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FORECASTING PRICES AND EXCESS RETURNS IN THE HOUSING MARKET

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ABSTRACT

The U. S. market for homes appears not to be efficient. A number of information variables predict housing price changes and excess returns of housing relative to debt over the succeeding year. Price changes observed over one year tend to continue for one more year in the same direction. Construction cost divided by price, the change in per capita real income, the change in adult population are all positively related to price changes or excess returns over the subsequent year.

The results are based on time-series cross section regressions with quarterly data 1970-1 to 1987-3 and for cities Atlanta, Chicago, Dallas, and San Francisco.

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In an earlier paper [Case and Shiller, 1989] we performed weak-form tests of the efficiency of the market for single family homes. The tests were based on a set of price indexes constructed from micro data on nearly 40,000 homes that were sold more than once during the period 1970-1986 in four metropolitan areas: Atlanta, Chicago, Dallas, and San Francisco (Alameda County).

In that paper we found evidence of positive serial correlation in real housing prices. A change in the log real price index in a given year and a given city tends to be followed by a change in the same direction the following year between 25 percent and 50 percent as large. We also found evidence of inertia in a crude measure of excess returns estimated for each of the four metropolitan areas.

This paper takes the analysis several steps further. First, we construct a more detailed estimate of excess returns to investment in single family homes in each of the four cities that is sensitive to changes in the market value of housing services and to changes in marginal personal income tax rates. Second, we perform strong-form efficiency tests by exploring the forecastability of excess returns and house prices with a number of forecasting variables.

<u>The WRS Index</u>

The biggest problem faced by analysts of the residential real estate market is a lack of good time series on house prices. The most commonly used series is the National Association of

Realtors' "median price of existing single family homes." While the NAR generates this series for a large number of metropolitan areas quarterly and for the U.S as a whole monthly, they are not useful for our purposes. First, they are only available since 1981. Second, changes in the median home price in an area depend both on changes in house prices and on changes in the mix of homes that happens to sell.

In an earlier piece (Case and Shiller [1987]) we discuss the problems associated with the NAR data and construct an alternative based on microdata using a technique that we call the Weighted Repeat Sales (WRS) method. The method used is a modification of one first proposed by Bailey, Muth and Nourse [1963] (hereafter BMN). The method uses observations on individual houses that sold more than once during the sample period. Specifically, the change in log price for each observation is regressed on a set of simple dummy variables. The dummys are set to -1 for the period of the first sale and to +1 for the period of the second sale and to 0 otherwise. The resulting coefficients are the values of the log price index (WRS_t). Baily, Muth and Nourse argued that if individual log house price changes differed from the city-wide log house price changes by an independent, identically distributed noise term, the BMN method produces the best linear unbiased estimate of the city-wide log price index.

In our earlier piece we argue that the house-specific component of the change in log price is not likely to be homoscedastic, but that the variance of the error is likely to increase with the interval between sales. Specifically we assumed that the log price P_{it} of the ith house at time t is:

(1) $P_{it} = C_t + H_{it} + N_{it}$

where C_t is the log of the city-wide level of housing prices at time t, H_{it} is a Gaussian random walk (where ΔH_{it} has zero mean and variance σ_h^2) that is uncorrelated with C_T and H_{jT} for all T and $i \neq j$, and N_{it} is an identically distributed normal noise term (which has zero mean and variance σ_N^2) and is uncorrelated with C_T and H_{jT} for all j and T and with N_{jT} unless i=j and t=T

In equation (1) N_{it} represents the truly random component of sales prices around true value resulting from random events in the search process, the behavior of real estate agents and other imperfections. H_{it} represents the individual drift in house value through time.

These assumptions led us to a three step weighted (generalized) least squares procedure. The BMN procedure was followed precisely in the first step, and the residuals were stored. The squared residuals from the first step regression were then used as the dependent variable in a second step with a constant term and the time interval between sales on the right hand side.[1] The constant term is the estimate of σ_N^2, and the

[1] Observations in which the time interval between sales is larger are likely to have larger errors. As a result, we used a weighted regression that downweighted the observations corresponding to large time intervals. As we mentioned in the text, the regression was run separately for each quarter using only information available in that quarter. For earlier quarters that meant that the coefficients were calculated with only a

coefficient on the time variable is the estimate of σ_h^2. In the third step, the first step was repeated after first dividing each observation by the square root of the fitted value in the second stage regression.

The above procedure was used to create two log price indexes, WRS_a and WRS_b. In each city, houses were randomly allocated to two samples, a and b, each with half the available observations, and the price indexes were estimated separately with these samples. In our regression results below, WRS_a was used for left-hand (dependent) variables, and WRS_b was used for right-hand (independent) variables. This method was adopted as a simple expedient to prevent the same measurement error from contaminating both sides of the equation; See Case and Shiller [1989].

<u>The Data</u>

Table 1 contains a list of the variables used in constructing the estimate of excess returns for each metro area and in the forecasting equations later in the paper. Most of the variables and their sources listed in Table 1 are self explanatory. Two, however, deserve some discussion: RENTINDEX, a metro area specific rent index, and MTR, a national marginal individual income tax rate series.

A critical component of the return to investment in owner

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small number of observations. In instances where the estimated coefficient of the interval between sales has the wrong sign, it was set to zero, and the procedure reduces to OLS in step three.

Table 1

VARIABLES USED IN THE ANALYSIS
All Variables are Quarterly
1970:1 to 1986:2

CODE	SERIES	SOURCE
RENT INDEX	CPI-RESIDENTIAL RENT INDEX (For each metro area)	BLS - Monthly Labor Review, adjusted to estimate mid-quarter
CPI	CONSUMER PRICE INDEX ALL ITEMS - URBAN	BLS - Data Resources, Inc.
NMTG	EFFECTIVE MORTGAGE RATE - EXISTING HOMES (CONVENTIONAL)	Federal Reserve System - Data Resources, Inc.
EMP	EMPLOYMENT, NON-AGRI. WAGE AND SALARY (For each metro area)	Data Resources, Inc. Monthly Labor Review
INCTOT	PERSONAL INCOME (For each metro area)	Data Resources, Inc. Survey of Current Business
PTAX	EFFECTIVE PROPERTY TAX RATE - Residential (For each metro area)	Constructed from Census of Governments and studies by the District of Columbia as reported annually in the Statistical Abstract of the United States
POP	POPULATION (For each metro area)	Data Resources, Inc. Current Population Survey
AGE1	PERCENT OF POPULATION BETWEEN 25 AND 34 (For each Metro area)	Data Resources, Inc. Dept. of Commerce, Bureau of the Census, "State Population and Household Estimates"
AGE2	PERCENT OF POPULATION BETWEEN 35 AND 44 (For each metro area)	Same as above

VARIABLES CONTINUED:

CODE	SERIES	SOURCE
AGEPOP	NUMBER OF PERSONS BETWEEN 25 AND 44 (For each metro area)	(AGE1+AGE2)*POP
HSTARTS	HOUSING STARTS - SINGLE FAMILY (For each metro area)	Data Resources, Inc. Dept. of Commerce Series C-25
TBILL	INTEREST RATE ON 90-DAY TREASURY BILLS	Economic Report of the President
MTR	MARGINAL INDIVIDUAL INCOME TAX RATE FOR MEDIAN HOME BUYERS (Annual)	Joint Center for Housing Studies, Harvard University Provided by D. Dipasquale and W. Apgar
CONCOST	CONSTRUCTION COST INDEX (For each metro area)	Engineering News Record

occupied housing is the value of housing services that accrues to owners. This "imputed rent" is in essence the dividend component of the return. Since there are no transactions involved, this component of the yield must be estimated and is problematic.

The only measures of rent available on a consistent basis for specific metropolitan areas is the "residential rent" component of the Consumer Price Index produced by the Bureau of Labor Statistics. In constructing the index, the BLS repeatedly surveys the same units from year to year to control for quality. A number of authors [Lowry (1982), Apgar (1987), DiPasquale and Wheaton (1989)] have criticized the index because it ignores the depreciation that takes place over time and thus tends to underestimate the level of rents controlling for unit quality. Some of the observed decline in real rents observed during the 1970's can be attributed to this downward bias. The BLS last year acknowledged the problem and introduced a correction into the index beginning in 1988. However, no changes were made in the historical series. Some of the authors critical of the series have suggested increasing rents annually from 0.5 to 0.9 percent [see again Lowry (1982), Apgar (1987) and DiPasquale and Wheaton (1989)].

For purposes of constructing an index of returns to homeowners, the unadjusted index that fails to adjust for depreciation is the appropriate one. It is reasonable to assume that owner occupied housing depreciates physically over time at about the same rate as rental housing with appropriate expenditures on maintenance and repair. The WRS index discussed above is based on repeat sales of the same unit. An individual home-owner will find that the market value of the housing services that he/she consumes declines slightly with the age of the unit as will the market value of the unit itself. Since this small decline is part of the net yield to owners, we do not adjust the BLS rent index nor do we attempt to wash depreciation out of our WRS index for purposes of estimating excess returns to investment.

Clearly, the decision to invest in owner occupied housing is likely to be influenced by the tax treatment of its yield. This has changed in complicated ways over time. First of all, net imputed rent has never been subject to taxation. Second, property tax payments and mortgage interest payments have always been deductible. The value of a deduction, of course, depends on the taxpayer's marginal tax rate.

The recent tax acts, specifically ERTA in 1981 and the Tax Reform Act of 1986, have changed the tax system in fairly dramatic ways. The change that has had the most significant direct effect on owner occupied housing has been the sharp decline in marginal tax rates, particularly at the top end.

The most significant indirect effect has come through the dramatic changes in depreciation rules, the ITC, and changes in passive loss rules that worked in favor of rental housing during the early 1980's and against it in the late 1980's.

To calculate excess returns to investment in owner occupied housing, a time series on an appropriate marginal tax rate is needed. The marginal tax rate series used was constructed by the researchers at the Joint Center for Housing Studies at Harvard. It was constructed by looking at the income profile of first time home buyers and calculating the average marginal rate, given the laws in effect in each year, for that group. The argument for using first time home buyers is that they can be thought of as the "marginal" investors. Those with higher incomes face higher marginal rates and will earn higher excess returns. Higher tax rates mean a lower opportunity cost of capital and lower net property taxes and mortgage payments.

It is important to note that both the rent series and the marginal tax rate series chosen will result in conservative estimates of excess returns.

Estimating Excess Returns

Two basic approaches were taken to estimating excess returns. The first assumes that the home is bought outright, no leveraging. The second assumes the purchase is financed with an 80% mortgage. Both are after-tax rates of return, and they assume that neither capital gains nor imputed rent are taxed.

$$EXCESS_A = \frac{\exp(WRS_{t+5})*PRICE70 + RENTINDEX_{t+1}*RENT70}{\exp(WRS_{t+1})*PRICE70}$$

$$- (1-MTR_{t+1})*PTAX_{t+1} - 1 - (1-MTR_{t+1})*TBILL_{t+1}/100$$

$$EXCESS_B = \frac{\exp(WRS_{t+5}) * PRICE70 + RENTINDEX_{t+1} * RENT70}{.2 * \exp(WRS_{t+1}) * PRICE70}$$

$$- [(1-MTR_{t+1}) * (PTAX_{t+1} + NMTG_{t+1} * .8)] / .2$$

$$- 5 - (1-MTR_{t+1}) * TBILL_{t+1} / 100$$

The WRS index is defined above, and RENTINDEX, PTAX, MTR and NMTG are defined in Table 1. Of critical importance are PRICE70, the base period house price, and RENT70, the base period value for imputed rent. The derivation of PRICE70 and RENT70 is shown in Table 2. The baseline house price is assumed to be the median value of owner occupied units from the 1970 Census. The figures are available separately for each of the four SMSA's. For the Alameda County series, we used the San Francisco/Oakland SMSA.

Estimating baseline imputed rent was not as easy. While median contract rent (which excludes utility payments) is available for each SMSA, rental units on average are smaller and of lower quality than owner occupied units. Thus, median rent will understate the market value of the housing services generated by the median owner occupied house. A rough correction was thus made to rent based on the number of rooms. The Census has data on the median number of rooms in renter occupied and owner occupied housing units and the median rent for each city was simply stepped up in proportion to the larger number of rooms in owner occupied units.

Table 2

A. BASELINE HOUSING PRICES AND RENT LEVELS:1970

CITY	(1) (PRICE70) MEDIAN VALUE OWNER OCCUPIED HOUSING UNITS[1]	(2) CONTRACT RENT[1] (Monthly)	(3) STEP-UP FACTOR[2]	(4) (RENT70) ESTIMATED ANNUAL RENT	(4)/(1) RETURN
ATLANTA	19,800	98	1.50	1764	.089
CHICAGO	24,300	116	1.40	1944	.080
DALLAS	16,600	110	1.50	1848	.111
SAN FRAN.	26,900	130	1.56	2424	.090

B. MEDIAN NUMBER OF ROOMS:1970

CITY	(1) OWNER OCCUPIED UNITS	(2) RENTER OCCUPIED UNITS	(1)/(2) STEP-UP FACTOR
ATLANTA	6.0	4.0	1.50
CHICAGO	5.6	4.0	1.40
DALLAS	5.6	4.0	1.40
SAN FRAN.	5.6	3.6	1.56

Sources: 1. U.S. Department of Commerce, Social and Economic Statistics Administration, Bureau of the Census, 1970 Census of Housing, Vol. 1, Parts 7, 12, 15, 45, Tables 1 and 14. Data are for Standard Metropolitan Statistical Areas.

2. Step-up factor based on Part B. below

The right hand column in table 2A shows the estimated baseline annual rent as a fraction of the median home value. Compared to the average dividend price ratio on common stocks, those numbers look high, but they are not unreasonable.

Excess Returns:1970-86

Tables 3 and 4 present the calculations for excess returns to investment in owner occupied housing ($EXCESS_A$ and $EXCESS_B$). Table 3 presents the figures assuming that the property was purchased outright ($EXCESS_A$). Table 4 assumes that the investment was leveraged with an 80 percent mortgage at the going rate of interest ($EXCESS_B$).

The tables confirm the conventional wisdom that housing was an exceptional investment throughout the decade of the 1970's in all cities. It should be kept in mind that these estimates are fairly conservative. The marginal income tax rate series and the rent price index both are, if anything, downward biased. Only the baseline imputed rent figure is at all suspect, and it could be halved and these series would show extraordinary excess returns.

As expected the highest returns are achieved in California and Texas during their respective price booms. The Dallas boom occurred in 1977-79, while the California boom was longer, starting in 1976 and running into early 1980. In both Dallas and Oakland, leveraged returns of over 100 percent per year were

Table 3

ESTIMATED EXCESS RETURNS FOR FOUR QUARTERS ENDING IN EACH QUARTER:1971-1986
($EXCESS_A$ -- NO LEVERAGE)

YEAR	ATLANTA	CHICAGO	DALLAS	SAN FRAN
1970.1	.	.	.	.
70.2	.	.	.	.
70.3	.	.	.	.
70.4	.	.	.	.
1971.1	0.09	0.06	0.09	0.09
71.2	0.09	0.02	0.10	0.11
71.3	0.17	0.04	0.09	0.14
71.4	0.05	0.08	0.17	0.06
1972.1	0.13	0.08	0.18	0.10
72.2	0.10	0.11	0.06	0.11
72.3	0.07	0.09	0.08	0.10
72.4	0.09	0.11	0.10	0.12
1973.1	0.09	0.12	0.09	0.11
73.2	0.16	0.11	0.17	0.10
73.3	0.12	0.10	0.16	0.15
73.4	0.14	0.10	0.03	0.14
1974.1	0.10	0.10	0.10	0.15
74.2	0.09	0.08	0.08	0.15
74.3	0.08	0.06	0.03	0.10
74.4	0.10	0.05	0.11	0.12
1975.1	0.03	0.06	0.07	0.13
75.2	0.00	0.05	0.08	0.11
75.3	0.01	0.04	0.08	0.11
75.4	-0.04	0.08	0.09	0.12
1976.1	0.03	0.10	0.05	0.14
76.2	0.02	0.10	0.04	0.17
76.3	-0.02	0.10	0.06	0.18
76.4	0.05	0.10	0.06	0.20
1977.1	0.06	0.11	0.15	0.27
77.2	0.06	0.13	0.16	0.32
77.3	0.09	0.14	0.17	0.33
77.4	0.04	0.18	0.21	0.29
1978.1	0.09	0.19	0.22	0.22
78.2	0.08	0.20	0.22	0.16
78.3	0.07	0.19	0.23	0.10
78.4	0.10	0.16	0.24	0.14
1979.1	0.13	0.13	0.28	0.09
79.2	0.15	0.10	0.23	0.16
79.3	0.19	0.07	0.23	0.20
79.4	0.19	0.05	0.15	0.16
1980.1	0.11	0.01	0.13	0.25
80.2	0.10	-0.03	0.14	0.16
80.3	0.07	-0.03	0.12	0.14
80.4	0.03	-0.06	0.13	0.12

YEAR	ATLANTA	CHICAGO	DALLAS	SAN FRAN
1981.1	0.02	-0.06	0.08	0.09
81.2	0.05	-0.01	0.07	0.10
81.3	0.04	-0.03	0.05	0.04
81.4	0.00	-0.06	0.00	-0.05
1982.1	-0.03	-0.09	0.00	-0.09
82.2	-0.04	-0.12	0.00	-0.07
82.3	-0.03	-0.05	0.01	-0.09
82.4	0.03	-0.01	0.04	-0.02
1983.1	-0.01	-0.01	0.01	-0.05
83.2	0.03	0.01	0.04	-0.05
83.3	0.05	0.02	0.08	0.02
83.4	0.04	-0.01	0.07	0.03
1984.1	0.07	0.00	0.07	0.04
84.2	0.04	0.01	0.06	0.02
84.3	0.03	-0.04	0.05	0.02
84.4	0.08	-0.01	0.07	0.03
1985.1	0.05	0.00	0.05	0.03
85.2	0.05	0.00	0.03	0.04
85.3	0.05	0.02	0.02	0.07
85.4	0.03	0.03	0.02	0.07
1986.1	0.06	0.03	0.03	0.08

Table 4
ESTIMATED EXCESS RETURNS FOR FOUR QUARTERS
ENDING IN EACH QUARTER: 1971-1986
(EXCESS B -- ASSUMING 80% MORTGAGE)

YEAR	ATLANTA	CHICAGO	DALLAS	SAN FRAN
1970.1	.	.	.	.
70.2	.	.	.	.
70.3	.	.	.	.
70.4	.	.	.	.
1971.1	0.42	0.28	0.42	0.41
71.2	0.39	0.06	0.47	0.49
71.3	0.79	0.13	0.39	0.65
71.4	0.13	0.28	0.73	0.21
1972.1	0.51	0.29	0.78	0.38
72.2	0.39	0.42	0.19	0.46
72.3	0.28	0.38	0.30	0.41
72.4	0.34	0.44	0.41	0.47
1973.1	0.33	0.47	0.30	0.40
73.2	0.68	0.46	0.72	0.41
73.3	0.51	0.40	0.67	0.65
73.4	0.61	0.39	0.05	0.63
1974.1	0.45	0.42	0.43	0.70
74.2	0.43	0.36	0.37	0.71
74.3	0.43	0.34	0.18	0.50
74.4	0.49	0.25	0.51	0.58
1975.1	0.12	0.24	0.32	0.59
75.2	-0.02	0.27	0.39	0.54
75.3	0.06	0.21	0.37	0.55
75.4	-0.28	0.34	0.39	0.53
1976.1	0.01	0.37	0.11	0.59
76.2	-0.01	0.41	0.11	0.75
76.3	-0.21	0.44	0.19	0.80
76.4	0.11	0.36	0.18	0.89
1977.1	0.19	0.43	0.64	1.22
77.2	0.16	0.52	0.66	1.46
77.3	0.31	0.56	0.70	1.54
77.4	0.06	0.76	0.92	1.33
1978.1	0.31	0.82	0.96	0.97
78.2	0.26	0.87	0.99	0.68
78.3	0.24	0.85	1.03	0.41
78.4	0.40	0.71	1.09	0.61
1979.1	0.55	0.56	1.30	0.35
79.2	0.66	0.43	1.04	0.69
79.3	0.89	0.28	1.08	0.91
79.4	0.89	0.23	0.73	0.76
1980.1	0.52	-0.01	0.60	1.24
80.2	0.46	-0.18	0.69	0.77
80.3	0.32	-0.19	0.55	0.65
80.4	0.16	-0.30	0.64	0.63

YEAR	ATLANTA	CHICAGO	DALLAS	SAN FRAN
1981.1	0.13	-0.29	0.40	0.46
81.2	0.13	-0.17	0.21	0.34
81.3	0.09	-0.24	0.14	0.12
81.4	0.03	-0.29	0.02	-0.22
1982.1	-0.11	-0.40	0.04	-0.41
82.2	-0.13	-0.55	0.06	-0.29
82.3	-0.15	-0.27	0.04	-0.46
82.4	-0.02	-0.18	0.04	-0.27
1983.1	-0.09	-0.09	0.01	-0.32
83.2	0.01	-0.06	0.07	-0.38
83.3	0.03	-0.12	0.18	-0.11
83.4	-0.01	-0.23	0.14	-0.05
1984.1	0.18	-0.14	0.19	0.06
84.2	0.06	-0.10	0.18	-0.05
84.3	0.06	-0.28	0.13	-0.01
84.4	0.26	-0.17	0.23	0.01
1985.1	0.15	-0.09	0.17	0.05
85.2	0.19	-0.05	0.09	0.15
85.3	0.16	0.02	0.03	0.27
85.4	0.04	0.02	-0.02	0.20
1986.1	0.19	0.02	0.03	0.26

achieved during those years. Home owners in Chicago and Atlanta also did extremely well during the 1970's.

Leveraged buyers suffered large percent losses during the early 1980's in every city but Dallas. Chicago had the longest string of negative returns from 1980-1984.

Several things can be seen in the raw numbers. First, if these numbers are correct, and there is no reason to believe that they are not, it is extraordinary that excess returns of this magnitude could persist for so long a period of time. In one sense, we now know a "buy rule" that will consistently earn an extraordinary return: simply buy housing. The puzzling question is, why didn't housing prices rise even further and more rapidly during the decade?

Second, it is apparent that there is a substantial degree of positive serial correlation in the data. Positive signs and negative signs are clumped and there are clear "waves" in the data. Some of this is, of course, explained by the fact that these are returns for the previous four quarters estimated with quarterly data so that the returns periods overlap. In our earlier paper (Case and Shiller [1989]), however, we concluded that positive serial correlation goes beyond what is explained by overlapping intervals. A substantial degree of quarter-to-quarter noise is still present.

<u>Estimation Procedure</u>

This section attempts to forecast house prices and excess returns using a set of forecasting variables. The tables below give time-series-cross-section regression results, pooling the four cities to reduce standard errors of the estimates. Pooling the data from the four cities can reduce standard errors because it increases the number of observations in the regression and also because it may increase the variance of independent variables by including cross-city variance. The estimated coefficients are ordinary least squares estimates $\beta = (X'X)^{-1}X'Y$ where the vector Y of observations of the dependent variable equals $[Y_1' \; Y_2' \; Y_3' \; Y_4']'$ where Y_i is the matrix of dependent variables for city i, and the matrix X of observations of independent variables equals $[X_1' \; X_2' \; X_3' \; X_4']'$ where X_i is the matrix of independent variables for city i. These are "stacked" ordinary-least squares regressions, effectively ordinary least squares regressions for each city but constraining all coefficients to be the same across cities. The estimated standard errors of the estimated coefficients take account of the overlap in one-year forecasts with quarterly data using a method of Hansen and Hodrick [1980] modified to allow time-series-cross-section regression. The estimated variance matrix of estimated coefficients was $(X'X)^{-1}X'SX(X'X)-1$ and S is an estimated covariance matrix of the error terms in the ordinary least squares regression. In composing S it was assumed that $\mathrm{cov}(u_{it1}, u_{jt2}) = c_{ij}L(t_1-t_2)$ where u_{it} is the error term for city i at time t and c_{ij} equals the sample variance of the residuals, $\mathrm{var}(Y - X\beta)$ if $i = j$, and equals the average covariance of contemporaneous residuals across all pairs of cities if i does not equal j. Thus two parameters were estimated to compose S. $L(t_1-t_2)$ equals $1 - |t1-t2|/4$ if $|t_1-t_2| < 4$, and zero otherwise, so that $L(t_1-t_2)$ is the degree of overlap between the forecast intervals.

Results

In discussing the results we will often refer to the "expected" sign of a coefficient. Of course, if markets are efficient there is no expected sign: all coefficients should be zero. If, however, information tends to be incorporated with a lag into housing prices, then the expected sign of a coefficient may tend to be the same as the expected effect on housing value of a change in the variable indicated.

We first observe [Table 5] that price changes show the positive serial correlation at short horizons and negative serial correlation at longer horizons that has been observed for other assets (Cutler, Poterba and Summers [1990]) and that is consistent with notion of "excess volatility" in prices. The results show that if prices once go up in a given year, they tend also to go up the next year, but by about a third as much. Moreover, this upward movement appears partly to be reversed in succeeding years, although the negative coefficients are not statistically significant.

The R^2s in Table 5 are not very high. We sought to improve our forecasting ability by including the other forecasting variables, in Table 6. Table 7 uses the forecasting variables to

Table 5
Regressions of Change in Log Real Price on Its Lagged Values
Dependent Variable is $W_{at+5} - W_{at+1}$

(W_a = log real price index estimated from a sample of homes)
(W_b = log real price index estimated from b sample of homes)

Regression No.		1	2	3	4	5
1	Constant	0.010 (1.033)	0.013 (1.244)	0.015 (1.354)	0.018 (1.596)	0.015 (1.581)
2	$W_{bt}-W_{bt-4}$	0.312 (2.863)	---	---	---	0.375 (3.341)
3	$W_{bt-4}-W_{bt-8}$	---	-0.014 (-0.120)	---	---	-0.095 (-0.911)
4	$W_{bt-8}-W_{bt-12}$	---	---	-0.124 (-1.025)	---	-1.118 (-1.142)
5	$W_{bt-12}-W_{bt-16}$	---	---	---	-0.110 (-0.895)	-0.029 (-0.283)
	R^2	0.108	0.000	0.018	0.014	0.162
	Nobs	229	213	197	181	181

Note: t-statistics (in parentheses) take account of overlap of observations of dependent variable and cross-section-time-series structure of data. Data are quarterly starting in 1970, first quarter. Data end in 1986, second quarter for each city except San Francisco, where data end in 1986, third quarter.

Table 6

Regressions of Price Change on Forecasting Variables
Dependent Variable is $W_{at+5} - W_{at+1}$

Regression No.	1	2 (dummied)	3	4	5
1 Const	-0.178 (-1.446)	-0.104 (-0.793)	-0.195 (-2.764)	-0.011) (-0.293)	-0.084 (-1.521)
2 $W_{bt}-W_{bt-4}$	0.237 (2.332)	0.197 (1.922)	0.202 (1.966)	---	---
3 $\frac{RENT}{P_b}$	-0.206 (-0.179)	-0.875 (-0.739)	---	0.353 (0.712)	---
4 $\frac{PAYMENT}{INCOME}$	-0.008 (-0.087)	-0.053 (-0.518)	---	---	---
5 CONCOST/P_b Atlanta	0.122 (1.708)	0.077 (0.483)	0.101 (2.229)	---	0.066 (1.420)
6 CONCOST/P_b Chicago	0.169 (2.053)	0.386 (2.545)	0.163 (2.688)	---	0.081 (1.522)
7 CONCOST/P_b Dallas	0.186 (1.820)	0.151 (1.277)	0.162 (2.629)	---	0.113 (1.843)
8 CONCOST/P_b San Fran	0.225 (2.501)	0.213 (1.953)	0.217 (3.197)	---	0.142 (2.287)
9 ΔEMP (%)	0.046 (0.547)	0.052 (0.615)	---	---	---
10 ΔINCOME (%)	0.310 (1.440)	0.297 (1.376)	0.276 (1.290)	---	---
11 ΔCONCOST (%)	-0.011 (-0.088)	0.002 (0.013)	---	---	---
12 ΔAGEPOP (%)	0.013 (2.002)	0.010 (1.479)	0.013 (2.158)	---	---
13 ΔMTR (%)	-0.839 (1.585)	-0.772 (-1.441)	-0.909 (-1.684)	---	---
14 HSTARTS/POP	-10.539 (-0.821)	-11.555 (-0.848)	---	---	---
R^2	0.336	0.361	0.329	0.009	0.163
Nobs	229	229	229	229	229

Table 7

Regressions of Excess Return A (No Leverage) on Forecasting Variables
Dependent Variable is $EXCESS_{Aa}$ from t+1 to t+5

Regression No.	1	2 (dummied)	3	4	5
1 Const	0.034 (0.240)	0.163 (1.076)	-0.126 (-1.092)	-0.037 (-0.736)	-0.159 (-2.169)
2 $EXCESS_{Ab}$	0.181 (1.949)	0.116 (1.263)	0.232 (2.256)	---	---
3 RENT / P_b	-1.876 (-1.471)	-1.848 (-1.400)	---	1.657 (2.463)	---
4 PAYMENT / INCOME	-0.303 (-2.659)	-0.383 (-3.184)	-0.184 (-1.908)	---	---
5 CONCOST/P_b Atlanta	0.137 (1.769)	0.218 (1.251)	0.097 (1.348)	---	0.184 (2.976)
6 CONCOST/P_b Chicago	0.168 (1.838)	0.357 (2.117)	0.168 (1.926)	---	0.201 (2.869)
7 CONCOST/P_b Dallas	0.247 (2.229)	0.286 (2.233)	0.174 (1.841)	---	0.288 (3.543)
8 CONCOST/P_b San Fran	0.258 (2.611)	0.132 (1.095)	0.243 (2.624)	---	0.300 (3.628)
9 ΔEMP (%)	0.090 (0.950)	0.095 (1.020)	---	---	---
10 ΔINCOME (%)	0.548 (2.330)	0.407 (1.721)	0.468 (1.982)	---	---
11 ΔCONCOST (%)	0.051 (0.363)	0.022 (0.159)	---	---	---
12 ΔAGEPOP (%)	0.013 (1.779)	0.015 (1.931)	0.015 (2.262)	---	---
13 ΔMTR (%)	-0.843 (-1.439)	-0.914 (-1.530)	-0.943 (-1.522)	---	---
14 HSTARTS/POP	-14.079 (-0.988)	-3.763 (-0.250)	---	---	---
R^2	0.559	0.582	0.536	0.109	0.281
Nobs	229	229	229	229	229

Table 8

Regressions of Excess Return B (Leveraged) on Forecasting Variables
Dependent Variable is $EXCESS_{Ba}$ from t+1 to t+5

Regression No.	1	2 (dummied)	3	4	5
1 Const	0.406 (0.667)	0.916 (1.393)	-0.312 (-0.619)	-0.253 (-1.052)	-0.839 (-2.372)
2 $W_{bt}-W_{bt-4}$	1.340 (2.496)	1.040 (1.985)	1.485 (2.585)	---	---
3 $\frac{RENT}{P_b}$	-11.926 (-2.384)	-10.345 (-1.848)	---	8.072 (2.476)	---
4 $\frac{PAYMENT}{INCOME}$	-1.838 (-4.220)	-2.116 (-4.600)	-1.250 (-3.147)	---	---
5 $CONCOST/P_b$ Atlanta	0.711 (2.041)	1.178 (1.470)	0.312 (0.943)	---	0.884 (2.964)
6 $CONCOST/P_b$ Chicago	0.817 (2.007)	1.632 (2.065)	0.608 (1.527)	---	0.966 (2.851)
7 $CONCOST/P_b$ Dallas	1.316 (2.704)	1.632 (2.065)	0.645 (1.469)	---	1.394 (3.557)
8 $CONCOST/P_b$ San Fran	1.301 (2.941)	1.574 (2.673)	0.991 (2.321)	---	1.452 (3.645)
9 ΔEMP (%)	0.650 (1.535)	0.634 (1.148)	---	---	---
10 ΔINCOME (%)	1.963 (2.053)	1.366 (1.419)	1.447 (1.410)	---	---
11 ΔCONCOST (%)	-0.025 (-0.043)	-0.166 (-0.286)	---	---	---
12 ΔAGEPOP (%)	0.056 (1.710)	0.072 (2.084)	0.066 (2.107)	---	---
13 ΔMTR (%)	-2.988 (-1.379)	-3.660 (-1.647)	-3.624 (-1.423)	---	---
14 HSTARTS/POP	-76.367 (-1.180)	-19.963 (-0.294)	---	---	---
R^2	0.615	0.640	0.571	0.109	0.284
Nobs	229	229	229	229	229

Notes to Tables 6, 7, and 8

Notes: $EXCESS_{Aa}$ and $EXCESS_{Ba}$ are from the expressions on page 9 using WRS_a, t+1 to t+5, . Row 2: $W_{bt}-W_{bt-4}$ is the lagged change in log price using price index b. $EXCESS_{Ab}$ is from the expression on page 9 using WRS_b t-4 to t. Row 3: $RENT/P_b$ is the ratio of rent on homes at time t to the price of homes at t. RENT, a measure of rent levels that is valid for comparison across cities, is for each city equal to RENTINDEX*RENT70/PRICE70. P_b is $\exp(WRS_b)$ (the WRS price index was in logs). Row 4: The variable PAYMENT is estimated mortgage payment divided by per capita personal income, time t. Rows 5-8: $CONCOST/P_b$ is a construction cost divided by price, time t. Since the construction cost index is the same for all cities in the base year by construction, it cannot be used for inter-city comparisons. Therefore, the variable for each city appears multiplied by a dummy which is 1.00 only for that city. Row 9: Percentage change in employment, t-4 to t, Row 10: Percentage change in real per capita income between t-4 and t, Row 11: percentage change in real construction costs, t-4 to t, Row 12: ΔAGEPOP is the percentage change in adult population (between ages of 25 and 44) t-4 to t, Row 13: Percentage change in marginal tax rate t-4 to t, Row 14: Housing starts, total for quarters t-4 through t divided by population at time t. See Table 1 for sources of data.

T-statistics (in parentheses) take account of overlap of observations of dependent variable and cross-section-time-series structure of data. Data are quarterly starting in 1970, first quarter. Data end in 1986, second quarter for each city except San Francisco, where data end in 1986, third quarter.

The regression reported in Column 2 included as well city dummies for Atlanta, Chicago, and Dallas. (Their coefficients are not reported here.)

forecast excess returns by definition A on page 9, and Table 8 uses them to forecast excess returns by definition B on page 9. There was only modest improvement in the R^2 for the price changes, to about a third. The excess returns are more forecastable since the real interest rate on the alternative asset which is used to compute the excess return is fairly forecastable.

Columns 1 and 2 in Tables 6, 7, and 8 use all of our forecasting variables as independent variables. Columns 3 and 4 drop some of the less significant variables to achieve a simpler forecasting relation. Columns 5, and 6 show some extremely simple forecasting relations.

The forecasting variables we consider include two that are measures of fundamental value relative to price: rent divided by price and construction cost divided by price. These are analogous to the dividend-price ratio that has been found to forecast stock market returns (see for example Shiller [1989] and Fama and French [1988]). When both of these are in the regression together (columns 1 and 2) the rent divided by price has the "wrong" sign, possibly reflecting multicollinearity problems. Thus, the rent divided by price was omitted from the regression reported in column 3. However, rent divided by price has a positive estimated coefficient (statistically significant in the excess returns regressions) in regressions with it as the only forecasting variable for price.

The estimated mortgage payment divided by per capita personal income variable is supposed to be an index of affordability, reflecting the difficulty that people face in carrying a mortgage. High values indicate that housing costs are out of line with income. Thus, we expect this variable to be negatively correlated with subsequent price changes. However, it was not significant in the regressions with price change as the dependent variable. It has the expected negative sign in the regressions with excess returns as dependent variable, though perhaps just because it is an interest-rate-related variable which is correlated with the interest rate on the alternative asset used to compute excess returns.

The change in employment and change in income variables had the expected positive signs, high values of these indicators of the strength of the economy portending price increases or high excess returns. But the former was never statistically significant, the latter only marginally so.

The change in construction cost variable was not statistically significant. The change in the adult population variable had a positive sign and was often significant. This variable is related to one used by Mankiw and Weil [1989] to predict housing prices in the United States.

Changes in the marginal income tax rate facing the marginal buyer is an important variable, but its sign in these equations is ambiguous ex ante. First, a decrease in marginal tax rate actually increases the cost of owning since property taxes and mortgage interest payments are deductible. But most statutory changes in marginal personal tax rates are accompanied with other tax provisions that may change the attractiveness of owner occupied housing relative to other assets. For example, the marginal rate reductions in the 1986 Tax Reform Act were accompanied by provisions that significantly curtailed the ability of taxpayers to shelter income with passive losses making home ownership one of the last commonly available "shelters." Similarly, ERTA of 1981 contained lower marginal rates and a host of provisions (ACRS, Safe Harbor Leasing, extension of the ITC, etc.) that changed the relative attractiveness of other assets. Finally, a cut in marginal rates may have an income effect; as disposable income increases, so will the demand for housing.

The variable MTR has a negative coefficient and is mildly significant in all equations. This indicates that the impact of marginal tax rates on the after tax cost of housing seems to be offset by other provisions. There is one other possible reason for the negative sign. The largest cut in tax rates (1981) took place at the same moment that interest rates were at extremely high levels (21% prime in the summer of 1981). These very high interest rates may have put sharp downward pressure on house prices at exactly the same moment that ERTA was cutting marginal rates.

Housing starts divided by population also has also the expected negative sign. High housing starts represent new supply on line that will tend to depress prices with a lag. However, the housing starts variable was not significant.

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