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

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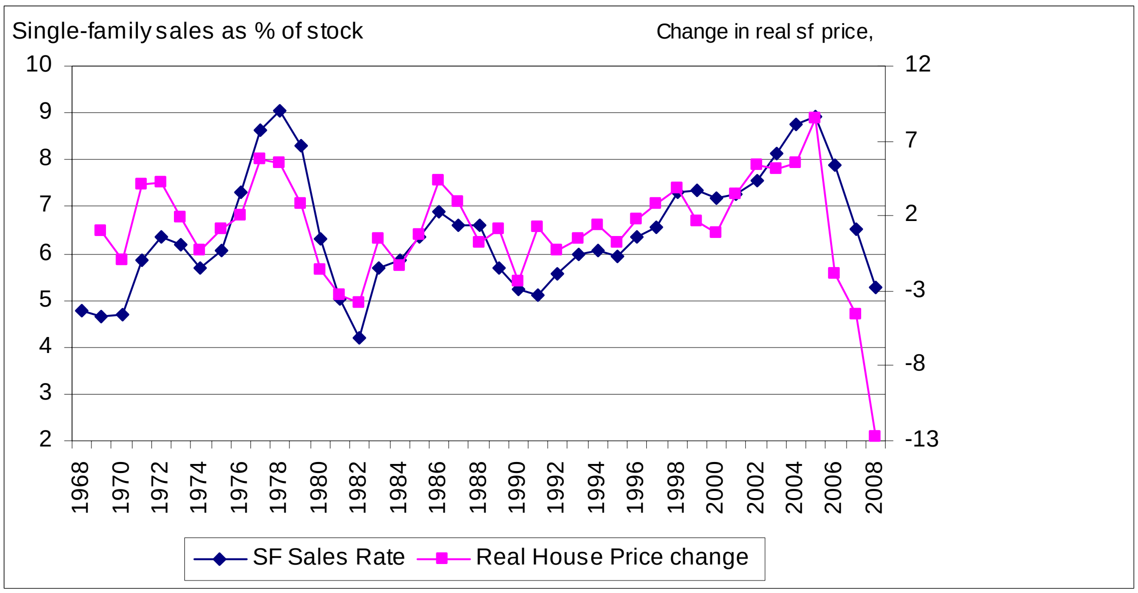
ABSTRACT

This paper examines the gross flows in the US housing market and finds that rent-to-own and own-to rent flows (which roughly offset each other) are each as large as own-to-own. A number of theories are reviewed which have focused solely on the own-to-own decision, suggesting that higher sales generate higher prices. Empirically we find that there exists a parallel relationship in which higher prices reduce sales. This most likely occurs as higher prices depress rent-to-own flows and increase own-to-rent flows. The result is not only a reduction in sales but an increasing inventory of for sale units. In effect the positive relationship between sales and prices is driven by own-to-own decisions, while the negative relationship between prices and sales results from decisions to enter or exit ownership. These two relationships receive extensive empirical support in a panel VAR of 101 MSA spanning 25 years. The two relationships together also provide a more complete picture of how the housing market operates.

I. Introduction.

As shown in Figure 1 below, there is a very strong positive correlation between housing sales (expressed as a percent of owner households) and the movement in housing prices. On the surface the relationship looks to be close to contemporaneous, and a number of authors have offered explanations for it.

Figure 1: US Housing Sales and Prices



One group of papers directly examines how owners trade housing in the presence of search frictions [Wheaton (1990), Berkovic and Goodman (1996), Lundberg and Skedinger (1999)]. As is true with most frictional market models (e.g. Pissarides 2000), increases in turnover (sales) tends to make trading easier and in a housing market composed of all owners this will increase prices. Increases in the inventory of units for sale (vacancy) have the opposite impact. Hence, in this camp the hypothesis is that positive shocks to sales will then increase prices while negative shocks will depress them.

Other authors have argued that when prices fall, homeowners have an aversion to selling at a loss - no matter how "rational" selling may be [Genesove and Mayer (2001), Englehardt (2003)]. It is unclear whether this micro-economic argument also extends to aggregate movements since those with a loss will most likely be recent movers and hence only a portion of the aggregate transaction market. Another group of papers comes to a similar conclusion, but through using liquidity and down payment constraints [Chan (2001), Stein (1995), Lamont and Stein (1999). The implication of both micro-economic theories is that after price declines, sales should be reduced, and following price increases, sales should recover.[1]

It is interesting that all of these arguments focus on the situation where existing owners wish to trade laterally into another (owned) house. What about renters? In this paper we demonstrate that the sales by owners who then become renters, and purchases by renters who then become owners – *each are almost as large a flow as sales by owners who stay owners (purchase another house).* In effect almost 50% of purchases are by households entering ownership and 50% of sales are by households exiting ownership. Thus any theory examining sales and price movements must surely incorporate these flows as well. In this paper we do so, both empirically and theoretically. Our approach involves the following contributions.

1). We provide some simple initial evidence in the aggregate data that when sales are high, so is homeownership, and so is the flow of rent-to-own relative to the opposing flow of own-to-rent.

2). When there are net entrants into ownership the inventory of units for sale is low and vice versa. Own-to-rent moves add to the inventory, while rent-to-own subtract. Own-to-own do both and hence have little impact on the inventory for sale. Thus if high prices cause sales (as well as net entrant flows and homeownership) to eventually fall, our argument is made. The simple correlation between prices on the one hand and these variables on the other (net flows, homeownership, inventory) is positive however (rather than negative), but then these are the relationships that are highly subject to simultaneity.

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[1] There have been a few recent attempts test whether the relationship between movements in sales and prices support one, or the other, or both theories. This is complicated by the fact that both theories predict positive relationships, just with different timing. 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.

3). Using a panel granger causality test across 25 years and more than a hundred metropolitan areas, we show unequivocally that in fact housing prices *negatively* "Granger cause" housing sales. This convincingly reinforces the view that the net flow into ownership is negatively impacted by housing prices.

4). In our panel analysis there also is overwhelming evidence that higher sales "Granger cause" rising prices. This reinforces the relationship posited in frictional search models wherein prices are driven positively by the ratio of sales/inventory (the inverse of sales duration or time-to-sale). Here we show that movement in and out of ownership creates a simultaneous relationship wherein higher prices dampen sales and increase the inventory. The full equilibrium solution to the operation of the housing market is the intersection of these two schedules.

Our paper is organized as follows. In section II we review gross housing flows from the AHS as well as the Census. This documents the magnitude of the rent-to-own and own-to-rent flows. We examine some survey questions in the census to try and ascertain why these gross flows into and out of ownership are so large (and generally close to offsetting). We also study the limited aggregate time series data to show the hypothesized relationships between sales, net ownership and inventory changes. In sections III – V we expand the range of our empirical analysis to a full panel study of the movement between sales and prices across 101 MSA covering the years from 1982-2006. Here 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 tests. Finally, in section VI we present a simple analytic framework for studying the equilibrium outcome of the two simultaneous relationships – and its comparative statics.

Our conclusion is that since the contemporaneous correlation between sales and prices is so strong, that the market must be driven largely by exogenous changes to net flows or homeownership. The most recent housing "bubble" and its painful "bust" provide an excellent example. Positive changes in mortgage availability shifted the (negative) price-sales schedule outward which then increases both prices and sales as the market moves along the (positive) search-based schedule. In the last two years, the reverse shift (from foreclosures) has done the opposite – dropping both sales and prices.

II. US Housing Flows.

Much of the literature discussed above investigates how existing homeowners behave as they try and sell their current home to purchase a new one. This flow is often referred to as "churn". To investigate how important a role such "churn" plays in the ownership market, we examined the 2000 Census, as well as the AHS for the same year. The Census has quite accurate counts of overall mobility, while the AHS goes into more detail about those households who did move – in particular their household and occupancy status in the previous year. Putting the two sources together we have created Figure 2, which estimates US gross housing flows in that year. The only other source of similar data is that reported by the National Association of Realtors (NAR) but it has no information on buyers or their previous status.

In the 2000 Census it appears that about 7,186,000 of US current owners moved over the previous 15 months, which annualizes to 5,186,000 or about 8% of owners. Annualized renter mobility is a far greater – at 30% or 11,040,000. The AHS interestingly shows that owners are almost as likely to have come from renting as from owning, but renters are 3 times more likely to have come from renting as owning. Since owners are twice as numerous as renters, these tenure change flows roughly offset each other. What is interesting is that own-to-rent as well as rent-to-own flows are virtually as large as own-to-own at least over a 15 month interval. In the 1990 Census, these patterns are also virtually identical. Focusing on the owned housing market, the Census and AHS allow us to account for virtually all of the events that add to the inventory of houses for sale (herein called LISTS) and all of those transactions that remove houses from the inventory (herein called SALES). The exception is new housing units. In 2000 roughly 800,000 units were delivered vacant to the for-sale market and these must be counted as additional LISTS. A problem with the Census data is that it does not totally reconcile with the NAR data on sales (which if anything undercounts - see later). Rent-to-own plus own-to-own is almost 700 thousand shy of known NAR sales. [2]

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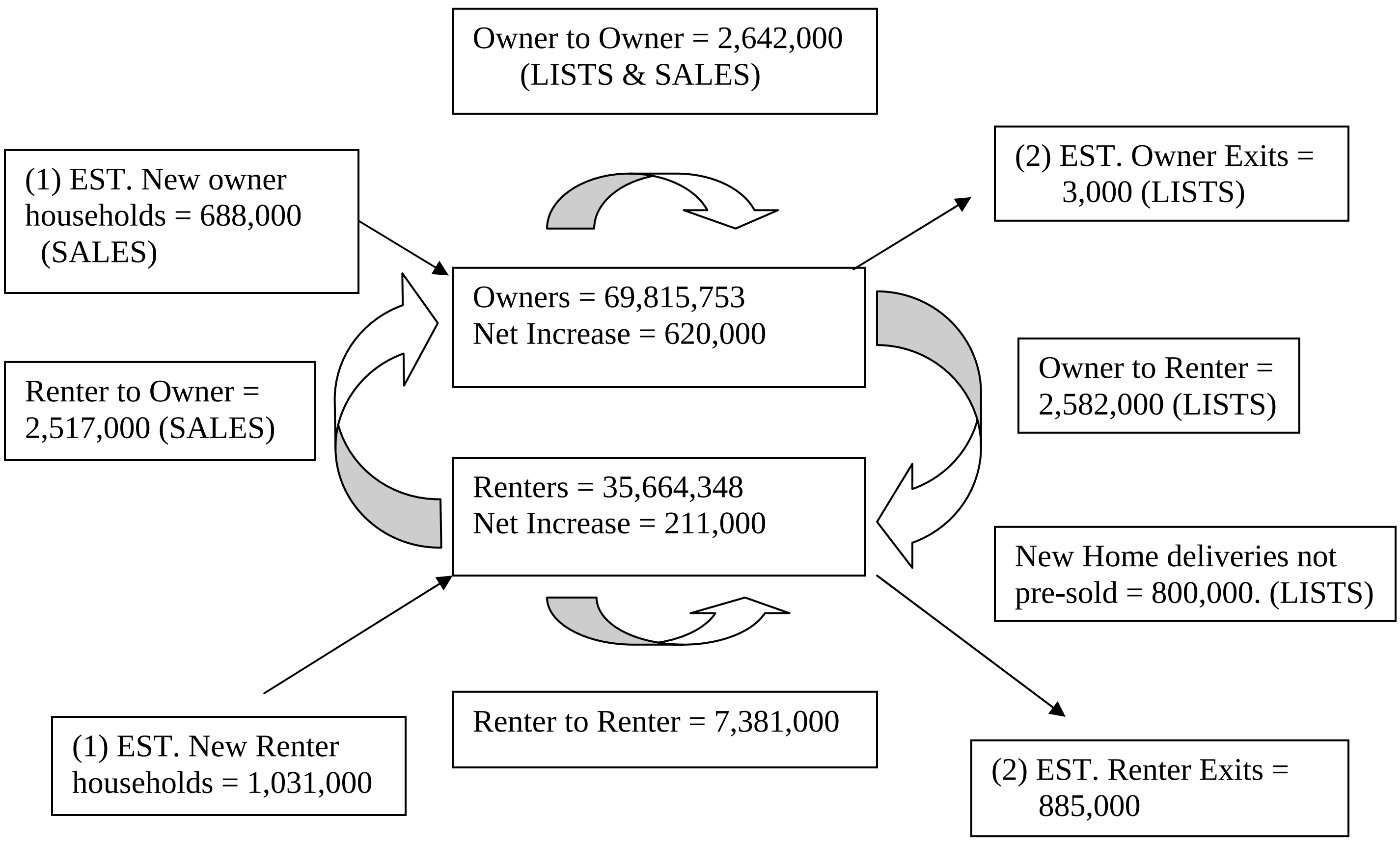
[2] NAR Single Family sales in 1999 were 5.4 million. To this figure we must also add roughly .5m of condominium sales.

The NAR also reports that in 1999-2000 the inventory of units for sale is nearly stable. Hence in that year total lists will have to roughly equal total sales. Using the following identities we can estimate the flow of *new households* into owning and renting, as well as the death or exist of households from each tenure category as residuals.

SALES = Owner-Owner + Renter-Owner + New Owner = 5.847million
LISTS = Owner-Owner + Owner-Renter + Owner Deaths + Deliveries = 6.027million

We also know that the formation and death of total households (including immigration) has been roughly: 0.8% deaths, 1.6% gross increases, with a net increase of around 0.8%. From the AHS we also can estimate the breakdown of household formation (those households not reporting previous headship) into tenure categories. More complicated, however, is the division of household "exits" by tenure. More than 80% of personal deaths occur in some institution, but the "death" of a household actually occurs with the move *into* the institution. In Figure 2 we have calculated household exits by tenure as a residual from the more solid data on moves, as well as gross and net household formation. It is interesting that the residual calculation suggests virtually all exits from "household" occur among renters.

Figure 2: US Housing Gross Flows (2000)



Notes:

(1): Total gross household formation is estimated to be 1,719,000 from known population birth/death rates and the known net increase in households. The AHS reports (average of 1999 and 2001), 40% of newly created households were owners and 60% renters.

(2). Total household exits are estimated to be 888,000 from known population birth/death rates and the net increases in households. Tenure allocations are calculated as a residual from new households, net tenure changes and known net increases.

What is most interesting to us is that more than 50% of SALES involve a purchaser that one would expect to be sensitive (negatively) to housing prices. When prices are high presumably new owner household formation may be discouraged as would be moves which involve changes from renting to owning. On the other side, more than 50% of LISTS involve decisions that should be positively sensitive to price. New deliveries certainly try to occur when prices are high, and such periods are also ideal times for owners to "cash out" and switch to renting. This presents a picture that is quite different from the models of Stein, Wheaton, or Berkovec and Goodman – where sales transactions are largely based on owners transitioning between different owned houses with little sensitivity to price – since households are both seller as well as buyer.

The gross flows of own-to-rent and rent-to-own are very large and must involve many type of housing "churn". The most common and simplistic view of the housing market has all households starting as renters and then permanently switching at some single point to owners. Were this so, in the steady state the rent-to-own flow would be no larger than new household formations – 1million or so and not 2.5 million. Similarly if owners remained so until death (or institutionalization) there would be little or no flow from own-to rent – only owner "deaths". Clearly there must be an enormous amount of economic or demographic change that leads households to move back and forth across tenure categories.

The Census survey of "reasons for moving" does provide some evidence to this effect. In Table 1 below we examine the "primary reason" given by each type of mover in the 2000 census. There are several distinguishing patterns. First lets examine the own to rent flow (versus the own to own). It is clear that owners move to renting primarily for job transfers and demographic changes (divorce, family changes, establish own household). The rent-to-own flow (versus rent-to-rent) is more like to involve a desire for more/better housing, establish an independent household (again) and a desire to own per se. One interesting hypothesis that cannot be evaluated with the AHS data is that many own-to-rent moves are actually linked with a later rent-to-own move, for example when households change cities. Such linked moves allow households to sell first, rent, and then buy. This is in contrast to the search models where owners must buy first, own two homes and then sell. Since the AHS unfortunately examines only a single (the most recent) move, we cannot directly determine the extent of such linked moves.

TABLE 1: Reasons for Moving (2000 AHS)

	Own to Own	Own to Rent	Rent to Own	Rent to Rent
New job or job transfer	10.23%	16.81%	7.34%	13.98%
To be closer to work/school/other	10.00%	8.62%	5.68%	11.85%
Other, financial/employment related	2.79%	4.74%	2.10%	4.16%
To establish own household	3.26%	6.47%	12.59%	5.83%
Needed a larger house or apartment	24.88%	2.16%	17.40%	12.77%
Married, widowed, divorced, or separated	8.60%	22.41%	4.90%	5.60%
Other, family/personal related	4.88%	8.62%	3.50%	5.78%
Wanted a better quality house (apartment)	14.88%	1.29%	9.70%	8.79%
Change from owner to renter OR renter to owner	1.86%	7.76%	20.80%	1.20%
Wanted lower rent or less expensive house to maintain	3.26%	2.16%	2.45%	6.20%

At the US aggregate level there is mixed evidence that the own-to-rent and rent-to-own flows are playing the role that we assert in the discussion above. Since there is high simultaneity between these variables we examine only simple correlations. In Table 2 we have calculated the difference between rent-to-own moves and own-to-rent moves (RO – OR) from the AHS in each year. We examine correlations between this and the designated variables. Consistent with our argument is that fact that sales are higher during periods of net owner entrants and that this is associated with years of higher home ownership and a low inventory of for-sale units. The only correlation at odds with the argument is that between real price levels and RO-OR. Here we hypothesized that when prices are high RO moves are less relative to OR, but the sample for these correlations is very short (1985-2007) and includes almost a decade during which homeownership grew rapidly and steadily due to subprime lending. Any reasonably serious examination of our hypothesis that there is a pair of opposing relationships between housing sales and housing prices must surely rely on a larger and more robust data set.

Table 2: 1985-2007 correlations

	RO-OR	Real House Price	Sales Rate	Inventory	Home ownership rate
RO-OR	1				
Real Housing Price Index	0.3923	1			
Sales Rate	0.8369	0.6282	1		
Inventory	-0.2811	0.738	0.0969	1	
Homeownership rate	0.7609	0.8309	0.8416	0.3277	1

III. Metropolitan Level Sales and Prices.

To more carefully study the relationship(s) between housing sales and housing prices we have assembled a large panel data base covering 101 MSA and the years 1980 through 2006. Examining data at the metropolitan level, however is also far more difficult and cannot rely on Census or AHS data. The latter is available only every decade and the AHS sample is too small to generate reliable flows at the MSA level.

For sales data, 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), but is available for each MSA over the full period from 1980 to 2006. To standardize the sales data, raw sales were compared with annual Census estimates of the number of total households in those markets. Dividing single family sales by total households we get a very crude sales rate for each market. In 1980 this calculated sales rate varied between 1.2% and 5.1% across our markets with a national average value of 2.8%. By contrast, in the 1980 census, 8.1% of owner occupied households had moved in during the last year. By 2000, the ratio of national NAR single family sales to total households had risen to 4.9%, while the Census owner mobility rate just inched up to 8.9%. Of course our crude calculated average sales rates should always be lower than the census reported owner mobility rates since the former excludes condo transactions and non-brokered sales. In addition we are dividing by total households rather than just single family owner-occupied households. Separate renter/owner single family household series at *yearly frequency* are just not available by metropolitan market.

The price data we use is the 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)]. These omissions could generate a significantly bias in the long term trend of the OFEHO series. 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 to purchase similar indices from CSW/FISERV, although they have most of the same methodological issues as the OFHEO data.

In Figures 3 and 4 we illustrate the yearly NAR sales rate data, along with the constant dollar OFHEO price series – both in levels and differences - for two markets that exhibit quite varied behavior, Atlanta and San Francisco. Over this time frame, Atlanta's constant dollar prices increase very little while San Francisco's increased almost 200%. San Francisco prices, however, exhibit far greater price volatility. Atlanta's average sales rate is close to 5% and *triples* over 1980-2006, while San Francisco's is almost half of that (2.6%) and increases by only 50%. These trends illustrate the typical range of patterns seen across our sample of 101 metropolitan areas. In appendix I we present the summary statistics for each market's price and sales rate series.

Figure 3: Atlanta

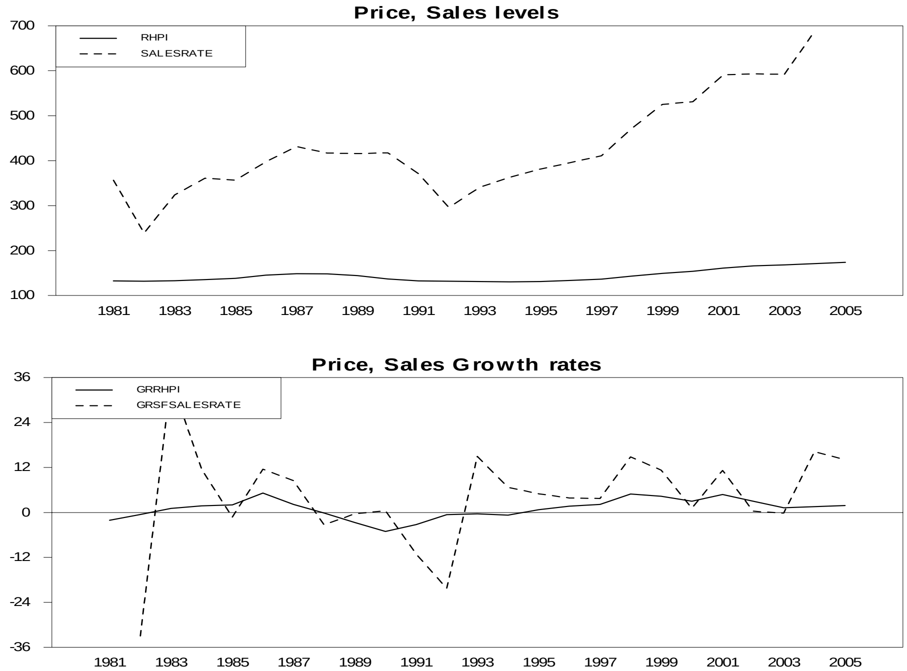
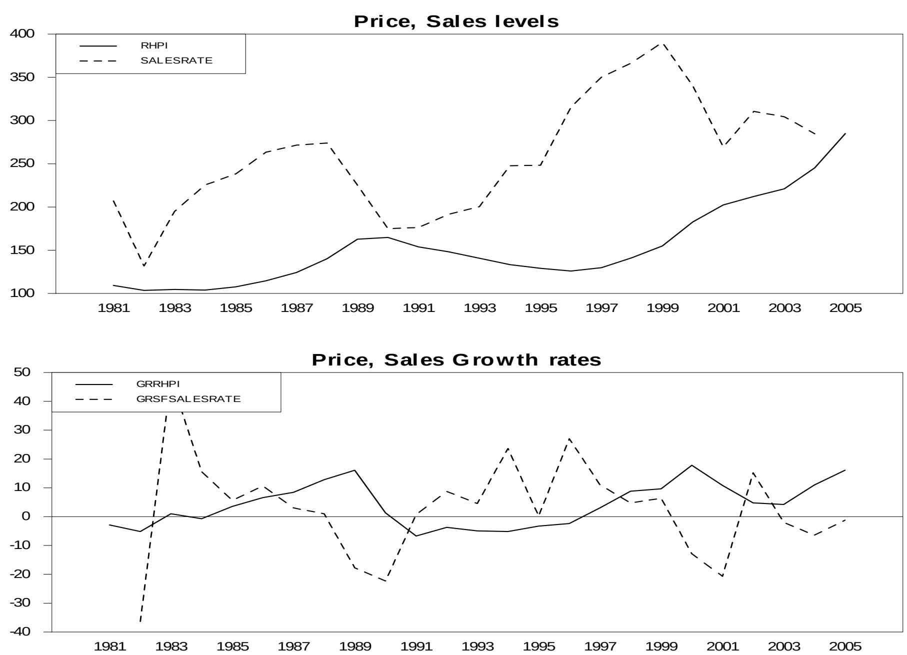


Figure 4:San Francisco



Given the persistent trends in both series it is useful and important to test for series stationarity. There are two tests available for use with panel data such as we have. In each, the null hypothesis is that *all* of the individual series have unit roots and are non stationary. Levin-Lin (1993) and Im-Persaran-Shin (2002) both develop a test statistic for the sum or average coefficient of the lagged variable of interest – across the individuals (markets) within the panel. The null is that all or the average of these coefficients is not significantly different from unity. In Table 3 we report the results of this test for both housing price and sale rate levels, as well as a 2nd order stationarity test for housing price and sales rate changes.

TABLE 3: Stationarity tests

RHPI (Augmented by 1 lag)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.10771	-18.535	0.22227	0.5879
First Difference	-0.31882	-19.822	-0.76888	0.2210
IPS test	T-Bar	W(t-bar)		P>T
Levels	-1.679	-1.784		0.037
First Difference	-1.896	-4.133		0.000

SFSALESRATE (Augmented by 1 lag)

Levin Lin's Test	Coefficient	T Value	T-Star	P>T
Levels	-0.15463	-12.993	0.44501	0.6718
First Difference	-0.92284	-30.548	-7.14975	0.0000
IPS test	T-Bar	W(t-bar)		P>T
Levels	-1.382	1.426		0.923
First Difference	-2.934	-15.377		0.000

With the Levin-Lin test we cannot reject the null (non-stationarity) for either house price levels or differences. In terms of the sales, we can reject the null for differences in sales rate differences, but not for levels. The IPS test (which is argued to have more power) rejects the null for house price levels and differences and for sales rate differences. In short, both variables would seem to be stationary in differences, but levels are more problematic and likely non-stationary.

IV. Panel Estimation Approaches.

Our panel approach uses a well-known application of Granger-type analysis. We will ask how significant lagged sales are in a panel model of prices which uses lagged prices and then several conditioning variables. The conditioning variables we choose are market area employment, and national mortgage rates. The companion model is to ask how significant lagged prices are in a panel model of sales using lagged sales and the same conditioning variables. This pair of model is shown (1)-(2).

$$P_{i,T} = \alpha_0 + \alpha_1 P_{i,T-1} + \alpha_2 S_{i,T-1} + \beta' X_{i,T} + \delta_\iota + \varepsilon_{\iota,\mathrm{T}} \qquad (1)$$

$$S_{i,T} = \gamma_0 + \gamma_1 S_{i,T-1} + \gamma_2 P_{i,T-1} + \lambda' X_{i,T} + +\eta_i + \varepsilon_{\iota,\mathrm{T}} \qquad (2)$$

In panel models, all of the estimation issues raised in time series continue to exist. In our case there is concern about the stationarity of both price and sales rate levels. This same concern is normally not present for differences. Hence we will need to estimate the model in first differences as well as levels – as outlined in equations (3) and (4).[3]

$$\Delta P_{i,T} = \alpha_0 + \alpha_1 \Delta P_{i,T-1} + \alpha_2 \Delta S_{i,T-1} + \beta' \Delta X_{i,T} + +\delta_\iota + \varepsilon_{\iota,\mathrm{T}} \qquad (3)$$

$$\Delta S_{i,T} = \gamma_0 + \gamma_1 \Delta S_{i,T-1} + \gamma_2 \Delta P_{i,T-1} + \lambda' \Delta X_{i,T} + +\eta_i + \varepsilon_{\iota,\mathrm{T}} \qquad (4)$$

In panel VAR models with individual heterogeneity there exists a specification issue. Equations (3) or (4) will have an error term that is correlated with the lagged dependent variables [Nickell, (1981)]. OLS estimation will 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 (the α and γ) may not be reliable. The bias problem

[3] In (3) and (4) the fixed effects are cross-section trends rather than cross section levels as in (1) and (2)

might not be serious in our case since we have 26 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.

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 the α_2 and γ_2) as the check of whether changes in sales "Granger cause" changes in price and vice versa.

In table 4 we report the results of equations (1) through (4) in each set of rows. The first column uses OLS estimation, the second the Random Effects IV estimates from Holtz-Eakin et al. 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 differences are indicated with the prefix GR.

Among the levels equations, we first notice that the two conditioning variables, the national mortgage rate and local employment can have the wrong signs – here in two cases. The mortgage interest rate in the OLS price levels equation and local employment in the IV sales rate equation are miss-signed. There is also an insignificant employment coefficient in the OLS sales rate equation (despite almost 2500 observations). Another troublesome result is that the price levels equation has excess "momentum" – lagged prices have a coefficient greater than one. Hence prices (levels) might 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. Interestingly, the two estimation techniques yield quite similar coefficients – as might be expected with a larger number of time series observations.

When we move to the results of estimating the equations in differences these issues all disappear. The lagged price coefficients are small, the price equations 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 significant. Furthermore in every sales rate or growth in sales rate equation, lagged prices (or its growth) are also always significant. Hence there is clear evidence of joint causality, *but 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. The impact is exactly the opposite of that predicted by theories of loss aversion or liquidity constraints, and completely consistent with our argument that entrants and exits from ownership drive sales.

TABLE 4: Sales-Price VAR

	Fixed Effects	E Holtz-Eakin estimator
<u>Levels</u>		
Real Price (Dependent Variable)		
Constant	-25.59144**	-12.47741**
	(2.562678)	(2.099341)
Real Price (lag 1)	1.023952**	1.040663**
	(0.076349)	(0.0076326)
Sales Rate (lag 1)	3.33305**	2.738264**
	(0.2141172)	(0.2015346)
Mortgage Rate	0.3487804**	-0.3248508**
	(0.1252293)	(0.1209959)
Employment	0.0113145**	0.0015689**
	(0.0018579)	(0.0003129)
Sales Rate (Dependent Variable)		
Constant	2.193724**	1.796734**
	(0.1428421)	(0.1044475)
Real Price (lag 1)	-0.0063598**	-0.0059454**
	(0.0004256)	(0.0004206)
Sales Rate (lag 1)	0.8585273**	0.9370184**
	(0.0119348)	(0.0080215)
Mortgage	-0.063598**	-0.0664741**
	(0.0069802)	(0.0062413)
Employment	-0.0000042	-0.0000217**
	(0.0001036)	(0.0000103)
<u>First Difference</u>		

GR Real Price (Dependent Variable)		
Constant	-0.4090542**	-0.49122**
	(0.1213855)	(0.1221363)
GR Real Price (Lag 1)	0.7606135**	0.8008682**
	(0.0144198)	(0.0148136)
GR Sales Rate (Lag 1)	0.0289388**	0.1826539**
	(0.0057409)	(0.022255)
GR Mortgage Rate	-0.093676**	-0.08788**
	(0.097905)	(0.0102427)
GR Employment	0.3217936**	0.1190925**
	(0.0385593)	(0.048072)
GR Sales Rate (Dependent Variable)		
Constant	0.7075247	1.424424**
	(0.3886531)	(0.3710454)
GR Real Price (Lag1)	-0.7027333**	-0.8581478**
	(0.0461695)	(0.0556805)
GR Sales Rate (Lag 1)	0.0580555**	0.0657317**
	(0.0183812)	(0.02199095)
GR Mortgage Rate	-0.334504**	-0.307883**
	(0.0313474)	(0.0312106)
GR Employment	1.167302**	1.018177**
	(0.1244199)	(0.1120497)

** 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) and collectively "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, we offer up Table 5 where price changes are tested against the level of sales (as a rate).

TABLE 5: Sales Price Mixed VAR

Differences and Levels	Fixed Effects	E Holtz-Eakin estimator
GR Real Price (Dependent Variable)		
Constant	-6.61475**	-1.431187**
	(0.3452743)	(0.2550279)
GR Real Price (lag 1)	0.5999102**	0.749431**
	(0.0155003)	(0.0141281)
Sales Rate (lag 1)	1.402352**	0.2721678**
	(0.0736645)	(0.0547548)
GR Mortgage Rate	-0.1267573**	-0.0860948**
	(0.0092715)	(0.0095884)
GR Employment	0.5059503**	0.3678023**
	(0.0343458)	(0.0332065)
Sales Rate (Dependent Variable)		
Constant	-0.0348229	0.0358686
	(0.0538078)	(0.0026831)
GR House Price (lag 1)	-0.0334235**	-0.0370619**
	(0.0024156)	(0.0026831)
Sales Rate (lag 1)	1.011515**	1.000989**
	(0.0114799)	(0.0079533)
GR Mortgage Rate	-0.0162011**	-0.0151343**
	(0.0014449)	(0.0014294)
GR Employment	0.0494462**	0.043442**
	(0.0053525)	(0.0049388)

** 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 (rather than rises). All conditioning variables are significant and correctly signed and lagged dependent variables have coefficients less than one.

V. 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 markets into two groups: so-called "coastal" cities that border both oceans 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 6. The second test is to divide the sample up by year – in this case we estimate separate models for 1980-1992 and 1993-2006. The year 1992 generally marks the bottom of the housing market from the 1990 recession. These results are depicted in Table 7. Both tests use just the differences model that seems to provide the strongest results in the previous section.

TABLE 6: Geographic Sub Panels

	Fixed Effects		E Holtz-Eakin estimator	
	Coastal MSA	Interior MSA	Coastal MSA	Interior MSA
GR Real Price (Dependent Variable)				
Constant	-0.6026028 (0.2974425)	-0.274607** (0.1132241)	-0.543562 (0.3332429)	-.338799** .1054476
GR Real Price (Lag 1)	0.7661637** (0.0255794)	0.7731355** (0.0178884)	0.855731** (0.0351039)	.7834749** .0171874
GR Sales Rate (Lag 1)	0.0608857** (0.0141261)	.0094349* (0.0054047)	0.3475212** (0.0573584)	.0799289** .0198759
GR Mortgage Rate	-0.106036** (0.023653)	-.0866954** (0.0092136)	-0.112101** (0.0278593)	-.0776626** .008816
GR Employment	0.5717489** (0.0978548)	.1978858** (0.0359637)	-0.0434497 (0.153556)	.1617733** .0381004
GR Sales Rate (Dependent Variable)				
Constant	2.098906** (0.7412813)	0.0396938** (0.4541917)	3.03388** (0.7426378)	0.8084169* (0.4261651)
GR Real Price (Lag1)	-0.8320889** (0.0637485)	-0.5447358** (0.0637485)	-0.9763902** (0.0798291)	-0.8519448** (0.0919725)
GR Sales Rate (Lag 1)	-0.0004387 (0.0352049)	0.0770193** (0.0216808)	-0.0350817 (0.0402424)	0.1111637** (0.0251712)
GR Mortgage Rate	-0.2536587** (0.0589476)	-0.3772017** (0.0369599)	-0.2390963** (0.0595762)	-0.3323406** (0.036746)
GR Employment	1.265286** (0.2438722)	1.172214** (0.1442662)	1.102051** (0.2223687)	1.03251** (0.1293764)

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 6, the results of Table 4 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 not only always significant, but 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 7). 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 7, the Holtz-Eakin estimates are now sometimes quite different than the OLS results.

TABLE 7: Time Subpanels

	Fixed Effects		E Holtz-Eakin estimator	
	1982-1992	1993-2005	1982-1992	1993-2005
GR Real Price (Dependent Variable)				
Constant	-2.63937** (0.2362837)	-0.1053808 (0.1453335)	-1.237084** (0.2879418)	-0.2731544 (0.1943765)
GR Real Price (Lag 1)	0.5521216** (0.0271404)	0.9364014** (0.0183638)	0.6752733** (0.0257512)	0.9629539** (0.0196925)
GR Sales Rate (Lag 1)	0.0194498** (0.0073275)	0.0363384** (0.0097935)	0.1622147** (0.0307569)	0.0874362 ** (0.0307703)
GR Mortgage Rate	-0.2315352** (0.0193262)	-0.0707981** (0.0116032)	-0.1432255** (0.0244255)	-0.0812995** (0.0163056)
GR Employment	0.6241497** (0.063533)	0.4310861** (0.0501575)	0.157348* (0.0910416)	0.3441402** (0.0493389)

GR Sales Rate (Dependent Variable)				
Constant	-6.269503** (0.9018295)	4.398222** (0.447546)	-4.898023** (0.8935038)	3.00473** (0.4587499)
GR Real Price (Lag1)	-0.8795382** (0.1035874)	-0.5704616** (0.0565504)	-1.080492** (0.1243784)	-0.4387881** (0.066557)
GR Sales Rate (Lag 1)	0.0056823 (0.027967)	-0.025242 (0.0301586)	-0.0035275 (0.0350098)	0.066557 (0.029539)
GR Mortgage Rate	-0.5636095** (0.0737626)	-0.1934848** (0.0357313)	-0.550748** (0.0819038)	-0.2720118** (0.0420076)
GR Employment	2.608423** (0.2424878)	0.4856197** (0.154457)	2.026295** (0.2237316)	0.7631351** (0.1325586)

Note:

a) Column labeled under 1982-1992 refer to the results using observations that span from 1982 to 1992.
b) Column labeled under 1993-2005 refer to the results using observations that span from 1993 to 2005.

VI. A More Complete Model of Housing Sales Volume and Prices

As discussed initially, 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 must always become sellers – there are no entrants or exits from the market. In such a situation prices are like “funny money” – when participants pay higher prices they also receiving more upon sale. It is only the transaction cost of owning 2 homes (during the moving/trade period) that grounds prices. If prices are high, the transaction costs can make moving so expensive as to erase whatever gains from moving were there originally. In this environment Nash-bargained prices move almost inversely to expected sales times - where the latter equals vacancy divided by the sales flow. In these models, both vacancy and sales churn are exogenous. Following Pissarides (2000) if the matching rate is exogenous or alternatively if the matching function is of specific form, sales time will be shorter with more churn and prices therefore higher. Hence greater sales cause higher prices. Similarly greater vacancy (inventory) raises sales times and causes lower prices.

There are also a series of paper’s which propose a relationship in which changes in prices will subsequently generate higher sales volumes. This again is a positive relationship between the two, 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 consumers are again moving from one house to another (market "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 to make the 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 will experience a loss when they sell tend to set higher reservations than those who will not experience a loss. With higher reservations, the market would see lower sales – if more and more sellers experience loss aversion as prices continue to drop. As long as prices are rising, however, the theory makes little prediction about what will happen to sales.

We have shown (hopefully convincingly) that the causal relationship between prices generating sales is actually negative – rather than positive. Our empirics are quite strong. To explain this, we have argued that actual flows in the housing market are remarkably large between tenure groups – and that a negative price-sales relationship makes sense in explaining these inter-tenure flows. Higher prices lead households to exit more than enter homeownership and lower housing prices do the reverse. When entrants exceed exits sales increase and the inventory declines. More exits than entrants will generate lower sales and a growing inventory.

Our empirical analysis also overwhelmingly supports the search-based models of house pricing. Here, a high sales/inventory ratio causes higher prices and a low ratio generates lower prices. Thus we arrive at a more complete equilibrium model of the housing market – as shown with the two schedules below in Figure 5.

FIGURE 5: Housing Market Equilibrium(s)

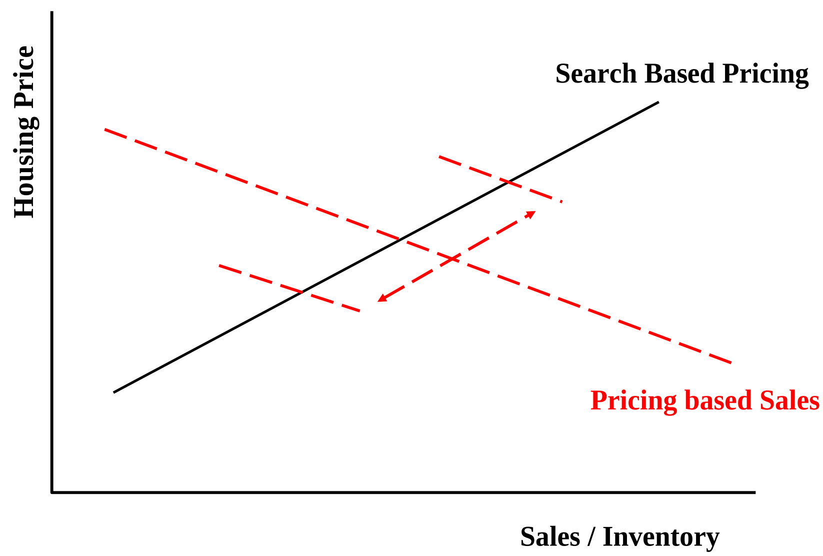


Figure 5 provides a compelling explanation for why in the data, the simple correlation between prices and sales is strongly positive. Over time it must be the “price based sales” schedule that is shifting up and down. Remember that this schedule is derived from the decision to enter or exit the ownership 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 and decreases the own-to-rent flow. 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 market fluctuations [Capozza, Hendershott, Mack (2004)].

Figure 5 also is useful for understanding the current troubles in the housing market. Rising foreclosures expand the rent-to-own flow and shift the “price based sales” schedule down (or in). This has decreased both sales and prices. Preventing foreclosures through credit amelioration would move the schedule upward again, but so could a countervailing policy of easing mortgage credit. This would most likely create an increase in the countervailing rent-to-own flow. It is interesting to speculate on whether there might be some policy that would shift the “search based pricing” schedule upward.

This would restore prices, although it would not increase sales. For example some policy to encourage interest-free bridge loans would certainly make it easier for owners to "churn". Likewise some form of home sales insurance might reduce the risk associated with owning two homes. That said, such policies would seem to be a less direct way of restoring home prices versus a stimulus to the "price-based-sales" schedule.

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APPRENDIX I

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.

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,

$$\Delta p_T = W_T B + V_T \quad (1)$$

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

$$\Delta p = WB + V \ . \quad (2)$$

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

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

which changes with T.

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

$$Z'\Delta p = Z'WB + Z'V \ . \quad (4)$$

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, $\tilde{\Omega}$, for Ω. $\tilde{B}$, the GLS estimator of the parameter vetor, is hence:

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

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