Are Housing Bubbles Contagious? A Case Study of Las Vegas and Los Angeles Home Prices

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ABSTRACT. *This paper asks whether speculative house-price pressure in an economic center can spill into related housing markets. In other words, are bubbles contagious? I develop a theoretical model that allows for speculative price appreciation to spread from one market to another. I estimate an error-correction model using quarterly housing data for Las Vegas and Los Angeles and fundamental market variables from 1978 Quarter 2 through 2008 Quarter 1. Las Vegas prices show significant persistence and adjust slowly to disequilibrium. Contagious price and income growth from the Los Angeles market sustained by naïve expectations contributed to the bubble that formed in Las Vegas.* (JEL C32, R21)

I. INTRODUCTION

Most economists agree that fundamental economic conditions including low mortgage interest rates, low unemployment, stable per capita income growth, and restricted supply in many areas contributed to the recent single-family home-price appreciation in the United States. However, evidence suggests that speculative pressures also played a role. Case and Shiller (2003) found evidence of a speculative bubble in many U.S. submarkets, including Los Angeles and San Francisco. They note that large swings in single-family home prices were poorly explained by fundamental forces such as income, population, interest rates, and housing starts. More importantly, price growth exhibited a persistence that was unrelated to fundamental forces, suggesting that speculative forces were at work.

Studies support the notion that housing-market bubbles originate from some precipitating event that acts to stimulate housing demand (Shiller 2000; Riddel 1999; Hu et al. 2006). According to this view, the initial price shock attracts speculative demand into the market. These demanders base their expectation of future price appreciation on past appreciation rather than market fundamentals, creating a mechanism whereby an initial price shock leads to continued appreciation in the future. This price-feedback mechanism amplifies the initial stimulus, and prices pull away from the underlying housing-market fundamentals. For example, Case and Shiller (2003) speculate that sharp declines in equity prices from the collapse of the "tech" bubble in 2000 may have led to a "flight to safety" as funds poured out of asset markets into the perceived safety of the housing market. Persistence in prices resulting from speculative demand amplified the initial price appreciation and a bubble formed.

Case and Shiller's "flight to safety" hypothesis exemplifies one possible source of an initial price shock that may lead to a price bubble. Another, yet unexplored, possibility is that housing-price bubbles spread from one market to another. Studies looking at cointegration among regional home prices have found that price appreciation initially begins in major urban centers (called the "urban core"), then spreads to peripheral markets (Meen 1996; Oikarinen 2006). Some authors speculate that this is because business cycles originate in the urban core, then spread to peripheral areas whose economies are strongly linked to the urban core. Under this hypothesis, expansion in economic activity in the urban core generates new income. The new income may be consumed or invested, with one of the investment choices being housing in the peripheral market. New investment in the peripheral housing market will tend to

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drive up prices, all else equal. If income growth accelerates in the peripheral market as the economic expansion spreads, this will also stimulate new housing demand and push prices higher. This pathway, where the price shock in the peripheral market is primarily driven by income, I denote as "income effects."

Another possible explanation for price cointegration among markets is what I will call "price effects." Under this scenario, housing investment in the peripheral market is a substitute for investment in the urban core's market. Consider investors who have enjoyed capital gains from speculation-based home-price appreciation in the urban core. As these investors search for more investment opportunities, they find that housing in the periphery is undervalued relative to the urban core. If these investors buy in the neighboring market, assuming that prices in the two markets should converge either in absolute price or the rate of appreciation, then new demand in the peripheral market is based on the expectation of future price appreciation and not market fundamentals. This paves the way for speculative forces to flow from the urban core to the peripheral market, pushing prices above the long-run expected price.

This paper explores the effects of price appreciation and income growth in an urban core on a peripheral housing market. If price and/or income effects in the urban core influence prices in peripheral markets, I call this "price contagion." For one, I ask whether speculative price pressure can spill into related markets. I also address the influence of income growth in the urban core on housing demand in peripheral housing markets. In other words, are bubbles contagious? I disaggregate home-price appreciation in peripheral markets into that caused by income effects and that driven by price effects, controlling for other variables that affect prices for single-family homes, such as construction costs, employment, and local income.

Using the Granger representation theorem, I show that the proposed model, based on a vector error correction model (ECM), is a substantial improvement over simple pairwise cointegration models of prices. I show that under standard regularity conditions for stable ECMs, if two markets are related through either price or income effects, then price appreciation in the peripheral market will be a function of past appreciation in the core market, past investment in the peripheral market, and national and regional economic variables that influence regional supply and demand.

I apply the model to a pair of cities that experienced rapid price appreciation and dramatic subsequent decline in homes prices: Los Angeles (the urban core) and Las Vegas (the peripheral market).[1] I estimate a multiple ECM that spans 1978 Quarter 2 (Q2) to 2008 Q1 using housing data for Las Vegas and Los Angeles and national and regional economic variables thought to influence housing demand and supply.

I find evidence that income and price contagion originating in Los Angeles contributed to the rapid appreciation in Las Vegas home prices observed from 2002 to 2006. More importantly, the results reveal a market subject to severe distortions. Las Vegas prices and stocks adjust very slowly to disequilibrium. The primary drivers of Las Vegas price changes are past changes in Las Vegas prices and changes in Los Angeles prices. While prices respond modestly to changes in income in Las Vegas and Los Angeles, other important market fundamental variables, such as mortgage interest rates and construction costs, are unrelated to price movements. Taken together, the results reveal a market that was for some time driven chiefly by irrational exuberance and virtually divorced from fundamental market forces.

[1] The Las Vegas economy is strongly linked to Los Angeles. More than 50% of employment in Las Vegas is linked to tourism, and 24% of those visitors come from southern California (GLS Research 2008). The housing markets are also strongly linked. In 2004, near the peak of the Las Vegas price bubble, 39.3% of homes were purchased by absentee owners (DataQuick 2004). Clauretie and Thistle (2007) found that 45% of absentee purchasers of homes in Clark County from 2000 to 2004 gave their home address in California. Of those, 39% reported addresses in Los Angeles County. During the boom years 2003–2006, there were frequent media reports that maintained that California investors had their eye on the Las Vegas residential real estate market. See for example Haddad (2006).

II. LITERATURE ON HOME-PRICE EXPECTATIONS AND REGIONAL PRICE RELATIONSHIPS

Two strands of the housing-market literature are particularly relevant to the current discussion. The first involves the manner in which people form expectations about future home-price movements. The second is the relationship between home prices in regions with strong economic ties. I discuss each strand of the literature in turn.

Price Expectation Formation

Several studies have found that forecast error in housing prices follows an autoregressive process (Case and Shiller 1989, 1990; Tirtiroglu 1992). Others have postulated more complex forms for the forecast error, including high-order autoregressive processes coupled with moving average processes (Dolde and Tirtiroglu 1997; Riddel 1999). If the forecast error is autoregressive of any order, this indicates that buyer's expectations of future price movements are based on past price movements. Riddel (1999) calls these types of investors "feedback traders." If the error follows an autoregressive process together with a moving average process, then buyers price forecasts are based on past price movements partially correcting for mistakes they have made in previous forecasts. From a time-series perspective, this is consistent with an adaptive expectations model.

Riddel (1999) contrasts feedback traders with fundamental investors who base price forecasts on expected economic conditions in the area. These investors combine information on fundamentals market variables such as income, interest rates, and demographics to form an expected future price. This type of investor would be more likely to purchase a home when prices are low relative to expected fundamentals and sell when prices are relatively high. Using a monthly dataset for Santa Barbara County, California, spanning 1983 through 1997, she decomposes housing demand into that originating from feedback traders and that attributable to fundamental traders. She finds significant evidence of feedback trading activity that is best described by an adaptive expectations model. She finds that virtually all of the price appreciation in the 18 months preceding the 1990 peak in Santa Barbara home price can be attributed to the activity of feedback traders.

Relationships between Regional Housing Markets

Several studies have used cointegration analysis and Granger causality tests to explore the relationships between regional home prices. Alexander and Barrow (1994) test for pairwise Granger causality between 65 pairs of regional housing markets in the United Kingdom. They find Granger causality in 26 of the 132 possible pairwise Granger causality tests. In a similar approach, Slade (2006) finds evidence of Granger causality between regional Australian housing markets.

Other studies have focused on cointegration between price pairs with mixed success. Alexander and Barrow (1994) test for pairwise cointegration between the same 65 price pairs in the United Kingdom. They find cointegration in 13 of the pairs. Cook (2005) tests for cointegration using the same 65 pairs but bases his tests on threshold autoregressive processes that allow for asymmetric adjustment to market disequilibrium. Allowing markets to adjust asymmetrically to disequilibrium only modestly improves the case for cointegration of price pairs. Of the 65 pairs of housing markets tested, only 20 are found to be cointegrated assuming a test based on a significance level of 10%. When 5% is used, only 14 price pairs are found to be cointegrated.

MacDonald and Taylor (1993) take a broader approach to analyzing cointegration in the U.K. housing market from 1969 Q1 through 1987 Q4. Using Johansen's multiple cointegration techniques, they find evidence of cointegration between prices in regional submarkets. Still, their analysis is restricted to prices, and they admit that their approach is a "black box" and makes no attempt to explain the regional home-price interactions.

Several studies indirectly address the causes of home-price cointegration. Meen (1996), Berg (2002), and Oikarinen (2006) find evidence to support the hypothesis that price appreciation originates in an urban core then diffuses to peripheral markets with strong economic ties to the urban core. Oikarinen (2006) studies the relationship between regional housing markets in Finland from 1987 to 2004, using a vector ECM that explicitly accounts for long-run relationships between regional home prices. He finds that housing-price appreciation originates in Helsinki (the center of Finish economic activity), then spreads to peripheral markets. Meen (1996) uses a panel of housing starts and prices in regional markets in the United Kingdom. He finds that house-price movements are unidirectional, spreading from urban centers to the periphery. Nevertheless, all of these studies focus on cointegration among prices and fail to acknowledge exogenous variables, such as interest rates, income, and construction costs, that may affect cointegration relationships and adjustments to long-run equilibrium prices. Larraz-Iribas and Alfaro-Navarro (2008) study Spanish housing markets over the recent rapid escalation in many regional prices. They find evidence of cointegration among regional prices, with physical proximity increasing the likelihood of price cointegration.

I argue below that models based solely on prices are atheoretical and, as such, fail to adequately capture the dynamics of prices within the market structure from which they originate. Most importantly, price-only models fail to account for other endogenous market variables that are necessary for cointegration. For one, the housing stock is a fundamental endogenous component of long-run equilibrium because prices likely rise and fall in accordance with the rate of new construction. Failing to include new construction therefore ignores an important source of variation in prices and a key factor in the long-run equilibrium relationship. Pairwise price models also ignore exogenous supply and demand variables such as mortgage interest rates, income, and general economic conditions that affect the rate of price adjustment. Including new construction and exogenous variables may reveal price cointegration that is missed when only prices are considered.

In the next section, I present a model that allows for exogenous growth in income and prices in the urban core to affect house prices at the periphery. I start with a simple, nondynamic model and then extend the model into a dynamic setting. The theoretical model makes it clear that any analysis of the interaction of regional house prices based solely on price is flawed and gives biased estimates of the relationship between prices. This is because pairwise price analysis ignores the important influence of net housing investment and regional exogenous determinants of house prices.

III. CONTAGION MODEL WITH REGIONAL INCOME AND PRICE EFFECTS

According to the motivating argument, house-price contagion arises from the dynamic relationships between an urban core and peripheral housing markets. The basic idea is that home price and income in the urban core affect demand for housing in the peripheral market, but peripheral income and prices are independent of demand in the urban core. These assumptions, which I test later, imply the simple nondynamic demand and supply equations for the peripheral market:[2]

$$Q_{\mathrm{DP},t} = \alpha_0 + \alpha_{\mathrm{P}} P_{\mathrm{P},t} + \alpha_{\mathrm{PU}} P_{\mathrm{U},t} + \alpha_{\mathrm{UI}} I_{\mathrm{U},t} + \mathbf{B}_{\mathrm{D}} \mathbf{X}_{\mathrm{D},t} + \varepsilon_{\mathrm{D},t}$$

$$Q_{\mathrm{SP},t} = \beta_0 + \beta_1 P_{\mathrm{P},t} + \mathbf{B}_{\mathrm{S}} \mathbf{X}_{\mathrm{S},t} + \varepsilon_{\mathrm{S},t}, \qquad [1]$$

where $Q_{\mathrm{DP},t}$ and $Q_{\mathrm{SP},t}$ equal quantity demanded and quantity supplied, respectively, in the peripheral market at time t, $P_{\mathrm{U},t}$ and $P_{\mathrm{P},t}$ are prices in the urban core and periphery, respectively, $I_{\mathrm{U},t}$ is income in the urban core, and $\mathbf{X}_{\mathrm{DP},t}$ is a vector of exogenous demand variables, such as the number of households and income in the peripheral market and mortgage interest rates. Note that by assumption, $P_{\mathrm{U},t}$ and $I_{\mathrm{U},t}$ are also exogenous variables in peripheral demand. Finally, $\mathbf{X}_{\mathrm{SP},t}$ is a vector of exogenous supply variables such as construction costs, and $\varepsilon_{\mathrm{D},t}$ and $\varepsilon_{\mathrm{S},t}$ are white noise disturbances.

Recall that the definition of contagion is

[2] Note that bold script represents a vector.

when income or price growth in the urban core causes changes in the peripheral price. The appearance of $P_{\mathrm{U},t}$ in the demand equation recognizes that the homes in the two markets are related goods. If $\alpha_{\mathrm{PU}} > 0$, then a home in the peripheral market is a substitute for one in the urban market, whereas $\alpha_{\mathrm{PU}} < 0$ implies the goods are complements. If prices in the core don't affect peripheral demand, then $\alpha_{\mathrm{PU}} = 0$. Further, if $\alpha_{UY} > 0$, then income growth in the urban core will cause prices and quantity demanded of homes to increase in the peripheral market, ceteris paribus.

For simplicity, I group all exogenous market variables into one vector: $\mathbf{X}_t = [\mathbf{X}_{\mathrm{D},t}, P_{\mathrm{U},t}, I_{\mathrm{U},t}, \mathbf{X}_{\mathrm{S},t}]$ and the $n = 2$ endogenous variables into the 1×2 vector $\mathbf{Y}t' = [P_{\mathrm{P},t}, Q_{\mathrm{P},t}]$. By definition, price growth in the urban core (ΔP_{U}) Granger-causes price growth in the peripheral market (ΔP_{P}) if lagged values of ΔP_{U} provide statistically significant information about the future of ΔP_{P}.[3] Urban price growth is an exogenous determinant of peripheral price growth if ΔP_{U} Granger-causes ΔP_{P}, but the reverse is not true. Likewise, urban income growth is an exogenous determinant of peripheral price growth if ΔI_{U} (growth in urban income) Granger-causes ΔP_{P}, but the reverse is not true (Hamilton 1994, 303).[4] Given these assumptions, the system has a dynamic component that takes the form of a nonstationary kth-order autoregressive process with exogenous market variables $\mathbf{X}_t$ and white-noise error vector $\boldsymbol{\eta}_t$.[5] The reduced form of the system may be written as (Hamilton 1994, 579)

$$\mathbf{Y}_t = \mathbf{B}\mathbf{X}_t + \sum_{i=1}^{k} \Pi_i \mathbf{Y}_{t-i} + \boldsymbol{\eta}_t. \qquad [2]$$

By the Granger representation theorem (Hamilton 1994, 581), this system's error-correction form is

$$\Delta\mathbf{Y}_t = \mathbf{B}\mathbf{X}_t + \boldsymbol{\Gamma}_0 \mathbf{Y}_{t-1} + \sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \Delta\mathbf{Y}_{t-i} + \boldsymbol{\eta}_t, \qquad [3]$$

where $\boldsymbol{\Gamma}_0$ is a 2×2 matrix given by $\boldsymbol{\Gamma}_0 = -\mathbf{I} + \sum_{i=1}^{k} \boldsymbol{\Pi}_i$ and $\boldsymbol{\Gamma}_s = -[\boldsymbol{\Pi}_{s+1} + \boldsymbol{\Pi}_{s+2} + \ldots + \boldsymbol{\Pi}_k]$ for $s = 1,2,\ldots,k$. The error-correction representation decomposes changes in the endogenous market variables into short-run dynamics arising from autoregressive forces in the endogenous variables given by $\sum_{i=1}^{k-1} \boldsymbol{\Gamma}_i \Delta\mathbf{Y}_{t-i}$, the deterministic exogenous effects given by $\mathbf{B}\mathbf{X}_t$, and changes driven by deviation from the long-run equilibrium, $\boldsymbol{\Gamma}_0 \mathbf{Y}_{t-1}$. The latter term can be written as $\boldsymbol{\Gamma}_0 \mathbf{Y}_{t-1} = -\boldsymbol{\Phi} A' \mathbf{Y}_{t-1}$, where vector $\mathbf{A}'\mathbf{Y}_{t-1}$ defines the cointegrating equation, and the matrix $\boldsymbol{\Phi}$ with elements ϕ_n defines the speed that the nth endogenous variable adjusts to disequilibrium arising from the cointegrating vector.

The estimable reduced-form equations for price and stock growth following [3] are

$$\begin{aligned}\Delta P_{\mathrm{P},t} = {} & \mathbf{B}_{\mathrm{P}}\mathbf{X}_t + \sum_{i=1}^{k-1} \gamma_{\mathrm{PP},i} \Delta P_{\mathrm{P},t-i} \\ & + \sum_{i=1}^{k-1} \gamma_{\mathrm{QP},i} \Delta Q_{\mathrm{P},t-i} + D_{\mathrm{PU}} \Delta P_{\mathrm{U},t} \\ & + D_{\mathrm{PI}} \Delta I_{\mathrm{U},t} + \phi_1 \mathbf{A}'\mathbf{Y}_{t-1} + \eta_{\mathrm{P}},\end{aligned} \qquad [4a]$$

$$\begin{aligned}\Delta Q_{\mathrm{P},t} = {} & \mathbf{B}_{\mathrm{Q}}\mathbf{X}_t + \sum_{i=1}^{k-1} \gamma_{\mathrm{PQ},i} \Delta P_{\mathrm{P},t-i} \\ & + \sum_{i=1}^{k-1} \gamma_{\mathrm{QQ},i} \Delta Q_{\mathrm{P},t-i} + D_{\mathrm{QU}} \Delta P_{\mathrm{U},t} \\ & + D_{\mathrm{QI}} \Delta I_{\mathrm{U},t} + \phi_2 \mathbf{A}'\mathbf{Y}_{t-1} + \eta_{\mathrm{Q}},\end{aligned} \qquad [4b]$$

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[3] Formally, ΔP_{U} does not Granger-cause ΔP_{P} if $MSE(\Delta P_{\mathrm{P},t+s} | \Delta P_{\mathrm{P},t}, \Delta P_{\mathrm{P},t-1}, \ldots)] = MSE[\hat{E}(\Delta P_{\mathrm{P},t+s} | \Delta P_{\mathrm{P},t}, \Delta P_{\mathrm{P},t-1}, \ldots, \Delta P_{\mathrm{U},t}, \Delta P_{\mathrm{U},t-1}, \ldots,)]$. In other words,is not informative about the future of ΔP_{P} (Hamilton 1994, 302–303).

[4] In a regional model, employment, income, the number of households, and wages could also be endogenous variables. For the following model development, I focus on prices and stocks to simplify the presentation. I discuss adding an additional endogenous variable, employment, in the presentation of the empirical model.

[5] A number of studies including those by Riddel (2000), Malpezzi (1999), and Alexander and Barrow (1994), as well as this study (as shown later) fail to reject the null hypothesis of nonstationarity in regional house-price indices for the United States and United Kingdom. Thus, I model house prices as nonstationary.

where

$$\gamma_{QQ,i} = \frac{\partial \Delta Q_{P,t}}{\partial \Delta Q_{P,t-i}},\ \gamma_{PQ,i} = \frac{\partial \Delta Q_{P,t}}{\partial \Delta P_{P,t-i}},$$

$$\gamma_{QP,i} = \frac{\partial \Delta P_{P,t}}{\partial \Delta Q_{P,t-i}},\ \gamma_{PP,i} = \frac{\partial \Delta P_{P,t}}{\partial \Delta P_{P,t-i}}$$

are elements of $\mathbf{\Gamma}_i$ from equation [3], and the error terms η_P and η_Q are uncorrelated white-noise processes. Equation [4a] has three components. The short-run dynamics, $\sum_{i=1}^{k-1} \gamma_{PP,i}\Delta P_{P,t-i} + \sum_{i=1}^{k-1} \gamma_{QP,i}\Delta Q_{P,t-i}$, describe the short-run effects of changes in peripheral prices and housing stocks on peripheral price appreciation. The contemporaneous effects of the exogenous variables on price changes are given by $\mathbf{B}_P\mathbf{X}_t + D_{PU}\Delta P_{U,t} + D_{PI}\Delta Y_{U,t}$. Finally, the cointegrating vector $\mathbf{A}'\mathbf{Y}_{t-1}$ determines the long-run equilibrium relationship between prices and stocks. If the system is in long-run equilibrium, $\mathbf{A}'\mathbf{Y}_{t-1} = 0$. When $\mathbf{A}'\mathbf{Y}_{t-1} \neq 0$, prices deviate from long-run expectations but return to equilibrium at rate ϕ_1. Large numbers for ϕ_1 and ϕ_2, in absolute value, suggest relatively rapid price and stock adjustment, respectively, whereas smaller values indicate a slower adjustment to equilibrium.

Notably, peripheral price adjustment is affected by past own-price and own-stock changes, changes in prices and income in the urban core, as well as the extent of deviation from the long-run equilibrium. If $\gamma_{QQ,i} \neq 0$ and/or $\gamma_{QP,i} \neq 0$, then excluding own-stocks from the ECM ignores the importance of new housing investment not only in the long-run equilibrium, but in the price-adjustment mechanism. This makes clear why simply looking at regional price pairs will give a biased view of price adjustment in the peripheral market. If price adjustment is a function of past stock adjustments, as we would expect if they arise from an underlying demand and supply model, then excluding stocks from the model will mean that price adjustments arising from new housing investment will be attributed to past price changes. This is classic omitted variable bias, where bias is observed in the coefficient of any variable included in the model that is correlated with the omitted variable.

An interesting question is what induces investment flows from the urban core into the peripheral real estate market. As stated earlier, others have found house prices that are best described by an adaptive expectations model. According to this finding then, investors in Los Angeles and Las Vegas both use backward-looking models to forecast prices in their home markets.

Buyers have information about past appreciation and depreciation in home prices in their local market, typically through watching changes in assessed value and/or sale prices for their own homes and those of their neighbors and friends. However, buyers' information sets are different when they attempt to forecast price movements outside their local housing market. Given that a Los Angeles investor decides to purchase in Las Vegas, it is difficult to assess the extent of his forecast error using standard time-series approaches. It could be he bases his forecast of nonlocal future price in a peripheral market on his forecast of local price under the assumption that the economic linkages should cause prices to rise at similar rates. It could also be that he forecasts price growth in the nonlocal but economically dependent market based on price differences between the local and nonlocal markets. Or, it could be that his is selling his home and using capital gains to purchase a similar, but cheaper home in Las Vegas with the intention of moving. Unfortunately, a time-series model is unable to distinguish between these three types of forecasts. Rather than trying to explicitly model nonlocals forecast of price, the paper explores the extent that income and home-price growth in Los Angeles influence home-price growth in Las Vegas.

Equations [4a] and [4b] clarify the mechanism by which a bubble can spill from one market into a related market. Assume a positive shock occurs to demand in the urban housing market in period $t = 0$. The demand shock causes prices in the urban market to rise. If $D_{PU} > 0$ and $D_{QU} > 0$ (the coefficients of urban price changes in the price and stock adjustment equations, respectively), this causes a contemporaneous increase in prices and stocks in the peripheral market. Since price growth in the peripheral market is a function of the previous period's own price and stock growth, positive autoregressive forces in the short-run dynamics amplify the initial shock, paving the way for peripheral-market price to rise well above the long-run equilibrium over a sustained period. If the markets are characterized by slow adjustment, it may take many periods for the price to completely reequilibrate the system. Whether a price bubble forms is determined by the speed the system adjusts to shocks and the sign and magnitude of the autoregressive short-run dynamics. A similar story can be told with respect to income shocks to the urban market. If $D_{\mathrm{PI}} > 0$ and $D_{\mathrm{QI}} > 0$, then income shocks in the urban core may initiate a bubble in the peripheral market in the same way.

Nevertheless, "bubble" prices that are high relative to the long-run equilibrium are not sustainable and must eventually fall to realign the market with long-run fundamentals. Depending on the short-run dynamics, the downward adjustment response may be amplified, forcing the previously elevated price to fall well below fundamentals. In other words, the bubble collapses, often bringing prices below the long-run expected price. In the absence of new shocks, price will eventually stabilize at a new long-run equilibrium that may be above or below the initial equilibrium, depending on any deterministic trends in the data.

Because of the complexity of the short and long-run dynamics implied by even relatively simple ECMs, it is difficult to predict the long-run effects of shocks on system variables by simply examining the system parameters. Instead, I rely on the impulse response functions to give insight into how price and quantity in the peripheral market respond to shocks over time. Impulse response functions help distinguish between scenarios where prices quickly return to their initial level following some perturbation and those where the perturbation effect persist in the form of higher (or lower) prices over time.

Define $\boldsymbol{\Omega}_{t+s}$ as a 2×1 orthogonalized impulse response matrix with elements (Hamilton 1994, 320)

$$\omega^{P_{\mathrm{P}}}_{P_{\mathrm{P},t+s}} = \frac{\partial E[P_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial P_{\mathrm{P},t}}, \qquad [5a]$$

$$\omega^{Q_{\mathrm{P}}}_{P_{\mathrm{P},t+s}} = \frac{\partial E[P_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial Q_{\mathrm{P},t}}, \qquad [5b]$$

$$\omega^{P_{\mathrm{P}}}_{Q_{\mathrm{P},t+s}} = \frac{\partial E[Q_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial P_{\mathrm{P},t}}, \qquad [5c]$$

$$\omega^{Q_{\mathrm{P}}}_{Q_{\mathrm{P},t+s}} = \frac{\partial E[Q_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial Q_{\mathrm{P},t}}. \qquad [5d]$$

For variables in logs, the impulse response $\omega^{k}_{n,t+s}$ gives the percent change in endogenous variable n at time $t+s$ for a 1% change in endogenous variable k at time t, given all past information about the endogenous variables.

To determine the long-run effect of a change in an *exogenous* variable on price, note that a 1% increase in an exogenous variable has a contemporaneous effect on the endogenous variables equal to the coefficient of the exogenous variable in the price or stock adjustment equation [4], namely, $\frac{\partial \Delta P_{\mathrm{P},t}}{\partial \Delta P_{\mathrm{U},t}} = D_{\mathrm{PU}}$. As a result, an increase in an exogenous variable affects the endogenous variables this period and is perpetuated according to the net impulse response function. Thus the effect of a change in an exogenous variable r at time t on an endogenous variable at time $t+s$ is

$$\Psi^{r}_{P_{\mathrm{P},t+s}} = \frac{\partial E[P_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial X_{r,t}} = B_{r,P}\, \omega^{P_{\mathrm{P}}}_{P_{\mathrm{P},t+s}} + B_{r,Q}\, \omega^{Q_{\mathrm{P}}}_{P_{\mathrm{P},t+s}}, \qquad [6a]$$

$$\Psi^{r}_{Q_{\mathrm{P}},t+s} = \frac{\partial E[Q_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial X_{r,t}} = B_{r,P}\,\omega^{P_{\mathrm{P}}}_{Q_{\mathrm{P}},t+s} + B_{r,Q}\,\omega^{Q_{\mathrm{P}}}_{Q_{\mathrm{P}},t+s}, \quad [6b]$$

where $B_{r,P}$ and $B_{r,Q}$ are the coefficients of exogenous variable r in the price and stock adjustment equations, respectively. For example, the effect of a shock to urban price on peripheral price is

$$\frac{\partial E[P_{\mathrm{P},t+s} \mid P_{\mathrm{P},t}, P_{\mathrm{P},t-1}, \ldots, Q_{\mathrm{P},t}, Q_{\mathrm{P},t-1}, \ldots]}{\partial P_{\mathrm{U},t}} = D_{\mathrm{PU}}\omega^{P_{\mathrm{P}}}_{P_{\mathrm{P}},t+s} + D_{\mathrm{QU}}\omega^{Q_{\mathrm{P}}}_{P_{\mathrm{P}},t+s}.$$

The impulse response functions for the exogenous variables are essentially marginal effects because they reflect changes to the endogenous variable at time $t + s$ for a change in an exogenous variable at t. The total effect of the change is given by the cumulative impulse response function, which is simply the sum of response through $t + s$. In other words, the long-run effect of an increase in some variable on the response variable is the accumulated short-run responses.

IV. THE EMPIRICAL MODEL

The Data

I model the period from 1978 Q2 through 2008 Q1, focusing on the response of Las Vegas single-family home prices and stocks to innovations in Los Angeles house price and income. In a regional economy, home prices, the housing stock, employment, income, and the number of households are typically thought of as endogenous. Unfortunately, the number of households is not available on a quarterly basis and is typically highly correlated with employment. Thus, I use the employment variable to capture both employment and household-formation effects in the housing market.

One of the challenges of error-correction modeling is choosing a parsimonious set of endogenous variables that contains enough market information to allow estimation of the cointegrating equations. I conduct a series of Granger causality tests to determine which of the variables can be treated as exogenous. I motivate the argument for contagion effects based on the idea that economic growth, in terms of growth in income and housing demand, originates in the urban core (Los Angeles) and spreads to peripheral markets (Las Vegas). If this is true, then we should find that Los Angeles income and home-price growth are exogenous determinants of Las Vegas housing price and stock growth. Granger causality tests, reported in Table 1, support this hypothesis. The tests indicate that growth in Los Angeles home prices Granger-cause growth in Las Vegas home prices (LV PRICE) and stocks (LV STOCK), but the reverse is not true. As a result, it is appropriate to include Los Angeles home prices as an exogenous variable in the Las Vegas ECM. The results also indicate that growth in Los Angeles income Granger-causes growth in Las Vegas prices and stocks, so it is appropriate to model Los Angeles income as exogenous.

The Granger causality tests also indicate that Las Vegas income growth Granger-causes Las Vegas price and stock growth, but the converse is not true. Thus, I assume Las Vegas income growth is an exogenous variable in the ECM. I treat employment as an endogenous variable because employment growth Granger-causes stock growth and vice versa.

With these considerations in mind, the endogenous variables in the model are Las Vegas single-family home prices, employment, and the stock of single-family homes. All variables are in natural logarithms. I average the Clark County, Nevada, monthly employment figures from the Bureau of Labor Statistics[6] (BLS) to estimate quarterly employment for Las Vegas (LV EMP). Quarterly home-price indices published by the Office of Federal Housing and Enterprise Oversight[7] (OFHEO) are used to represent single-family home prices for Las Vegas (LV PRICE). The Las Vegas housing stock (LV STOCK) is con-

[6] www.bls.gov.
[7] www.fhfa.gov.

TABLE 1
Pairwise Granger Causality Tests for Exogeneity

Null Hypothesis:[a]	F-Statistic	Probability
D(LV PRICE) does not Granger-cause D(LA PRICE)	1.8028	0.1696
D(LA PRICE) does not Granger-cause D(LV PRICE)	7.1595	0.0012
D(LV STOCK) does not Granger-cause D(LA PRICE)	1.0491	0.3868
D(LA PRICE) does not Granger-cause D(LV STOCK)	2.7076	0.0354
D(LV STOCK) does not Granger-cause D(LA INCOME)	0.9317	0.4495
D(LA INCOME) does not Granger-cause D(LV STOCK)	4.9369	0.0012
D(LV PRICE) does not Granger-cause D(LA INCOME)	0.7961	0.4536
D(LA INCOME) does not Granger-cause D(LV PRICE)	2.8318	0.0631
D(LV PRICE) does not Granger-cause D(LV INCOME)	1.1975	0.3058
D(LV INCOME) does not Granger-cause D(LV PRICE)	4.9791	0.0085
D(LV STOCK) does not Granger-cause D(LV INCOME)	1.6159	0.2043
D(LV INCOME) does not Granger-cause D(LVSTOCK)	5.0706	0.0081
D(LV EMP) does not Granger-cause D(LV STOCK)	5.2004	0.0070
D(LV STOCK) does not Granger-cause D(LV EMP)	19.7986	0.0000
D(LV PRICE) does not Granger-cause D(LV EMP)	1.3475	0.2644
D(LV EMP) does not Granger-cause D(LV PRICE)	1.0962	0.3380
D(LV STOCK) does not Granger-cause DLVPRICE	20.8939	0.0000
D(LV PRICE) does not Granger-cause D(LV STOCK)	2.9979	0.0234

[a] D( ) denotes the one-period lag operator. Number of lags equals 4.

TABLE 2
Tests for Unit Root: Endogenous Variables

Variable	ADF Statistic	Lag Length[a]	p-Value	Conclusion
LV PRICE[b]	−0.0281	9	0.9533	Nonstationary
LV STOCK[c]	−1.9500	12	0.6205	Nonstationary
LV EMP[c]	−0.9270	2	0.9483	Nonstationary

Note: The series is inferred to be an I(0) process if the null hypothesis of an I(1) process is rejected using a 5% decision rule. ADF, augmented Dickey-Fuller.
[a] Lag length determined using the Schwartz information criterion.
[b] Test based on series with a constant term.
[c] Test based on series with a constant and deterministic trend.

structed by taking the U.S. Census estimate of the number of homes in 1980 Q1 and adding or subtracting census-based quarterly housing starts for single-family homes each quarter. If this method over- or underpredicts stocks for the next census (1990 or 2000), I rebase the number in the census years and subtract or add the same number of new homes from each of the noncensus years so that the census count matches the estimated count in the census years. Augment Dickey-Fuller tests, summarized in Table 2, indicate that all of these series are nonstationary with a constant and deterministic trend.

Exogenous variables include a national construction cost index (BUILDCOST), national mortgage interest rates (MORT), Los Angeles house price (LA PRICE), estimated quarterly per capita income for the Las Vegas MSA (LV INCOME), and estimated quarterly per capita income for the Los Angeles Metropolitan Statistical Area (MSA) (LA INCOME). All exogenous variables save the mortgage interest rate are in natural logarithms, and all prices and interest rates are adjusted using the GDP deflator. MORT is the national average rate on conventional 30-year mortgages taken from the St. Louis Federal Reserve's ALFRED database.[8] BUILDCOST is the Bureau of Labor Statistics' construction cost index. Quarterly income and population

[8] http://research.stlouisfed.org.fred2/.

TABLE 3
Trace Test and Maximum Eigenvalue Tests for the Number of Cointegrating Relations; Endogenous Series: LV PRICE, LV EMP, and LV STOCK 1978 Q2–2008 Q1

Hypothesized No. of CEs	Trace Statistic	*p*-Value	Max. Eigenvalue Statistic	*p*-Value
None	53.0483	0.0000	35.9763	0.0002
At most 1	17.0721	0.0287	15.7852	0.0285
At most 2	1.2869	0.2566	1.2869	0.2566

Note: Trace test and maximum eigenvalue tests suggests two cointegrating vectors at the 0.05 level of significance. Statistics based on a linear deterministic trend and lag length equal to 4. CE, cointegrating equations.

TABLE 4
Two Cointegrating Equations: Coefficients and Adjusted *t*-Statistics in Brackets

Dependent Variable	LVPRICE	LVEMP	LVSTOCK	CONSTANT
μ_{price}	1	0	2.4622	−35.4298
	—	—	[3.008]	—
μ_{emp}	0	1	−0.8064	3.7252
	—	—	[−12.2982]	—

data are not available at the county or MSA level. I construct the per capita income series for the Los Angeles and Las Vegas MSAs based on annual personal income available from the Bureau of Economic Analysis and U.S. Census annual population estimates.[9] I transform annual personal income and population into quarterly series based on the assumption that population and income grow at a constant quarterly rate. Given the quarterly estimates, per capita income for each of the MSAs is calculated as the ratio of personal income to population.

Results

I establish the number of cointegrating vectors based on Johansen's (1988) multiple-error correction techniques. Table 3 gives the results of the trace and maximum eigenvalue tests for the number of cointegrating relationships assuming the set of endogenous variables is LV PRICE, LV EMP, and LV STOCK. Both tests agree that two cointegrating relationships are present in the data. Table 4 reports those relationships for the Phillip's normalization with the coefficients of LV PRICE and LV EMP normalized to 1 and 0, respectively, in the first cointegrating equation, and 0 and 1, respectively, in the second equation. The estimated coefficients are significant at the 5% level. Levin, Lin, and Chu's test (not reported) rejects the null hypothesis of a common unit root in the two cointegrating vectors.

The first cointegrating equation defines a negative long-run relationship between price and stocks, so that an increase in stocks is correlated with a decrease in prices in the long run. Thus, a positive value for the disequilibrium error suggests that prices are high relative to stocks and must eventually fall to reequilibrate the system. The second equation defines the long-run relationship between Las Vegas stocks and employment. Las Vegas stocks are positively related to employment, thus a positive value for the disequilibrium error suggests that stocks are low relative to the long-run equilibrium employment, suggesting that either employment must fall or stocks rise to reequilibrate the system.

In initial model runs, I included the log of the quarterly average of the S&P500 stock index and real U.S. GDP as exogenous variables in the ECM. Neither variable was significant, so they are excluded from the final model.

Table 5 reports the speeds of adjustment and their *t*-statistics for the stock, price, and employment adjustment equations of the ECM, as well as the coefficients and *t*-statistics for the exogenous variables. The model results for the short-run dynamics ($\gamma_{PP,i}$, $\gamma_{PQ,i}$, $\gamma_{QQ,i}$, and $\gamma_{QP,i}$ from equations [4a] and [4b]) are reported in the Appendix. The Akaike information criterion and the Schwartz information criterion support a lag length of four quarters.

The response of prices to disequilibrium and contemporaneous changes in the exogenous variables are given in Column 1 of Table 5. Over 80% of the variation in prices is explained by the regression. Notably, price adjusts very slowly to price disequilibrium (μ_{price}): only 1.5% of the adjustment takes

[9] www.bea.gov and www.census.gov.

TABLE 5
Error-Correction Model: Coefficients with t-Statistics in Brackets

Error Correction:	D(LV PRICE)	D(LV STOCK)	D(LV EMP)
Speeds of adjustment			
μ_{price}	−0.0152 [−4.9583]	−0.0031 [−1.6529]	−0.0106 [−2.5638]
μ_{emp}	−0.2018 [−5.3659]	−0.0005 [−0.0203]	−0.1032 [−2.0495]
Exogenous variables			
C	0.0082 [1.5737]	0.0094 [2.5672]	0.0167 [2.3966]
LA PRICE	0.6391 [8.1631]	0.1110 [2.0170]	0.0764 [0.72890]
MORT	−0.0100 [−0.3398]	0.0017 [0.0827]	0.0717 [1.8221]
BUILDCOST	0.0163 [0.1203]	0.1411 [1.4841]	0.0293 [0.1620]
LV INCOME	−0.1123 [−0.8805]	−0.0672 [−0.7495]	−0.2121 [−1.7721]
LA INCOME	0.0392 [0.2837]	0.1681 [1.7318]	0.0292 [0.1578]
R^2	0.8175	0.8073	0.5269
Adjusted R^2	0.7633	0.7501	0.3865
Akaike information criterion	−6.0116	−6.7174	−5.4279
Schwarz information criterion	−5.4329	−6.1386	−4.8491

Note: Only adjustment coefficients and coefficients of exogenous variables reported. See Appendix for coefficients and t-statistics for the short-run dynamics.

place in the first quarter.[10] The response is faster to changes in μ_{emp}, with about 20% of disequilibrium offset in the first quarter. Of the exogenous variables, only the Los Angeles home-price index is shown to have a significant contemporaneous impact on Las Vegas home-price growth.

The stock responses to disequilibrium and contemporaneous changes in the exogenous variables are given in Column 2 of Table 5. Stocks adjust to disequilibrium in μ_{price}, but not μ_{emp}. Stock adjustment is strikingly slow: after one quarter, less than 1% of the adjustment takes place. Of the exogenous variables, Los Angeles prices, the construction cost index, and Los Angeles income growth gave a statistically significant (p-value < 0.1) and positive contemporaneous impact on stock growth.

Employment responds slowly to price and employment disequilibrium at a rate of 1% and 10% each quarter, respectively. An increase in mortgage interest rates has a positive and statistically significant impact on employment. The other exogenous variables are not statistically significant. The equation has much less explanatory power than the price or stock adjustment equations. This is not particularly surprising, since the aim of the model is to predict house price and stock changes, and I therefore do not include variables, such as wages, that might improve the equation significantly but would have no theoretical impact on the housing market.

As noted before, the ECM is a dynamic model, and the marginal effects of a change in any of the model variables are not fully captured by the model coefficients. Unlike a static model, the coefficients of the exogenous variables give only the contemporaneous impact of the exogenous variable on the endogenous variable. To understand the long-run effects of a change in any of the model variables on prices and stocks, we turn to the impulse response functions. For the sake of brevity, I focus the discussion on the impulse response functions for price and stock changes. Discussing the employment impulse response functions adds nothing to the discussion of the housing market but would take considerable additional space.

[10] It would be interesting to compare these adjustment rates to those of other studies, but this is difficult. It is very unusual for studies to address housing markets using ECM models, and when they do, they typically rely on a very limited set of variables such as stocks and prices only. The one study that is comparable in that it includes a comprehensive set of housing-market variables is by Riddel (2004). Although the estimation technique is quite different, that study found that national housing prices adjust at a rate of 65% annually to price disequlibrium. This is much faster than the current model. However, the national home-price series does not contain any marked periods of bubble activity like that observed in Las Vegas.

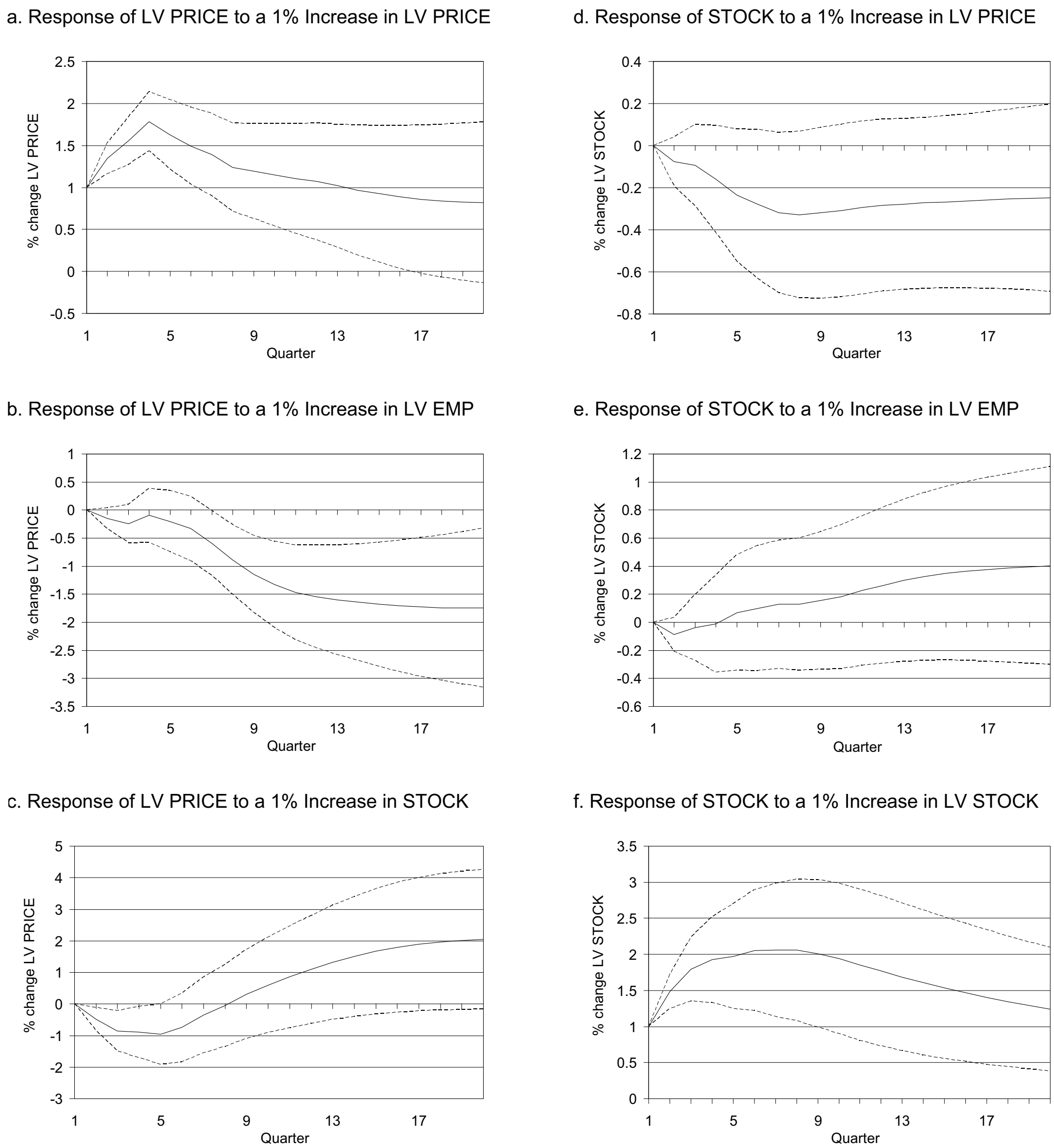

FIGURE 1
Impulse Response Functions and 95% Confidence Intervals for Endogenous Variables: Response of Endogenous Variable at Time s to a 1% Increase in the Endogenous Variable at Time $t = 0$

Impulse Response Functions for Endogenous Variables

Recall that the impulse response functions for the endogenous and exogenous variables $\omega^k_{n,t+s}$ and $\Psi^r_{n,t+s}$, respectively, give the percent change in the endogenous variable n at time $t + s$ for a 1% increase in endogenous variable n or exogenous variable r, respectively, at time t. Figure 1a–c shows the impulse response function for Las Vegas house price at time s for a 1% increase in each of the three endogenous variables at $t = 0$. The graphs also give a 95% confidence interval

TABLE 6
Cumulative Response of Las Vegas Prices and Stocks after Eight Quarters to a 1% Increase in Each Variable

	Response Variable (2-Year Response)[a]	
1% Change In:	%Δ LV PRICE	%Δ LV STOCK
LA PRICE	6.83%	NS
LA INCOME	NS	2.41%
BUILDINDEX	NS	NS
MORT	NS	NS
LV PRICE	11.43%	NS
LV INCOME	NS	2.57%
LV STOCK	−0.49%	14.35%
LV EMP	−1.47%	NS

[a] NS indicates impulse response not significant for $\alpha = 0.05$.

around each of the functions.[11] Table 6 gives cumulative impulse responses for endogenous and exogenous variables.

The everywhere positive price impulse response function in Figure 1a indicates significant persistence in the Las Vegas home-price series. A 1% increase in Las Vegas home prices in the initial period leads to 1.75% increase in the fourth quarter following the shock. The marginal effect declines then stabilize at just under 1% after 4 years, at which point the confidence interval widens enough that the effect is statistically indistinguishable from a zero marginal effect. The cumulative effect of a 1% increase in price after 2 years is an 11.43% increase in price.

According to Figure 1b, a 1% increase in employment has a negative impact on future prices in Las Vegas, but the effect is not statistically significant until the sixth quarter after the initial employment shock. At first blush, this is somewhat counterintuitive. As employment grows, so do the number of households, which should cause prices to increase. However, the impulse response functions are reduced form responses and, as such, include demand and supply forces. On the supply side, more employment means more construction workers. During the period under study, 12% of total nonfarm employment was in the construction sector. More construction workers reflect the expansion in residential construction activity. All else equal, new construction tends to lower prices. Apparently, this effect is dominating the upward pressure on prices resulting from new household formation, giving a net effect that is negative.

According to Figure 1c, a 1% increase in stocks has a negative and significant effect on Las Vegas prices for four quarters. The cumulative response indicates that a 1% increase in stocks at time $t = 0$ will result in a 2.23% drop in price after a year, when the marginal effect becomes statistically insignificant.

Figure 1d–f gives the impulse response function for Las Vegas stocks at time s for a 1% increase in each of the three endogenous variables at $t = 0$. Las Vegas housing stock does not exhibit a statistically significant response to changes in either prices or employment. However, Figure 1f indicates statistically significant persistence in the Las Vegas housing stock series. A 1% increase in stocks at $t = 0$ generates a marginal stock increase of just over 2% after 2 years, after which the response falls modestly to about 1.25% after 4 years.

Impulse Response Functions for Exogenous Variables

The response of Las Vegas home prices to innovations in the exogenous variables are given in Figure 2 together with 95% confidence intervals. Shocks to the construction cost index (BUILDINDEX), Los Angeles income, and mortgage interest rates do not result in statistically significant changes in Las Vegas prices.

Of primary interest in this analysis is the price contagion effect. The impulse response function for Los Angeles price supports the hypothesis that growth in Los Angeles prices contributed to Las Vegas price growth. The marginal effect is at just over 1% in the fourth quarter following the shock, then declines and stabilizes at around 0.8%. The cumulative effect after 4 years is a 13.2% increase in Las Vegas home prices. Las Vegas prices exhibit a lagged response to changes in Las Vegas

[11] Note that cointegration implies that a new equilibrium will be achieved in response to system shocks, and the inherent nonstationarity of the variables means that shocks permanently affect endogenous variables. Thus, unlike those for stationary variables, the impulse response functions do not necessarily go to zero over time.

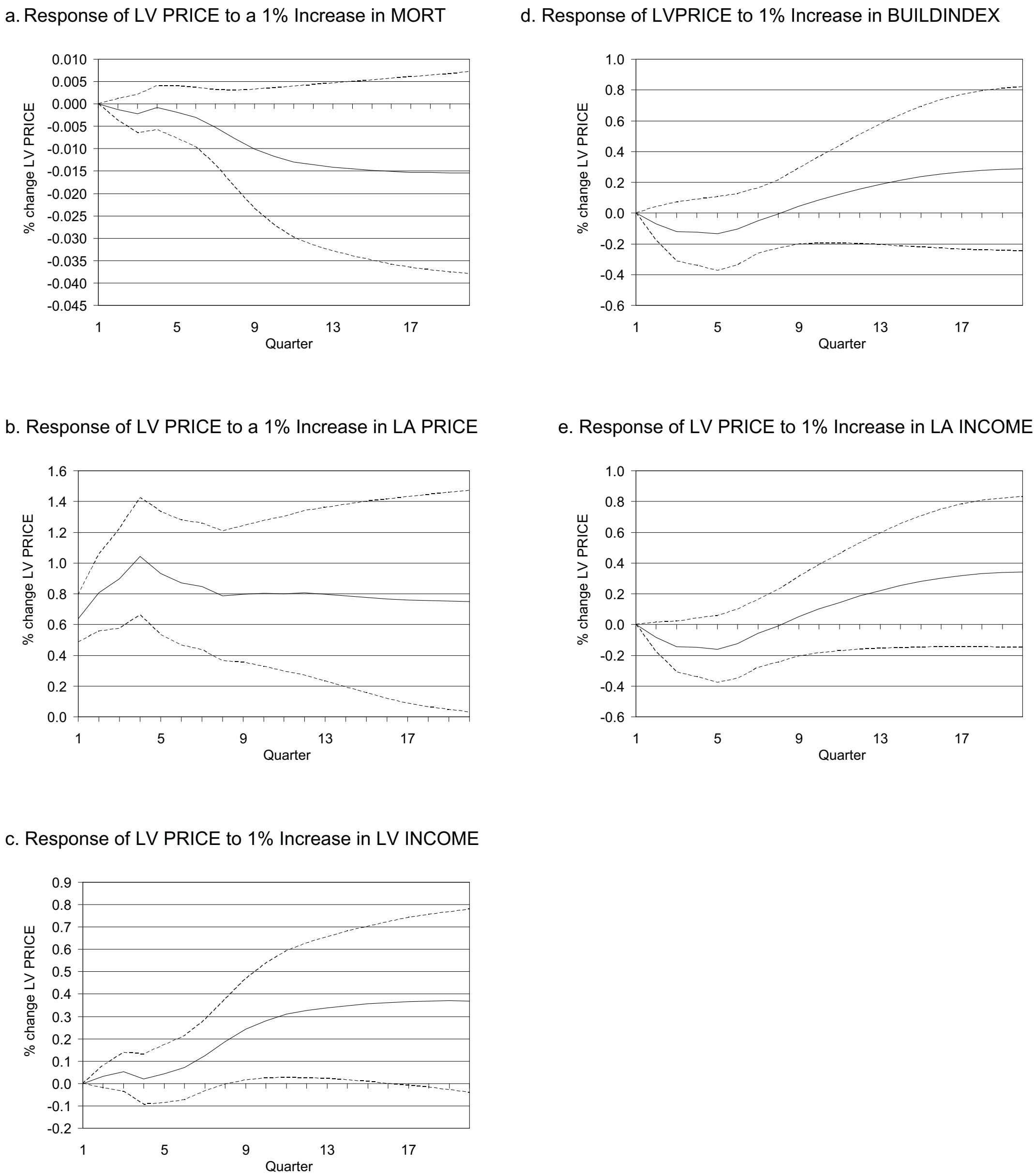

FIGURE 2
Price Impulse Response Functions and 95% Confidence Intervals for Exogenous Variables: Response of Las Vegas Housing Price at Time s to a 1% Increase in the Exogenous Variable at Time $t = 0$

income. The impulse response function is positive and significant for quarters 8 through 15. Accordingly, a 1% increase in Las Vegas income at $t = 0$ will cause prices to rise by slightly more than 0.25% after 2 years. The total cumulative effect of the income change after the fifteenth quarter is a 2.6% increase in Las Vegas home price.

As expected, a 1% positive shock to the mortgage interest rate is followed by a decline in Las Vegas price. The effect is significant over the 5 years of the calculated impulse response functions. The cumulative effect of a 1% increase in the mortgage rate is a roughly 1% drop in price over 4 years.

Figure 3 gives the response of stock to a 1% increase in each of the five exogenous variables. All of the stock impulse response functions are statistically insignificant, save for the Los Angeles income. According to Figure 3e, a 1% increase in Los Angeles income leads to a roughly 0.3% increase in stocks in each of the subsequent eight quarters, after which the marginal effect becomes insignificant.

V. THE LAS VEGAS BUBBLE: HOUSE-PRICE CONTAGION OR FUNDAMENTALS?

The model results suggest a housing market that is subject to serious distortions. Notably, these distortions go well beyond slow adjustment to equilibrium that others have explained by inefficiencies associated with high transaction and search costs and asymmetric information about home quality (DiPasquale and Wheaton 1994; Riddel 2004). Rather, this model suggests that prices adjust extremely slowly (less than 2% per quarter) to price disequilibrium. Although price adjustment to employment disequilibrium is faster, it is still quite slow, with a quarterly adjustment rate of 20%. Taken together, these findings indicate that prices are not responding to fundamental forces.

The question is: What *is* driving price movements if not fundamental market forces? The answer lies in the impulse response functions. Notably, the primary driver of current price movements is past price movements. This provides strong evidence of the feedback trading effect described by Riddel (1999). Apparently, a substantial share of Las Vegas housing demand over the period of study came from feedback traders, in other words, those who forecast price based on past price and ignored fundamental forces, such as costs, population, income, and mortgage interest rates.

Very little of the price adjustment is driven by fundamental market variables. Las Vegas prices don't respond to changes in income, mortgage interest rates, or construction costs, and the response to changes in stocks and employment (the model's proxy for the number of households) is very small. In fact, the second-most influential determinant of Las Vegas price growth is Los Angeles price growth. The long-run (2-year) effect of a change in Los Angeles prices on Las Vegas prices is 4.6 times the impact of a change in employment and 14 times that of a change in stocks. Of course, Los Angeles prices should affect Las Vegas prices if homes in the two regions are substitutes, but it is difficult to believe that market forces alone would lead to such substantial substitution effects that dominate the effects of other market variables.

Stock growth also appears to ignore normally important market fundamental variables. Adjustment to price disequilibrium is inordinately slow, with less than 1% adjustment per quarter. Even more importantly, stocks do not respond at all to employment disequilibrium. Like prices, most of the growth in stocks can be explained by positive autoregressive forces. The only other statistically significant determinant of stock growth is income growth in Los Angeles. This is highly suggestive of construction that is based on past construction rates and not fundamental forces such as household formation. Construction that is motivated by autoregressive forces rather than fundamentals can lead to overbuilding, thereby setting the stage for the subsequent collapse in the price bubble.

The finding that stocks do not respond to prices is interesting, but not inconsistent with the notion of a bubble. The data suggest that firms simply could not build fast enough to satisfy speculative demand. Construction did, in fact, increase dramatically during the bubble years 2003–2005. For example, in 2005, just prior to the collapse of the bubble, 39,000 permits were issued for single-family units in Clark County. During that year, the population grew by an estimated 68,675 people. Assuming 2.5 people per household, this would suggest new demand of roughly 27,400 new households (Metropolitan Research Association 2008). A back of the envelope calculation suggests that this was more than 11,000 units in excess of those warranted by changes in

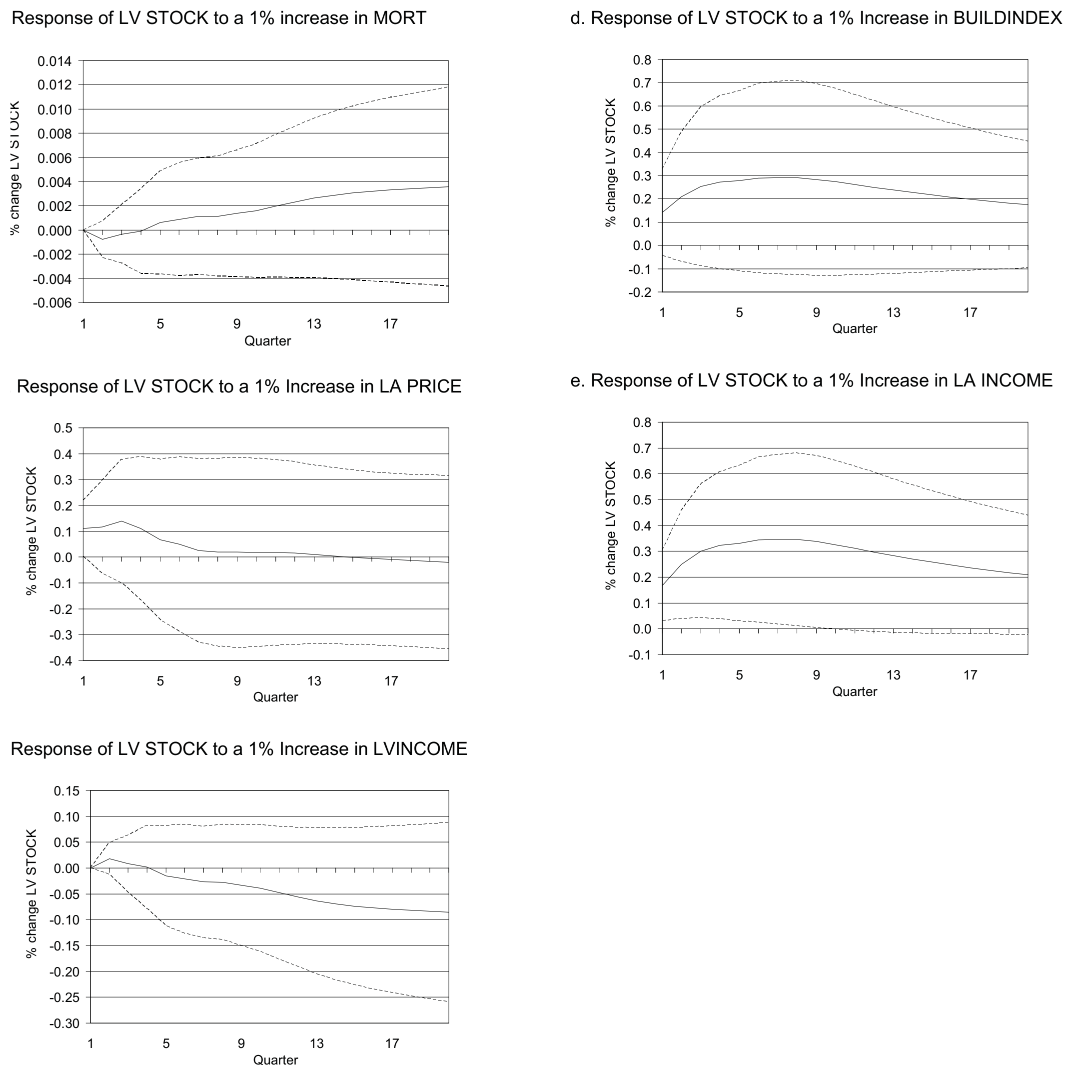

FIGURE 3
Stock Impulse Response Functions and 95% Confidence Intervals for Exogenous Variables: Response of Las Vegas Housing Stock at Time s to a 1% Increase in the Exogenous Variable at Time $t = 0$

fundamental demand conditions. In 2004, price growth peaked at an annual rate of 53%. Although builders overbuilt in retrospect, it was not possible for them to build enough homes to dampen price growth, since prices were growing at such an astounding rate.

One obvious interpretation is that the model simply has little explanatory power, and adding more years of data would show that market forces are, in fact, significant. But this interpretation is not supported by the relatively high R-square, which shows that over 80% of the stock and price variations is explained by the model. The other interpretation is that fundamental forces were not operative over the period of study. Instead, feedback trading dominated the market, pushing prices well above the long-run equilibrium values.

Market forces failed to establish a new fundamental equilibrium. Instead, rapid price appreciation encouraged builders to expand new construction, and the stock of units quickly overshot the long-run fundamental-based demand for units.

To investigate whether speculative activity undermined the effect of fundamental forces on price and stock changes, I reestimated the model excluding the "bubble" years 2002 Q1–2008 Q1. The results are interesting and generally support the conclusion that speculation dominated fundamental forces during the bubble years. For one, the speeds of adjustment are markedly faster in the truncated model, with price adjusting to price equilibrium at 10% per quarter, 10 times faster than in the full model, and stocks adjusting to price disequilibrium at 16% per year compared to less than 1% per year observed in the full model. Also, Los Angeles income does not have a contemporaneous positive impact on price growth as it does in the full model. Finally, the coefficient of Los Angeles price in the price adjustment equation is significantly larger in the full model compared to the truncated model. Taken together, these results suggest that speculative activity was active in the market during the bubble years. Prices became less responsive to market forces, and the Los Angeles home price and income became more influential in determining price and stock growth.

Given the magnitude of price persistence revealed by the model, it is not surprising that once price growth began to accelerate, a bubble formed quite rapidly. But what was the precipitating event that led to the initial shock to Las Vegas home prices? The insignificance of the S&P 500 index in initial model runs argues against a "flight to safety" argument for Las Vegas. In other words, the decline in equity values observed in 2000–2001 did not lead investors to divert funds into the "safe haven" of the Las Vegas housing market. Rather, the model results point squarely to Los Angeles investors as at least one source of the initial market distortion. Of primary importance is the strong, positive, and persistent relationship between growth in Los Angeles and Las Vegas home prices. I posit that persistently high prices in Los Angeles increased home equity, freeing up investment funds to flow into the Las Vegas market. This investment represented new demand over and above long-run expectations, causing prices in Las Vegas to rise. Further evidence of this scenario is the model finding that growth in Los Angeles income generated new construction activity in Las Vegas. Of course, new construction should dampen the upward pressure on prices. However, new construction was not sufficient to offset the upward pressure on prices caused by speculative demand.

Price persistence means that if prices drop, investors forecast more depreciation. This has played out in the Las Vegas market. By the end of 2005, prices were well above those attributable to fundamental forces as defined by the cointegrating vectors. The initial drop in price was modest, but price depreciation accelerated in subsequent quarters as investors, still relying on the previous period's price change to forecast future price changes, scrambled to take capital gains and/or cut their losses. As a result, prices dropped by over 13% between 2005 Q4 and 2008 Q1.

VI. CONCLUSION

In this paper, I develop a simple dynamic model that explains how a housing-price bubble in one regional housing market could spill into a neighboring market. Other studies have tested for regional house-price correlation using cointegration among price pairs in neighboring markets with limited success. I argue that any long-run equilibrium arising from supply and demand relationships must include stocks, since prices rise and fall, in the long run, largely in response to excess demand or supply. As a result, pairwise or other price-only-based tests for cointegration exclude important information about the price adjustment mechanism.

I estimate an ECM based on Johansen's (1988) multiple error correction techniques, seeking to determine the relative influence of fundamental variables and contagion from shocks to Los Angeles home prices and income on prices and the housing stock in Las Vegas. The model results show significant persistence in Las Vegas home prices. Key economic variables such as historically low interest rates and rising construction costs are shown to have far less influence on prices and stock growth than appreciation in Las Vegas prices and income and prices in Los Angeles. The model results also suggest that a substantial portion of the rapid appreciation and subsequent declines in home prices experienced in Las Vegas from 2001 to 2008 can be traced back to activity in the Los Angeles market. This provides empirical support for the contagion hypothesis.

APPENDIX

TABLE A1

Short-Run Dynamics for the Error-Correction Model, Coefficients of Lagged Endogenous Variables and *t*-Statistics in Brackets

Variable	D(LVPRICE)	D(LV EMP)	D(LVSTOCK)
D(LV PRICE(−1))	0.3611 [3.9370]	−0.0164 [−0.13335]	−0.0728 [−1.1291]
D(LV PRICE(−2))	0.0622 [0.6492]	−0.0661 [−0.51527]	0.0458 [0.6796]
D(LV PRICE(−3))	0.1386 [1.4972]	0.0429 [0.34610]	−0.0557 [−0.8572]
D(LVPRICE(−4))	−0.2924 [−3.5792]	0.0703 [0.64303]	−0.0077 [−0.1344]
D(LV EMP(−1))	0.0545 [0.59161]	−0.0596 [−0.48324]	−0.0864 [−1.33447]
D(LV EMP(−2))	0.0874 [0.91463]	0.2406 [1.87949]	0.0661 [0.98332]
D(LV EMP(−3))	0.4278 [4.55753]	−0.2467 [−1.96280]	0.0318 [0.48158]
D(LV EMP(−4))	0.0975 [0.95136]	0.0221 [0.16076]	0.0294 [0.40840]
D(LV STOCK(−1))	−0.6132 [−3.49573]	0.1335 [0.56841]	0.4922 [3.99344]
D(LV STOCK(−2))	−0.0583 [−0.33315]	−0.1431 [−0.61094]	0.0607 [0.49376]
D(LV STOCK(−3))	0.1289 [0.74091]	0.2050 [0.88034]	−0.0585 [−0.47841]
D(LV STOCK(−4))	−0.1720 [−1.15940]	−0.2155 [−1.08490]	0.0183 [0.17527]

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