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The Co-Movement of Housing Sales and Housing Prices: Empirics and Theory

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The co-movement of Housing Sales and Housing Prices: Empirics and Theory

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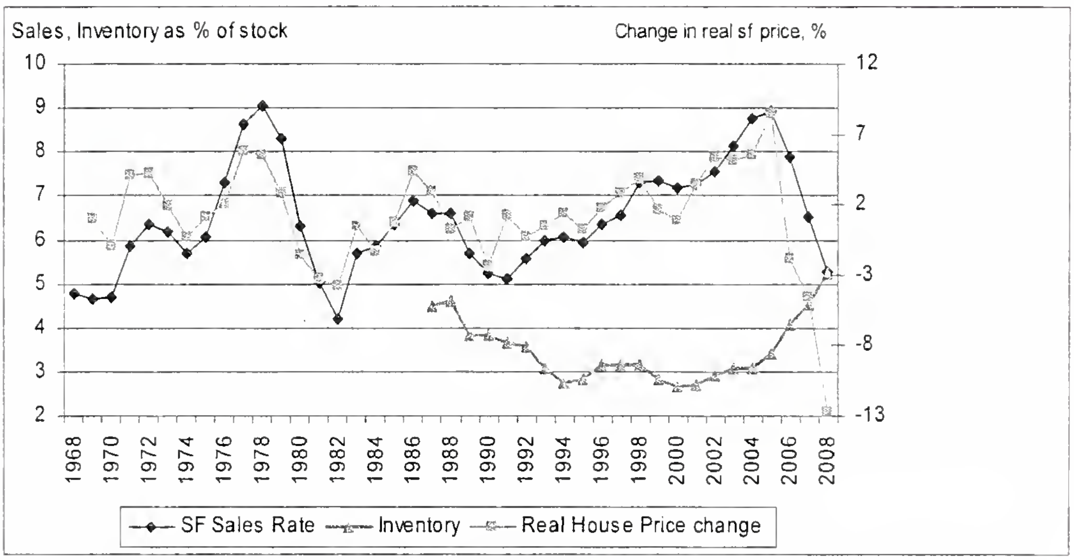
ABSTRACT

This paper examines the strong positive correlation that exists between the volume of housing sales and housing prices. We first examine gross housing flows in the US and divide sales into two categories: transactions that involve a change or choice of tenure, as opposed to owner-to-owner churn. The literature suggests that the latter generates a positive sales-to-price relationship, but we find that the former actually represents the majority of transactions. We develop a simple model of these inter-tenure flows which suggests they generate a negative price-to-sales relationship. This runs contrary to a different literature on liquidity constraints and loss aversion. Empirically, we assemble two data bases to test the model: a short panel of 33 MSA covering 1999-2008 and a long panel of 101 MSA spanning 1982-2006. Our results from both are strong and robust. Higher sales "Granger cause" higher prices, but higher prices "Granger cause" both lower sales and a growing inventory of units-for-sale. These relationships together provide a more complete picture of the housing market – suggesting the strong positive correlation in the data results from frequent shifts in the negative price-to-sales schedule.

I. Introduction.

As shown in Figure 1 below, there is a strong positive correlation between housing sales (expressed as a percent of all owner households) and the movement in housing prices (R^2=.66). On the surface the relationship looks to be close to contemporaneous. There is also a somewhat less obvious negative relationship between prices and the shorter series on the inventory of units for sale (R^2=.51). A number of authors have offered explanations for these relationships, in particular that between prices and sales.

Figure 1: US Housing Sales, Prices, Inventory



On the one hand, there is a growing literature of models describing home owner "churn" in the presence of search frictions [Wheaton (1990), Berkovec and Goodman (1996), Lundberg and Skedinger (1999)]. In these models, buyers become sellers – there are no entrants or exits from the market. In such a situation if participants pay higher prices, they also receive more upon sale. It is the transaction cost of owning 2 homes (during the moving period) that actually determines price levels. The greater transaction costs accompanying high prices can make trading expensive enough to erase the original gains from moving. In this environment Nash-bargained prices move almost inversely to sales duration – equal to the vacant inventory divided by the sales flow. In these models, both the inventory and sales churn are exogenous. Following Pissarides (2000) if the matching rate is exogenous or alternatively of specific form, greater sales churn will shorten duration and lead to higher prices. Similarly greater vacancy (inventory) raises sale duration and causes lower prices.

There also are a series of papers which propose that negative changes in prices will subsequently generate lower sales volumes. This again is a positive relationship between the two variables, but with opposite causality. The first of these is by Stein (1995) followed by Lamont and Stein (1999) and then Chan (2001). In these models, liquidity constrained owners are again moving from one house to another ("churn") and must make a down payment in order to purchase housing. When prices decline consumer equity does likewise and fewer households have the remaining down payment necessary to make a lateral move. As prices rise, equity recovers and so does market liquidity. Relying instead on behavior economics, Genesove and Mayer (2001) and then Englehardt (2003) show empirically that sellers who would loose some equity upon sale set higher reservations than those who would not. With higher reservations, the market as a whole should see lower sales if more and more sellers experience loss aversion as prices continue to drop.

In this paper we try to unravel the relationship between housing prices and housing sales, and in addition, the inventory of housing units for sale. We accomplish the following:

1). First, we carefully examine gross housing flows in the AHS for the 11 years in which the survey is conducted and find there are more purchases of homes by renters or new households than there are by existing owners. Hence the focus on own-to-own trades does *not* characterize the *majority* of housing sales transactions.

2). We also examine which flows add to the inventory of for-sale units (called LISTS) and which subtract (called SALES). Own-to-own moves, for example do both. We show with a simple model of tenure choice flows that higher prices should generate more LISTS, lower SALES, and hence a larger inventory. When prices are low, the reverse happens.

3). This leads us to hypothesize a very specific form of joint causality between sales and prices. Own-to-own churn generates a positive schedule between sales and prices as suggested by frictional market theory. At the same time, inter-tenure transitions should lead to a negative schedule. In equilibrium, the overall housing market should rest at the intersection of these two schedules.

4). To test these ideas we first assemble a US panel data base of 33 MSA from 1999-2008. The shortness of the panel is due to limited data on the for-sale inventory. An estimated panel VAR model perfectly confirms our hypothesized relationships. Sales positively drive subsequent prices while prices negatively drive subsequent sales and also positively increase the inventory.

5). We also assemble a longer panel of 101 MSA from 1982 to 2006 on just sales and prices. Using a wide range of model specifications and tests of robustness we find again that sales *positively* "Granger cause" subsequent housing price movements, while prices *negatively* "Granger cause" subsequent housing sales. These joint relationships are exactly as our model suggests when owner churn is combined with inter-tenure moves.

Our paper is organized as follows. In section II we set up an accounting framework for more completely describing gross housing flows from the 2001 AHS. This involves some careful assumptions to adequately document the magnitude of all the inter tenure flows relative to within tenure churn and to household creation/dissolution. In Section III, we develop a simple stylized model of the inter tenure flows to illustrate how they can generate a negative relationship between prices and sales and a positive relationship between prices and inventory. We present our hypothesized *pair* of relationships between prices and the sales/inventory ratio. In section IV we test these ideas with a short panel data base (33 MSA) that covers the inventory as well as prices and sales. In sections V through VII we present an analysis of a longer panel data set between just sales and prices across 101 MSA covering the years from 1982-2006. Here again we find conclusive evidence that sales positively "Granger cause" prices and that prices negatively "Granger cause" sales. Our analysis is robust to many alternative specifications and subsample tests. We conclude with some thoughts about historic market fluctuations as well as the outlook for US house prices and sales.

II. US Gross Housing Flows: Sales, Lists, and the Inventory.

Much of the theoretical literature on sales and prices investigates how existing homeowners behave as they try and sell their current home to purchase a new one. This flow is most often referred to as "churn". To investigate how important a role "churn" plays in the ownership market, we closely examine the 2001 American Housing Survey. In "Table 10"of the Survey, respondents are asked about the tenure of the residence they previously lived in – for those that moved during the last year. The total number of moves in this question is the same as the total in "Table 11" – asking about the previous status of the current head (the respondent). In "Table 11" it turns out that 25% of current renters moved from a residence situation in which they were *not* the head (leaving home, divorce, etc.). The fraction is a smaller 12% for owners. What is missing is the joint distribution between moving by the head and becoming a head. The AHS is not strictly able to identify how many current owners moved either a) from another unit they owned b) another unit they rented or c) purchased a house as they became a new or different household.

To generate the full set of flows, we use information in "Table 11" about whether the previous home was headed by the current head, a relative or acquaintance. We assume that all current owner-movers who were also newly created households - were counted in "Table 10" as being part of a previous owner household. For renters, we assume that all renter-movers that were also newly created households were counted in "Table 10" in proportion to renter-owner households in the full sample. Finally, we use the Census figures that year for the net increase in each type of household, and from that together with the data on moves we are able to identify household "exits" by tenure. Gross household exits occur mainly through deaths, institutionalization (such as to a nursing home), or marriage.

Focusing on just the owned housing market, the AHS also allows us to account for virtually all of the events that add units to the inventory of houses for sale (herein called LISTS) and all of those transactions that remove units from the inventory (herein called SALES). There are two exceptions. The first is the net delivery of new housing units. In 2001 the Census reports that 1,242.000 total units were delivered to the for-sale market. Since we have no direct count of demolitions[1] we use that figure also as net and it is counted as additional LISTS. The second is the *net* purchases of 2nd homes, which must be counted as additional SALES, but about which there is simply little data[2]. In theory, LISTS – SALES should equal the change in the inventory of units for sale. These relationships are depicted in Figure 2 and can be summarized with the identities below (2001 AHS values are included).

SALES = Own-to-Own + Rent-to-Own + New Owner [+ 2nd homes] = 5,281,000
LISTS = Own-to-Own + Own-to-Rent + Owner Exits + New homes = 5,179,000
Inventory Change = LISTS – SALES
Net Owner Change = New Owners – Owner Exits + Rent-to-Own – Own-to-Rent
Net Renter Change = New Renters – Renter Exits + Own-to-Rent – Rent-to-Own

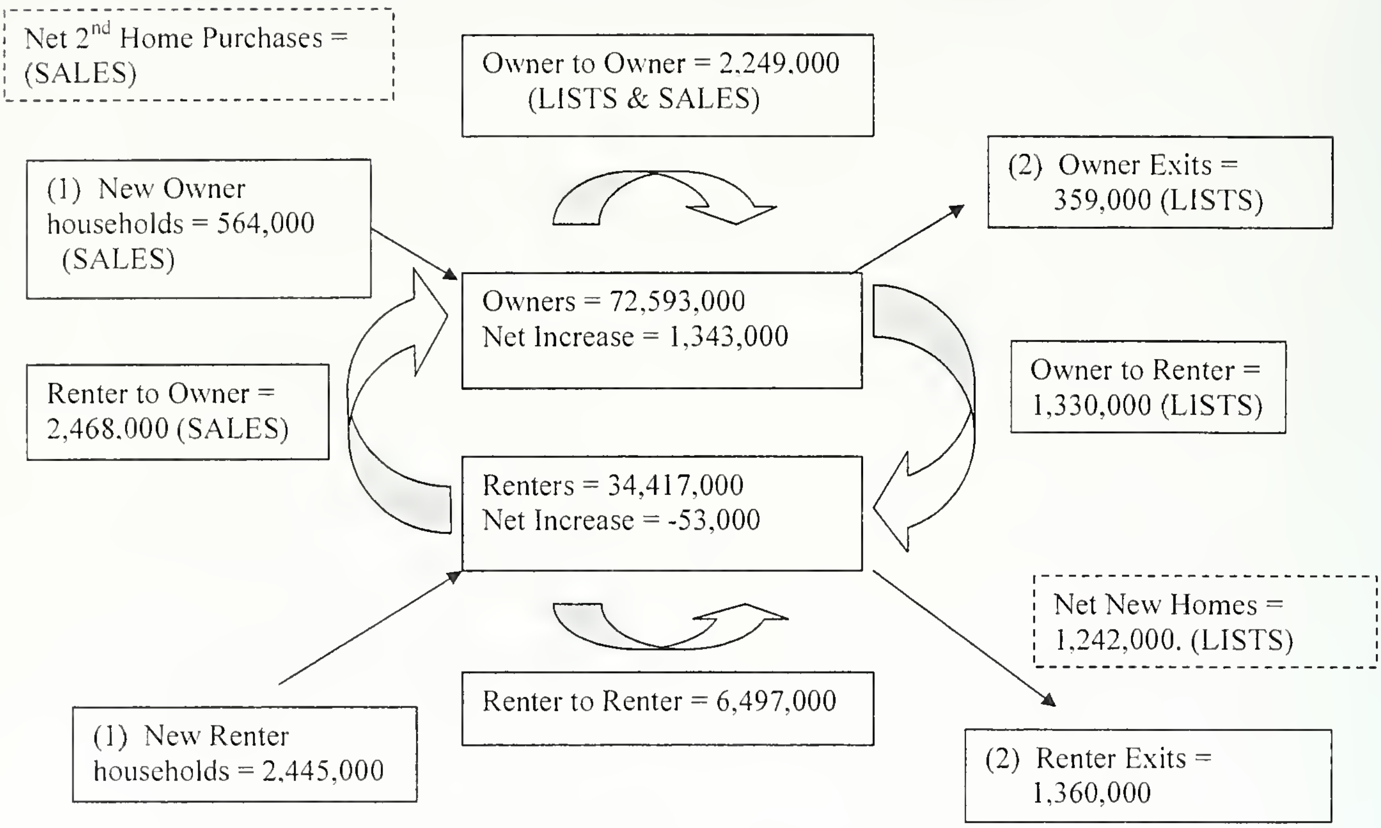
(1)

The only other comparable data is from the National Association of Realtors (NAR), and it reports that in 2001 the inventory of units for sale was nearly stable. The NAR however reports a slightly higher level of sales at 5,335,000 existing units.[3] This small discrepancy could be explained by repeat moves within a same year since the AHS asks only about the most recent move. It could also represent 2nd home sales which again are not part of the AHS move data.

What is most interesting to us is that almost 60% of SALES involve a buyer who is not transferring ownership laterally from one owned house to another. So called "churn" is actually a *minority* of sales transactions. The various flows between tenure categories also are the more critical determinant of change-in-inventory since "churn" sales leave the inventory unaffected. .

[1] The growth in stock between 1980-1990-2000 Censuses closely matches summed completions suggesting negligible demolitions over those decades. The same calculation between 1960 and 1970 however suggests removal of 3 million units.
[2] Net second home purchases might be estimated from the product of: the share of total gross home purchases that are second homes (reported by Loan Performance as 15.0%) and the share of new homes in total home purchases (Census, 25%). This would yield 3-4% of total transactions or about 200,000 units. There are no direct counts of the annual change in 2nd home stocks.
[3] The AHS is a repeat sample of housing units and excludes moves into new houses. Thus we compare its move number to NAR sales (both single and multi family) of existing units.

Figure 2: US Housing Gross Flows (2001)



In Figure 2, most inter-tenure SALES would seem to be events that one might expect to be negatively sensitive to housing prices. When prices are high presumably new created owner household formation is discouraged or at least deflected into new renter household formation. Likewise moves which involve changes in tenure from renting to owning also should be negatively sensitive to house prices. Both result because higher prices simply make owning a house less affordable. At this time we are agnostic about how net 2nd home sales are related to prices.

On the other side of Figure 2, many of the events generating LISTS should be at least somewhat positively sensitive to price. New deliveries certainly try to occur when prices are high, and such periods would be appropriate for any owners who wish or need to "cash out", consume equity or voluntarily choose to switch to renting. At this time we are still seeking a direct data source which investigates in more detail what events tend to generate the own-to-rent moves. Thus the flows in and out of homeownership in Figure 2 suggest that when prices are high sales are likely to decrease lists increase and the inventory grow.

The AHS has been conducted only semi-annually until recently and also has used consistent definitions of moving only since 1985. In Appendix III we calculate the flows for each of the 11 AHS surveys between 1985 and 2007. The flows are remarkably stable, although there exists some year to year variations. In all years, own-to-own moves ("churn") are less than the sum of new owners plus rent-to-own moves. Since the 2001 survey, the AHS calculated values for LIST-SALES have increased significantly. This is consistent with the growing national for-sale inventory reported in the NAR data over this period.

III. A stylized model of inter-tenure flows.

Here we assume that the total number of households T is fixed with $H \leq T$ being home owners. Those not owning rent at some fixed (exogenous) rent – hence we largely ignore the rental market. The total stock of units available for ownership U(p) is assumed to depend positively on price (long run supply) and with fixed rents we ignore rental supply. In this situation the inventory of units for sale is the difference between the owner stock and owner households: $I = U(p) - H$.

Households flow out of ownership at some constant rate α which could represent unemployment, foreclosure, or other economic shock. Rental households purchase units out of the owner stock (become owners) at some *rate* s(p) which we presume depends negatively on price. High prices (relative to the fixed rent) make ownership less appealing, but in general renters wish to become owners because of some assumed advantage (a tax subsidy for example). – hence the purchase rate is always positive.

The equations below summarize both flows (time derivatives) and steady state values (denoted with *). In Equation (2) the stable homeownership rate depends negatively on prices and the constant economic shock rate. When prices generate a sales rate equal to the economic shock rate, homeownership is 50%. Equation (3) cleanly divides up the inventory change into the same two categories from our more detailed flow diagram: LISTS-SALES. Here LISTS are stock change (new construction) plus own-to-rent flows (economic shocks) while SALES are rent-to-own flows. The equilibrium level of SALES is in (5), and the equilibrium inventory in (4). The latter must be constrained positive.

$$dH/dt = s(p)[T-H]-\alpha H, \qquad H^* = \frac{s(p)T}{\alpha + s(p)} \tag{2}$$

$$dI/dt = dU/dt - dH/dt = [dU/dt + \alpha H] - s(p)[T-H], \tag{3}$$

$$I^* = U(p) - H^* = \frac{\alpha U(p) + s(p)U(p) - sT}{\alpha + s(p)} \geq 0 \tag{4}$$

$$s(p)[T-H^*] = \frac{\alpha s(p)T}{\alpha + s(p)} \tag{5}$$

In (6) we derive comparative statics which show that as prices increase, the steady state value of the inventory grows and the steady state level of SALES decreases – as hypothesized about the flows which were diagramed in Figure 2.

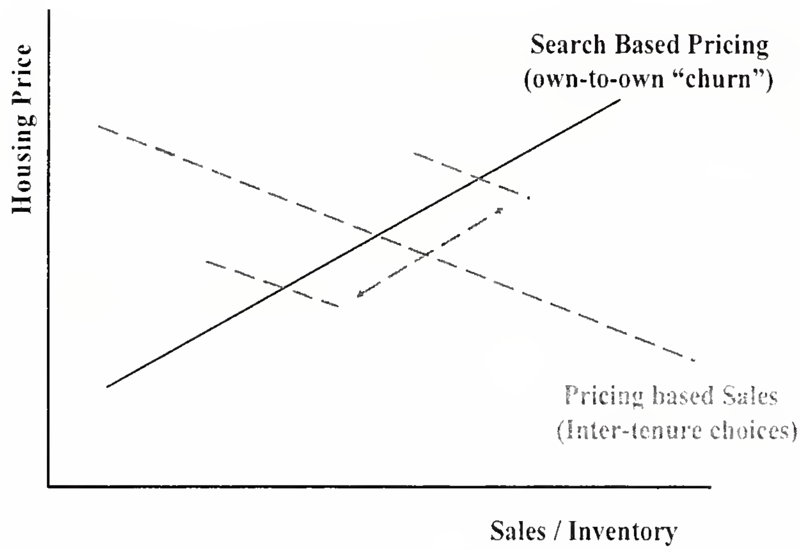
$$dI^*/dp = dU/dp - dH^*/dp = dU/dp - \frac{Tds/dp}{\alpha + s}[1 - \frac{s}{\alpha + s}] \geq 0 \tag{6}$$

$$d\,(s(p)[T-H^*])/dp = \frac{\alpha Tds/dp}{\alpha + s}[1 - \frac{s}{\alpha + s}] \leq 0$$

Again, the conclusions above follow from the *assumptions* that long run stock is positively related to price and the sales *rate* is negative related to price. Thus this simple model of inter-tenure flows establishes a negative relationship between housing prices and Sales/Inventory ratios. Alternatively, there should be a positive relationship between prices and duration.

With this new schedule between prices and duration we are now ready to better describe the full set of relationships in the owner market between sales, prices and the inventory. We combine this new schedule with a positive schedule between prices and the Sale/Inventory ratio – created from the various models of own-to-own decisions. In these latter models it is sales that are determining prices, while with the model in (2)-(5) above it is prices that are determining sales. At a more complete equilibrium (in the ownership market) sales, prices and the inventory all rest at the intersection of the two schedules shown in Figure 3. Figure 3 presents a more complete picture of the housing market than the models of Stein, Wheaton, or Berkovec and Goodman – since it accounts for the very large role of inter-tenure mobility as well as for owner churn.

FIGURE 3: Housing Market Equilibrium(s)



The out-of-equilibrium dynamics of this model are also appealing and seem in line with economic intuition as well. Consider a permanent increase in economic shocks (α). Using (2)- (5), owner households decline, and the inventory increases. Sales however also increase and so the impact on duration is technically ambiguous. Within a wide range of reasonable parameter values however, we can show that the sale/inventory ratio declines with greater α – the net shift in the price-to-sales schedule is therefore inward.[4] The new equilibrium then results in lower prices with a lower sales/inventory ratio as well (a higher duration). If we shift the s(p) schedule up (e.g. a greater tax subsidy) the number of owners increases, the inventory drops, and sales increase. This leads to an

[4] A sufficient condition is for the number of renters [T-H] to exceed the for-sale inventory.

unambiguous rightward shift in the price-to-sales schedule with a corresponding rise in equilibrium Sales/Inventory (drop in duration). Prices of course rise as well.

The next task is to see if we can empirically identify the relationships in Figure 3. For this, we examine two several panel data bases with different degrees of richness. The first data base is shorter and covers only 33 MSA. Its advantage is that it includes data on the inventory for sale by market – a series which the NAR has collected only recently. The second data base is much longer, covers 101 MSA, but includes only information on sales and prices.

IV. A Short Panel Analysis of Metropolitan Sales, Prices and Inventory.

Carefully constructed series on house prices are available from the late 1970s or early 1980s and for a wide range of metropolitan areas. The price data we use is the deflated OFHEO repeat sales series [Baily, Muth, Nourse (1963)]. This data series has recently been questioned for not factoring out home improvements or maintenance and for not factoring in depreciation and obsolescence [Case, Pollakowski, Wachter (1991), Harding, Rosenthal, Sirmans (2007)]. That said we are left with what is available, and the OFHEO index is the most consistent series available for most US markets over a long time period. The only alternative is CSW/FISERV, and it is available for far fewer metropolitan areas that in turn are disproportionately concentrated in the south and west.

In terms of sales, the only consistent source is that provided by the National Association of Realtors (NAR). The NAR data is for single family units only (it excludes condominium sales at the MSA level), but is available for each MSA over a period from 1980 to the present. The more limiting data series is that on the inventory of housing units for sale. Here the NAR distributes MSA data only from 1999 or later. We have been able to put together all three series since 1999 for 33 MSA, and Figures 4 through 6 depict the 33 series for each variable. The patterns are quite discernable and in Appendix I we present summary statistics for each market.

In Figure 4 we clearly see all house prices rising and then falling since 1999. The sample almost evenly divides between market where this movement is very pronounced and those with only the slightest of changes. In terms of the inventory, Figure 5 shows that over the first half of the sample the inventory was roughly constant. After 2004 it rises and falls in a pattern again similar to prices. Both the Prices and Inventory are raw series and exhibit little seasonality. As for sales, in Figure 6 we see a little bit of the same "hump shaped" pattern, but it seems weaker. What is more problematic with the sales data is the strong pattern of seasonality in each series – seasonality that varies by specific market in many cases.

Figure 4: Quarterly Real House Prices (33 MSA), 1999-2008

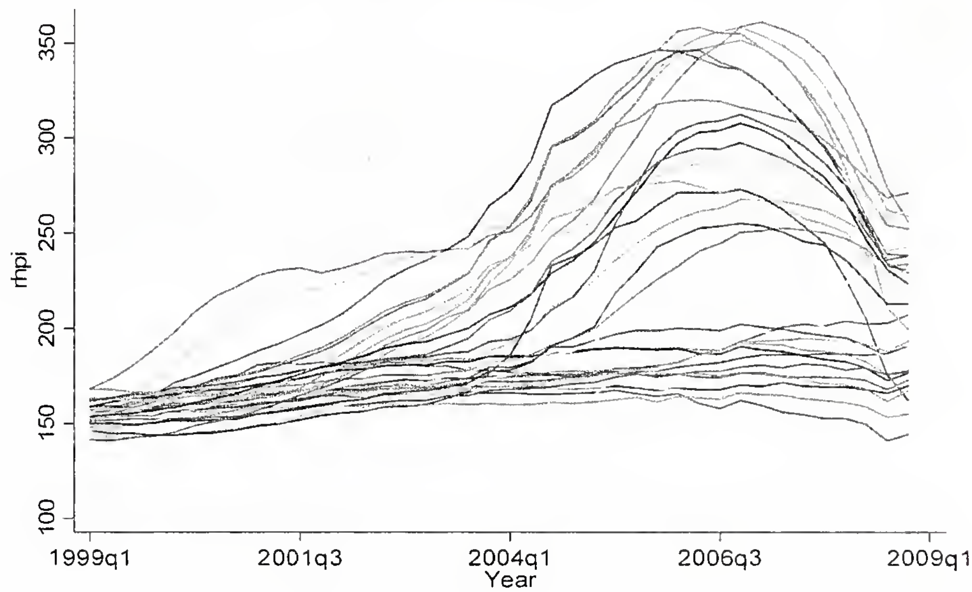
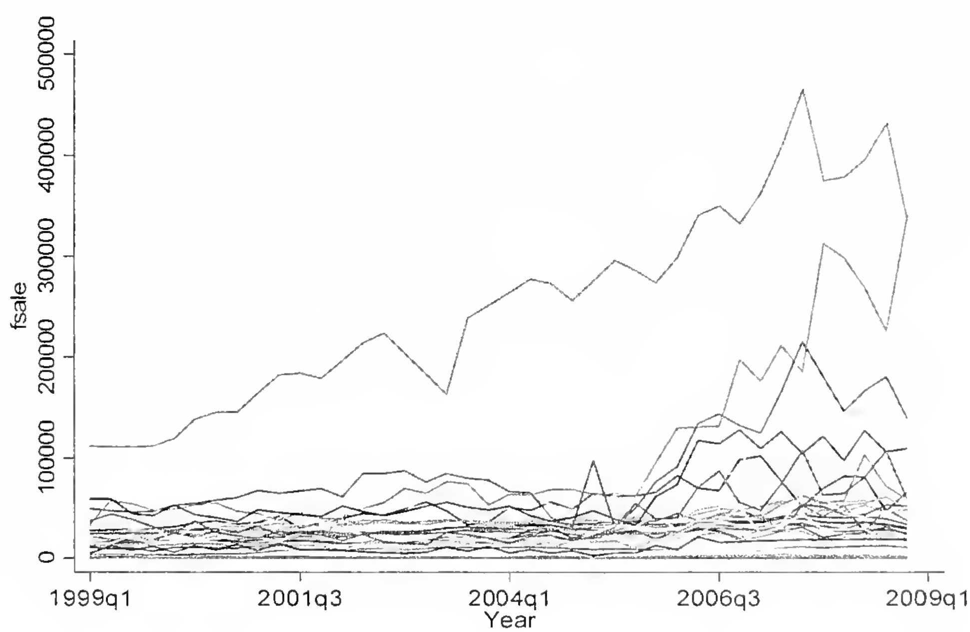
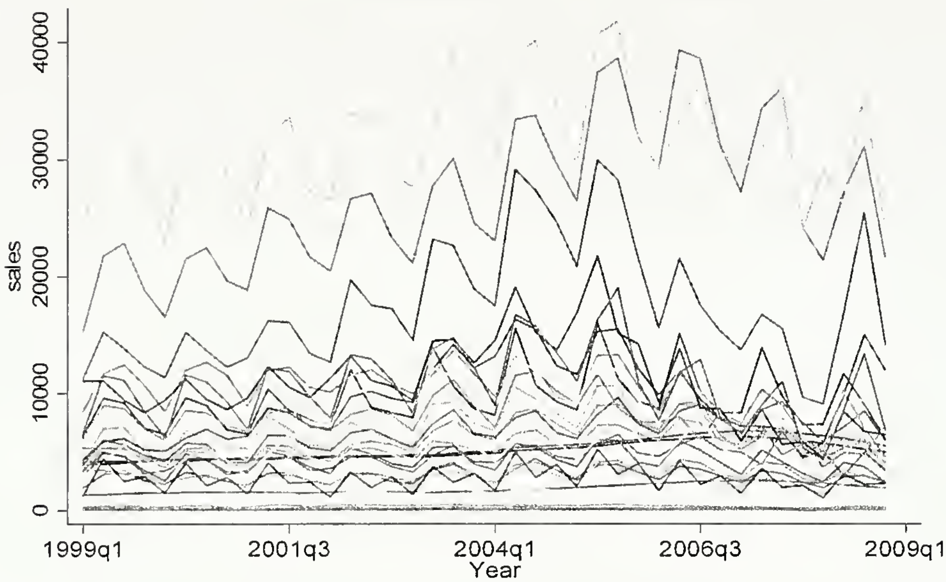


Figure 5: Quarterly For-Sale Inventory (33 MSA), 1999-2008



: **Figure 6: Quarterly House Sales (33 MSA), 1999-2008**



These observations suggest that a panel VAR is an appropriate instrument to test the relationships between prices, sales and inventory. In the VAR we will have each variable depending on lagged values of itself and the other variables. If the panel is of order one, we also can use each coefficient as an effective test of "Granger causality". Before turning to such a model, however, we need to examine each series to see if they are stationary. There are two tests available for use with panel data and in each, the null hypothesis is that sum or average of *all* the individual series have unit roots and are non stationary. In Levin-Lin (LL, 1993) the null has no constant (or drift) while in Im-Persaran-Shin (IPS, 2002) the null includes a constant to allow for drift. In Table 1 we report the results of this test for housing prices, sales and inventory – in levels. With the possible exception of prices, where we can be confident only at the 7% level, the non-stationary null is rejected and we should be on solid grounds undertaking our proposed VAR.

Table 1: Stationary tests, Short Panel

Inventory (Augmented by 1 quarters)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.17706	-11.214	-2.43396	0.0075
IPS test		T-Bar	W(t-bar)	P>T
Levels		-2.482	-2.114	0.017

Sale (Augmented by 1 quarters)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.02848	-3.924	-3.85330	0.0001
IPS test		T-Bar	W(t-bar)	P>T
Levels		-3.221	-10.906	0.000

Price (Augmented by1 quarters)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.00977	-6.168	-6.05783	0.0000
IPS test		T-Bar	W(t-bar)	P>T
Levels		-1.750	-1.477	0.070

In panel VAR models with individual heterogeneity there exists a specification issue: the error term can be correlated with the lagged dependent variables [Nickell, (1981)]. OLS estimation can yield coefficients that are both biased and also that are not consistent in the number of cross-section observations. Consistency occurs only in the number of time series observations. Thus estimates and any tests on the parameters of interest may not be reliable. These problems might not be serious in our case since we have 32 quarterly time series observations (more than many panel models). To be on the safe side, however, we also estimated the equations following an estimation strategy by Holtz-Eakin et al. As discussed in Appendix II, this amounts to using 2-period lagged values of sales and prices as instruments with GLS estimation.

A final concern with our VAR is the handling of seasonality. Here we propose 2 adjustments. In Tables 2 and 3, we report results using quarterly seasonal effects interacted with the cross section fixed effects. This effectively allows each MSA to have its own set of seasonal influences. Our second approach is to change all of the lags in the

VAR to 4-periods rather than one. In effect we are asking how our vector of variables relates to the vector lagged a year previous rather than a quarter ago. The year-over-year VAR results are presented in Table 4. [5]

Table 2: OLS Quarterly Panel VAR, 1999-2007, 33 MSA
(1 quarter lags, interactive seasonal effects)

$$P_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(234.97)}{.9795975} \; P_{it-1} - \underset{(-8.98)}{.000594} \; I_{it-1} + \underset{(7.78)}{.0004185} \; S_{it-1}, \quad R^2 = 0.9889$$

$$I_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(79.82)}{.9350479} \, I_{it-1} + \underset{(5.15)}{37.95637} \; P_{it-1} - \underset{(-0.86)}{.0819372} \; S_{it-1}, \quad R^2 = .9660$$

$$S_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(87.66)}{.9209442} \; S_{it-1} - \underset{(-1.64)}{1.332737} \; P_{it-1} - \underset{(-3.14)}{.0040554} \; I_{it-1}, \quad R^2 = .9950$$

Table 3: Holtz-Eakin Quarterly Panel VAR, 1999-2007, 33 MSA
(1 quarter lags, interactive seasonal effects)

$$P_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(427.85)}{.9819115} \; P_{it-1} - \underset{(-22.31)}{.000966} \; I_{it-1} + \underset{(19.79)}{.0006243} \; S_{it-1},$$

$$I_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(62.09)}{.8864376} \; I_{it-1} + \underset{(7.64)}{94.84853} \; P_{it-1} - \underset{(-0.57)}{.0774326} \; S_{it-1}$$

$$S_{it} = \alpha_i + \sum_{j=1}^{4} \alpha_i Q_j + \underset{(76.60)}{.916015} \; S_{it-1} - \underset{(-1.84)}{2.864344} \; P_{it-1} - \underset{(-2.30)}{.0041445} \; I_{it-1},$$

Table 4: OLS Quarterly Panel VAR, 1999-2007, 33 MSA
(4 quarter lags)

$$P_{it} = \alpha_i + \underset{(63.1)}{.833} \; P_{it-4} - \underset{(-8.5)}{.0021} \; I_{it-4} + \underset{(9.4)}{.0012} \; S_{it-4}, \quad R^2 = .885$$

$$I_{it} = \alpha_i + \underset{(53.2)}{.921} \; I_{it-4} + \underset{(10.1)}{93.9} \; P_{it-4} + \underset{(3.3)}{.032} \; S_{it-4}, \quad R^2 = .945$$

$$S_{it} = \alpha_i + \underset{(52.3)}{.819} \; S_{it-4} - \underset{(-6.7)}{102.6} \; P_{it-4} - \underset{(-2.4)}{.072} \; I_{it-4}, \quad R^2 = .98$$

[5] We do not report the cross section fixed effects, nor their interactive terms. Many, but not all are significant. Standard errors are shown in parenthesis below each coefficient.

In examining Tables 2-4, we find that our three hypotheses are validated in every case. First, the inventory negatively impacts price while sales has a positive effect. Interestingly the coefficient on inventory is always slightly larger than sales. In a log model if the ratio (duration) were all that mattered the coefficients would be identical in magnitude. In this linear model they are close enough to suggest a similar conclusion. Secondly, the inventory responds quite significantly (and positively) to prices. Thirdly, prices negatively impact subsequent sales. All of these effects are statistically significant, but the price impact on Sales shows up more strongly in the 4-quarter lag model. In the 1-quarter model it is at the threshold of significance. In all respects, the results fully support equations (2)-(5) and the pair of relationships in Figure 3. Duration negatively "Granger causes" subsequent prices to decline. Price then positively "Granger causes" the inventory to grow, and likewise for sales to decline. The first VAR equation validates the upward schedule in Figure 3 while the second and third combine to yield the downward schedule.

In comparing the different models we note that the Holtz-Eakin estimation does increase the coefficients a bit and reduce standard errors – relative to the OLS results. We did not undertake Holtz-Eakin estimation for our 4-quarter lag model. The OLS 4-quarter results are expectedly different. Inventory and sales, for example, have an impact on prices 4 periods hence that is roughly 4 times their impact in the 1-quarter model. Similarly, prices impact sales 4 quarters hence with much greater impact than from just 1 quarter back.

V. A long-Panel of Metropolitan Sales and Prices.

It is possible to test the just the relationship between prices and sales over a much longer time horizon – if we ignore the inventory. [6] For this we assemble a larger panel data base covering 101 MSA and spanning the years 1982 through 2006. This panel was

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[6] There have been a few recent attempts test the relationship between movements in sales and prices. Leung, Lau, and Leong (2002) undertake a time series analysis of Hong Kong Housing and conclude that stronger Granger Causality is found for sales driving prices rather than prices driving sales. Andrew and Meen (2003) examine a UK Macro time series using a VAR model and conclude that transactions respond to shocks more quickly than prices, but do not necessarily "Granger Cause" price responses. Both studies are hampered by limited observations.

purposely structured to be annual so as to avoid the seasonality of the shorter panel, while still maintaining plentiful time-based degrees of freedom.

Over this longer period, many metropolitan areas almost doubled their housing stock so we decided to standardize the sales data to eliminate some of the trend. Raw sales were compared with yearly Census estimates of the number of total households in those markets. Dividing single family sales by total households we get an estimated sales rate for each market in each period. Using sales rates also eliminated much of the cross section variation in the raw number of sales. In a similar manner we set the real price level in each market to 100 in the base year. These re-scaling of the data will help make the cross section fixed effects smaller in the estimated VAR models.

In Figures 7 and 8 we illustrate the constant dollar OFHEO price series along with the yearly NAR sales rate data, for all 101 of our markets. Over this time frame, the price series vary widely across markets, with some areas experiencing long term although episodic increases (e.g. San Francisco) while others are almost totally constant (e.g. Dallas). As for the sales rates, virtually every market has a slow gradual trend in sales rates, with the sample average increasing from 3% to 5.6 %. In appendix III we present the summary statistics for each market's price and sales rate series.

The data in Figure 4 through 6 for the short panel showed no obvious trends; prices, sales and the inventory generally rise and then fall. The longer term series in Figures 7 and 8 may have more persistent trends and so again we need to test for whether the series are stationary. In Table 5 we report the results of both Levin-Lin and IPS tests for both housing price and sale rate levels.

With both tests the null for house prices is rejected at high confidence levels, but with the IPS test the null hypothesis for the sales rate is quite likely to hold. Given the steady trends seen in Figure 8, this of course seems reasonable. To be on the safe side, then we estimate our long term sales-price VAR in differences as well as levels.

Figure 7: Annual House Prices (101 MSA), 1982-2006

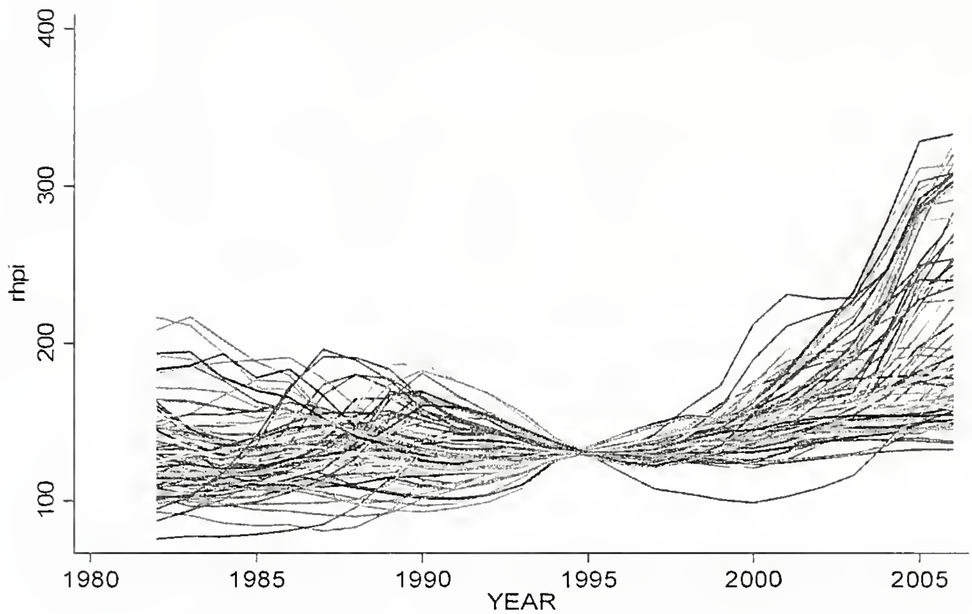

Figure 8: Annual Sales Rate (101 MSA), 1982-2006

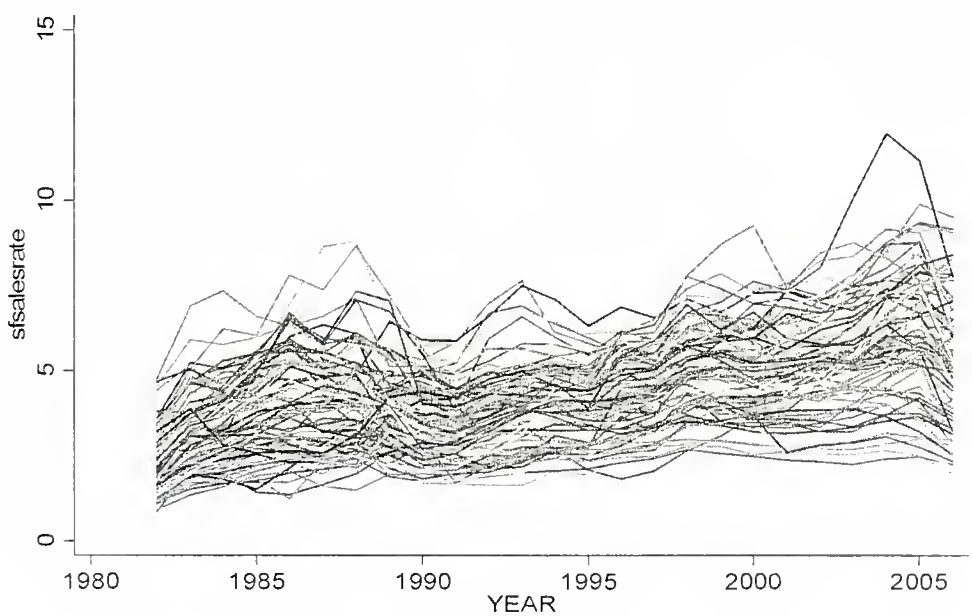

TABLE 5: Stationary tests, Long Panel

RHPI (Augmented by 1 lag, no constant for Levin lin's test)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.06626	-12.158	-11.71630	0.0000
IPS test		T-Bar	W(t-bar)	P>T
Levels		-1.772	-2.729	0.003

SFSALESRATE (Augmented by 1 lag, , no constant for Levin lin's test)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.03852	-5.550	-5.34822	0.0000
IPS test		T-Bar	W(t-bar)	P>T
Levels		-1.339	1.847	0.968

Since economies change more in the longer term, we decided to include several conditioning variables. The conditioning variables we choose are market specific employment, and the national mortgage rate. The resulting 2-variable VAR in levels is shown in (7), while in (8) the companion model is presented in first differences.[7]

$$P_{i,t} = \alpha_0 + \alpha_1 P_{i,t-1} + \alpha_2 S_{i,t-1} + \beta' X_{i,t} + \delta_i + \varepsilon_{i,t} \qquad (7)$$

$$S_{i,t} = \gamma_0 + \gamma_1 S_{i,t-1} + \gamma_2 P_{i,t-1} + \lambda' X_{i,t} + \eta_i + \varepsilon_{i,t}$$

$$\Delta P_{i,t} = \alpha_0 + \alpha_1 \Delta P_{i,t-1} + \alpha_2 \Delta S_{i,t-1} + \beta' \Delta X_{i,t} + \delta_i + \varepsilon_{i,t} \qquad (8)$$

$$\Delta S_{i,t} = \gamma_0 + \gamma_1 \Delta S_{i,t-1} + \gamma_2 \Delta P_{i,t-1} + \lambda' \Delta X_{i,t} + \eta_i + \varepsilon_{i,t}$$

We estimate each model using both OLS and also applying the previously discussed estimation strategy by Holtz-Eakin et al. From either estimates, we conduct a "Granger" causality test. Since we are only testing for a single restriction, the t statistic is the square root of the F statistic that would be used to test the hypothesis in the presence of a longer lag structure (Greene, 2003). Hence, we can simply use a t test (applied to

[7] In (6) the fixed effects are cross-section trends rather than cross section levels as in (5)

the α_2 and γ_2) as the check of whether changes in sales "Granger cause" changes in price and/or whether prices "Granger cause" sales.

In table 6 we report the results of equations (7) and (8) in each set of rows. The first column uses OLS estimation, the second the Random Effects IV estimates from Holtz-Eakin et al. Interestingly, the two estimation techniques yield quite similar coefficients – as might be expected with a larger number of time series observations and data rescaling to reduce the cross section effects.The first set of equations is in levels, while the second set of rows reports the results using differences. In all Tables, variable names are self evident and variable differences are indicated with the prefix GR. Standard errors are reported in parenthesis.

Among the levels equations, we first notice some anomalies. The mortgage interest rate in the price levels equation is always of the wrong sign, and the employment coefficient in the OLS sales rate equation is insignificant (despite almost 2500 observations). A more troublesome result is that the price levels equation has excess "momentum" – lagged prices have a coefficient greater than one. Hence prices (levels) can grow on their own without necessitating any increases in fundamentals, or sales. We suspect that these two anomalies are likely the result of the non-stationary feature to both the price and sales series when measured in levels. When we move to the results of estimating the equations in differences all of these issues disappear. The lagged price coefficients are less than one so the price equations are stable in the 2^{nd} degree, and the signs of all coefficients are both correct – and highly significant.

As to the question of causality, in every price or price growth equation, lagged sales or growth in sales is always significantly positive. Furthermore in every sales rate or growth in sales rate equation, lagged prices (or its growth) are also always significant. There is clear evidence of joint causality, *and the effect of lagged prices on sales is always of a negative sign.* Holding lagged sales (and conditioning variables) constant, a year after there is an increase in prices – sales fall. This is the opposite of that predicted by theories of loss aversion or liquidity constraints, but fully consistent with the role played by tenure choices in Figure 2 and our simple model of these flows.

TABLE 6: Annual Sales-Price VAR, 1982-2006

	Fixed Effects	E Holtz-Eakin estimator
Levels		
Real Price (Dependent Variable)		
Constant	-28.20461** (2.324949)	-37.72296** (1.832941)
Real Price (lag 1)	1.074879** (0.0064924)	1.03659** (0.0048785)
Sales Rate (lag 1)	3.402427** (0.1735139)	4.759024** (0.1520669)
Mortgage Rate	0.4064326** (0 .0989936)	0.7712427** (0.0762181)
Employment	0.0085368** (0.0014575)	0.0114432** (0.0018957)
Sales Rate (Dependent Variable)		
Constant	2.263661** (0.1623367)	2.473909** (0.1699231)
Real Price (lag 1)	-0.0071418** (0.0004533)	-0.0077217** (0.0004666)
Sales Rate (lag 1)	0.8484933** (0.0121154)	0.665248** (0.013338)
Mortgage	-0.0615272** (0.0069121)	-0.0766582** (0.0068535)
Employment	0.0000882 (0.0001018)	0.0011314** (0.0001835)
First Difference		
GR Real Price (Dependent Variable)		
Constant	-0.3966703** (0.1231288)	-0.9248402** (0.1219398)
GR Real Price (Lag 1)	0.7570639** (0.014764)	0.6737341** (0.013863)
GR Sales Rate (Lag 1)	0.0293207** (0.0058313)	0.0546765 ** (0.0061649)
GR Mortgage Rate	-0.0901117** (0.0099561)	-0.1197969** (0.009232)
GR Employment	0.3123949** (0 .0394799)	0.5318401 ** (0.0414994)
GR Sales Rate (Dependent Variable)		
Constant	0.8397989** (0.393873)	1.526134** (0 .4334829)
GR Real Price (Lag1)	-0.7050644** (0.0472282)	-1.106593** (0.0578562)
GR Sales Rate (Lag 1)	0.0544417** (0.0186536)	-0.02252 (0.0188086)

GR Mortgage Rate	-.3251265** (0.0318483)	-0.3078643** (0.031781)
GR Employment	1.134269** (0.126291)	1.391958** (0.1463704)

** indicates significance at 5%.

We have experimented with these models using more than a single lag, but qualitatively the results are the same. In levels, the price equation with two lags becomes dynamically stable in the sense that the sum of the lagged price coefficients is less than one. As to causal inference, the sum of the lagged sales coefficients is positive, highly significant, and passes the Granger F test. In the sales rate equation, the sum of the two lagged sales rates is virtually identical to the single coefficient above and the lagged price levels are again significantly negative (in their sum). Collectively higher lagged prices "Granger cause" a reduction in sales. We have similar conclusions when two lags are used in the differences equations, but in differences, the 2nd lag is always insignificant.

As a final test, we investigate a relationship between the *growth* in house prices and the *level* of the sales rate. In the search theoretic models sales rates determine price levels, but if prices are slow to adjust, the impact of sales might better show up on price changes. Similarly the theories of loss aversion and liquidity constraints relate price changes to sales levels. While the mixing of levels and changes in time series analysis is generally not standard, this combination of variables is also the strong empirical fact shown in Figure 1. In Table 7 price changes are tested for Granger causality against the level of sales (as a rate).

TABLE 7: Annual Sales-Price Mixed VAR, 1981-2006

Differences and Levels	Fixed Effects	E Holtz-Eakin estimator
GR Real Price (Dependent Variable)		
Constant	-6.698605** (0.3568543)	-11.10693** (0.4174099)
GR Real Price (lag 1)	0.5969905** (0.015889)	0.4286127** (0.0156827)
Sales Rate (lag 1)	1.424051** (0.0760102)	2.340478** (0.0912454)
GR Mortgage Rate	-0.1230876** (0.009451)	-0.1573441** (0.0086482)
GR Employment	0.4987545** (0.0349922)	0.7781044** (0.0373462)

Sales Rate (Dependent Variable)		
Constant	-0.0458271** (0.0541373)	0.283191** (0.0642588)
GR House Price (lag 1)	-0.0328973** (0.0024105)	-0.0355432** (0.002961)
Sales Rate (lag 1)	1.01549** (0.0115313)	0.9482599** (0.0139037)
GR Mortgage Rate	-0.0156137** (0.0014338)	-0.0132519** (0.0013497)
GR Employment	0.0462483** (0.0053086)	0.7280071** (0.1643153)

** indicates significance at 5%

In terms of causality, these results are no different than the models estimated either in all levels or all differences. One year after an increase in the *level* of sales, the *growth* in house prices accelerates. Similarly, one year after house price *growth* accelerates the *level* of home sales falls. All conditioning variables are significant and correctly signed and lagged dependent variables have coefficients less than one.

VII. Long Panel Tests of Robustness.

In panel models it is always a good idea to provide some additional tests of the robustness of results, usually by dividing up either the cross section or time series of the panel into subsets and examining these results as well. Here we perform both tests. First we divide the MSA markets into two groups: so-called "coastal" cities that border either ocean, and "interior" cities that do not. There are 31 markets in the former group and 70 in the latter. The coastal cities are often felt to be those with strong price trends and possibly different market supply behavior. These results are in Table 8. The second test is to divide the sample up by year – in this case we estimate separate models for 1981-1992 and 1993-2006. The year 1992 generally marks the bottom of the housing market from the 1990 recession. These results are depicted in Table 9. Both experiments use just the differences model that seems to provide the strongest results from the previous section.

TABLE 8: Geographic Sub Panels, 1982-2006

	Fixed Effects		E Holtz-Eakin estimator	
	Coastal MSA	Interior MSA	Coastal MSA	Interior MSA
GR Real Price (Dependent Variable)				
Constant	-0.4766326** (0.272633)	-0.3510184 (0.130979)	-1.188642** (0.2669406)	-0.721773** (0.1290227)
GR Real Price (Lag 1)	0.7340125** (0.0271992)	0 .77654** (0.0173987)	0.6845244** (0.0255055)	0.6926984** (0.0162093)
GR Sales Rate (Lag 1)	0.0615042** (0.0133799)	0.016732** (0.0061206)	0.089245** (0.0135344)	0.0373314** (0.0064948)
GR Mortgage Rate	-0.0885175** (0.0214447)	-0.0908632** (0.0107238)	-0.1275495** (0.020204)	-0.1119299** (0.0098326)
GR Employment	0.413934** (0.0864868)	0.2599301** (0.0422536)	0.6823408** (0.0890938)	0.4168953** (0.0438112)
GR Sales Rate (Dependent Variable)				
Constant	1.01577** (0 .71945)	0.726351** (0.4706282)	0.9888512** (0.7707496)	1.406108** (0.5151945)
GR Real Price (Lag1)	-0.7510799** (0.0717759)	-0.680113** (0.0625164)	-0.9596828** (0.0802649)	-1.092057** (0.0775659)
GR Sales Rate (Lag 1)	-0.0111514 (0.0353082)	0.0786527** (0.0219922)	-0.0686389** (0.0362451)	.013674 (0.0219432)
GR Mortgage Rate	-0.3092647** (0.0565903)	-0.3335691** (0.0385322)	-0.3139948** (0.057035)	-.3112734** (0.0383706)
GR Employment	1.265646** (0.2282296)	1.097809** (0.1518239)	1.651104** (0.2580107)	1.285375** (0.1738679)

Note:

a) *- 10 percent significance. **- 5 percent significance.
b) MSAs denoted coastal are MSAs near the East or West Coast (see Appendix I).
c) MSAs denoted interior are MSAs that are not located at the East or West Coast.

In Table 8, the results of Table 6 hold up remarkably strong when the panel is divided by region. The coefficient of sales rate (growth) on prices is always significant although so-called "costal" cities have larger coefficients. In the equations of price (growth) on sales rates, the coefficients are always significant, and the point estimates are very similar as well. The negative effect of prices on sales rates is completely identical across the regional division of the panel sample. It should be pointed out that all of the instruments are correctly signed and significant as well.

The conclusion is the same when the panel is split into two periods (Table 9). The coefficients of interest are significant and of similar magnitudes across time periods, and all instruments are significant and correctly signed as well. The strong negative impact of prices on sales clearly occurred during 1982-1992 as well as over the more recent period from 1993-2006. With fewer time series observations in each of the (sub) panels in Table 9, the Holtz-Eakin estimates are now sometimes more different than the OLS results.

TABLE 9: Time Subpanels, 101 MSA

	Fixed Effects		E Holtz-Eakin estimator	
	1982-1992	1993-2006	1982-1992	1993-2006
GR Real Price (Dependent Variable)				
Constant	-2.648239** (0.2403419)	0.1098486 (0.1788311)	-2.597524** (0.2538611)	0.0280176 (0.1594721)
GR Real Price (Lag 1)	0 .5533667** (0.0273908)	0.7581002** (0.0202692)	0.7006637** (0.0281998)	0.709152** (0.0212374)
GR Sales Rate (Lag 1)	0.0204875** (0.0074812)	0.0631485** (0.0114222)	0.0273097** (0.0081297)	0.0373903** (0.0114339)
GR Mortgage Rate	-0.2309851** (0.0195574)	-0.113025** (0.0148469)	-0.2164034** (0.0174734)	-0.1088119** (0.0120579)
GR Employment	0.6215331** (0.0644479)	0.3634738** (0.0594376)	0.5073589** (0.0719732)	0.5722806** (0.0629962)
GR Sales Rate (Dependent Variable)				
Constant	-6.077011** (0.9073653)	3.339319** (0.494379)	-4.553209** (1.017364)	4.601864** (0.5978177)
GR Real Price (Lag1)	-0.8804394** (0.1034087)	-0.7628642** (0.0560344)	-0.9065855** (0.1359358)	-0.8880742** (0.0738609)
GR Sales Rate (Lag 1)	0.0053538 (0.0282439)	-0.0100386** (0.0315767)	0.0706683 (0.0302102)	-0.0313258 (0.035461)
GR Mortgage Rate	-0.5534765** (0.0738353)	-0.3843505** (0.0410443)	-0.5593403** (0.0731325)	-0.2695104** (0.0383087)
GR Employment	2.564815** (0.2433108)	0.7280071** (0.1643153)	1.88701** (0.293079)	0.5015754** (0.2095683)

Note:

a) Column labeled under 1982-1992 refer to the results using observations that span those years..
b) Column labeled under 1993-2006 refer to the results using observations that span those years.

VII. Conclusions

We have shown that the "Granger causal" relationship from prices-to-sales is actually negative – rather than positive. Our empirics are quite strong. As an explanation, we have argued that actual flows in the housing market are remarkably large between tenure groups – and that a negative price-to-sales relationship makes sense as a reflection of these inter-tenure flows. Higher prices lead more households to choose renting than owning and these flows decrease SALES. Higher prices also increase LISTS and so the inventory grows. Conversely, when prices are low, entrants exceed exits into ownership, SALES increase, LISTS decline and so does the inventory.

Our empirical analysis also overwhelmingly supports the positive sales-to-price relationship that emerges from search-based models of housing churn. Here, a high sales/inventory ratio causes higher prices and a low ratio generates lower prices. Thus we arrived at a more complete description of the housing market at equilibrium – as shown with the two schedules in Figure 3.

Figure 3 offers a compelling explanation for why in the data, the simple price-sales correlation is so overwhelmingly positive. Over time it must be the "price based sales" schedule that is shifting up and down. Remember that this schedule is derived mainly from the decision to enter or exit the ownership market. Easy credit availability and lower mortgage rates, for example would shift the schedule up (or out). For the same level of housing prices, easier credit increases the rent-to-own flow, decreases the own-to-rent flow, and encourages new households to own. SALES expand and the inventory contracts. The end result of course is a rise in both prices as well as sales. Contracting credit does the reverse. In the post WWII history of US housing, such credit expansions and contractions have indeed tended to dominate housing market fluctuations [Capozza, Hendershott, Mack (2004)].

Figure 3 also is useful for understanding the current turmoil in the housing market. Easy mortgage underwriting from "subprime capital" greatly encouraged expanded homeownership from the mid 1990s through 2005 [Wheaton and Nechayev, (2007)]. This generated an outward shift in the price-based-sales schedule. Most recently, rising foreclosures have expanded the rent-to-own flow and shifted the "price based sales" schedule back inward. This has decreased both sales and prices. Preventing foreclosures through credit amelioration programs theoretically would move the schedule upward again, but so could any countervailing policy of easing mortgage credit.

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APPENDIX I: Sales, Prices, Inventory statistics for Short Panel

Market Code	Market	Average Yearly Change in Real Price Index (%)	Average Inventory	Average number of Sales
1	Dallas	0.009389	27765.84	5060.031
2	Houston	0.021073	29451.69	5274.866
3	Austin	0.030452	8341.217	1846.983
4	Los Angeles	0.075205	169724.2	37086.3
5	San Francisco	0.054361	59371.1	18343.6
6	San Diego	0.060397	85596.3	14763.43
7	Riverside	0.064271	65512.58	16534.53
8	Oakland	0.05496	34458.68	10413.08
9	Ventura	0.059364	14962.35	5015.75
10	Orange County	0.066787	68704.65	14874.88
11	Akron	-0.00793	21536.21	2954.509
12	Atlanta	0.011579	251270.1	26648.73
13	Baltimore	0.064701	30307.89	6897.025
14	Columbus	0.000834	46261.74	6603.301
15	Honolulu	0.067967	7333.894	1394.813
16	Kansas City	0.010979	52400.64	9495.937
17	Las Vegas	0.042517	19149.78	9506.009
18	Louisville	0.00919	30180.93	4507.799
19	Memphis	-0.00037	38817.24	5431.602
20	Miami	0.082253	97230.11	9453.403
21	Milwaukee	0.024482	21320.8	4223.433
22	Nashville	0.018568	34115.38	6109.578
23	New York	0.063452	67426.32	31415.68
24	Oklahoma City	0.017577	32241.31	6680.985
25	Omaha	0.003076	16562.51	3143.348
26	Phoenix	0.054298	89985.95	17518.8
27	Portland	0.042035	42870.82	8640.185
28	Providence	0.056324	18498.47	2737.019
29	Richmond	0.04495	24590.73	5294.051
30	St. Louis	0.023095	29147.49	9707.496
31	Tampa	0.055711	66049.88	11035.37
32	Tucson	0.047889	10922.03	2710.266
33	Washington DC	0.068125	39808.88	10710.28

APPENDIX II

Let $\Delta p_T = [\Delta P_{1T},....., \Delta P_{NT}]$' and $\Delta s_T = [\Delta S_{1T},....., \Delta S_{NT}]$', where N is the number of markets. Let $W_T = [e, \Delta p_{T-1}, \Delta s_{T-1}, \Delta X_{i,T}]$ be the vector of right hand side variables, where e is a vector of ones. Let $V_T = [\varepsilon_{1T},..., \varepsilon_{NT}]$ be the N x 1 vector of transformed disturbance terms. Let $B = [\alpha_0, \alpha_1, \alpha_2, \beta_1, \delta_1]$' be the vector of coefficients for the equation.

Therefore,

$$\Delta p_T = W_T B + V_T \tag{1}$$

Combining all the observations for each time period into a stack of equations, we have,

$$\Delta p = WB + V . \tag{2}$$

The matrix of variables that qualify for instrumental variables in period T will be

$$Z_T = [e, \Delta p_{T-2}, \Delta s_{T-2}, \Delta X_{i,T}], \tag{3}$$

which changes with T.

To estimate B, we premultiply (2) by Z' to obtain

$$Z'\Delta p = Z'WB + Z'V . \tag{4}$$

We then form a consistent instrumental variables estimator by applying GLS to equation (4), where the covariance matrix $\Omega = E\{Z'VV'Z\}$. Ω is not known and has to be estimated. We estimate (4) for each time period and form the vector of residuals for each period and form a consistent estimator, $\widetilde{\Omega}$, for Ω. $\widetilde{B}$, the GLS estimator of the parameter vetor, is hence:

$$\widetilde{B} = [W'Z(\widetilde{\Omega})^{-1} Z'W]^{-1} W'Z(\widetilde{\Omega})^{-1} Z'\Delta p . \tag{5}$$

The same procedure applies to the equation wherein Sales (S) are on the LHS.

APPRENDIX III: Sales, Prices Statistics for long Panel

Market Code	Market	Average GRRHPI (%)	Average GREMP (%)	Average SFSALES RATE	Average GRSALES RATE (%)
1	Allentown*	2.03	1.10	4.55	4.25
2	Akron	1.41	1.28	4.79	4.96
3	Albuquerque	0.59	2.79	5.86	7.82
4	Atlanta	1.22	3.18	4.31	5.47
5	Austin	0.65	4.23	4.36	4.86
6	Bakersfield*	0.68	1.91	5.40	3.53
7	Baltimore*	2.54	1.38	3.55	4.27
8	Baton Rouge	-0.73	1.77	3.73	5.26
9	Beaumont	-1.03	0.20	2.75	4.76
10	Bellingham*	2.81	3.68	3.71	8.74
11	Birmingham	1.28	1.61	4.02	5.53
12	Boulder	2.43	2.54	5.23	3.45
13	Boise City	0.76	3.93	5.23	6.88
14	Boston MA*	5.02	0.95	2.68	4.12
15	Buffalo	1.18	0.71	3.79	2.71
16	Canton	1.02	0.79	4.20	4.07
17	Chicago IL	2.54	1.29	4.02	6.38
18	Charleston	1.22	2.74	3.34	6.89
19	Charlotte	1.10	3.02	3.68	5.56
20	Cincinnati	1.09	1.91	4.87	4.49
21	Cleveland	1.37	0.77	3.90	4.79
22	Columbus	1.19	2.15	5.66	4.61
23	Corpus Christi	-1.15	0.71	3.42	3.88
24	Columbia	0.80	2.24	3.22	5.99
25	Colorado Springs	1.20	3.37	5.38	5.50
26	Dallas-Fort Worth-Arlington	-0.70	2.49	4.26	4.64
27	Dayton OH	1.18	0.99	4.21	4.40
28	Daytona Beach	1.86	3.06	4.77	5.59
29	Denver CO	1.61	1.96	4.07	5.81
30	Des Moines	1.18	2.23	6.11	5.64
31	Detroit MI	2.45	1.42	4.16	3.76
32	Flint	1.70	0.06	4.14	3.35
33	Fort Collins	2.32	3.63	5.82	6.72
34	Fresno CA*	1.35	2.04	4.69	6.08
35	Fort Wayne	0.06	1.76	4.16	7.73
36	Grand Rapids MI	1.59	2.49	5.21	1.09
37	Greensboro NC	0.96	1.92	2.95	7.22

38	Harrisburg PA	0.56	1.69	4.24	3.45
39	Honolulu	3.05	1.28	2.99	12.66
40	Houston	-1.27	1.38	3.95	4.53
41	Indianapolis IN	0.82	2.58	4.37	6.17
42	Jacksonville	1.42	2.96	4.60	7.23
43	Kansas City	0.70	1.66	5.35	5.17
44	Lansing	1.38	1.24	4.45	1.37
45	Lexington	0.67	2.43	6.23	3.25
46	Los Angeles CA*	3.51	0.99	2.26	5.40
47	Louisville	1.48	1.87	4.65	4.53
48	Little Rock	0.21	2.22	4.64	4.63
49	Las Vegas	1.07	6.11	5.11	8.14
50	Memphis	0.46	2.51	4.63	5.75
51	Miami FL	1.98	2.93	3.21	6.94
52	Milwaukee	1.90	1.24	2.42	5.16
53	Minneapolis	2.16	2.20	4.39	4.35
54	Modesto*	2.81	2.76	5.54	7.04
55	Napa*	4.63	3.27	4.35	5.32
56	Nashville	1.31	2.78	4.44	6.38
57	New York*	4.61	0.72	2.34	1.96
58	New Orleans	0.06	0.52	2.94	4.80
59	Ogden	0.67	3.25	4.22	6.08
60	Oklahoma City	-1.21	0.95	5.17	3.66
61	Omaha	0.65	2.03	4.99	4.35
62	Orlando	0.88	5.21	5.30	6.33
63	Ventura*	3.95	2.61	4.19	5.83
64	Peoria	0.38	1.16	4.31	6.93
65	Philadelphia PA*	2.78	1.18	3.52	2.57
66	Phoenix	1.05	4.41	4.27	7.49
67	Pittsburgh	1.18	0.69	2.86	2.75
68	Portland*	2.52	2.61	4.17	7.05
69	Providence*	4.82	0.96	2.83	4.71
70	Port St. Lucie	1.63	3.59	5.60	7.18
71	Raleigh NC	1.15	3.91	4.06	5.42
72	Reno	1.55	2.94	3.94	8.60
73	Richmond	1.31	2.04	4.71	3.60
74	Riverside*	2.46	4.55	6.29	5.80
75	Rochester	0.61	0.80	5.16	1.01
76	Santa Rosa*	4.19	3.06	4.90	2.80
77	Sacramento*	3.02	3.32	5.51	4.94

78	San Francisco CA*	4.23	1.09	2.61	4.73
79	Salinas*	4.81	1.55	3.95	5.47
80	San Antonio	-1.03	2.45	3.70	5.52
81	Sarasota	2.29	4.25	4.69	7.30
82	Santa Barbara*	4.29	1.42	3.16	4.27
83	Santa Cruz*	4.34	2.60	3.19	3.24
84	San Diego*	4.13	2.96	3.62	5.45
85	Seattle*	2.97	2.65	2.95	8.10
86	San Jose*	4.34	1.20	2.85	4.55
87	Salt Lake City	1.39	3.12	3.45	5.72
88	St. Louis	1.48	1.40	4.55	4.82
89	San Luis Obispo*	4.18	3.32	5.49	4.27
90	Spokane*	1.52	2.28	2.81	9.04
91	Stamford*	3.64	0.60	3.14	4.80
92	Stockton*	2.91	2.42	5.59	5.99
93	Tampa	1.45	3.48	3.64	5.61
94	Toledo	0.65	1.18	4.18	5.18
95	Tucson	1.50	2.96	3.32	8.03
96	Tulsa	-0.96	1.00	4.66	4.33
97	Vallejo CA*	3.48	2.87	5.24	5.41
98	Washington DC*	3.01	2.54	4.47	3.26
99	Wichita	-0.47	1.43	5.01	4.39
100	Winston	0.73	1.98	2.92	5.51
101	Worcester*	4.40	1.13	4.18	5.77

Notes: Table provides the average real price appreciation over the 25 years, average job growth rate, average sales rate, and growth in sales rate.
* Denotes "Costal city" in robustness tests.