs provides similar results.

Based on estimated VECM coefficients, this paper further constructs generalized impulse response functions to visualize the dynamic relationships among home sales, house prices, and GMP. The functions reveal that, a one standard deviation shock to home sales (log differences) increases the growth rate of per capita GMP by up to 11% of its standard deviation, and increases the growth rate of house prices by up to 9.5% of its standard deviation. Further, GMP responds more promptly to the home sale volume shock than house prices, which is consistent with the notion that economic production increases as a response to increases in home sales, anticipating higher house prices in the future.

This paper finds additional evidence for the economic impact of home sales in contemporaneous panel regressions of GMP on home sales: home sales have positive and significant coefficients. Further, decreases in home sales have greater coefficients than increases, and home sales have greater coefficients in MSAs with less elastic housing supply. Both findings are consistent with the causation from home sales to GMP.

This paper makes a few original contributions to the literature on the economic impact of the housing market. First, this paper seems the first to hypothesize and test the effects of *anticipated* future house price changes on economic production. Second, this paper is the first to substantiate the Granger causality between home sales and economic production, and to visualize the dynamic interactions. Third, this paper provides novel evidence for the asymmetric effects of increases and decreases in home sales on economic production, and stronger effects of home sales in markets with less elastic housing supply.

This paper significantly differs from and provides a few important extensions to Miller, Peng and Sklarz (2010) (MPS), which also analyzes the impact of the housing market on economic production. First, the two papers have different research questions. The MPS focuses on quantifying and comparing the wealth effect and the collateral effect of house price changes, and does not include home sales in its analysis. This paper focuses on the economic effects of home sales and use house price changes as control variables, but does not distinguish the wealth effect from the collateral effect. Second, the key econometric models are different: while the MPS analyses are based on contemporaneous regressions, the main model in this paper is the VECM. Finally, the two papers provide distinctive economic insights. The MPS highlights the importance of the relative magnitude of the wealth effect and the collateral effect, while this paper is the first to argue that anticipated future house price changes, which home sales help capture, can affect economic production.

This rest of this paper is organized as follows. Section 2 describes the data. Section 3 estimates the VECM and conducts Granger causality tests, which substantiates the causation from home sales to GMP. Section 4 constructs impulse response functions based on the VECM results, and visualizes the economic effects of home sales. Section 5 finds a contemporaneous relationship between home sales and GMP, and provides additional evidence for the economic effects of home sales. Section 6 concludes.

Data

The data used in this paper comprise 372 out of the 379 MSAs in the U.S. (the U.S. Census Bureau 2007 definitions), covers the first quarter of 1981 to the second quarter of 2008. Two points are worth noting regarding the data. First, the data are unbalanced panel, as some variables miss observations in early quarters of the sample period. Second, the Census definitions and boundaries of some MSAs have changed during the sample period. However, both

Economy.com and OFHEO, the two main data sources in this paper, recalculate historical values of the variables used in this analysis whenever the MSA definitions change. As a result, the time series of the variables are geographically consistent.

The first main variable in this analysis, the quarterly per capita GMP, is constructed by dividing the quarterly GMP (seasonally adjusted annual values) with quarterly population, both of which are estimated by Moody's economy.com. Economy.com estimates quarterly population series using Census data and migration flows among MSAs (the data source being the IRS). It estimates the GMP in a MSA with using a two step procedure. 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 (the data being from the BLS):

$$\hat{e}^i_{j,t} = \frac{w^i_{j,t}}{w^i_t} E_{j,t} . \tag{1}$$

In equation (1), for MSA i and quarter t, $\hat{e}^i_{j,t}$ is the estimated "weighted" productivity for Supersector industry j; $w^i_{j,t}$ is the employment in industry j; w^i_t is the total employment; and $E_{j,t}$ is the U.S productivity for industry j. The second step estimates the GMP using the "weighted" productivity for each industry and the total MSA employment:

$$GMP^i_t = w^i_t \times \sum_{j=1}^{J} \hat{e}^i_{j,t} , \tag{2}$$

with J being the total number of industries. From equations (1) and (2), it is clear that Economy.com essentially estimates the GMP by adding up the products of industry employment and the U.S. industry productivity across industries: $GMP^i_t = \sum_{j=1}^{J} w^i_{j,t} E_{j,t}$.

It is possible that Economy.com may measure MSA industry productivity and thus the GMP with error. This paper assumes that the measurement error is not correlated with home sales, which means that MSAs with industries that have higher productivity than the same industries in other MSAs do not necessarily have more or fewer house sale transactions. Therefore, readers should take this assumption into account when interpreting the empirical results provided here. Note that this assumption is not testable, because the test is not about the correlation between home sales and estimated GMP, but about the correlation between home sales and the unobserved measurement *error* in GMP, which is $\sum_{j=1}^{J} w_{j,t}^{i}\left(E_{j,t} - e_{j,t}^{i}\right)$ with $e_{j,t}^{i}$ being the unobserved true productivity for industry j in MSA i.

The second main variable in this analysis, the quarterly single family home price index, comes from two sources: the OFHEO and Standard and Poor's. The OFHEO estimates quarterly house price indexes using repeat sales regressions, which control for time invariant attributes of houses that enter into the sample at least twice. Consequently, the OFHEO indexes are considered superior to median or mean sale prices in the sense that changes in OFHEO indexes capture changes in market valuations instead of changes in attributes of traded houses. This paper matches MSAs in the OFHEO house price index file with MSAs in the data file from Economy.com, using the name and the state of each MSA, and obtains 372 matched MSAs.

Note that the OFHEO indexes likely over-represent houses that use Freddie Mac/Fannie Mae financing since the indexes are constructed with transactions financed with conforming mortgages only. In this sense, the OFHEO indexes seem inferior to the Case-Shiller home price indexes provided by Standard and Poor's, which are constructed using all transactions. However, the Case-Shiller indexes cover only 20 metropolitan areas, which is about 5% of all the MSAs. Further, the Case-Shiller indexes start no earlier than 1987:1, which further reduces the sample size. Therefore, this paper mainly analyzes the OFHEO indexes, and uses the Case-Shiller indexes for robustness checks.

This paper matches the metro areas of the Case-Shiller indexes with the MSAs in Economy.com data by identifying the MSAs in Economy.com of which the names contain the Case-Shiller metro names. For example, "Los Angeles, LA" of the Case-Shiller metros is matched with "Los Angeles-Long Beach-Glendale, CA" in Economy.com data. This leads to 19 matches. The only Case-Shiller metro area that cannot be matched with Economy.com MSAs is "Phoenix, AZ". Note that this match criterion is not perfect and might generate errors. Therefore, readers should interpret the results with caution.

The third main variable, the volume of home sales, is measured by house transaction turnover. Note that it is not appropriate to use the number of house transactions to measure home sales, as some MSAs simply have more total stock of houses than others and thus likely have more transactions. Instead, it seems more sensible to use housing turnover, which is the number of transactions divided with the total number of houses as a measure of home sales activity. This paper constructs the housing turnover by dividing the sales of existing single family houses (seasonally adjusted annual values) with the number of households, both of which are provided by Economy.com. Economy.com compiles the house sale data using data from the National Association of Realtors and Bureau of Census, and estimates the number of households using Census data. The number of households is used to proxy the stock of existing single family houses. While these two numbers may not be always equal and thus there are likely measurement errors in the constructed house transaction turnover, the possible attenuation bias due to the measurement errors does not seem to weaken our results. If the coefficient estimate of the turnover is statistically significant despite the attenuation bias, the true coefficient should be even more significant.

This analysis includes three MSA level control variables: the median household income (seasonally adjusted annual rate), population, and the unemployment rate (seasonally adjusted). All variables are provided by Economy.com, which compiles unemployment rates using Bureau of Labor Statistics data, and estimates median household income using Bureau of Economic Analysis data. Both the level and the changes of these control variables are likely correlated with both GMP growth and changes in house prices and house transaction turnover. For example, Ortalo-Magné and Rady (2004) suggest that changes in household income affect not only the economy but also the housing market. In addition, changes in population often concur with changes in the industrial structure in a MSA, which may indicate changes in productivity and economic production. Changes in population also often concur with migrations, which may affect the dynamics of house prices (see, e.g. Gabriel, Mattey and Wascher (1999) for direct evidence). Finally, the unemployment rate may capture 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.

The panel VECM in this paper uses quarterly dummies to control for all national level macroeconomic variables. The contemporaneous panel regressions not only use a multifactor error structure (Pesaran (2006)) to control for common variables, but also directly include the following quarterly financial and macroeconomic variables: the national average 30-year conventional mortgage interest rate, the GDP (seasonally adjusted annual rates), the marker risk premium of the stock market, the term spread, and the credit spread. The market risk premium of the stock market equals the total return on the value-weighted stock market portfolio minus the corresponding quarterly return on U.S. Treasury securities from the CRSP. The term spread is the difference between the annual yields on 10-year and 1-year Treasuries. The credit spread is the difference between the annual yields on Moody's BAA and AAA rated corporate bonds. All the variables are retrieved from the FRED database, except the market risk premium of the stock market, which is downloaded from Kenneth French's website.

The large panel data set used in this paper has a few important advantages. First, the panel data allow us to control individual heterogeneity (fixed effects) and help increase the power of tests due to a large number of observations. Second, being all in the U.S., the MSAs are homogenous in the sense that they are subject to similar if not identical monetary policies, political environment, federal tax codes, and financial market conditions.[2] Consequently, obtained results would be more reliable since some heterogeneity is difficult to address statistically. Finally, the sample period in this paper covers both economic expansions and recessions, so this analysis is less likely biased in the sense that the effects it captures are transitory or only valid in booming or busting economies. Despite the advantages, note that the MSAs in our sample are not independent and are likely related to each other economically. As a result, readers should be cautious in interpreting the results, for they pertain to *open* economies.

This analysis processes the data by first converting per capita GMP, OFHEO and Case-Shiller house price indexes, median household income, and GDP into real terms, using quarterly CPI obtained from the FRED database, and then calculating the first order log differences of main variables. This paper chooses to work on log differences of the main variables instead of the levels because OFHEO home price indexes are all normalized to be 100 in 1995:1, and thus house price index *levels* are not comparable across MSAs.[3] Note that while GMP, house price indexes, and house transaction turnover are in log differences, both the level and the log differences of the MSA level control variables are included in the analysis. To illustrate the

[2] It is worth noting that there could be significant variation in the state tax code, bankruptcy laws, political environment, etc. across MSAs.
[3] For example, all MSAs have the same price index level in 1995:1 but the actual average house prices certainly differ across MSAs.

temporal behavior of the three main variables and the three MSA level control variables, Figures 1 to 6 plot the 25%, 50%, and 75% percentiles of across MSA distributions of per capita GMP (dollars), house transaction turnover (percentage points), log differences of the OFHEO house price index, population (1,000 people), the median household income ($1,000), and unemployment rates (percentage points) in the sample period.

Table 1 reports summary statistics for the log differences of the six variables. Panel A reports the respective mean, median, and standard deviation of log differences of the variables across MSAs. Panel B reports respective autocorrelations up to the fourth lags, and Panel C reports correlations among the variables. All reported numbers are across MSA averages. The table also reports the t-statistics with the null hypotheses being that the distributions of the statistics have zero means. Note that house transaction turnover significantly correlates with per capita GMP, which is consistent with a positive effect of home sales on economic production. However, both GMP and house turnover are significantly correlated with many control variables, which highlights the importance of including the control variables in our analysis.

A Vector Error Correction Model and Granger Causality

This paper uses a panel Vector Error Correction Model to analyze the possible effects of home sales on economic production at the MSA level. The model treats per capita GMP, house prices, and home sales as endogenous variables, and controls for MSA fixed effects, quarterly fixed effects, and MSA level economic and geographic exogenous variables. The VECM allows the test of Granger causality among the three endogenous variables. While Granger causality is essentially a lead-lag relationship, the Granger causality from home sales to GMP is apparently consistent with the effects of home sales on GMP. The VECM also allows the construction of impulse response functions, which help visualize the dynamic relationships among the GMP, house prices, and home sales.

Assume that the per capita GMP, house prices, and home sales are functions of quarterly dummy variables that capture national exogenous variables (quarterly fixed effects), MSA dummy variables that capture time invariant individual heterogeneity (MSA fixed effects), MSA level exogenous variables and lagged endogenous variables.

$$Y_{i,t} = Q_t + MSA_i \times t + \sum_{s=1}^{k} B_s Y_{i,t-s} + CX_{i,t} + u_{i,t} \tag{3}$$

In equation (3), $Y_{i,t}$ is a 3 by 1 vector with the three elements being the per capita GMP ($GMP_{i,t}$), the house price level ($HP_{i,t}$), and the house transaction turnover ($TO_{i,t}$), all in logs, for the i th MSA in quarter t; Q_t is a 3 by 1 time dependent intercept term, which captures effects of time varying national level variables on the three endogenous variables; MSA_i is a 3 by 1 vector of MSA dummy variables; $X_{i,t}$ is a n by 3 matrix of MSA level exogenous variables with the 3 by n coefficient matrix denoted by C; $u_{i,t}$ is a 3 by 1 vector of errors. Note that equation (3) describes short-term dynamic relationship among the three endogenous variables.

Further assume that the three endogenous variables (in logs) have a simple long-term equilibrium relationship as follows.

$$GMP_{i,t} = k + \rho_1 HP_{i,t} + \rho_2 TO_{i,t} + v_{i,t} \tag{4}$$

The long-term relationship is substantiated with cointegration tests at the MSA level using OFHEO house price indexes.[4] Note that equation (4) essentially imposes constraints on coefficients in equation (3), and thus introduces an error correction term in the first order difference model of (3).

[4] In the cointegration tests, the three variables are measured with their deviations from their own MSA means, to overcome the problem that OFHEO house price indexes are not comparable across MSAs.

Since the OFHEO house price indexes are normalized to 100 in 1995:1 and thus do not directly measure house price levels, this paper estimates the first order difference of (3).

$$y_{i,t} = q_t + MSA_i + \sum_{s=1}^{k} B_s y_{i,t-s} + Cx_{i,t} + e_{i,t-1} + \varepsilon_{i,t} \tag{5}$$

The lower case variables in (5) are the first order differences of corresponding upper case variables in (3), and $e_{i,t-1}$ is the error correction term that equals the error $v_{i,t-1}$ in (4). This paper estimates the panel VECM in (5) and tests the hypothesis that house transaction turnover Granger causes both house prices and per capita GMP. This hypothesis is consistent with the notion that economic production, which is measured with per capita GMP, would increase with home sales, anticipating higher future house prices and thus higher future consumption.

To determine an appropriate specification of the VECM, this paper first conducts augmented Dickey–Fuller tests for the three endogenous variables in (5) for each MSA, and finds no evidence of unit roots. The error correction terms for each MSA are constructed using one quarter lagged residuals of the regressions of per capita GMP on house price index and house transaction turnover. This regression is run for each MSA separately since house price indexes are not comparable across MSAs. Note that the deviations of house price index levels from actual house price levels are captured by coefficients and thus do not affect the residuals. Finally, AIC is used to detect optimal lag order for each MSA, and about 90% of the MSAs have optimal lag order smaller than or equal to 5. Consequently, we let the lag order of the panel VECM in (5) be 5. Each of the three equations in the VECM also includes contemporaneous levels and the first order differences of population, median household income, and the unemployment rate (all in log) as MSA level control variables. National level variables, such as GDP, the term spread, and the credit spread, are not included because the quarterly dummy variables completely capture their effects.

Table 2 reports the coefficient estimators of lagged dependent variables in each of the three equations. First, the results substantiate the causation from house prices and home sales to per capita GMP. Specifically, in the equation with $gmp_{i,t}$ being the dependent variable, lagged house price and house transaction turnover both have positive coefficients, and some of the coefficients are statistically significant at the 1% level. Second, the results also indicate that per capita GMP affects house prices and transactions. Specifically, coefficients of lagged per capita GMP changes are all positive in the equations of house prices and home sales, and some are statistically significant at the 1% level. Third, lagged house prices significantly affect house transactions, and lagged house transactions also significantly affect house prices, with the signs and the magnitude of the coefficients being consistent with the results in Clayton, Miller and Peng (2010), which analyze the interactions between house prices and home sales.

Table 3 reports results of formal Granger causality tests, which are essentially F tests. For example, the null hypothesis that house transaction turnover does not affect per capita GMP requires that all coefficients of lagged house transaction turnover are not significantly different from 0 in the equation with per capita GMP being the dependent variable. The results in Table 3 provide strong evidence for Granger causality among all three variables. Particularly, per capita GMP and home sales Granger cause each other, which provides strong evidence for positive effects of house transactions on future per capita GMP.

This paper further conducts robustness analyses, using the Case-Shiller house price indexes instead of the OFHEO house price indexes to estimate the panel VECM model and conduct the Granger causality tests. While the Case-Shiller house price indexes might be more accurate than the OFHEO indexes, they cover only 19 metropolitan areas and have shorter sample periods. As a result, the sample size of the robustness analyses is only about 6% of the sample size of the

VECM based on OFHEO indexes. Table 4 reports the sample period for each of the 19 metro areas covered by the Case-Shiller indexes.

Table 5 reports the estimation results, and Table 6 reports the Granger causality tests. It is clear that the results still provide strong evidence for effects of home sales on per capita GMP. Specifically, coefficients of lagged house transaction turnovers are all positive and three of them are significant at 1% level in the regression with GMP being the dependent variable. The F-tests corroborate this result: house transaction turnover Granger causes per capita GMP at the 5% level, and Granger causes house prices at the 1% level. However, note that per capita GMP no longer Granger causes home sales or house prices, and house prices no longer Granger causes per capita GMP and home sales. To analyze if this is due to the smaller sample size and thus less statistic power of the robustness analyses or differences between the OFHEO indexes and the Case-Shiller indexes, this paper repeats the robustness analyses using the OFHEO indexes for the same 19 metro areas covered by the Case-Shiller indexes and finds similar results. Therefore, the disappearing Granger causality does not seem to be due to the differences between the OFHEO indexes and the Case-Shiller indexes, but likely due to the much smaller sample size of the robustness analyses.

Impulse Response Analysis

This paper uses impulse response functions to visualize the effects of changes in home sales on per capita GMP and house prices. The impulse response functions are constructed for the VECM in (5), and thus illustrate the effects of a shock to the log difference of house transaction turnover on the growth rates (log differences) of per capita GMP and OFHEO house price indexes. In this paper, impulse response functions based on log differences seem more sensible than level-based functions for the following reasons. First, this paper estimates the difference model in (5) instead of the level model in (3). Therefore, the impulse response functions constructed with the estimated coefficients clearly measure "impulses" and "responses" in log differences. Second, log differences of per capita GMP and house prices are stationary while the levels are apparently non-stationary. Changes in stationary variables seem to provide a cleaner description of responses than changes in non-stationary variables, because the later depends on the time varying means and standard deviations of the non-stationary variables. Finally, to construct level based impulse response functions, extra conversions and assumptions regarding the non-stationary GMP and house prices are needed. These assumptions might introduce additional errors and thus do not seem desirable. From now one, discussions in this section use per capita GMP, house prices, and house transaction turnover to refer to their log differences.

The impulse response functions are constructed from 1,000 rounds of simulations. Each round consists of the following steps. The first step generates a transient one-quarter shock to the system. The shock is a 3 by 1 vector, with the three elements representing respective shocks to per capita GMP, the house price index, and house transaction turnover. While the impulse response functions intend to illustrate the effects of a shock to house transaction turnover on per capita GMP and house prices, house transaction turnover is not only endogenous but also correlated with per capita GMP and house prices (see Table 1 for correlations). As Koop, Pesaran and Potter (1996) and others point out, it is inappropriate to assume a shock on one endogenous variable while keeping other endogenous variables fixed. To generate the vector of shock, the shock to house transaction turnover is fixed in all 1,000 rounds of simulations, and equals its sample standard deviation (average across MSAs). The shocks to per capita GMP and house price index are randomly generated from a conditional multivariate Normal distribution, which is inferred from a joint Normal distribution of per capita GMP, house price index, and house transaction turnover. The vector of the means and the covariance matrix of the joint Normal distribution are estimated using corresponding sample moments (averages across MSAs).

The conditional distribution is conditional upon the value of house transaction turnover being its standard deviation.

The second step randomly generates coefficients for the VECM model using the estimated means and covariance of the coefficients, assuming Normality. This is to take into account the fact that the coefficients are estimated, and thus may contain errors. It is clear that this step further introduces randomness into the responses of GMP and house prices. The third step lets all exogenous variables, lagged endogenous variables, and the error correction term be 0, and then let the endogenous variables equal the shock generated in the first step. The future values of all three endogenous variables are then calculated by repeatedly plugging into the VECM the coefficients generated in the second step and lagged endogenous variables. In this iterated calculation, exogenous variable remain 0. Therefore, the values of endogenous variables measure net effects of the initial shock, which are deviations from the values of the endogenous variables that are determined by exogenous variables. In this process, the lagged error correction term also remains 0 to simplify the calculation. This has no real effect on the result, since the coefficient of the lagged error correction term is not significantly different from 0 (the estimated coefficient is about 0.00007 with standard deviation being about 0.001). Further, to construct the error correction term, this paper must construct the levels of non-stationary GMP, house prices, and house transaction turnover, which involves more assumptions and calculations. The insignificant effect of the error correction term on the impulse response functions does not seem to justify the possible errors due to the extra assumptions and calculations.

The above three steps are repeated for 1,000 times, and generate 1,000 time series of per capita GMP, house prices, and house transaction turnover, from which the mean and standard deviation of each endogenous value in each quarter after the shock can be calculated. Figure 7 plots the mean values of per capita GMP in the 20 quarters after the shock. In this figure, per capita GMP values are all normalized with their standard deviation, which is estimated with the sample moment, so 1 means one standard deviation. This figure shows that a one standard deviation increase in the house transaction turnover increases per capita GMP by about 11% of its standard deviation in the next quarter. The effect diminishes gradually and becomes about 0.5% of the standard deviation after about 4 quarters. Note that the illustrated effects are on log differences; therefore, while the illustrated effects diminish eventually, the effects on the levels are permanent. This figure also plots the confidence bands that cover one standard deviation above and below the mean values, which indicates that per capita GMP increases with about 67% probability in the quarter after the shock, even with the estimation errors in coefficients taken into account.

Figure 8 plots the mean values and the same confidence band for the house price index in the 20 quarters after the shock. The figure shows that the house price first increases, reaches the highest value (9.5% of its standard deviation) in the third quarter after the shock, and then decreases gradually over time, becoming less than 1% 18 quarters after the shock. This is consistent with the conventional notion that increases in home sales predict house price increases. Contrasting Figures 7 and 8 indicates that per capita GMP responds more promptly to the home sale shock than house prices, in the sense that per capita GMP reaches the highest value one quarter after the shock while house prices reaches the highest value three quarters after the shock. This seems to suggest that the economic production increases as a response to increases in home sales, anticipating higher house prices in the future. It is also apparent that the effects on house prices last much longer than the effects on per capita GMP, which suggests that the housing market adjusts slowly to shocks.

Contemporaneous Analysis

The causation from home sales to GMP may also imply a contemporaneous relationship between them, which is analyzed with the following linear multifactor error structure panel model.

$$gmp_{i,t} = \alpha_i + \beta \times to_{i,t} + \rho \times x_{i,t} + \gamma \times z_t + u_{i,t} \quad (6)$$

In equation (6), for MSA i, α_i is a MSA fixed effect; $gmp_{i,t}$ and $to_{i,t}$ are respectively the log differences of per capita GMP and house transaction turnover between quarter $t-1$ and t; $x_{i,t}$ is a vector of MSA level control variables, including the log difference of house price index, the level (in log) and the log differences of population, median household income, and the unemployment rate; z_t is a vector of national level control variables, including the national average 30-year conventional mortgage interest rate, the GDP growth rate (log difference), the stock market risk premium, the level and first order difference of the term spread and the credit spread; and $u_{i,t}$ is the error term.

Assume that the error term includes latent variables as well as an idiosyncratic error.

$$u_{i,t} = \gamma_i' f_t + \varepsilon_{i,t} \quad (7)$$

In equation (7), f_t is a vector of latent variables that might affect both GMP and house transactions in all MSAs, such as the availability of mortgage financing and consumer confidence. CD (Cross-section Dependence) tests based on Pesaran (2004) provide strong evidence of the existence of such common factors.[5] $\varepsilon_{i,t}$ is an idiosyncratic error, which is assumed to be distributed independently of $x_{i,t}$ and f_t and across MSAs. Note that equation (7) allows the coefficients of the latent variables to differ across MSAs. As a result, the latent factors include variables that affect some but not all MSAs since the MSAs that are not affected by these factors simply have zero coefficients. The error structure in equation (7) is more flexible than time dummies in controlling for latent variables. While a time dummy (time fixed effect) is able to control for common variables that affect *all* MSAs *to the same extent*, the multifactor error

[5] The test statistics are not reported but all significant at the 1% level.

structure is able to control variables that affect MSAs *differently*. Further, if the multifactor error structure is the correct specification, $\gamma_i' f_t$ is expected to take away the explanatory power of observed national level control variables z_t.

We estimate four specifications of the multifactor error structure model in (6) and (7). All four specifications include log differences of house price indexes and national level control variables. The first specification does not include MSA fixed effects nor MSA control variables. The second specification includes MSA fixed effects but not MSA control variables. The third specification includes MSA control variables but not MSA fixed effects. The fourth specification includes both MSA fixed effects and MSA control variables. The coefficient estimators of the multifactor error structure model are called Common Correlated Effects estimators (CCE estimators). According to Pesaran (2006), the CCE estimators are constructed using regressions augmented with cross-sectional averages of all dependent and independent variables. Pesaran (2006) proves that the cross-sectional averages of all dependent and independent variables span the same space of and thus control for the latent common factors.

Table 7 reports the estimation results. First, the coefficient of house transaction turnover is positive and significant at the 1% level in all four specifications, which indicates that home sales help explain economic production. Second, while the coefficients of house transaction turnover are statistically significant, they are much smaller than the coefficients of home prices. For example, in the fourth and the most exhaustive specification the coefficient for house transaction turnover is 0.003, while the coefficient for house prices is 0.054. This indicates that house transactions have less explanatory power than home prices. Third, while not reported in the table, all national level control variables are insignificant. This is expected because Pesaran (2006) proves that cross-sectional averages of all dependent and independent variables, which are augmented to the regressions to generate the CCE estimators, span the same space as all national level explanatory variables do, including both observed and unobserved ones. Further, it is worth noting that the same regressions using Case-Shiller indexes and their 19 metro areas have insignificant results. To analyze if the insignificant results are due to the smaller sample size and the shorter sample period of the Case-Shiller indexes or the differences between the OFHEO indexes and the Case-Shiller indexes, this paper runs the same regressions using the OFHEO indexes for the same 19 metro areas, which also provide insignificant results. Therefore, the insignificant results are possibly due to the small sample size and the short sample period of the Case-Shiller indexes.

Direction of Causation in Contemporaneous Analysis

The significant coefficients of home sales in Table 7 do not necessarily suggest causation from home sales to GMP, as these two variables affect each other. This paper tests two hypotheses to help identify the causation from home sales to GMP in the contemporaneous regressions. The first hypothesis pertains to the asymmetric coefficients of increases and decreases in home sales. The causation from home sales to GMP predicts that decreases in home sales have larger coefficients, while the causation from GMP to home sales predicts the opposite.

If GMP changes affect housing demand, and thus affect home sales, GMP decreases would reduce house transactions more than GMP growth increases house transactions. House prices tend to be stickier in busting (declining) markets than in booming markets due to liquidity constraints, the default option, and loss aversion of home sellers (See Cauley and Pavlov (2002), Chan (2001), Engelhardt (2003), Genesove and Mayer (1997, 2001), and Lamont and Stein (1999), among others for empirical evidence). As a result, decreasing housing demand due to decreasing GMP is better captured by decreasing house transactions than decreasing house prices. On the other hand, increasing housing demand due to increasing GMP is better captured by increasing house prices than increasing home sales. Therefore, if GMP changes cause home sale changes, negative changes in house transactions would have a smaller coefficient in regressions of GMP on house transactions. Figure 9 visualizes this asymmetry.

Driven by the same phenomenon that home sales adjust more quickly than house prices in busting markets, decreasing home sales due to negative demand shocks would predict lower future house prices (see Clayton, Miller and Peng (2010) and Wheaton and Lee (2009) for direct evidence). Anticipating lower future house prices and thus lower expected wealth of homeowners, economic production would decrease. In booming markets, house prices adjust quickly so home sales have weaker predictive power for future house prices, and thus have weaker effects on the economy. Therefore, if home sale changes cause GMP changes, negative changes in house transactions would have a larger coefficient in regressions of GMP on house transactions.

While the coefficient of home sales likely captures causation in both ways, a larger coefficient for decreases in home sales would substantiate the causation from home sales to GMP. Empirically, we separate positive changes of house transaction turnover from negative changes, and re-estimate the model in (6) and (7). Specifically, we replace the log differences of house transaction turnover $to_{i,t}$ with $to.positive_{i,t}$ and $to.negative_{i,t}$. $to.positive_{i,t}$ ($to.negative_{i,t}$) equals $to_{i,t}$ if it is positive (negative) and 0 otherwise. Table 8 reports the estimation results. In the last two specifications, decreases in house transaction turnover have greater and more significant coefficients than increases. For example, in the fourth specification, which includes MSA fixed effects and local control variables, the coefficient for decreases in house transaction turnover is 0.004 and significant at 1% level, while the coefficient for increases is 0.002 and insignificant. This is consistent with the causation from house transactions to GMP.

The second hypothesis this paper tests relates to possible heterogeneity in the coefficients of home sales across MSAs. If home sales cause GMP, MSAs with less elastic housing supply may have larger coefficients of home sales. As Wheaton (1990) and Novy-Marx (2009) suggest, in MSAs with less elastic housing supply, changes in house transactions more likely lead to changes in future home values since housing supply (housing units) do not adjust quickly.

The literature provides a variety of measurements for housing supply elasticity, such as the geographical constraint measure developed by Saiz (2010), the Wharton Residential Land Use Regulatory Index from Gyourko, Saiz and Summers (2008), and the measure from Maclennan and Malpezzi (2001) and Goodman and Thibodeau (2008). Unfortunately, the metro area boundaries in these papers are based on older definitions of MSAs and thus are not consistent with the MSA boundaries in this paper. For instance, Saiz (2010) uses 1999 MSA, PMSA, and NECMA definitions and provides supply elasticity measures for 95 MSAs with population more than 500,000. Gyourko, Saiz and Summers (2008) use 1999 definitions/boundaries and provide the Wharton Residential Land Use Regulatory Index for 293 out of 337 metro areas. Differences between the MSA definitions in this paper and the older definitions include not only the new MSAs but also possibly mismatched definitions and boundaries of all other MSAs due to the fact that new MSAs are often created by splitting old MSAs.

To avoid possible matching errors due to the inconsistency in the MSA boundaries, this paper does not use the housing supply measurements based on older definitions of MSAs. Instead, this paper follows Clayton, Miller and Peng (2010) and constructs a simple in-sample measure of the long term housing supply elasticity. The measure is the log ratio of population growth to the appreciation of the house price index over the available sample period for each MSA. Holding constant the increases in population in a market, the greater is the house price appreciation, the "tighter" is the market and the lower is the long term supply elasticity. Figure 10 plots the histogram of the long term supply elasticity across the 372 MSAs.

This paper estimates the third and the fourth specifications of the model in equation (6) and (7) for the top 20% tight and loose markets respectively. Table 9 lists the top 20% (75) tight and loose MSAs. The list seems consistent with conventional wisdom regarding housing supply elasticity: the top tight markets include many MSAs on the coasts, and the top loose markets include most inland MSAs such as 19 MSAs in Texas. Table 10 reports the estimation results, and shows significant differences in the coefficient of house transaction turnover between tight and loose markets. Specifically, the coefficients of house transaction turnover are significantly (5% level) positive for the tight markets, but are insignificant for the loose markets. This result helps substantiate the causation from house transactions to GMP.

Conclusion

A growing literature provides supporting empirical evidence for the economic impact of realized house price changes, motivated by the notion that house price changes affect the expected life time wealth of home owners and the collateral values of homes. If house prices indeed have the wealth and the collateral effects, not only realized house price changes but also anticipated house price changes should have an economic impact. This paper hypothesizes that changes in home sales, which is shown to be able to help predict future house price changes, affect the economic production. Using a large panel data set that covers 372 MSAs from 1981:1 to 2008:2, this paper estimates a panel VECM model that includes per capita GMP, house prices, and home sales as endogenous variables. The estimation results and Granger causality tests provide strong evidence for the impact of home sales on GMP and the predictive power of home sales on house prices. This paper uses impulse response functions to visualize the dynamic interactions among the three variables, which confirms the economic effects of home sales. Contemporaneous panel regressions of per capita GMP on home sales also provide supporting evidence. The results indicate a significant and positive correlation between home sales and GMP. Moreover, the coefficient of house transaction turnover is larger and more significant for decreases in house transactions, and in MSAs within more supply constrained housing markets, which is consistent with causation from home sales to per capita GMP. Overall, this paper provides strong evidence for the economic impact of *anticipated* house price changes in addition to the effects of realized house price changes.

We thank the editor Crocker Liu and two anonymous referees, and Toni Whited and participants of the 2008 AREUEA conference for constructive comments. All errors are ours.

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