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Local Market and National Components in House Price Appreciation

Authors
Gyourko, Joseph
Voith, Richard

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Joseph Gyourko
Associate Professor
Finance Department and
Real Estate Unit
The Wharton School
University of Pennsylvania
and
Visiting Associate Professor
Anderson Graduate School
of Management
UCLA

Richard Voith
Senior Economist
Federal Reserve Bank
of Philadelphia
and
Visiting Assistant Professor
Finance Department and
Real Estate Unit
The Wharton School
University of Pennsylvania

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ABSTRACT

We analyze real home price appreciation using a long time series (1971-1989) and large cross section (56 metro areas). Our findings yield important new insights into two outstanding issues in real estate finance and economics. The first deals with the implications for investment opportunities in housing across metro areas. A striking result is that we cannot reject the null hypothesis of equal appreciation rates across locales. **A priori**, the results are suggestive of equal expected appreciation across the different local markets. However, it is noteworthy that we find significant serial correlation in some local appreciation series. This is consistent with previous findings by Case and Shiller (1989), and suggests that prescient market timers might have been able to make money in selective markets.

We then consider the potential implications of equal appreciation rates across cities for housing market equilibrium in light of the fact that price levels do materially differ across metro areas. We argue that equal real appreciation rates starting from different price levels imply increasingly divergent prices of local traits in terms of foregone consumption. Without special assumptions with respect to income and productivity differentials across locations or to local trait income elasticities of demand, the appreciation rates in high housing price level areas ultimately have to fall below those in low price level areas. Preliminary evidence indicates that higher priced areas tend to have significantly lower appreciation rates (controlling for local fixed effects and a common, time-varying effect).

I. INTRODUCTION

Most people understand the adage that the three most important factors in real estate are 'location, location, and location'. Casual observation of home prices over the past decade has reinforced the insight of that aphorism. At various times in the 1980's, we have simultaneously seen booming housing markets in the northeast and in California, with deteriorating markets in the southwest. A few of the local markets have boomed so remarkably that some economists are searching for evidence of bubbles (e.g., Case (1986) for the Boston market) while others have begun more broadly-based investigations into the efficiency of the single-family housing market (e.g., Case & Shiller (1989) and Linneman (1986)).

We step back from detailed analyses of bubbles and market efficiency in order empirically to examine longer-term national components in real house price growth as well as the extent of intermetropolitan area differences in home price appreciation rates. We also investigate the persistence of local trends in housing prices, working with a time series cross section of annual data on existing home prices for 56 metropolitan statistical areas (MSAs) over the 1971-1989 time period.

Housing price appreciation rates have differed across local markets. However, we treat annual appreciation rates as realizations of random processes and analyze whether house buyers could reasonably have expected differential capital gains across MSAs. Our empirical work decomposes the variance in housing price appreciation rates into common (national) time-varying components and city-specific fixed effects, while allowing for local persistence in price movements. We then investigate whether the intermetropolitan variance in house price appreciation rates is the result of different realizations from a common national distribution or of draws from city-specific distributions of appreciation rates.

Our results confirm that the national economy strongly influences local housing markets.

We report significant common national effects which vary with time. The importance of the national component is highlighted by the fact that we cannot reject the hypothesis that there are no city-specific fixed effects in house price appreciation rates over this two decade period. From an investment perspective, the data suggest equal expected returns across housing markets.[1] Our results also contrast with, but do not contradict, the suggestion of price bubbles by Case (1986). In fact, we find evidence of unequal persistence in appreciation rates across MSAs, suggesting that there are periods of local housing price appreciation which diverge from national trends.

In light of the fact that price **levels** vary dramatically across localities, not being able to reject equal trend appreciation rates across metropolitan areas has potentially important implications for housing market equilibrium. An investment in a home is inextricably linked to the purchase of a given local amenity and service bundle. Thus, housing market equilibrium must result in equal household utility across localities as well as satisfy asset market arbitrage conditions. Starting from unequal prices, equal growth rates imply increasingly divergent absolute price levels. An interesting question is whether similar appreciation rates across MSAs can continue in the face of differing local house price levels. With binding budget constraints and steady amenity package differentials, we argue tht it typically is not feasible for high housing

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[1]Because our appreciation measure captures only the change in capital value and not the housing service flow (an implicit dividend), we caution that the term 'return' is loosely used. However, we suspect it is accurate in spirit. Assuming you reside in a single location and actually consume the housing service flow only in that jurisdiction, you still could have invested in homes in other areas. It is probably the case that if a home you own in another jurisdiction is appreciating rapidly, then the same is occurring to the rent you could charge a tenant.

Finally, while we do not find differences in appreciation rates across cities, we do find evidence of differences when cities are grouped by the nine Census divisions. This is due to high average appreciation rates in the Pacific division. The remaining eight groups do not have significantly different division-specific components to their house price appreciation rates. See the discussion at the end of Section II for more on this.

price areas to appreciate indefinitely at the same rate as the low housing price areas unless special assumptions are made with respect to productivity differentials or local trait income elasticities. We find that higher price levels are significantly negatively correlated with appreciation rates when we control for local fixed effects and a time-varying common component. Our results imply that a 10% higher real house price is associated with a 1.4% lower average real appreciation rate.

II. EMPIRICAL SPECIFICATION AND RESULTS

To investigate the relative importance of common national impacts and metropolitan area differences on house price appreciation, we estimate an equation similar in spirit to that used by labor economists in their analyses of spatial differences in unemployment.[2] We postulate that the housing price appreciation rate in city i at time t (APP_{it}) is a long-run, MSA-specific trend (α_i) plus random fluctuation (ϵ_{it}):

$$1) \quad APP_{it} = \alpha_i + \epsilon_{it}.$$

The fluctuations reflect national movements in housing prices (β_t) and MSA-specific persistence in errors (ρ_i), so that

$$2) \quad \epsilon_{it} = \beta_t + \rho_i \epsilon_{it-1} + \upsilon_{it},$$

where υ_{it} is white noise. (1) and (2) yield:

[2]See Marston (1985) for example. Voith & Crone (1988) use a similar equation in their study of intermetropolitan area vacancy rate differentials.

$$3) \quad APP_{it} = \alpha_i(1-\rho_i) + \beta_t + \rho_i APP_{it-1} + \upsilon_{it}.$$

To estimate (3) we construct a set of dichotomous MSA dummy variables (C_i where i indexes the MSAs) and a vector of time dummies for each year (T_t where t denotes the time period) and estimate the following non-linear equation:

$$4) \quad APP_{it} = \alpha_i(1-\rho_i)C_i + \beta_t T_t + \rho_i APP_{it-1}C_i + \upsilon_{it}.$$

The vector of coefficients (α_i) captures any city-specific fixed effects in appreciation. The vector of coefficients on the time dummies (β_t) captures the annual national movements in housing price appreciation. Persistence in appreciation away from trend is reflected in the one-period serial correlation parameter, (ρ_i).[3,4]

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[3]We use annual data in the empirical work and experimented with specifications including an added lag term. We could always reject the joint significance of all lags beyond the first year. We work over annual intervals for a variety of reasons. As Case and Shiller (1987) have noted, there is a relatively high amount of quarterly variation and seasonality in the National Association of Realtors (NAR) data we use. Given that our interests are focused on longer-term and persistent local and national components in metropolitan area house prices, using annual data abstracts from various short-term influences on local prices. We could have used quarterly data to address some of these long-term issues. However, only a complex lag structure, requiring a dramatic increase in the number of parameters to be estimated, could hope to capture the longer-term dynamics we find in annual house price appreciation. Even with the annual data, we have to estimate 131 parameters.

[4]It should be noted that the specification in (4) does not arise from any particular structural model. For example, one might think that current house price appreciation is a function of past construction cost increases. Solving out might (conceptually) lead to a specification with lagged dependent variables as in (4). In that situation, introducing the lagged dependent variable could result in inconsistent OLS parameter estimates. This is not an issue here because our goals are more modest in that we wish to ascertain the relative importance of a city-specific component in house prices. This paper does not deal with underlying structural causes for such differences.

Estimation of (4) provides tests of several hypotheses. They include:

(a) $\beta_t = 0$ for all t; a finding of joint insignificance indicates that there is no time-varying common component to price appreciation; it would imply that only local factors are of significance;

(b) $\alpha_i = \alpha_j$ for all $i \neq j$; joint equality indicates that there are no significant differences in appreciation rates across metropolitan areas, controlling for MSA-specific differences in persistence of appreciation trends;

(c) $\rho_i = \rho_j$ for all $i \neq j$; joint equality indicates that persistence in appreciation does not differ across MSAs; $\rho_i < 0$ implies that there is mean reversion in the area's appreciation rate; $\rho_i > 0$ implies that there is persistence in the growth rates of housing prices and that a high (low) appreciation rate tends to be followed by another high (low) rate the next period;

(d) $\rho_i = 0$; joint insignificance indicates that there is no persistence in the price series; shocks to local markets are completely absorbed within a single period t and do not spill over into the next period;

(e) $\alpha_i = \alpha_j$, $\rho_i = \rho_j$ for all $i \neq j$; joint equality of both the MSA-specific appreciation and persistence terms indicates that there are no local effects on appreciation.

The time-series cross-section data used to estimate (4) are derived from the National

Association of Realtors' (NAR) series on median existing home sales prices.[5] Four annual house price appreciation series are constructed by using first quarter over first quarter levels, second quarter over second quarter levels, etc.[6] We have eighteen annual observations on price appreciation per local area beginning with the 1971-1972 time period and ending with the 1988-1989 time period. The 1989 data were available only through mid-1989. The cross section is composed of observations on fifty-six of the largest MSAs. Table 3 includes a listing of each metro area. All house price data are deflated by the appropriate national consumer price index (CPI) value.

Comparison Across Metropolitan Areas

Table 1 presents the results for the various hypothesis tests for each annual series for the 1971-1989 period, and for the 1971-1979 and 1980-1989 subperiods. We examine the subperiods to see if significant intermetropolitan area differences exist over a decade horizon, but

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[5]It is noteworthy that Case & Shiller (1987) have criticized price series such as the NAR data. First, the NAR data do not control for quality growth in the housing stock. Case & Shiller (1987) use a 'weighted repeat sales' method to construct housing price indexes for four cities (Atlanta, Chicago, Dallas, and San Francisco). They also provide information on the standard errors of their constructed price indexes. The NAR provides no such information. We use the NAR data because of its easy availability, low cost, and its broad cross section (which is particularly important for our purposes). Quality-adjusted price data for the vast majority of our cities simply are not available on a consistent basis. We have examined U.S. Bureau of the Census' quality-adjusted data for new homes at the regional level. We comment more extensively on the implications of quality growth for our results at the end of this section.

[6]We also constructed a smoother annual series by averaging the quarterly median prices for each year. This smoother averaged series eliminates a great deal of quarter-to-quarter variation in the NAR data and, hence, abstracts from seasonal and other short-term influences on housing price movements. We suspect that employing a smoother series makes it more likely to observe city-specific effects in the data because substantial idiosyncratic quarterly fluctuations are eliminated. However, this series is not ideal because it introduces serial correlation in the error structure (see Working (1960)). However, results using this series are not materially different from those reported below in that we still do not observe city-specific fixed effects in appreciation.

dissipate as the sample becomes longer.

The data indicate a strong time-varying common component to house price appreciation in each series. At better than a 99% confidence level, we can always reject the hypotheses that $\beta_t=\beta_s$ and $\beta_t=0$. Controlling for this common component and local persistence, the data do not reject the null that there are **no** differences in existing home price appreciation rates across MSAs (i.e., $\alpha_i=\alpha_j$ for all $i\neq j$)at standard confidence levels. This is true for the entire sample period and for both subperiods. These findings are in marked contrast to the implications for office markets from Voith & Crone (1988). They report significant differences in local vacancy rates and they found that some cities did not appear to cycle about a natural vacancy rate.[7]

The data reject the hypotheses of no persistence ($\rho_i=0$) and of equal persistence across cities ($\rho_i=\rho_j$) for the full 1971-1989 period, except in the third quarter over third quarter series. This pattern does not hold over the subperiods. We find no evidence of persistence in appreciation rates in the 1970's, but the 1980's show evidence of persistence which differs across cities in two of the four series. Finally, the joint hypothesis of equal local appreciation and equal local persistence ($\alpha_i=\alpha_j$, $\rho_i=\rho_j$) cannot be rejected in two of the four series for either the whole sample or the 1980's. In the 1970's we cannot reject this hypothesis at standard confidence levels in any case.

The estimates for the common time-varying component of housing appreciation are shown in Table 2. Only the results using first quarter data are presented, but results using other series are qualitatively similar. These estimates clearly illustrate that the 1970's were much more

[7]We examined the housing price appreciation cycles for the same cities as in Voith & Crone (1988). Our data showed that the observed local appreciation rates cycle about the estimated market-specific rates (which, as just noted, are not statistically significantly different from one another) in all cases. This pattern continues to hold even over the shorter eight year time period analyzed in Voith & Crone (1988).

favorable to housing than the 1980's have been. Relative to 1989 (the omitted year), the average common real appreciation rate over the whole sample period was 0.73 percent, 1.91 percent over 1971-1979, and -0.64 percent for 1980-1989. In the 1980's the only significant positive real appreciation rate was in the 1986-1987 period. Because the NAR data are not quality-adjusted, these coefficients overstate the common real appreciation in home prices.[8] Census data indicate that **new** home quality growth has been extremely high in the 1980's (just over 3% per annum for the U.S.). Average quality growth in the much larger stock of **existing** homes covered by the NAR undoubtedly is much smaller, but any such growth amplifies how poor real housing price appreciation has been in the 1980's.

Casual inspection also shows that housing performs poorly during recessions as evidenced by the negative coefficients for the 1974-1976 and 1980-1983 period. Titman's (1982) model suggests that real owner-occupied home prices are higher when anticipated inflation is higher, which is another reason for the relatively high appreciation of the 1970's. While this issue is beyond the scope of the paper, future research should use the longer time series on prices now available to investigate the determinants (including unanticipated as well as anticipated inflation) of the influential time-varying common component.

Table 3 presents the estimates of the MSA-specific parameters, α_i and ρ_i, over the full 1971-1989 period. Again, we show results for first quarter over first quarter data only. Differences across series for any MSA with a significant α_i or ρ_i in at least one series are shown in Table 4. Table 3 also shows the average median price over the same period as well as the compound annual appreciation rate for each MSA (which is based on their beginning and ending price

[8]Note that this does not imply that total returns would also be overstated. Higher-quality homes provide an enhanced service flow, resulting in greater utility for the tenant and higher rent for the landlord.

levels).

The only two statistically significant α_i's are for San Francisco and San Jose, further emphasizing the dominance of national trends. No other metropolitan areas have significant α_i's in any of the other series. Pairwise comparisons for all possible pairs of the α_i's reveal that only 50 of 1540 pairs, or 3.2 percent, were significantly different from one another. Almost all of these involved either San Francisco or San Jose.

Despite our inability to reject the equality of the α_i's, casual observation of the gross compound annual appreciation rates reveals some large **ex post** differences. Note that different gross compound annual appreciation rates do not necessarily imply different expected appreciation rates because different realizations from the same distribution could result in unequal **ex post** appreciation rates. Also, these gross appreciation rates are larger, on average, than the estimated MSA-specific fixed effects (α_i). This result is expected given the positive average national effects.

Turning to the serial correlation estimates, the first quarter data reject the hypothesis of joint equality of the ρ_i's, but only eight MSAs that have ρ_i's that are significantly different from zero. Rather than focus on only the first quarter estimates of ρ_i, Table 4 presents the estimated values for ρ_i for every MSA in which ρ_i is individually significant in at least one of the annual series. For the most part, the estimates are consistent across series. Nine of the seventeen MSAs had significantly negative ρ_i's in at least one of the series, suggesting mean reversion in prices.[9] Most of these MSAs are in the north central area of the country with the exception of Phoenix, San Jose, and Washington, D.C.

The six MSAs exhibiting persistence as opposed to mean reversion in price appreciation fall

[9]Two additional MSAs, Baton Rouge and San Francisco, had significantly negative ρ_i's but also had significantly positive ones as well.

into one of two categories. One group, including Ann Arbor, Los Angeles, New York, and Providence, experienced relatively high appreciation over the sample period, with compound annual appreciation rates over 3.4 percent. The remaining two, Houston and Oklahoma City, are linked to natural resource industries that were battered in the 1980's. Each of these cities had compound annual appreciation rates near zero for the sample period.

The finding of significant serial correlation parameters for some MSAs is consistent with previous findings by Case (1986) and Case & Shiller (1989). Serial correlation implies both forecastability and deviation from national trends over prolonged periods of time. Thus, while there appear to be no long-term local fixed effects on appreciation, it appears that investors could achieve differentially higher capital gains in some MSAs because of the high levels of persistence in those markets, especially in the 1980's. Note that if investors did not take advantage of the persistence in trends, that is, if their timing of entrance into the 1980's housing markets was random, they would have expected no differences in appreciation across MSAs.

Comparison across Census Divisions

That we cannot reject the null of no MSA-specific fixed effects in appreciation rates suggests that price trends are drawn from a common distribution rather than from separate MSA-specific distributions. However, it might be argued that the power of our test is low if certain groups of MSAs are affected by similar economic forces, and thereby have more similarity to other MSAs within their group than with localities outside their group. This argument is similar in spirit to that made by MacKinlay (1987) in his study of tests of the capital asset pricing model. In our case, the distributions of the appropriate test statistics for a comparison across two MSAs may be similar so that we could not hope to distinguish between trends in their appreciation rates without vastly superior data. However, if, say, regionally-grouped MSAs are similar to one another because they face the same regionally-based economic fundamentals, then grouping across MSAs and testing across groups may be appropriate and should lead to tests with greater power.

Towards this end, we grouped our MSAs by Census division and estimated a version of (4) such that each MSA within one of the nine Census divisions was restricted to have the same local fixed effect. The nine Census divisions listed in Table 6 make up the four Census regions. As shown in Table 5, the appropriate F-tests document that we can generally reject the null hypothesis of equality across the nine Census division α_i's for the whole sample and for both subsamples at conventional confidence levels. We can also reject equality of ρ_i for the whole sample and for the 1980's, though the results for the 1970's subsample are mixed. Further tests on the divisional regression reveal, however, that there are no significant differences across eight of the nine divisions. The Pacific division alone has significantly higher appreciation rates. Thus for most of the country, the MSA and division regressions yield similar conclusions.[10]

Morevoer, examining the Census data on quality growth in **new** homes across regions suggests that the NAR series on existing homes may overstate how different the Pacific division is. The Census data indicate that new home quality growth in the West has been extremely high, particularly in the 1980's. (The Census data show much less regional variation in quality growth in the 1970's.) While the new home quality index for the U.S. has grown by an average 3.16% per year from 1980-1989, the analogous figure for the Western region is 4.23%. While the quality growth in the existing home stock must have been smaller, some of the higher appreciation in the NAR data for the West appears due to relatively large increases in housing

[10]The F-statistic for equality of the division fixed effects, excluding Pacific is 1.33, with an associated rejection probability of .77. The F-statistic for equal serial correlation parameters is 1.31, with a rejection probability of .76.

quality over the decade. Further, quality growth in new homes in the Northeast has been the slowest of all regions in the 1980's, averaging only 1.03% per year. With quality-adjusted data, it is possible that appreciation in the Northeast, not the West, would appear to be drawn from a different distribution.[11]

We remain agnostic as to whether it is more appropriate to compare MSAs resulting in a test of potentially weaker power, or to group MSAs by census division. The reason is that the grouping may be due to what Lo and MacKinlay (1990) term "data snooping" which can artificially increase the size of test statistics. Lo and MacKinlay (1990) note that a potential problem arises when groupings of observations are not made randomly, but are based on some empirical trait (e.g., firm capitalization value for stocks or regional location for MSAs) which is correlated with returns. If there is no economic theory for why the correlation exists, this pre-test analysis essentially uses the "snooped" properties of the data to influence the size of test statistics which are then employed in analyzing the same data. With respect to home price appreciation, if it is only in an **ex post** sense that one notices growth rates are regionally correlated, then regional aggregation clearly is data-snooping. **Ex ante**, if theory would have allowed the researcher to predict the regionally-based correlation, then the grouping would not appear to bias the size of resulting test statistics. Both researchers and investors need to satisfy themselves that they are not data-snooping when grouping across MSAs.[12]

[11]It is worth noting that, with rapid quality changes, the weights used to construct the index can vary materially over only a few years. Megbolugbe (1989) documents the changes in weights from 1980 to 1988 and presents Laspeyres and Paasche versions of the Census' quality-adjusted price series.

[12]Undoubtedly, there are other groupings which would lead to rejection of equal α_i's. The same question needs to be answered whatever the grouping strategy.

III. PRICE LEVELS, APPRECIATION, AND LONG-RUN EQUILIBRIUM

While we cannot reject equal real house price appreciation rates across MSAs, real house price levels do vary widely across MSAs. As shown in the first column of Table 3, mean house prices over the two decade sample period vary from a low of $40,189 in Portland (ME) to a high of $132,569 in Ann Arbor. At any point in time, we would expect differences in housing price levels due to different amenity packages across localities (see Roback (1982)). The varying house price levels partially reflect the price that consumers are willing to pay for the amenity bundle associated with any particular MSA, holding house quality constant.

In equilibrium, housing prices should adjust to levels appropriate to consumers being indifferent across locations.[13] However, equal real housing price appreciation given different initial real house price levels implies increasingly divergent real housing price levels, and hence, divergent local trait prices. The extent of the divergence of housing price levels over time is shown in Figure 1 which plots the cross-sectional standard deviation of real housing price levels over the sample period. This has nearly tripled from $13,063 in 1972 to $36,025 in 1989.[14]

The compensating differential model, however, implies that the extent of price level differences in terms of foregone goods is bounded by differences in amenity levels across locations, given a finite amenity price. While it comes as no surprise that sunny southern

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[13]As noted earlier though, asset arbitrage occurs on total risk-adjusted returns, rather than on absolute price levels. This highlights the dual investment and consumption aspects involved in analyzing housing markets. A full utility model might allow one to combine the asset and consumption arbitrage conditions.

[14]In a study of Japanese land prices, Boone and Sachs (1989) present a simple neo-classical growth model which generates continually diverging land price levels. Their results are dependent upon a firm not being able to substitute away from local land in production. While that may be a reasonable assumption for most Japanese producers in an international setting, we do not believe it is accurate for firms or workers within a given country. Also, their model does not focus on the consumption aspect of housing.

California has higher priced homes than Omaha, the two areas cannot indefinitely continue to have the same real price appreciation rates as long as the sunshine differential remains constant and the price of sunshine does not increase markedly. Note that local trait prices can increase faster in one location than in another if residents of the high-priced area enjoy commensurately higher income growth which is tied to greater productivity in the region. In that case, the price of the local traits in terms of foregone consumption could be the same across regions. However, we suspect that long-run differentials in productivity growth across locations in the U.S. are not likely to be observed. Congestion, as well as factor market constraints, vitiate productivity increases in high growth areas in the long run.

The following stylized example illustrates how the divergence in relative trait prices can occur. Assume no non-housing inflation and normalize the price of other goods consumption to one. Let the initial absolute and relative existing home prices be \$100,000 in area 1 and \$10,000 in area 2. Now assume an equal 10% housing appreciation rate for both areas. Home prices are now \$110,000 and \$11,000, respectively, in areas 1 and 2. Relative prices have increased proportionately across locations. However, a homebuyer in area 1 had to forego \$10,000 worth of other goods consumption while a homebuyer in area 2 only had to forego \$1000 worth of other goods consumption. The differential in absolute cost increases with time. Ultimately, this cost exceeds the finite willingness-to-pay for the amenity differential.[15] At some point, the appreciation rate in the high price level area has to fall below that for the low price level area to

[15]Note that this would not hold if the income elasticity for the trait was always at least one, no matter what the cost of the trait in terms of the nonhousing composite good. Stated differently, if we introduce rising income into our simple example, it is technically feasible for a household in the high price area to continue consuming the same (or a greater) precentage of its budget in terms of housing. However, its preferences must be such that it wishes to do so even in the face of a potentially very high relative local trait price. With reasonable assumptions about diminishing marginal utility of local traits, we suspect that the requisite high income elasticities are not likely to be maintained in the face of rising relative trait prices.

keep the divergence in price levels within the bounds implied by differences in amenity packages (subject to our previous caveat about income differentials). The binding nature of this constraint may be becoming evident for some MSAs. Firms which produce for a national market and are located in the Los Angeles, San Francisco, and New York City areas appear to be having increasing trouble paying even their skilled employees enough to buy a home and to enjoy a reasonable level of nonhousing consumption.

To further investigate this issue, we estimated the following equation:

$$(5) \qquad \log(P_{it}/P_{it-1}) = a_i + B_t + \gamma \log(P_{it-1}) + \epsilon_{it}$$

where the a_i are city fixed effects, B_t are national effects and P_{it} is the real price of housing in period t.[16] If the implicit trait prices of housing are an important factor in consumers' location decisions, we would expect MSAs with higher price levels to have, **ceteris paribus**, lower appreciation rates. We therefore expect a negative coefficient on $\log(P_{i,t-1})$. We do find a negative value for γ of -0.139 with a standard deviation of 0.021. The full regression is available upon request. While this finding must be viewed as only suggestive, the elasticity implies that, on average, an area with 1% higher real home prices is associated with about a 0.14% lower appreciation rate.[17] (The mean appreciation rate for our sample is about 1.5% per annum.)

[16]This equation resembles that used by Barro and Martin (1989) in their paper examining convergence in economic growth rates. Note that a_i and B_t are not equivalent to α_i and β_t, respectively.

[17]The caution about the result being only suggestive in nature is for two reasons. With respect to the high statistical significance found, analysis of univariate time series has demonstrated that standard errors effectively are biased downward when there is a unit root (e.g., see Dickey (1975)). We are not aware of an analogous finding for pooled series such as ours, but a similar bias well could occur. Second, ordinary least squares estimates of parameters such as γ in (5) are likely to be consistent only as t becomes very large (e.g., see Anderson &

The negative coefficient implies that price levels ultimately will converge. However, this may occur over very long time periods because initially the absolute increase in price level resulting from a modest appreciation of a high priced house may be greater than high appreciation on a low priced house.

IV: SUMMARY

Using a large panel data set on median existing housing prices spanning eighteen years, we have found strong evidence of national movements in housing price appreciation, but virtually no MSA-specific fixed effect in the appreciation series. There appears to be persistence in price appreciation trends, and that persistence appears to vary locally, especially in the 1980's. The finding of equal appreciation rates across MSAs is in contrast to the wide divergence in price levels occurring over the sample period. Equilibrium in the housing services and amenities market suggests that continued equal appreciation, and hence divergence of price levels, can continue only if there are corresponding increases in MSA-specific income levels to offset the increasing implicit trait prices. Initial evidence suggests that, holding other factors constant, higher price levels are associated with lower appreciation rates. Finally, there are significant differences in trend growth rates across Census divisions but those differences appear to be solely due to a "California effect" which itself may partially reflect rapid increases in housing quality. Before arguing for the existence of regionally-specific housing markets, we urge that researchers satisfy themselves that they are not engaging in pretest data snooping.

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TABLE 1
HYPOTHESIS TESTS

Null (H_0)	Number of Restrictions	Probability of Rejection			
		Q1/Q1	Q2/Q2	Q3/Q3	Q4/Q4
Annual Data 1971-1989					
$\alpha_i = \alpha_j$	55	0.071	0.005	0.227	0.400
$\beta_t = \beta_s$	16	0.999	0.999	0.999	0.999
$\beta_t = 0$	17	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	55	0.994	0.999	0.616	0.967
$\rho_i = 0$	56	0.996	0.996	0.613	0.986
$\alpha_i=\alpha_j,\ \rho_i=\rho_j$	110	0.965	0.998	0.515	0.899
Annual Data 1971-1979					
$\alpha_i = \alpha_j$	55	0.745	0.540	0.578	0.667
$\beta_t = \beta_s$	7	0.999	0.999	0.999	0.999
$\beta_t = 0$	8	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	55	0.055	0.500	0.020	0.001
$\rho_i = 0$	56	0.895	0.874	0.811	0.920
$\alpha_i=\alpha_j,\ \rho_i=\rho_j$	110	0.588	0.902	0.432	0.257
Annual Data 1980-1989					
$\alpha_i = \alpha_j$	55	0.842	0.053	0.441	0.363
$\beta_t = \beta_s$	7	0.999	0.999	0.999	0.999
$\beta_t = 0$	8	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	55	0.972	0.997	0.545	0.795
$\rho_i = 0$	56	0.971	0.998	0.515	0.821
$\alpha_i=\alpha_j,\ \rho_i=\rho_j$	110	0.998	0.997	0.667	0.923

TABLE 2 NATIONAL APPRECIATION RATES ANNUAL DATA 1971-1989			
YEAR	APPRECIATION RATE	YEAR	APPRECIATON RATE
1971-72	3.35*	1980-81	-3.78*
1972-73	3.38*	1981-82	-2.80*
1973-74	1.42*	1982-83	-1.50*
1974-75	-2.48	1983-84	-0.31
1975-76	-0.75	1984-85	-0.26
1976-77	3.39*	1985-86	1.35
1977-78	5.63*	1986-87	2.85*
1978-79	2.08	1987-88	-0.63
1979-80	1.14	1988-89	omitted[a]

*Significant at the 5 percent level.
[a]All national appreciation rates are measured against the 1988-89 rate.

TABLE 3 HOUSING PRICES, CITY-SPECIFIC APPRECIATION AND PERSISTENCE ANNUAL DATA 1971-1989				
CITY	AVERAGE MEDIAN HOUSING PRICE	COMPOUND ANNUAL APPRECIATION RATE	CITY-SPECIFIC APPRECIATION RATE (α_i)	PERSISTENCE (ρ_i)
Akron	$53,020	-0.96	-1.54	0.05
Albany	69,688	1.13	0.69	-0.06
Ann Arbor	132,569	4.44	4.68	0.54*
Baltimore	59,157	2.99	2.41	-0.01
Baton Rouge	54,848	1.11	1.52	-0.31*
Birmingham	59,901	1.27	0.82	-0.33
Boston	110,273	2.55	2.04	0.20
Buffalo	45,641	1.39	1.04	-0.25
Chicago	73,228	1.24	0.66	0.10
Charlotte	66,917	0.70	0.18	-0.09
Charleston	61,483	0.07	-0.51	0.03
Cincinnati	56,909	0.49	-0.08	0.00
Cleveland	67,062	-1.33	-1.91	0.10
Columbus, OH	55,744	0.90	0.34	-0.22
Dallas	70,174	1.31	0.81	-0.02
Denver	72,946	1.25	0.57	0.19
Des Moines	51,886	0.26	-0.13	-0.35
Detroit	53,022	0.46	-0.10	0.21
El Paso	58,714	-0.99	-1.16	-0.29
Fort Lauderdale	71,860	0.45	-0.17	0.09
Grand Rapids	48,509	0.05	-0.38	-0.32
Houston	65,269	0.24	-2.08	0.67*
Indianapolis	49,729	1.41	0.84	-0.05
Jacksonville	49,441	1.50	1.25	-0.28
Kansas City	51,491	1.93	1.41	0.02
Las Vegas	80,082	-0.59	-1.20	0.14
Los Angeles	111,687	4.07	4.23	0.61*
Louisville	51,491	-0.62	-0.90	-0.24

*Significant at the 5 percent level.

TABLE 3 (cont'd.)
HOUSING PRICES, CITY-SPECIFIC APPRECIATION AND PERSISTENCE
ANNUAL DATA 1971-1989

CITY	AVERAGE MEDIAN HOUSING PRICE	COMPOUND ANNUAL APPRECIATION RATE	CITY-SPECIFIC APPRECIATION RATE (α_i)	PERSISTENCE (ρ_i)
Memphis	$60,497	0.66	0.22	-0.12
Miami	72,867	0.39	-0.24	0.13
Milwaukee	66,108	0.15	-0.38	-0.08
Minneapolis	70,466	0.60	0.21	-0.24
Nashville	62,013	1.07	0.78	-0.27
New York	97,208	3.44	2.99	0.24
Oklahoma City	53,395	-0.16	-3.40	0.64*
Omaha	51,252	0.23	0.09	-0.35
Orlando	62,339	1.30	0.95	-0.22
Philadelphia	66,755	2.33	1.90	-0.03
Phoenix	61,911	2.01	1.97	-0.33
Portland, ME	40,189	2.83[a]	2.56	-0.39
Portland, OR	53,938	2.22	1.65	0.15
Providence	63,496	4.36	3.88	0.38*
Rochester, NY	55,106	0.99	0.61	-0.39
San Francisco	122,084	5.29	5.00*	-0.06
San Jose	106,457	4.32[b]	3.96*	-0.21
San Diego	92,557	4.53	4.12	0.46
Salt Lake City	61,877	0.48	-0.04	0.11
San Antonio	55,496	0.17	-0.48	0.10
St. Louis	65,671	-0.04	-0.52	0.05
Syracuse	50,555	1.62	1.19	-0.34
Tampa	52,341	1.32	0.80	-0.24
Toledo	50,735	-0.39	-0.65	-0.54*
Tulsa	59,004	0.10	-0.16	-0.14
Washington, D.C.	93,944	1.93	2.76	-0.30*
West Palm Beach	80,769	1.03	0.50	0.05

*Significant at the 5 percent level.
[a]Data for 1971-1986.
[b]Data for 1971-1987.

TABLE 4 COEFFICIENTS FOR SELECTED CITIES				
CITY	Q1/Q1	Q2/Q2	Q3/Q3	Q4/Q4
FIXED EFFECTS				
San Francisco	5.00*	2.21	3.90*	3.83*
San Jose	3.96*	2.17	3.46	3.26*
PERSISTENCE				
Ann Arbor	0.54*	0.18	0.33	0.17
Baton Rouge	-0.31*	-0.32*	0.36*	-0.22
Buffalo	-0.25	-0.09	-0.42*	-0.41*
Columbus, OH	-0.22	-0.62*	0.03	-0.20
Grand Rapids	-0.32	-0.12	0.02	-0.39*
Houston	0.67*	0.50*	0.41	0.38
Indianapolis	-0.05	-0.63*	-0.32	-0.01
Los Angeles	0.61*	0.67*	0.31	0.33
New York	0.24	0.50*	0.04	0.33*
Oklahoma City	0.64*	0.63*	0.17	0.55
Omaha	-0.35	-0.53*	0.09	0.00
Phoenix	-0.33	-0.23	-0.38*	-0.17
Providence	0.38*	0.31*	0.35*	0.47*
San Francisco	-0.06	0.68*	-0.38*	-0.59*
San Jose	-0.21	0.13	-0.03	-0.44*
Toledo	-0.54*	-0.44*	-0.27	-0.47*
Washington, D.C.	-0.30*	0.31	-0.08	-0.24

*Significant at the 5 percent level.

TABLE 5
HYPOSTHSIS TESTS (DIVISIONS)

Null (H_0)	Number of Restrictions	Probability of Rejection			
		Q1/Q1	Q2/Q2	Q3/Q3	Q4/Q4
Annual Data 1971-1989					
$\alpha_i = \alpha_j$	8	0.968	0.542	0.996	0.996
$\beta_t = \beta_s$	16	0.999	0.999	0.999	0.999
$\beta_t = 0$	17	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	8	0.851	0.999	0.998	0.966
$\rho_i = 0$	9	0.853	0.999	0.989	0.982
$\alpha_i=\alpha_j$, $\rho_i=\rho_j$	16	0.999	0.999	0.999	0.999
Annual Data 1971-1979					
$\alpha_i = \alpha_j$	8	0.999	0.995	0.999	0.999
$\beta_t = \beta_s$	7	0.999	0.999	0.999	0.999
$\beta_t = 0$	8	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	8	0.636	0.872	0.930	0.919
$\rho_i = 0$	9	0.999	0.998	0.999	0.999
$\alpha_i=\alpha_j$, $\rho_i=\rho_j$	16	0.999	0.999	0.999	0.999
Annual Data 1980-1989					
$\alpha_i = \alpha_j$	8	0.999	0.940	0.997	0.991
$\beta_t = \beta_s$	7	0.999	0.999	0.999	0.999
$\beta_t = 0$	8	0.999	0.999	0.999	0.999
$\rho_i = \rho_j$	8	0.926	0.999	0.994	0.960
$\rho_i = 0$	9	0.904	0.999	0.994	0.937
$\alpha_i=\alpha_j$, $\rho_i=\rho_j$	16	0.999	0.999	0.999	0.999

TABLE 6
HOUSING PRICES, DIVISION-SPECIFIC APPRECIATION AND PERSISTENCE
ANNUAL DATA 1971-1989

DIVISION	AVERAGE MEDIAN HOUSING PRICE	COMPOUND ANNUAL APPRECIATION RATE	DIVISIONAL APPRECIATION RATE (α_i)	PERSISTENCE (ρ_i)
East North Central	$64,758	0.97	-0.17	0.02
East South Central	58,476	0.64	0.05	-0.23
Middle Atlantic	64,159	2.04	1.21	-0.04
Mountain	69,204	0.70	0.17	-0.01
New England	73,049	3.24	2.81	0.16
Pacific	97,149	4.36	3.39*	0.14
South Atlantic	66,672	1.17	0.75	-0.14*
West North Central	55,024	0.56	0.00	-0.16
West South Central	59,557	0.29	-0.28	-0.06

*Significant at the 5 percent level.

FIGURE 1

Cross-Sectional Standard Deviation of Real Price Levels Over Time