THE EFFICIENCY OF THE SINGLE-FAMILY HOUSING MARKET: AN EMPIRICAL STUDY

By

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This study examines the efficiency of the single-family housing market. It is hypothesized that *ex post* returns to housing are weak-form efficient; thus, investors cannot construct *ex ante* investment rules based on historical information that consistently yield above-normal risk-adjusted returns. Previous work is extended by examining the effects of time-varying marginal tax rates, unexpected inflation, and return volatility in the analysis of after-tax excess returns.

After-tax excess housing returns are found to be negatively correlated with marginal tax rates and positively correlated with unanticipated inflation. Furthermore, the data suggest that inflationary shocks were autocorrelated during the late 1970s and early 1980s, and not anticipated by either the bond or housing markets. When returns are

adjusted for both time-varying marginal tax rates and unanticipated inflation, after-tax excess returns are no longer significantly autocorrelated in three of the four markets studied. While persistent return "momentum" exists, when all transaction costs are considered, no investment rule can be constructed to consistently exploit any remaining serial dependence present in the markets. Hence, the hypothesis of a weak-form efficient market cannot be rejected.

Finally, limited preliminary evidence is found to suggest that expected after-tax excess returns to housing are weakly correlated (positively) with expected volatility. This provides provisional support that return volatility (risk) is priced by housing market participants. However, modeling expected volatility using an ARCH-M model reveals no evidence of time-varying risk premia in housing returns.

CHAPTER 1
INTRODUCTION

Purpose and Contribution of the Study

This study evaluates the efficiency of the single-family housing market using conventional methodologies reported in the finance literature. Although there exists a plethora of literature that models and tests the efficient market hypothesis (EMH), it is largely confined to corporate security market analysis. Research in the area of real estate market efficiency is quite limited. Few, if any, definitive answers to the efficiency question have emerged.

Previous studies by Rayburn, Devaney and Evans (1987) and Guntermann and Smith (1987) indicate that real returns to housing are not autocorrelated, hence, the hypothesis of a weak-form efficient market could not be rejected on that basis. However, Case and Shiller (1989) find that after-tax excess returns to housing (annual returns to housing in excess of the after-tax one-year treasury bill yield) are autocorrelated.[1] They indicate that home buyers, indifferent to time of purchase, can construct an investment

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[1]Case and Shiller find only limited evidence of autocorrelated real price changes in housing and indicate that single-family housing prices are not forecastable.

rule that yields above-normal returns. They suggest that real treasury bill yield information is not incorporated in the decision to invest in housing and on that basis reject the notion of a weak-form efficient market.

It is well known that any test of market efficiency is a joint test of some market equilibrium model. Fama (1976) demonstrates that even a serially correlated return series is consistent with market efficiency if equilibrium returns are serially correlated as well.[2]

Adopting the weighted repeat-sales housing price indices constructed by Case and Shiller (1987), this study examines whether their results are sensitive to alternative specifications of what a "fair" or equilibrium expected excess return is for single-family housing. This study explores the possibility that findings of serial correlation in the Case and Shiller study may be the result of serial

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[2]Serial correlation of stock returns is generally taken as evidence of market inefficiency because it seems reasonable to assert that the variance of price changes due to new information arrival is large relative to the variance in expected equilibrium returns over time. Hence, if new information arrives randomly, autocorrelation in returns is expected to be low. Serial correlation in treasury bill returns, however, is not judged to be a violation of weak-form efficiency because the treasury bill returns are expected to follow the level of inflation, which is a highly autocorrelated series. A discussion of the conditions necessary for market efficiency appears in Chapter 2.

correlation in home buyer marginal tax rates, *ex post* inflationary shocks, and return volatility.[3]

Significance of Housing Market Efficiency

The single-family housing market is a prime component of the U.S. economy. A substantial portion of aggregate U.S. wealth, housing is the most widely held of all investment assets. As shown in Figure 1-01, the estimated value of single-family housing approximates $5 trillion. In addition, nearly one-third of all income property, valued at $2.5 trillion (debt and equity), consists of multi-family residential property.

At $3 trillion, the estimated value of single-family home equity is approximately equal to that of corporate equity. The remaining $2 trillion value of single-family housing mortgages approaches the size of the U.S. Treasury debt.[4]

Furthermore, the largest portion of many individuals' investment portfolios is their leveraged equity in their home. Efficient capital allocation within this market segment could strongly affect economic conditions.

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[3]Case and Shiller assumed these factors do not change (are constant) or randomly change during the sample period. These assumptions are detailed later in Chapter 6.

[4]Statistics quoted are 1987 estimates of the Board of Governors of the Federal Reserve and Salomon Brothers, Inc.

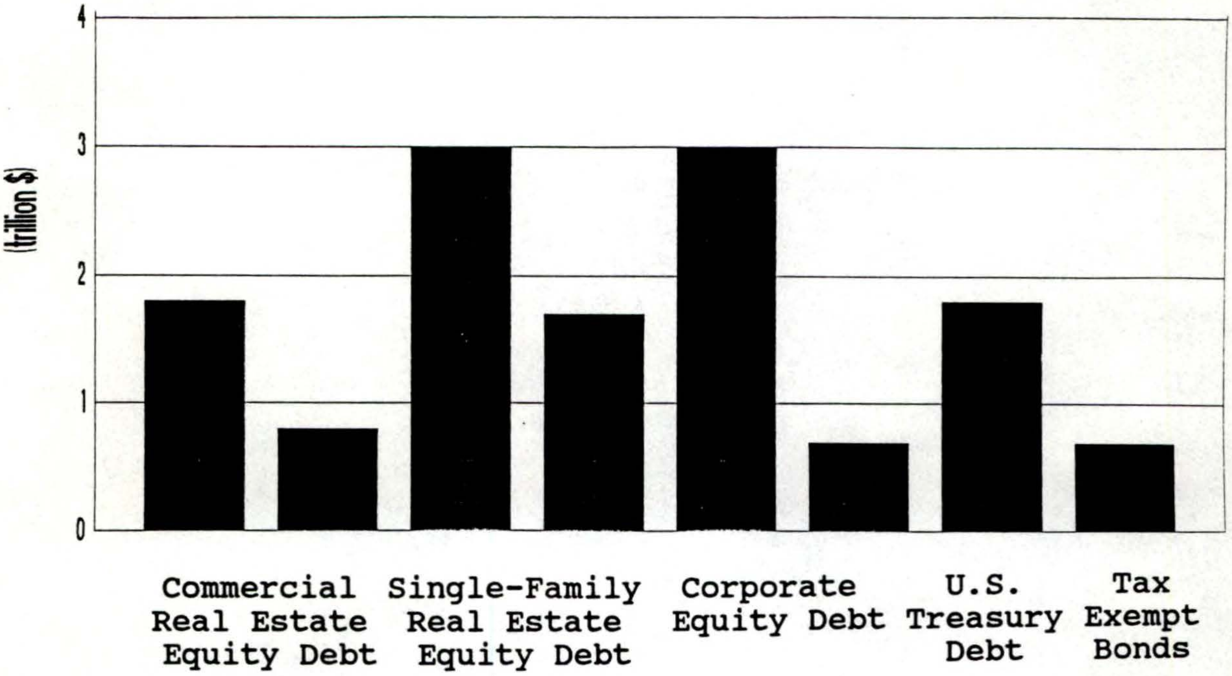

FIGURE 1-01 Compostion of the U.S. Capital Market
(December 31, 1986)
Source: Hartzell, Lepcio, Fernald, and Jordan (1987)

In addition to issues of general economic welfare, the efficiency of the market is important to the development of real estate economic theory. It is fundamental to fields such as valuation, investment, and urban economics.

Valuation theory assumes that all buyers and sellers are well informed. Grossman and Stiglitz (1976), in reference to capital markets, have suggested that in an efficient market, it is the activity of informed individuals that serves to set prices for the entire market. The market prices aggregate information so that all participants (both informed and uninformed) become informed by observing prices.

If single-family homes within a market are viewed as substitutes, then "the result of this process, which is the competitive process, is an arraying of all the sites {properties} in a hierarchy of land {property} values" (Ratcliff 1949, p.366).[5] Individual property value movements are not purely idiosyncratic. Under this condition, it is plausible that the activity of informed individuals in the real estate market highly influences, if not sets, the most probable selling price of homes in each market. While it is recognized that real estate market participants are not price-takers, findings of market efficiency strengthen the notion that informed individuals substantially influence market prices. This strengthens the concept of using the direct sales comparison approach in real estate valuation and may lead to alternative valuation methods.

Institutional investment in single-family housing occurs almost solely as whole mortgages or mortgage-backed securities (MBS). Savings and loans, commercial banks, government agencies, pension funds, and insurance companies, along with others who manage large portfolios of assets, collateralized by housing, have an interest in the stability and predictability of the housing market. Understanding the degree of housing market efficiency lends valuable support

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[5]Parenthetical insertions are by this author.

to this effort. Support of the EMH indicates that unanticipated returns to housing are random and normally distributed, an assumption central to much investment theory.

Finally, the concept of real estate market efficiency is central to evaluating the economic effects of local government policies such as zoning, growth controls, and property taxation. For example, the assertion that real estate markets are inefficient supports proponents of suburban growth restrictions.

Real estate inefficiencies, proponents argue, result in abnormally high risk-adjusted returns to developers of suburbs. This overstimulates investment in suburban real estate, resulting in excessive suburbanization--a misallocation of resources. The existence of efficient real estate markets, especially the single-family housing market, gives provisional strength to the argument that historical suburbanization trends may be allocationally efficient. It suggests that development patterns may arise primarily from individual locational preferences responding to demographic and technical change. However, it is important to note that informationally efficient markets do not necessarily imply allocational efficiency.

CHAPTER 2
THE EFFICIENT MARKET HYPOTHESIS

Types and Forms of Market Efficiency

The concept of efficiency is central to understanding both the positive and normative economics of a market. Economists commonly describe efficiency as the allocation of resources such that no reallocation can occur that increases the utility of some without decreasing the utility of others. A new allocation of goods in an economy is said to be "Pareto-preferred" to another if all economic agents are at least as well off or better off than under the prior allocation.[1] Thus, efficiency is a relative, not absolute, concept.

The financial economics literature distinguishes between three distinct, but interrelated, types of efficiency in capital markets: informational, allocational, and operational.[2] An "informationally efficient" market, as defined by Fama (1970), implies that all market prices fully and instantaneously reflect all relevant information.

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[1]This concept of efficiency is attributed to Vilfredo Pareto, a nineteenth-century Italian economist and sociologist.

[2]This study focuses on informational efficiency. Unless otherwise noted, the use of the term efficiency refers to informational efficiency.

In such a market, prices serve as an accurate signal for the allocation of scarce capital.[3] Funds flow to the most productive investments, those yielding the highest marginal risk-adjusted return, maximizing net social benefits.

Funds continue to flow to productive investments until the marginal risk-adjusted return on the least profitable investment equals the opportunity cost of borrowing. Subsequently, when the risk-adjusted marginal rate of return is equated for all borrowers and lenders, the market is said to be "allocationally efficient." Fama states that

> in general terms, the ideal is a market in which prices provide accurate signals for resource allocation: that is, a market in which firms can make productive-investment decisions, and investors can choose among the security prices that represent ownership of a firm's activities under the assumption that security prices at any time "fully reflect" all available information. A market in which prices always "fully reflect" available information is called "{informationally} efficient" (Fama (1970), p.383).

Consistent with the traditional concept of Pareto efficiency, all economic agents (borrowers and lenders) are better off in an efficient capital market.

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[3]It is important to differentiate between the allocational efficiency defined in the finance literature and that (Pareto-efficiency) defined in the mainline economics literature. Whereas the finance literature describes allocational efficiency on an after-tax (corporate) basis, Pareto-efficiency of investment requires that risk-adjusted marginal returns be equated before-tax. It maintains that the substitution effect resulting from taxes creates allocational inefficiencies. Thus, full Pareto-efficiency requires an efficient tax structure, an optimal level of public goods and services, efficient private goods and factors, no externalities, and labor resources that are efficiently distributed across regions.

Finally, "operational efficiency" considers the degree of transaction costs associated with the flow of funds. Perfect operational efficiency would require a market with no transaction costs. Informational efficiency is a necessary condition for both allocational and operational efficiency.

In the single-family housing market, informational efficiency implies that the distribution of market prices accurately reflects the expected risk-adjusted marginal return associated with the bundle of housing services at each location. This suggests that the market prices both private and public goods and services embedded in the purchase of real estate assets. As in the financial markets, this distribution of prices is useful in directing investment and consumption decisions to maximize welfare. If the housing market is informationally efficient, firms and households will choose to invest in and consume an optimal amount of housing services, given the current state of nature.[4]

Fama defines three "forms" of informational efficiency: strong-form, semistrong-form, and weak-form. Each form is based on a different notion of exactly what type of information is relevant in the phrase "all market prices

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[4]This condition assumes that housing capital is fluid and mobile, that negative externalities and congestible housing services are priced by the market, and that future conditions are accurately anticipated. In addition, it assumes investors have identical preference sets.

fully and instantaneously reflect all available relevant information."

Strong-form efficiency states that investors cannot consistently earn above-normal risk-adjusted returns using any or all information, whether private or public. This, of course, is highly restrictive, suggesting that even insider information is reflected in current prices. Grossman and Stiglitz (1976) have shown that such efficiency is impossible since costless information is both a sufficient and necessary condition in order for prices to fully reflect all available information. When information is costly, as in both financial and real estate markets, prices cannot reflect all information since those who incur costs gathering it would not be fully compensated. However, weaker forms of efficiency are not precluded by the existence of costly information. Investors acquiring costly information in such environments could be fairly compensated with returns commensurate with the cost of the information acquired.[5]

The information subset relevant for semistrong-form efficiency consists of all available current public

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[5]The cost of acquiring new information should equal the added return benefits such that the net present value, including entrepreneurial profit, will equal zero. Cornell and Roll (1981) argue that the average investor who uses costly security analysis data will outperform investors who use less information; but, their net returns will be identical. Some empirical evidence has been found to support this position.

information. This would include newly released information such as government economic statistics, investment advisory data, or information in newspapers or news services.

For weak-form efficiency the information set is further partitioned to include only publically available historical information, such as past asset prices or returns. Hence, by definition, weak-form efficient markets indicate that investors cannot construct investment strategies based on past returns alone that consistently yield above-normal risk-adjusted returns. Weak-form efficiency is a necessary condition for both strong and semistrong-form efficiency.

Conventional Tests of Weak-Form Efficiency

Random-Walk Theory and Autocorrelated Returns

Conventional tests of efficiency reported in the finance literature are generally based on the assumption that market equilibrium conditions can be expressed in terms of expected returns. The earliest references to this concept have been attributed to Bachelier (1900) and are now referred to as expected return theory or the "fair-game" model. Bachelier hypothesized that speculation should be a "fair-game"--in the long-run (aggregate) excess profits to speculation should be zero. Formally, this is defined

$$\epsilon_{j,t} = r_{j,t} - E(r_{j,t} | \Phi_{t-1}), \tag{2-01}$$

where

$$r_{j,t} = (P_{j,t} - P_{j,t-1}) \;/\; P_{j,t-1}, \tag{2-02}$$

and $P_{j,t}$ is the actual price of asset j at the end of period t; $r_{j,t}$ is the one-period return of asset j in t; $E(r_{j,t}|\Phi_{t-1})$ is the expected one-period return of asset j in t given the available relevant information set, Φ, at end of period t-1; $\epsilon_{j,t}$ is the residual (abnormal) return in t; and $\epsilon_{j,t}$ and $r_{j,t}$ are randomly distributed variables at t. Taking (mathematical) expectations of (2-01) yields,

$$E(\epsilon_{j,t}) = E[r_{j,t} - E(r_{j,t}|\Phi_{t-1})]. \tag{2-03}$$

The fair-game model is based on the behavior of average residuals, not the entire distribution.[6] In other words, on average, across a large number of observations, the expected return is equal to the average actual return. Therefore, by definition, the expected residual is equal to zero -- a fair-game,

$$E(\epsilon_{j,t}) = 0. \tag{2-04}$$

Modifications of the fair game model include the martingale, the submartingale, and the random-walk models. The martingale and the submartingale represent only minor modifications of the fair-game model. The random-walk model,

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[6]Fama (1970) indicates that the model properties are implied by the assumptions that (1) the conditions of market equilibrium can be stated in terms of expected returns, and (2) the relevant information set, Φ_t, is fully utilized by the market in forming equilibrium expected returns.

however, substantially strengthens the conditions necessary for efficiency.[7]

Samuelson (1965) established that randomly fluctuating returns represent a sufficient condition for weak-form efficiency.[8] Others reasoned that, if price changes can be assumed to be attributable to new relevant information priced by the market, and there exists no reason to believe that this information enters the market other than randomly, then it is reasonable to assume that in an efficient market one-period returns are statistically independent of one another (such as Alexander, 1961). This indicates that not only should the expected residuals be zero, but also residual returns should follow a random-walk.[9] Although somewhat imprecise, this describes an enhanced return generating process from that of the fair-game model.

Random-walk theory, with regard to returns, explicitly requires that a series of successive one-period returns is

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[7]The martingale is a fair game model where the expected price next period is the same as the price this period, $E(P_{j,t+1}|\Phi_t)=P_{jt}$. The submartingale simply says that the expected price next period is greater than the price this period, $E(P_{j,t+1}|\Phi_t)>P_{jt}$. Tests of these models typically compare the models' abnormal returns with those of the buy-and-hold strategy.

[8]In analyzing commodity futures contracts, Samuelson indicates that the expected price change of such contracts from period to period will be zero and is deemed a "fair-game."

[9]Prices may be hypothesized to follow a random-walk if price changes are independent and identically distributed. If price changes are expected to be non-zero prices are said to follow a random-walk with "drift".

both independent and identically distributed. Unlike the fair-game model, it is not based solely on average returns or average residuals. The probability density function of returns, $f(\ )$, must be the same for all t. Hence, the random-walk model can be described as

$$f(r_{j,t}) = f(r_{j,t} \mid \Phi_{t-1}), \tag{2-05}$$

where $r_{j,t}$ is the one-period return on asset j in time t. This says, that each one-period return can be characterized as an independent variable, where the conditional probability distribution (conditioned on the information set) is identical to the marginal probability distribution. Thus, the information set is independent of the sequence of returns.[10] Indicating,

$$E(\epsilon_{j,t}) = E(\epsilon_{j,t} \mid \Phi_{t-1}) = 0, \tag{2-06}$$

since all past relevant information is priced by the market, consistent with weak-form efficiency. In addition, all of the moments of the return distribution (i.e. the mean, variance, skewness, and kurtosis) must be the same with or without the relevant information set for all t.

Under these conditions, if we assume that the number of underlying transactions is sufficiently large in generating each one-period return, then by the Central Limit Theorem

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[10]This condition is described later to be consistent with the rational expectations hypothesis.

and (2-03) we can expect returns to be normally and independently distributed, $(\epsilon_{j,t}) \sim N(0,\sigma^2)$.

Previous empirical work indicates that stock returns do not meet all the distributional properties of the random walk.[11] However, it appears unlikely that investors can exploit any of the distributional anomalies. In addition, it is highly improbable that risk is constant over time; the variance of the return distribution is likely to be time dependent.[12] Therefore, applying strict distributional random-walk model criteria for tests of efficiency is widely viewed as quite restrictive.

Since returns are required to be independent, the random-walk model requires no autocorrelation of returns for all lags, including one-period lags. However, significant one-period autocorrelated returns are not inconsistent with the fair-game model.[13] This then, for purposes of testing efficiency, is viewed as the primary statistical difference between the random-walk model and the fair-game model.

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[11]Both Fama (1970) and Copeland and Weston (1988) provide a good review of this evidence.

[12]French, Schwert, and Stambaugh (1987) indicate that the volatility of stock returns is time-varying and autocorrelated. While random-walk theory assumes that the volatility of asset returns is constant, random changes in volatility should have no adverse affect on tests of efficiency.

[13]Copeland and Weston (1988) show that even though the residual, ϵ_t, is a fair-game variable, the conditional expectation of returns for t can depend on the return observed for t-1. Thus, the autocorrelation of one-period returns need not be zero in a fair-game model.

Together, the random-walk of independent returns and Equation (2-06), have important implications which form the basis of most efficient market tests.

Conventional tests of weak-form efficiency take the form of: (1) tests of autocorrelation, (2) trading strategies, or (3) runs tests, to evaluate the stochastic characteristics of the return time series. All represent testable derivatives of the expected returns theory. Consistent with the random-walk literature, studies involving tests of autocorrelation, most commonly tests of first-order autocorrelation, are often applied. Findings of autocorrelated returns indicate that future returns may be predicted from historical returns. Such findings reject the random-walk model hypothesis. However, to rejection the efficient market hypothesis (EMH) additionally requires that trading strategies, which account for transaction costs, be identified that exploit any serial dependence.

Trading Strategies

Like most theoretical models, the fair-game and random-walk models are limited in their ability to describe fully the necessary conditions of the efficient market hypothesis. The formal notation of this concept takes on a general notion that is not necessarily consistent with market efficiency. While the return generating process described by a random-walk implies independent and identical return distributions, variations of distributions are nearly impossible to exploit; hence, they do not necessarily imply a rejection of the EMH. In order for market efficiency hypotheses to be rejected, random-walk inconsistencies must be exploitable by market participants.

The existence of predictable patterns (serial correlation) of residual returns potentially represents relevant information. If this information can be used by investors to construct trading strategies that yield excessive profits, then the market is considered inefficient.[14] Only if the strategies consistently outperform a buy-and-hold option is the EMH definitively rejected.

If predictable return patterns exist but transaction costs are too high for the information to benefit investors they are rational not to use it. Consequently, under the latter scenario, the hypothesis of a weak-form efficient market cannot be rejected.

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[14]With respect to security returns, Alexander (1961) examines a number of trading rules (strategies) to determine whether they could outperform a buy-and-hold strategy. Of the strategies put forth the y% filter rule was examined rigorously. It was used to exploit the presence of positively correlated returns. This rule suggests that if the security price rose y%, the investor should buy-and-hold the security until its price moved down at least y% from the next high, then the investor should sell and go short. The short position is held until the price again rises y% above the low, at which time the investor covers the short position and buys long. Alexander concludes that such filters cannot beat the buy-and-hold position supporting the EMH.

Interestingly, the very use of a known pattern will serve to eliminate it. Weak-form efficiency does not rule out the fact that patterns in returns may exist over very short periods of time. For instance, if research by investors indicates that on Friday prices rise and on Monday prices fall, informed investors will sell at the end of Friday, forcing prices down, and buy at the end of Monday, raising prices. This, of course, will mitigate any long term "weekend effect".[15]

Joint Hypotheses

It is important to note that all tests of market efficiency are tests of joint hypotheses. While conducting primary tests of the EMH, constructed actual returns and expected equilibrium return models are typically applied. As with hedonic modeling, if a component either of the return or expected return series is misspecified or ignored, the primary test may yield spurious results. Thus, the equilibrium model is implicitly tested, as well as the explicit testable hypothesis.

For example, assume a dividend stream is a significant component of a return series of a particular asset and the dividend series is autocorrelated. Neglecting to include it in the construction of a return series could produce

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[15]French (1980) found significant evidence of a weekend effect in security returns using data from 1953-1977. However, it is not exploitable when transactions costs are considered.

findings that the market is inefficient (efficient) when, in fact, it is efficient (inefficient).

Rational and Adaptive Expectations

Price Extrapolation

The rational expectations hypothesis (REH), posits that prices are formed on the basis of the expected future returns, including components of appreciation and dividends adjusted for taxes. As introduced by Muth (1961), the REH requires both expected prediction residuals (errors) to equal zero and that prediction residuals are uncorrelated with the information set that is available to the forecaster at the time of the forecast. If the residuals are correlated with the information set it implies that the forecaster has not used all available information. Maddala (1988) formalizes this as

$$y^{*}_{t} = E(y_t | \Phi_{t-1}), \qquad (2\text{-}07)$$

where y^{*}_{t} is the predicted value of y_t in period t. Maddala indicates that the left-hand side of (2-07) should be interpreted as the "subjective" expectation and the right-hand side the "objective" expectation conditional on the data available when the expectation is formed. Both the descriptive and the formal definitions are consistent with random-walk theory (see Eq. 2-06); hence, the EMH implies rational rather than adaptive expectations on the part of investors.

Adaptive expectations hypotheses indicate that economic agents revise their expectations upward or downward, based on the most recent errors. Formally, this may be modeled

$$E(r_{j,t+1}) - E(r_{j,t}) = \Gamma \; [r_{j,t} - E(r_{j,t} | \Phi_{t-1})] \qquad (2.08)$$

with $0 < \Gamma < 1.$

This indicates that market participants adapt gradually to new information.[16] This is in contrast to the EMH where "all relevant information is fully and instantaneously reflected in the market".

The popularity of technical analysis in security markets is evidence that many investors think the market is adaptive, that it overreacts (underreacts) to new information leading to predictable corrections (resistance targets).

Similarly, the real estate market is often assumed to follow an adaptive pricing process. Models of expected housing returns found in the literature often extrapolate recent experience. In these models, high appreciation rates in period t raise the likelihood of high appreciation rates

[16]Other models of adaptive expectations such as distributive and polynomial lagged processes are well known and often applied.

in period t+1, thus, implying returns are positively correlated.[17]

Furthermore, at times reporters allege that some regional markets are highly overpriced, forecasting an impending long-term "bear" real estate market in these regions.[18] Here, market prices are believed to react slowly to excess supply and demand conditions. If true, investors may be able to take advantage of the predictable cyclical nature of these real estate markets.

Arbitrageurs

Arbitrageurs play an important role in maintaining market efficiency in both national and global security markets. The trading activity of both arbitrageurs and informed traders (those with better estimates of future states) serves to bring securities prices quickly into equilibrium. While it is impossible for investors to consistently earn abnormal risk-adjusted returns in an efficient market, short-term excessive returns are not

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[17]Since inflation, especially long-term inflation rates, are positively autocorrelated and returns to real estate are generally positively correlated with inflation, it is likely that such models largely capture the component of housing return due to general inflation.

[18]In March of 1990, *The Wall Street Journal* indicated that both the real estate markets in the northeast and west were highly overvalued and long-run "bear" markets were widely predicted. In the short-run this information may be self-fulfilling; however, if this information is used by market participants it should serve to mitigate long-run forecasts.

inconsistent with weak-form efficiency. In fact, the possibility of such returns is a necessity of market efficiency, since it serves as an incentive for price correcting arbitrage. Arbitrage trading serves to inform uninformed investors through the pricing mechanism.

The absence of any formal arbitrage activity in real estate markets may justify arguing that the market is inefficient. Real estate markets tend to be local or regional rather than national, futures contracts in real estate equities do not exist, and short selling is not available. All this curtails or perhaps eliminates arbitrage activity. Nevertheless, informed corporate, institutional, and individual investors actively participate in both debt and equity investment in real estate. Investors, including realty groups, systematically analyze regional real estate markets for mispriced properties and hold portfolios of rental housing on the basis of expected (abnormal) returns. If informed buyers and sellers participate and if agents disburse relevant information, prices may adjust quickly to new information in active markets. The lack of formal arbitrage activity does not preclude the notion of market efficiency.

Real Estate Market Efficiency

Historical View

Market imperfections. Historically, the real estate market has been widely perceived and casually asserted to be inefficient.[19] This is not surprising, since the market suffers high transaction and information costs. These costs result from a market distinguished by highly heterogeneous products cleared through a negotiated market rather than an auction market. Therefore, it is difficult for investors to identify and take advantage of market mispricing opportunities which would serve to correct the market. Nonetheless, the existence of transaction and information costs are not sufficient to preclude market efficiency.[20]

In addition, other real estate market imperfections such a barriers to entry, indivisible assets, and limited liquidity have been suggested as potential sources of market inefficiency.[21] Liu, Grissom, and Hartzell (1988) theorize

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[19]Jaffe and Sirmans (1984), in discussing real estate investment decisions, offer evidence that the real estate market is perceived to be inefficient. However, they indicate that scholarly work on the issue is limited.

[20]Analyzing security markets, Fama (1970) indicates that all available relevant information can be reflected in the market price despite the existence of high transaction costs. In addition, Copeland and Weston (1988) show that the existence of information costs is not inconsistent with weak-form market efficiency.

[21]Roulac (1978) provides a discussion of these issues. It is important, though, to distinguish between perfect markets and efficient markets. Of the two concepts, market efficiency is much less restrictive. For example, it does

that market imperfections are priced. Gau (1985) and others have argued that alternative forms of ownership and financing serve to mitigate market limitations to some extent. Unfortunately, the influence of such market innovations on market efficiency is not clear.

Abnormal Returns. Empirical work reporting excessive returns to real estate is often cited as evidence that real estate markets are inefficient.[22] Wendt and Wong(1965), Coyne, Goulet and Piconni (1980), Kaplan (1985), as well as others, indicate that returns on nonsecuritized real estate investments have consistently outperformed alternative asset groups, such as stocks and bonds, on both a risk-adjusted or unadjusted basis. Bruggeman, Chen and Thibodeau (1984), Miles and McCue (1984), and Hartzell, Hekman, and Miles (1987) find that returns on risk-adjusted commingled real estate funds (CREFs) dominate stock and bond returns, in addition to providing a superior inflation hedge. Finally, Burns and Epley (1982) find similar results for the returns on real estate investment trusts (REITs).

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not require markets to be frictionless or assets to be infinitely divisible. It simply requires market imperfections to be fully and instantaneously reflected in the market price. Hence, the EMH does not require conditions consistent with a perfect market.

[22]While the majority of the work cited in this section analyzes income properties (securitized and unsecuritized), housing returns are viewed as substantially correlated to other real estate assets. Unfortunately, research which relates housing market returns to other capital asset groups is quite limited.

In contrast, Smith and Shulman (1976) find no abnormal returns for REITs while Kapplan and Schwartz (1988) find stock and bond returns to substantially outperform those of real estate limited partnerships (RELPs). Finally, Ibbotson and Siegel (1984) indicate returns, unadjusted for risk, on direct real estate investments fall between those of stocks and bonds, as shown in Table 2-01 below.

TABLE 2-01
U.S. Capital Market Total Annual Returns
(1947 - 1982)

Asset Group	Arithmetic Mean Return	Geometric Mean Return	Standard Deviation
Common Stocks:			
NYSE	12.25%	10.90%	17.16%
AMEX (1960-1982)	9.33	6.66	23.30
OTC	15.17	13.05	21.41
Corporate Bonds:			
Long-Term	3.40	2.99	9.71
Intermediate-Term	4.84	5.03	6.41
Real Estate:			
Farms	11.32	11.09	7.21
Residential Housing	7.50	7.43	3.89
Commerical (1960-1982)	8.10	8.03	3.78
U.S. Government Securities:			
U.S. Treasury Notes:	4.38	4.49	4.85
U.S. Treasury Bonds:	2.99	3.31	8.57
Municipal Bonds:			
Long-Term:	2.21	1.65	11.12

Source: Ibbotson and Siegel (1984)

Liu, Grissom, and Hartzell (1988) have relaxed the standard assumptions of the capital asset pricing model (CAPM) to explicitly recognize limited marketability, high transactions costs, and a mildly segmented market. They suggest factors unique to the market imperfections of real estate, such as non-market risk and cost components, are priced by the market. These results are consistent with investment theory, past empirical findings, and the EMH. They argue that excessive risk-adjusted returns to real estate are an artifact of misspecified modeling. Unfortunately, infrequent transactions pose a difficultly to modeling and conducting extensive tests of directly held real estate investments.

<u>The Single-Family Housing Market</u>

Tests of efficiency conducted on returns to single-family housing are marked by an additional difficulty. Unlike most real estate assets, single-family housing is viewed as a consumption good as well as an investment. In this case consumption characteristics of housing assets may dominate the purchase decision. If consumers have heterogeneous housing preferences, tests of efficiency using individual house returns are likely to be judged invalid. Assuming preferences are not time dependent, aggregating constant-quality housing returns across a large number of transactions will reduce this limitation to some extent.

It is widely recognized, however, that many home buyers consider housing to be a major investment and price it with respect to its investment characteristics. In doing so, the home buyer considers both its implicit rental stream (dividends to housing) and the expected future selling price. While the dynamics of this process are poorly understood, these buyers form "subjective" estimates of the expected future value.[23] In this case, the price is formed as the "objective" present value of the expected after-tax "cash flows". Applications of investment motivated housing models are widely applied in the literature.[24]

In contrast to directly held real estate investment markets, single-family housing transactions occur frequently, and comparative price data are readily available. In addition, economic factors impacting future housing prices, such as mortgage rates or employment rates, are widely published and often appear in local newspapers.

Speculative "Bubbles"

Speculative "bubbles" have been alleged in some housing markets. At times, persistent price rises caused by speculative buying in regional markets appear to suggest such conditions. In the 1980s, markets in the northeast and California experienced substantial price increases while markets in Texas declined. Case (1986) examines large reoccurring price increases in the Boston market and Gyouko and Voith (1990) indicate evidence of a "California effect".

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[23]This terminology is consistent with Maddala (1988) in discussing rational expectations (see Eq. 2-06).

[24]Among other places, examples of this appear in the tenure choice literature such as Hendershott and Shilling (1982).

Speculative price "bubbles" suggest upward (downward) price inertia, implying positively autocorrelated returns and market inefficiency.[25] While some supporting evidence exists, rigorous study of the rationale for such conditions is viewed as preliminary.

[25]Hypotheses of rational speculative "bubbles" do exist in security markets; however, such theory is quite preliminary at this time.

CHAPTER 3
LITERATURE REVIEW

Principal Studies

Research on the efficiency of the housing market is an emerging topic in at least two areas of the real estate literature, investment and housing economics.[1] Each area has developed largely independent of the other. To date, no comprehensive effort has been made to review the full body of the literature.

Single market studies by Linneman (1986) and Rayburn, Devaney and Evans (1987), and Case and Shiller (1989) and inter-market studies by Guntermann and Smith (1987) and Hamilton and Schwab (1985) are the principal attempts to analyze the efficiency of the single-family housing market. These studies are described in detail in the following sections. In addition, a complete listing of the real estate market efficiency literature appears in Table 3-01 following a review of the principal work.

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[1]Works by Roulac (1978), and Jaffe and Sirmans (1984), among other provide a basis for relating the EMH to real estate. In addition, empirical work on the efficiency of income properties by Gau (1984, 1985) provided substantial impetus for testing the efficiency of the single-family housing market.

Linneman (1986)

In what has been described as "a first step in testing the efficiency of the single-family housing market," Linneman conducted a semistrong-form efficiency test of the Philadelphia market. Using a sample of 3,776 owner-occupied residential properties, the owners' contemporaneous assessments of value in 1975 were regressed on nearly fifty detailed structural and neighborhood characteristics considered to be "current public information". In essence, the study described a trading rule that examined whether undervalued houses (those with negative residuals) in 1975 represented arbitrage opportunities in 1978, *ex post*.

The findings indicate abnormal increases to the subsequent values of the undervalued properties.[2] However, when transaction costs were considered, assumed to be twelve percent, an insignificant number of properties were judged to be profitable arbitrage candidates. Thus, the hypothesis of a semistrong-form efficient market could not be rejected.

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[2]This result may be spurious. Since property values are owner-assessments, they can be assumed to be normally distributed across the actual "market" value. If it is assumed that the total variance remains constant between periods, it is likely that the residuals would have mean reverting tendencies. That is, properties with negative (positive) residuals would subsequently redistribute themselves toward the mean when reappraised, while those properties with zero residuals would randomly change to nullify any "abnormal returns". Thus, consistent with Linneman's findings, "undervalued" properties would appear to yield excessive returns and "overvalued" properties would yield less than normal returns.

The Linneman study is quite limited because it examines one market over a single period. In addition, is assumed implicit rental income to owner-occupied housing to be invariant and proportionate to home value. Consequently, the distribution of implicit rents and risk are assumed equal across properties of equal value. Finally, returns were estimated using each owner's assessment of value rather than actual transaction prices, potentially biasing the results.[3]

Rayburn, Devaney, and Evans (1987)

Rayburn, Devaney, and Evans (RDE) test a weak-form efficiency hypothesis from a sample of 140,000 residential transactions in Memphis, Tennessee, occurring from 1970 to 1984. Ten submarkets were delineated based on the Memphis City Planning Commission planning districts. From actual transaction prices a monthly series of mean prices was constructed on a square foot basis from a minimum of ten transactions for each of the ten submarkets.

Using log transformations of the time series and taking the first differences, subjective tests of autocorrelation

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[3]The use of subjective value estimates as a proxy for market value is likely to be biased. Since the housing market is heterogeneous, the implicit assumption that all homeowners have identical utility functions, equally valuing different housing characteristics, is highly restrictive. In addition, the estimates are severely contaminated with "private" information. Linneman examines the bias, arguing that it is insignificant. Yet, he concedes that the use of actual transaction price data would strengthen the study.

at various lags were performed. A time pattern of returns was revealed in seven of the ten submarkets.[4] However, applying Alexander's filter rule, a buy-and-hold strategy could not be outperformed for the period sampled, assuming transactions costs as low as two percent. Therefore, the weak-form efficiency hypothesis for each submarket could not be rejected. It was, however, rejected for some markets during the period 1970 to 1975. Reasons for this anomaly could not be explained.

Like Linneman, RDE's single market efficiency test was based solely on the change in prices; no explicit examination of the dividend stream to housing was conducted. It is particularly interesting to note also that no explanation was offered as to why submarkets appeared to yield different degrees of efficiency.

Case and Shiller (1989)

Case and Shiller examine four major metropolitan markets. Using actual transaction price data, repeated sales were identified and a citywide weighted repeat-sales (WRS) index was constructed. Case and Shiller argue that the use of repeat-sales price data for the construction of

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[4]As indicated later, Case and Shiller [1989] suggest the use of first differences of logged indices may induce spurious correlations.

an index is an improvement over other methods, controlling for quality changes in the sample.[5]

Weak-form efficiency tests were performed on the citywide index and on returns which included city-specific rental indices from the U.S. Bureau of Labor Statistics to proxy implicit rent. Consistent with previous work, they find that real changes in house prices are not predictable. However, Case and Shiller find that after-tax excess annual returns are positively autocorrelated and conclude that the single-family housing market is inefficient.

Assuming round-trip transaction costs to be zero, they indicate it is possible to construct a trading rule that yields abnormal returns to home buyers who are unconstrained as to time of purchase.[6] They suggest the most likely reason for the apparent inefficiencies is that "predictable" real interest rates are not priced by the market.

Nevertheless, their study is subject to rather strong assumptions. They assume that individual marginal income tax rates and property tax rates are constant during the study period. Additionally, they implicitly assume that treasury bill yields accurately predicted the volatile

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[5]A description of the WRS price index methodology appears in Appendix A, including a discussion of its strengths and limitations.

[6]They consider only first-time home buyers and home owners who want to trade up and therefore argue that transaction costs can be ignored since the purchase will be made anyway. In addition, they assume all costs associated with waiting to purchase are zero.

inflation rates of the late 1970s and early 1980s and that the risk premia to single-family housing are constant throughout this time period.

Guntermann and Smith (1987)

Guntermann and Smith, in a study of 57 metropolitan areas, test the housing market for weak-form efficiency using FHA-based housing price data from 1968 to 1982. They test for efficiency across major metropolitan markets rather than within each market, making their study distinctly different from the previously cited studies. In a second test of efficiency, they proxy implicit rental streams by including a city-specific rental variable computed from the rental component of the CPI for 27 of the metropolitan areas.

To their advantage, the use of FHA-based price data imposes a degree of homogeneity on their study. Annually, FHA financing has accounted for ten to forty percent of the value of all new loans on one- to four-family houses in the United States. These homes are largely confined to the middle segment of the U.S. housing market due to statutory maximum mortgage limits. Furthermore, FHA financing insures that the property is in good condition by meeting FHA specifications and complying with standard building code regulations.

Using a portfolio approach, Guntermann and Smith compare the correlation of aggregate unanticipated returns to the aggregate unanticipated returns of subsequent years.[7] Findings indicate no relationship of returns for lags of one to three years and a weak negative relationship in returns for lags of four to ten years. However, a trading rule based on ranked decile groupings of annual returns for the areas could not be constructed to outperform the buy-and-hold strategy after transactions costs were considered. Thus, the hypothesis of an efficient housing market could not be rejected.

While Guntermann and Smith's approach is elegant and their findings interesting, the applicability is quite limited. Even if findings had indicated that the housing market was inefficient, it would be difficult for investors or homeowners to exploit the market failure, since investors would need to be able to hold city indexed portfolios of housing subject to homestead property taxes across multiple cities. Alternatively, highly mobile homeowners could earn abnormal returns only by identifying and trading mean return properties that were substitutable across cities.

It is unreasonable to assume that investors, development firms, or individual homeowners would be able to

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[7]In using this approach they note that significant autocorrelation of actual returns is not inconsistent with market efficiency. *Ex post* actual returns may exhibit significant serial correlation over some time intervals even though expected returns would not. They point out that serial correlation of raw returns over long time intervals would be expected even in an efficient market due to such factors as changes in inflation rates and changes in tax rates.

exploit such market inefficiencies. A more pertinent question is whether investors (development firms, landholding companies, or homeowners) can earn abnormal risk-adjusted returns by sequentially investing in housing on the basis of publically available information within a particular metropolitan area.

Additionally, the FHA-based price data have their deficiencies. At least three important problems with the data can be identified. First, the maximum mortgage limit truncates the data base. Thus, there may be systematic moves in the index brought about by periodic statutory changes to the limit. With respect to the overall housing price index, Greenlees (1982) indicates that truncation causes the FHA price index to be biased downward prior to mortgage ceiling adjustments, and biased upward following ceiling increases. Second, homes within the sample that experience abnormal returns may subsequently move out of the sample again because of the truncated data. Finally, findings of efficiency in this market segment may not generalize to housing market segments using other types of financing. Segmented by value, location, and quality, FHA financed homes may appreciate in value at rates different from other market segments.

Additional Research

In other studies of housing market efficiency, Hamilton and Schwab (1985), using FHA-based price data, indicate that households failed to incorporate past appreciation in their housing return expectations during the period from 1976 to 1979. However, by incorporating more time periods Guntermann and Smith show that these findings cannot be generalized to include other periods. Finally, Essayyad, Marx, and Green (1987) also used FHA data to conduct conventional random-walk tests following Fama's (1970) methodology. They report findings of weak-form efficiency consistent with those of Guntermann and Smith. Unfortunately, their findings are severely limited by the length of the data series. Table 3-01 provides a listing of the real estate EMH empirical literature.

TABLE 3-01
Tabular Survey of the Real Estate EMH Empirical Literature

PANEL A Single-Family Residential Market Studies, (City-Specific)

Case and Shiller (1989)

Test Markets:	Atlanta, GA; Chicago, IL; Dallas, TX; San Francisco, CA.
Test Period:	1970:1-1986:2; 62 Quarters.
Index:	Quarterly index constructed using weighted-repeat sales methodology.
Test Form:	Weak-form efficiency.
Data:	Repeat sales were identified for each market from a data base of individual sales of single-family homes maintaine by the SREA Market Data Center in Atlanta, GA and the California Market Data Cooperative in Glendale, CA.
Major Results:	Hypothesis of weak-form efficient market is rejected since excess returns are found to be autocorrelated. Findings indicate buyers indifferent to time of purchase could earn abnormal returns. Suggest trends in real interest rates do not appear to be incorporated in prices.

Rayburn, Devaney and Evans (1987)

Test Market:	Memphis, TN; subdivided into 10 single-family housing submarkets.
Test Period:	1970:1-1984:12; 180 months.
Index:	Monthly mean price per square foot in each submarket, based on a minimum of ten transactions.
Test Form:	Weak-form efficiency.
Data:	Submarket transactions were identified from a data base of 140,000 sales maintained by the Chandler Company in Memphis, TN.
Major Results:	Hypothesis of a weak-form efficient market could not be rejected. Evidence of autocorrelated returns were found in seven of the ten submarkets. However, an investment rule could not be constructed to outperform the buy-and-hold strategy. No seasonal patterns were found to the rate of change in the index.

Skantz and Strickland (1987)

Test Market:	Houston, TX.
Test Period:	1977 and 1981.
Index:	Cross-sectional and time-series analysis using using ordinary least squares regression model.
Test Form:	Semistrong-form efficiency.
Data:	Actual transaction data on 183 houses located in two Houston-area subdivisions. Each transaction included price, square footage, lot size, financing, month of sale, age of home, location data.
Major Results:	Findings are consistent with a semistrong-form efficient single-family housing market. The market reacted quickly and rationally to new publicly available information.

Linneman (1986)

Test Market:	Philadelphia, PA.
Test Period:	1975 and 1978.
Index:	Cross-sectional analysis using hedonic model and a one-period modified capital asset pricing model.
Test Form:	Semistrong-form efficiency.
Data:	Annual Housing Survey, using owners assessment of value as well as structural and neighborhood variables.
Major Results:	Hypothesis of a semistrong-form efficient market cannot be rejected when transaction costs are included.

PANEL B Single-Family Residential Market Studies, (U.S. Market)

Gunterman and Smith (1988)

Test Market:	U.S. single-family housing market using 57 metropolitan areas.
Test Period:	1967:1-1982:1; 15 years.
Index:	Annual mean price per square foot for each metropolitan area for existing homes using FHA financing.
Test Form:	Weak-form efficiency.
Data:	FHA 203(b) data available in publications of the U.S. government.
Major Results:	Hypothesis of a weak-form efficient market could not be rejected. Evidence of autocorrelated returns found, however, trading rule cannot be constructed to outperform buy-and-hold strategy.

TABLE 3-01 -- continued

Essayyad, Marx and Green (1987)
Test Markets: 73 U.S. metropolitan areas.
Test Period: 1966-1985; 19 years
Index: Annual mean price per square foot for each metropolitan area for existing homes using FHA financing.
Test Form: Weak-form efficiency.
Data: FHA 203(b) data available in publications of the U.S. government.
Major Results: Hypothesis of a weak-form efficient market could not be rejected.

Hamilton and Schwab (1985)
Test Markets: U.S. single family housing market, using 49 metropolitan areas.
Test Period: 1974-1976.
Index: Cross-sectional analysis using hedonic model.
Test Form: Rational expectations.
Data: Annual Housing Survey. Estimates of appreciation are based on changes in each cities average sales price of existing single family homes insured under the FHA 203(b) program.
Major Results: Results do not support the rational expectations hypothesis. Households fail to accurately incorporate past appreciation in housing into their future expectations of housing appreciation.

PANEL C Multi-Family Residential Market Studies, (City-Specific)

Gau (1985)
Test Market: Vancouver, B.C.; apartment market.
Test Period: 1971-1980; 120 months.
Index: Price per square foot using a single, but different, property transaction each month within a defined submarket.
Test Form: Semistrong-form efficiency; event study.
Data: Apartment transaction data [See Gau (1984)]
Major Results: Could not reject semistrong-form efficiency hypothesis. Investors are unable to earn abnormal risk-adjusted returns on the basis of new government policy information or interest rate movements.

Gau (1984)
Test Market: Vancouver, B.C.; apartment and commercial markets
Test Period: 1971-1980; 120 months.
Index: Price per square foot using a single, but different, property transaction each month within a defined submarket.
Test Form: Weak-form efficiency.
Data: Transaction data including price, date, location, use, square footage, and the number of units.
Major Results: Could not reject the hypothesis of a weak-form efficient market. Study reports a general absence of significant autocorrelation and support for the fair game model as a representation of price behavior in real estate markets.

PANEL D Multi-Family Residential Market Studies, (U.S. Market)

Isakson and McInish (1988)
Test Market: 18 U.S. metropolitan areas.
Test Period: 1978-1983; 5 years.
Index: Examined the income component of multi-family housing returns for autocorrelation. Did not include the capital appreciation component.
Test Form: Weak-form efficiency.
Data: Average income and operating expenses of apartments reported by the Institute of Real Estate Management (IREM).
Major Results: Hypothesis of a weak-form efficient market could not be rejected for lags of one year; however, the hypothesis was rejected for lags of two and three years.

TABLE 3-01 -- continued

PANEL E	Urban Land Market Studies, (City-Specific)
Davies (1977)	
Test Market:	London, Canada
Test Period:	1966-1973; 93 months.
Index:	Monthly median sales price of vacant lots.
Test Form:	Semistrong-form efficiency.
Data:	Land transactions reported in Teela Digest.
Major Results:	Evidence indicates that builders form their expectations about future prices and costs primarily on the basis of the rate of change in the profitability of building and selling houses four months prior to the current period. However, evidence suggests that the adoption of restrictive land use policies triggered an immediate and sustained speculative increase in prices due to increased costs or uncertainty.
Adams, Milgram, Green, and Mansfield (1968)	
Test Market:	Northeast Philadelphia.
Test Period:	1945-1962, 14 years.
Index:	Cross-section hedonic model with time trend dummy.
Test Form:	Residual analysis with macroeconomic variables.
Data:	1111 land transactions and associated land characteristics, such as access, condition, and zoning.
Major Results:	Interest rates were inversely related to regression residuals and positively related to construction activity and general economic conditions. Land development trends were anticipated and priced by the market

PANEL F	Commercial Market Studies, (City-Specific)
McIntosh and Henderson (1989)	
Test Market:	Dallas-Fort Worth, Texas.
Test Period:	1979-1985, 73 months.
Indices:	Three indices were constructed: (1) price per square foot of net rentable area, (2) gross income multiplier, and (3) net income multiplier, each using a single, but different, property transaction each month.
Test Form:	Weak-form efficiency
Data:	127 office transactions and associated characteristics, such as address, vacancy rates, financing, land area, age, condition, and parking.
Major Results:	Hypothesis of a weak-form efficient market could not be rejected.

PANEL G	Real Estate Securities Market Studies, (Foreign Market)
Locke (1985)	
Test Market:	British and Australian income producing real estate markets.
Test Period:	1980s, Specific period not by study.
Indices:	Indices were constructed using the ratio of the change in monthly prices of the property investment trusts.
Test Form:	Weak-form efficiency
Data:	Companies listed on each respective country's stock exchange which invested primarily in real estate were identified and aggregated to form an index of real estate company returns.
Major Results:	Hypothesis of a weak-form efficient market could not be rejected.

Summary of the Literature

In summary, research in the area of real estate market efficiency is quite limited and findings have been mixed, largely due to insufficient data. Linneman (1986) indicates that the hypothesis of a semistrong-form efficient housing market cannot be rejected. Rayburn, Devaney and Evans (1987), and Guntermann and Smith (1988) report that the housing market is consistent with weak-form efficiency. In addition, these three studies indicate that trading rules cannot be constructed that consistently yield abnormal risk-adjusted marginal returns to housing net of transaction costs.

In contrast, Case and Shiller (1989) reject the hypothesis of a weak-form efficient housing market. Although, they find real house prices to be unpredictable, after-tax excess returns are found to be positively autocorrelated. Case and Shiller indicate that buyers indifferent to time of purchase can construct a trading rule that yields abnormal returns. However, their study is subject to some rather strong assumptions. They assume home buyer marginal tax rates and risk-premia are constant and although evidence exists to the contrary, inflation expectation errors by investors are assumed to be random during the unprecedented inflation of the 1970s.

While important, these findings are partial and preliminary. Data limitations are such that there exists little hope of definitively answering the efficiency question by conducting comprehensive studies on the total U.S. housing market. Until such time that adequate data become available, it appears that a piecewise research agenda is required to resolve definable market specific efficiency issues.

CHAPTER 4
GENERAL HYPOTHESES

This study hypothesizes that the single-family housing market is weak-form efficient. This implies that investors (home buyers) cannot construct *ex ante* investment rules, on the basis of historical information alone, that consistently yield above-normal risk-adjusted returns. The study, in examining both real returns and after-tax excess returns, relaxes the strong assumptions of Case and Shiller (1989) and tests the robustness of their results.

More specifically, the assumptions that home buyer marginal tax rates are constant, inflationary shocks are stochastic, and the risk-premia to housing do not vary are relaxed. It is speculated that the findings of autocorrelated after-tax excess returns, reported by Case and Shiller, are largely induced by the autocorrelated attributes of these factors. Therefore, when *ex post* returns are constructed using time-varying marginal tax rates and are adjusted for unexpected inflation and time-varying return volatility, abnormal returns (deviations from expected returns) are hypothesized to follow a random-walk. Thus, consistent with the EMH, there exists no information in the historical housing return series that can be used to consistently earn excessive profits.

Expressed formally, the model defined earlier (see Eq. 2-01), is restated in terms of housing returns such that,

$$E(\epsilon_{n,t+1}) = E[R_{n,t+1} - E(R_{n,t+1}|\Phi_t)] = 0, \qquad (4\text{-}01)$$

and

$$R_{n,t+1} = (P_{n,t+1}-P_{n,t})/P_{n,t}, \qquad (4\text{-}02)$$

where $R_{n,t+1}$ is the actual nominal return in period t+1 to housing, $E(R_{n,t+1}|\Phi_t)$ is the expected nominal return in period t+1 to housing given the information set, Φ, in period t, and $\epsilon_{n,t+1}$ is the abnormal return, which is assumed to be normally distributed with mean, zero, and variance, σ^2.

As described previously, if the market is efficient differences between the actual *ex post* nominal return series and the expected nominal return series can be characterized as a random-walk. Findings of autocorrelation in the predicted errors suggest that not all available historical information has been used in setting prices, and provisionally rejects the EMH.

Expressing Equation 4-01 in terms of after-tax excess returns yields

$$E(\epsilon_{t+1}) = E[(R_{n,t+1}-(1-\tau)RF_{t+1})-E((R_{n,t+1}-(1-\tau)RF_{t+1})|\Phi_t)] = 0, \qquad (4\text{-}03)$$

where RF_{t+1} is the risk-free rate in period t+1, τ is the home buyer's marginal tax rate, and the difference between the actual nominal return and the after-tax risk free rate is defined as the after-tax excess return.[8] From this expression it can be seen that if income tax rates are time-varying, they must be included in the construction of actual returns, since they are imbedded in the formation of home buyers return expectations. Thus, assuming marginal tax rates to be constant may induce spurious autocorrelation in the prediction errors.

Additionally, tests of efficiency assume that market inflation expectations are accurately imbedded in both the risk-free rate and housing return expectations. If inflation prediction errors are random, they will have no impact on any subsequent tests of efficiency. However, there is evidence that during the volatile 1970s and early 1980s, markets in general systematically underestimated price level changes, resulting in autocorrelated inflationary shocks.[9] Since this condition appears to be an enigma confined to the time period studied, tests of efficiency over this sample period may be biased and judged

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[8]Income taxes on returns to housing are assumed to be zero. Since implicit rent is not taxed and capital gains due to appreciation can be avoided by continuously rolling over gains on principal residences, this assumption is reasonable. Additionally, consistent with this and other work, RR = MAX [MB, (1-t)TB] + RP. This states that the required (expected) return on housing, RR, is equal to the maximum rate between municipal bills and the after-tax treasury bill, plus some risk premium. As shown later, assuming the marginal tax rate of the median priced home buyer, the after-tax rate on the treasury bill is the dominant security in determining the required rate of return on housing.

[9]A discussion of this evidence is found in Chapter 8.

myopic. It is therefore argued that the after-tax excess returns to housing should be adjusted for unexpected inflation and the Case and Shiller (1989) results examined to determine their sensitivity to such conditions.

Using Equation (4-03), the general condition can be expressed to explicitly consider the impact of inflationary shocks such that,

$$E(\epsilon_{t+1}) = E[(R_{n,t+1}-(1-\tau)RF_{t+1}-R_{ui,t+1})-E((R_{n,t+1}-(1-\tau)RF_{t+1})|\Phi_t,\Phi_{\pi,t})] = 0, \quad (4\text{-}04)$$

where $R_{ui,t+1}$ is the after-tax excess return attributable to unexpected inflation, and $\Phi_{\pi,t}$ denotes the information available in t to form inflationary expectations in t+1.

Still, (4-04) assumes that the risk-premium associated with housing returns is constant. Recent research provided some evidence of time-varying risk-premia in security markets; however, such tests of risk-premia have yet to be conducted in real estate markets.[10] If changes in risk are random or constant, the removal of the return component due to risk is unnecessary and tests of efficiency are not biased. Omitting conditions of autocorrelated risk-premia however, will induce spurious conditions of autocorrelated returns and bias the subsequent tests of efficiency.

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[10]A discussion of time-varying risk-premia is found Chapter 9.

Expressed to include volatility expectations, Equation (4-04) can be explicitly modeled by simply stating

$$E(\epsilon_{t+1}) = E[(R_{n,t+1}-(1-\tau)RF_{t+1}-R_{ui,t+1})-E((R_{n,t+1}-(1-\tau)RF_t)\,|\,\Phi_t,\Phi_{\pi,t},\Phi_{\sigma^2,t})] = 0, \qquad (4\text{-}05)$$

where $\Phi_{\sigma^2,t}$ denotes all historical information on risk in the housing market available in t.

This model formally states this study's general hypothesis; when after-tax excess returns to housing are constructed considering time-varying marginal tax rates, unexpected inflation, and return volatility, abnormal returns are hypothesized to follow a random-walk with mean equal to zero, $(E(\epsilon_t)=0)$.[11]

[11] Alternatively, this condition can be formalized as follows: Assume R = MAX [MB,(1-t)TB] + RP + e,

where R is the actual nominal return to housing, MB is the yield on municipal bills, (1-t)TB is the after-tax yield on treasury bills, RP is the risk premium to housing and e is a random error term. Separating yields:

$$r + r_{ei} + r_{ui} = (1-t)(tb+ei) + RP + e$$

where r is the expected real return to housing, r_{ei} is the expected return to housing due to inflation, r_{ui} is the unexpected return to housing from inflation, tb is the real treasury bill yield, and ei is the expected inflation rate imbedded in the treasury bill. Under this condition expected returns equal actual nominal returns less returns due to unexpected inflation, $E(R) = R - r_{ui}$. Rearranging, we get: $E(R_e) = R - (1-t)(TB) - r_{ui} = RP + e$,

where $E(R_e)$ is the expected excess return, and e is a random variable. Therefore,

$$E(e) = E[(R-(1-t)(TB)-r_{ui})-E(R_e)] = 0,$$

where E(e) is a random variable with mean zero and variance σ^2, assuming the risk-premium is random.

CHAPTER 5
TESTABLE HYPOTHESES

To test the hypothesis that the housing market is weak-form efficient, conventional methodologies reported in the finance literature are employed. Historical single-family housing returns are assumed to constitute the relevant information set and are hypothesized to be instantaneously captured in the future return expectations of home buyers. Therefore, in a weak-form efficient single-family housing market, home buyers cannot predict future abnormal returns to housing on the basis of past price or return information and consequently, cannot consistently "time the market".

If home buyers expect a positive return on their investment, future price movements should follow a random-walk return generating process with "drift". Consistent with conventional tests of the random-walk model, *ex post* returns to housing are constructed, tested for autocorrelation and the findings compared to previous findings. In addition, the test methodology in Chapter 8 examines the random-walk characteristics of any abnormal returns (return prediction errors) to housing.

Hypotheses of the EMH are primarily tested by examining the first-order autocorrelation characteristics of each relevant return series. Regression coefficients of the alternative return series are estimated and evaluated using the general linear model,

$$R_{(t)} = \beta_0 + \beta_1 R_{(t-L)} + u_{(t)}, \qquad (5\text{-}01)$$

where $R_{(t)}$ is the relevant *ex post* return in period t and L denotes the number of periods lagged. The expected return that is unrelated to the previous return is captured in the intercept term, β_0. Since home buyers expect a positive return in housing assets, the intercept is postulated to be positive when testing real housing returns and zero if testing abnormal returns to housing.

The slope of the regression, β_1, measures the correlation between the return in period t-L and that in period t, given β_0. Under the hypothesis that the market is weak-form efficient, $\beta_1 = 0$, should not be rejected. A significant positive or negative slope coefficient suggests that all historical return information is not used by investors. Under this scenario it may be possible, given transaction costs, to devise an investment rule based on this information to earn excessive profits. The error term, $u_{(t)}$, is assumed to be independent and normally distributed with mean zero and variance, σ^2.

In addition to testing annual returns to housing, the random-walk characteristics of the quarterly returns are briefly examined. Autocorrelation functions of the quarterly returns are estimated and examined at various lags. If there is no significant autocorrelation at important lags (first or fourth) then it is concluded that there is no autoregressive or moving average structure in the time series. In addition, the results of a chi-square test, the Q-statistic, for no time pattern in the series are reported for each quarterly return series.[1] If the Q-statistic is close to zero, we accept the null hypothesis that there exists no time pattern in the return series. This result is consistent with the a weak-form efficient market.

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[1]If the autocorrelation function, ACF, is defined as,

$$ACF_L = \text{covariance}\ (R_{(t)}, R_{(t-L)})\ /\ \text{variance}\ (R_{(t)}),$$

the Q-statistic is composed of the first L residual autocorrelations $ACF_1, \ldots ACF_L$, such that,

$$Q = T\ \Sigma\ ACF^2_{(L)}, \quad \text{for } L=1 \text{ to } L.$$

Hence, the Q-statistic is the sum of squared independent normal random variables, with mean 0 and variance 1/T. Box and Pierce (1970) have shown that it is distributed as chi-square.

CHAPTER 6
TEST METHODOLOGY, BASE MODEL, AND DATA

This chapter describes the methodology used to test the housing market EMH. An overview of the entire test methodology is first outlined. Following this, the initial model (base model) and data are described in detail. Initial tests of the EMH are then conducted using the base model and the results reported. The chapter concludes with a discussion of the base model assumptions and their potential impact on the initial test findings.

Overview of the Test Methodology

The test methodology employed is comprised of four principal steps. In each step, an assumption of the preceding model is relaxed and a time-series of housing returns constructed. Conventional tests of the EMH outlined in Chapter 5 are applied to each of the constructed return series and the results evaluated.

In the first step, real returns and after-tax excess returns to single-family housing are estimated for four U.S. metropolitan markets (Atlanta, Chicago, Dallas, and San Francisco) using methodologies similar to Case and Shiller

(1989).[1] To insure the accuracy of the constructed data, tests of autocorrelation, replicating Case and Shiller, are conducted on each return series. These initial tests provide baseline statistics used to compare and evaluate the subsequent tests of weak-form efficiency conducted in the following phases under alternative model assumptions. The results of these initial tests are reported later in this chapter.

In the second step, using data from Ling (1990) and other secondary sources, a series of time-varying home buyer marginal tax rates are estimated for each city. After-tax excess returns to housing are then reconstructed using the city-specific home buyer time-varying marginal tax rates rather than a constant marginal tax rate, as assumed previously by Case and Shiller. Tests of autocorrelation are conducted on both the annual and quarterly return data. These procedures and the data used are discussed in Chapter 7.

After-tax excess returns to housing are adjusted for unexpected inflation in the third step. In this phase a step-ahead autoregressive integrated moving average (ARIMA) model is used to model annualized inflation expectations on a quarterly basis. The values forecasted by the ARIMA model are compared to survey data available every other quarter to

[1]Base model after-tax excess returns to housing use the same procedure and are identical to those estimated by Case and Shiller.

determine their validity. The after-tax excess returns are then regressed on the unexpected inflation series and the estimated regression coefficients evaluated. The estimated intercept and slope coefficient of the regression are interpreted as the expected return to housing and the after-tax excess return component to housing associated with unexpected inflation, respectively. Additionally, the residual is defined as the abnormal after-tax excess return to housing. Abnormal after-tax excess return series are constructed for each city and their random-walk characteristics statistically examined. This phase of the EMH test is described in Chapter 8.

Step four, discussed in Chapter 9, examines the impact of time-varying return volatility on expected returns. An autoregressive conditional heteroskedastic (ARCH) model is used to estimate expected after-tax excess returns adjusted for return volatility. If the risk-premium to housing is time-varying we would anticipate expected volatility (risk) to be positively correlated with expected after-tax excess returns and unexpected volatility to be negatively correlated. In this model, if volatility is shown to be time-varying and predictable, the random-walk properties of the model residuals, defined as abnormal returns, are examined and subsequently used to test the hypothesis that the single-family housing market is weak-form efficient.

Description of the Base Model

Real Returns to Housing

Housing returns are constructed using both the house appreciation component and the dividend (implicit rent) stream.[2] Using the quarterly WRS housing price indices for each market reported by Case and Shiller (1987), *ex post* annual real returns for sample j in quarter t, $R_{r,j\_(t)}$, are defined as

$$R_{r,j\_(t)} = \exp[WRS_{j\_(t+4)} - WRS_{j\_(t)}] + \frac{C_j[D_{(t)}+D_{(t+1)}+D_{(t+2)}+D_{(t+3)}]}{\exp[WRS_{j\_(t)}]} - [CPI\text{-}U_{(t+4)}/CPI\text{-}U_{(t)}]. \qquad (6\text{-}01)$$

$WRS_{j\_(t)}$ is the nominal WRS price index (in logs) constructed from sample j in quarter t; C_j is a constant used to calibrate the model; $D_{(t)}$ is the quarterly city-specific residential rent component of the consumer price index CPI-U; and $CPI\text{-}U_{(t)}$ is the quarterly city-specific consumer price index, net of shelter.

In Equation (6-01), the annual housing return due to the house price appreciation component is captured by the

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[2]Real changes in house prices were analyzed by Case and Shiller (1989). In contrast, this study includes the return component due to implicit rent and analyzes the predictability of total real returns to housing. If it is assumed that implicit rent is not taxed and capital gains can be avoided, these returns can be interpreted as real after-tax returns.

first term, $\exp[WRS_{j\_(t+4)} - WRS_{j\_(t)}]$, while returns due to the dividend stream to housing are captured by the second, $(C_j[D_{(t)}+D_{(t+1)}+D_{(t+2)}+D_{(t+3)}])$ over $\exp[WRS_{j\_(t)}]$. The nominal return to housing is then adjusted for changes in the general price level, $[CPI\text{-}U_{(t+4)}/CPI\text{-}U_{(t)}]$, to derive real returns to housing.

C_j is initially set so that the average dividend to price ratio is five percent during the study sample period, hence,

$$\left[\sum_{t=1}^{n} [C_j\,[D_{(t)}+D_{(t+1)}+D_{(t+2)}+D_{(t+3)}]/\exp(WRS_{j\_(t)})]\right] / n = 0.05 \tag{6-02}$$

for each city.[3]

<u>After-Tax Excess Returns to Housing</u>

Excess returns, are defined as the realized return (actual) on a risky asset in excess of the contemporaneous risk-free interest rate. Thus, $R_e = R_a - R_f$, where R_e denotes the excess return, R_a is the realized return on a risky asset, and R_f is the risk-free rate of return. After-

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[3]If C_j is set so that the assumed dividend to housing ratio differs from the actual ratio and the actual ratio is not constant, subsequent tests of the random-walk properties of the time series will be biased. Since sizable returns due to price appreciation are commonly anticipated during the sample period, the five percent dividend to price ratio, assumed by Case and Shiller, appears quite high. However, the impact of this assumption on tests of autocorrelation is found to be minor.

tax excess returns are defined simply as excess returns constructed on an after-tax basis. For housing, the after-tax excess return for sample j in period t, $R_{e,j\_(t)}$, is formally defined,

$$R_{e,j\_(t)} = \exp\,[WRS_{j\_(t+4)} - WRS_{j\_(t)}] + \frac{C_j[D_{(t)}+D_{(t+1)}+D_{(t+2)}+D_{(t+3)}]}{\exp\,[WRS_{j\_(t)}]} - 1 - [(1-\tau_{(t)})TB_{(t)}/100], \quad (6\text{-}03)$$

where $\tau_{(t)}$ is the home buyer marginal personal income tax rate and $TB_{(t)}$ is the one-year treasury bill rate. Here it is assumed that neither capital gains nor dividends to housing are subject to income taxes; however, interest on risk-free securities is considered taxable.[4]

Description of the Data

To estimate the models, quarterly WRS price indices, available for the metropolitan areas of Atlanta, Chicago, Dallas, and San Francisco, were obtained directly from Karl Case. Constructed using a weighted repeat-sales methodology

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[4]Adjusting after-tax housing returns to account for property tax changes and services received from the local government would be appropriate to model, however, adequate data do not appear to be available at this time. Therefore, any subsequent tests of autocorrelation assume that year-to-year changes in these variables are not serially dependent. In addition, base model marginal tax rates, τ, are assumed to be constant at thirty percent, consistent with Case and Shiller.

from a data base consisting of nearly one million recorded transactions (221,876 Atlanta, 397,183 Chicago, 211,638 Dallas, 121,909 San Francisco), each city index extends from the first quarter of 1970 and through the second quarter of 1986.[5]

The underlying transactions data for Atlanta, Chicago and Dallas, as well as the transactions data from before 1979 for San Francisco, were originally obtained from the Society of Real Estate Appraisers' Market Data Center in Atlanta, Georgia. The transaction data for San Francisco from 1979 through 1986 were obtained from the California Market Data Cooperative in Glendale, California.

To estimate the dividend to housing, city-specific residential rent indices are obtained from the U.S. Bureau of Labor Statistic (BLS) - Consumer Price Index Division in Washington D.C. General inflation for each city is estimated using the city-specific CPI-U, net-of-shelter, price level index from the BLS. One-year and three-month treasury bill yields as reported in the *Federal Reserve Bulletin* are used as a proxy for the annual and quarterly risk-free rate.

A description of the data used to construct home buyer marginal tax rates and inflationary expectations in the second and third steps of the analysis is deferred until

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[5]Sales transaction data for San Francisco were available for the third quarter of 1986. The empirical results for this market include this additional period.

Chapters 7 and 8. Table 6-01 below, summarizes the data used in testing the base model.

TABLE 6-01
Summary of the Data Used to Estimate the Base Model

Series	Dates Used	Freq	Source
Weighted Repeat Sales Housing Return Indices (Samples A,B, and T).	70:1-86:2	Q	Karl Case & Robert Shiller.
Yield on one-year treasury securities (Nominal).	70:01-86:06	M	*Federal Reserve Bulletin*.
Yield on three-month treasury securities (Nominal).	70:01-86:06	M	*Federal Reserve Bulletin*.
CPI-U, Net of Shelter (City-Specific).	70:1-86:2	Q	Bureau of Labor Statistics (BLS).
CPI-U, Residential Rent (City-Specific).	70:1-86:2	Q	Bureau of Labor Statistics (BLS).

Note: M and Q indicate that the data series used were available monthly or quarterly, respectively. Samples A and B refer to half-sample indices for each city, while sample T refers to the city index estimated from the complete sample of sales transactions.

Base Model EMH Test Results

Annual real returns and annual after-tax excess returns to single-family housing are first estimated on a quarterly basis from 1970:1 to 1986:2 for Atlanta, Chicago, and Dallas and from 1970:1 to 1986:3 for San Francisco. Later, quarterly returns are estimated for each series over the same sample period on a quarterly basis. Statistics of real quarterly returns for each city are reported in Table 6-02.

Consistent with previous work, real housing returns are estimated to be similar to that of common stock during the study period, while the variances of housing returns are less than that of stock. In addition, the regional nature of the real estate market is evidenced by low cross-correlation statistics, although real returns in each of the cities studied are positively correlated.

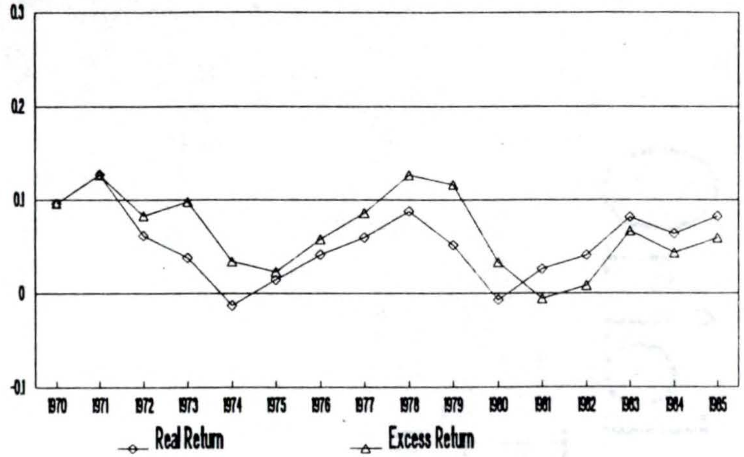
However, estimated annual real returns and after-tax excess returns, shown in Figure 6-01, yield some common characteristics. Real and after-tax excess returns appear highly correlated with the period of rising inflation in the 1970s and the subsequent decline in inflation in the early 1980s. In addition, after-tax excess returns, in each market, are less than real returns for the periods after mid-1980.

To insure the accuracy of the initially constructed return series and to provide baseline test statistics, tests of first-order autocorrelation are conducted that replicate those reported by Case and Shiller (1989). Commonly, tests of first-order autocorrelation involve regressing a time series of returns on the same time series lagged one period. However, Case and Shiller have shown that when using the housing return model above, spurious serial correlation may result from such tests. Alternatively, they indicate that

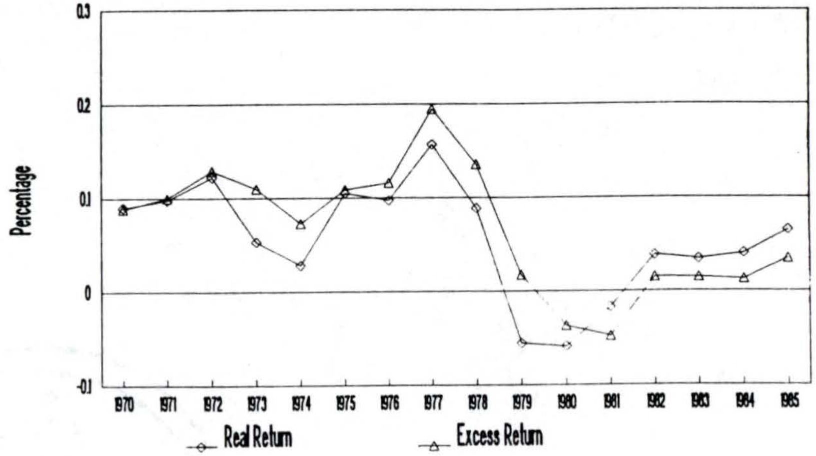
TABLE 6-02 Summary Statistics of Capital Market Real Quarterly Returns (1970:1 - 1986:1)

Statistic	Common Stocks	Small Stocks	L.Term Corp. Bonds	L.Term Gov. Bonds	I.Term Gov. Bonds	Real T-Bill Rates	Change in CPI-U	Real Returns (ATL)	Real Returns (CHI)	Real Returns (DAL)	Real Returns (OAK)
Mean	.013	.031	.009	.008	.007	.003	.016	.013	.013	.018	.023
Std. Dev.	.091	.143	.071	.072	.043	.009	.010	.028	.023	.027	.026
Skewness	-.389	.204	.424	.629	.534	.365	.392	.502	-.859	.038	-.084
Kurtosis	3.618	3.344	3.934	4.147	3.822	2.626	2.734	4.864	3.416	2.914	5.044
Cross-Correlations											
Common Stocks	1.000										
Small Stocks	.800	1.000									
LT Cor. Bonds	.607	.380	1.000								
LT Gov. Bonds	.545	.331	.963	1.000							
IT Gov. Bonds	.487	.300	.955	.954	1.000						
Real T-Bill	.305	.157	.441	.470	.529	1.000					
CPI-U Infl.	-.413	-.238	-.517	-.511	-.536	-.731	1.000				
Real Housing Returns (ATL)	.212	.238	.216	.145	.142	.087	-.159	1.000			
Real Housing Returns (CHI)	.143	.017	.321	.283	.235	-.026	-.338	.066	1.000		
Real Housing Returns (DAL)	-.084	-.027	-.065	-.041	-.023	.010	-.172	.013	.142	1.000	
Real Housing Returns (OAK)	.292	.325	.267	.248	.190	-.216	-.148	.153	.174	.029	1.000

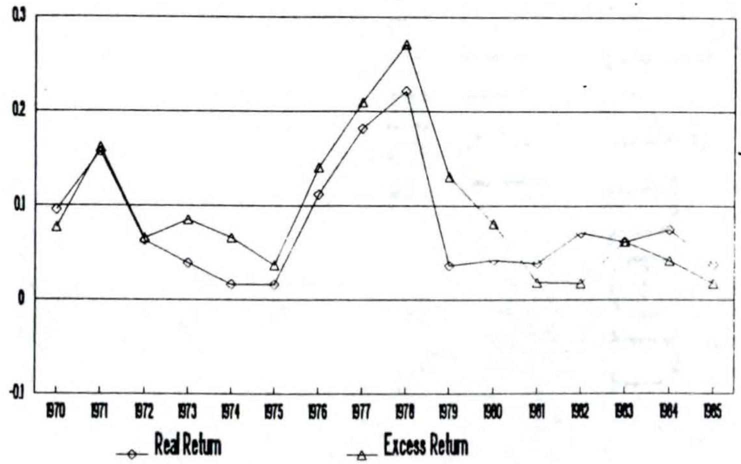
Stock, bond, and series as reported in *SBBI 1989 Yearbook*, Ibbotson Associates.



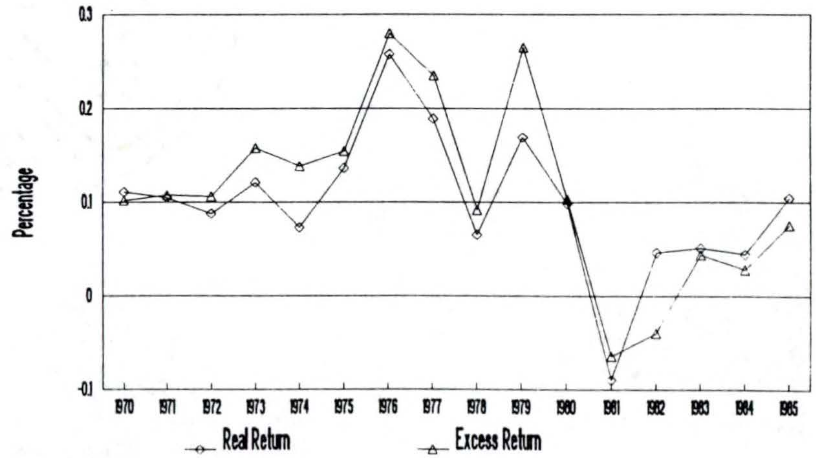
Atlanta



Chicago



Dallas



San Francisco

FIGURE 6-01 Real and Excess Annual Returns to Housing such correlation is avoided by constructing two independent indices for each city and regressing the changes in one index on the lagged changes of the second index. Consequently, they randomly divide the underlying transaction data for each city into two half-samples and estimate WRS price indices for each sample. Each of the resultant indices (two half-sample indices for each city) were obtained from Karl Case and are used to estimate returns to housing (real and after-tax excess returns) using Equations 6-01 and 6-03.

Annual real returns and annual after-tax excess returns are tested for first-order autocorrelation by adjusting Equation (5-01) to consider returns generated from each of the half-sample indices. Modifying (5-01) yields,

$$R_{r,j\_(t)} = \beta_0 + \beta_1 R_{r,k\_(t-L)} + u_{(t)}, \qquad (6\text{-}04)$$

and

$$R_{e,j\_(t)} = \beta_0 + \beta_1 R_{e,k\_(t-L)} + u_{(t)}, \qquad (6\text{-}05)$$

where j and k indicate that the respective return is constructed from each half-sample index and L is the number of periods lagged. To increase the sample size, annual returns are calculated on a quarterly basis and L set equal to four quarters. This eliminates the possible influence of seasonality in this test.[6] Equations (6-04) and (6-05) are

[6]Case and Shiller (1987) report that while second quarter price changes tend to be high and third quarter changes low, the difference is small, and only in Chicago is seasonality significant at the five percent level.

estimated using ordinary least squares (OLS) regression. To correct for the overlapping characteristics of the resulting data series, the estimated coefficient standard errors are adjusted using the methodology suggested by Hansen and Hodrick (1980).

The results of these regressions are reported in Tables 6-03 and 6-04. For each city, the contemporaneous regression of one return series (constructed from half-sample j) on the other return series (constructed from half-sample k) is first reported. If the indices were measured without error, the estimated β_1 coefficient and the coefficient of determination, R^2, would both be equal to 1.0. However, the errors-in-variables characteristics should result in the estimated coefficients being somewhat less than 1.0.

In this set of regressions the estimates of β_1, for each of the cities, are found to be significantly positive. In addition, the estimated β_1 coefficient for the cities of Atlanta, Chicago, and Dallas is close to the anticipated level of 1.0. However, both the estimated β_1 coefficient and the R^2 statistic reported for San Francisco are relatively small, suggesting that the series may contain substantial measurement error.[7] This result is slightly different from that reported by Case and Shiller. It is suspected that the difference is due to differences in the

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[7]This is later evaluated using quarterly return data.

TABLE 6-03

Regression of Annual Real Returns
on
Annual Real Returns Lagged (Full Sample Period)

$$R_{r\_j(t)} = \beta_0 + \beta_1 R_{r\_k(t-L)} + u_{(t)}$$

City Parameters	Period	Obs. Root MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 $adjR^2$
Atlanta	1970:1-1985:2				
j=A, k=B, L=0		62	0.012	0.768	.658
		0.028	(1.180)	(6.175)	.652
j=A, k=B, L=4		58	0.039	0.185	.045
		0.045	(2.716)	(1.090)	.028
j=B, k=A, L=4		58	0.038	0.211	.045
		0.049	(2.501)	(1.089)	.028
Chicago	1970:1-1985:2				
j=A, k=B, L=0		62	0.003	0.948	.832
		0.027	(0.291)	(8.291)	.829
j=A, k=B, L=4		58	0.024	0.540	.274
		0.058	(1.184)	(2.334)	.261
j=B, k=A, L=4		58	0.028	0.452	.222
		0.057	(1.428)	(2.045)	.208
Dallas	1970:1-1985:2				
j=A, k=B, L=0		62	0.028	0.724	.439
		0.048	(1.723)	(4.477)	.429
j=A, k=B, L=4		58	0.055	0.328	.099
		0.059	(2.493)	(1.558)	.083
j=B, k=A, L=4		58	0.044	0.296	.117
		0.053	(2.334)	(1.759)	.101
San Francisco	1970:1-1985:3				
j=A, k=B, L=0		63	0.049	0.488	.293
		0.069	(1.878)	(2.803)	.282
j=A, k=B, L=4		59	0.050	0.451	.253
		0.073	(2.090)	(2.800)	.240
j=B, k=A, L=4		59	0.070	0.300	.074
		0.090	(2.234)	(1.320)	.058

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE 6-04

Replication of Case and Shiller (1989) Statistics
Regression of After-Tax Annual Excess Returns on
After-Tax Annual Excess Returns Lagged

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters	Period	Obs. Root MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A, k=B, L=0		62	0.011	0.825	.688
		0.029	(0.960)	(6.229)	.683
j=A, k=B, L=4		58	0.039	0.350	.136
		0.048	(2.163)	(1.780)	.120
j=B, k=A, L=4		58	0.040	0.328	.113
		0.049	(2.199)	(1.610)	.097
Chicago	1970:1-1985:2				
j=A, k=B, L=0		62	0.004	0.933	.847
		0.027	(0.427)	(8.593)	.844
j=A, k=B, L=4		58	0.017	0.707	.481
		0.051	(0.879)	(3.531)	.472
j=B, k=A, L=4		58	0.020	0.664	.452
		0.052	(1.040)	(3.391)	.442
Dallas	1970:1-1985:2				
j=A, k=B, L=0		62	0.026	0.794	.589
		0.048	(1.589)	(5.714)	.582
j=A, k=B, L=4		58	0.045	0.551	.285
		0.063	(1.778)	(2.654)	.272
j=B, k=A, L=4		58	0.032	0.529	.291
		0.061	(1.381)	(2.965)	.279
San Francisco	1970:1-1985:3				
j=A, k=B, L=0		63	0.042	0.604	.459
		0.071	(1.585)	(3.758)	.450
j=A, k=B, L=4		59	0.047	0.552	.381
		0.100	(1.722)	(3.512)	.370
j=B, k=A, L=4		59	0.056	0.509	.205
		0.079	(1.551)	(2.200)	.191

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

calibrating of the model and the inclusion of the entire return series.

Statistics are reported for the regression of annual real returns to housing constructed with sample A on the one-year lagged real returns constructed from sample B in Table 6-03. The estimated β_1 coefficients from these regressions are consistently positive, suggesting that there is evidence of inertia in real returns. This is consistent with the findings of Case and Shiller with respect to changes in real house prices. Still, β_1 coefficients significant at the five percent level are found only in the Chicago market, the largest of the four markets. While some inertia does exist in each of the real return series, from these results it appears unlikely that excessive profits are consistently predictable or available.

Comparing these results with those reported by Case and Shiller indicate that including the implicit rental stream has little impact on the random-walk properties of the real return series in all the markets studied, with the exception Chicago. In Chicago, the first-order autocorrelation properties of the real return series are stronger than that of the changes in real housing price series. The reason for this condition is not clear. Table 6-03 indicates that, in the four markets studied, real returns become more predictable as the size of the market increases. It is speculated that in larger markets, such as Chicago, competing real estate assets are more difficult to identify. City-wide information, while available, is cumbersome and difficult to assimilate; hence, transaction costs increase, resulting in decreased efficiency.

In contrast, Table 6-03 shows that after-tax excess returns to housing are found to exhibit substantial serial dependency in all of the markets studied. The hypothesis that $\beta_1 = 0$, is rejected at a 95 percent confidence level for the markets of Chicago, Dallas, and San Francisco. Although the hypothesis can not be rejected for Atlanta substantial return inertia exists. These findings replicate those reported by Case and Shiller (1989), confirming the underlying constructed after-tax excess return series.

In addition to tests using annual returns to housing, quarterly real returns and quarterly after-tax excess returns in each city are examined for first-order autocorrelation. The results of these tests are reported in Tables 6-05 and 6-06.

Regression of real quarterly returns on itself indicates that the two half-sample price indices are best matched in the Atlanta and Chicago markets. The data are most homogenous in Atlanta. The Dallas indices are less similar and remarkably the two Oakland indices are statistically independent. This suggests that quarterly returns in that market are measured with significant error or that the returns are highly unpredictable.

TABLE 6-05

Regression of Quarterly Real Returns
on
Quarterly Real Returns Lagged (Full Sample Period)

$$R_{r\_j(t)} = \beta_0 + \beta_1 R_{r\_k,(t-L)} + u_{(t)}$$

City Parameters	Period	Obs. Root MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1986:1				
j=A, k=B, L=0		65	0.006	0.522	.409
		0.022	(1.998)	(6.601)	.400
j=A, k=B, L=1		64	0.016	-0.233	.081
		0.027	(4.292)	(-2.338)	.066
j=B, k=A, L=1		64	0.017	-0.314	.066
		0.034	(3.752)	(-2.088)	.051
Chicago	1970:1-1986:1				
j=A, k=B, L=0		65	0.006	0.662	.307
		0.023	(1.672)	(5.288)	.296
j=A, k=B, L=1		64	0.007	0.503	.181
		0.025	(1.990)	(3.701)	.168
j=B, k=A, L=1		64	0.012	0.030	.001
		0.023	(3.715)	(0.273)	-.015
Dallas	1970:1-1986:1				
j=A, k=B, L=0		65	0.014	0.220	.047
		0.034	(2.925)	(1.770)	.032
j=A, k=B, L=1		64	0.016	0.161	.027
		0.034	(3.354)	(1.309)	.011
j=B, k=A, L=1		64	0.017	-0.032	.002
		0.029	(4.053)	(-0.305)	-.015
San Francisco	1970:1-1986:2				
j=A, k=B, L=0		66	0.026	-0.162	.039
		0.038	(4.995)	(-1.612)	.024
j=A, k=B, L=1		65	0.019	0.158	.038
		0.038	(3.623)	(1.566)	.022
j=B, k=A, L=1		65	0.005	0.767	.386
		0.038	(1.033)	(6.287)	.376

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters.

TABLE 6-06
Regression of After-Tax Quarterly Excess Returns
on
After-Tax Quarterly Excess Returns Lagged

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters	Period	Obs. Root MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1986:1				
j=A, k=B, L=0		65	0.007	0.517	.370
		0.022	(2.323)	(6.085)	.360
j=A, k=B, L=1		64	0.017	-0.136	.025
		0.027	(4.522)	(-1.276)	.010
j=B, k=A, L=1		64	0.019	-0.219	.035
		0.032	(4.176)	(-1.494)	.019
Chicago	1970:1-1986:1				
j=A, k=B, L=0		65	0.008	0.526	.261
		0.022	(2.625)	(4.721)	.250
j=A, k=B, L=1		64	0.007	0.555	.297
		0.021	(2.376)	(5.121)	.286
j=B, k=A, L=1		64	0.011	0.250	.064
		0.023	(3.073)	(2.053)	.049
Dallas	1970:1-1986:1				
j=A, k=B, L=0		65	0.015	0.252	.070
		0.034	(3.198)	(2.182)	.055
j=A, k=B, L=1		64	0.017	0.224	.057
		0.033	(3.508)	(1.944)	.042
j=B, k=A, L=1		64	0.017	0.088	.009
		0.032	(3.603)	(0.762)	.000
San Francisco	1970:1-1986:2				
j=A, k=B, L=0		66	0.026	-0.033	.002
		0.041	(4.656)	(-0.341)	.000
j=A, k=B, L=1		65	0.022	0.165	.045
		0.040	(3.877)	(1.722)	.030
j=B, k=A, L=1		65	0.004	0.844	.424
		0.040	(0.734)	(6.812)	.415

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters.

Regressing real quarterly returns using the full sample period leads to widely differing results in each of the cities. In Atlanta, real returns are negatively correlated to the previous quarter's return, while in Dallas we see reversed signs depending on the regression specification.

Looking at quarterly after-tax excess returns, correlation patterns are much more consistent across the cities. While β_1 is not significantly different from zero in Atlanta, Dallas and Oakland, Chicago demonstrates significant patterns of a first-order autoregressive process.

In Tables 6-07 and 6-08, autocorrelation functions of the real and after-tax excess quarterly return series are estimated at lags up to eight quarters and a chi-square test for no time pattern in returns is reported. These tables indicate that when the full sample is used, real quarterly returns for Chicago and quarterly after-tax excess returns for both Chicago and San Francisco exhibit some pattern in their return structure. However, real quarterly returns in Atlanta, Dallas, and San Francisco and quarterly after-tax excess returns in Atlanta and Dallas are found consistent with random "white-noise".

Base Model Assumptions

The return methodology described previously depends largely on four primary assumptions. It is assumed that (1)

TABLE 6-07 Autocorrelation Function of Real Quarterly Returns

Series	p_1	p_2	p_3	p_4	p_5	p_6	p_7	p_8	CHI-SQ (12)	PROB.
PANEL A (1970:1-1986:1)										
Real Returns										
Atlanta	-0.394	0.207	0.088	-0.045	0.086	0.041	-0.026	-0.114	17.01	0.149
Chicago	0.228	0.201	0.460	0.177	0.194	0.128	-0.100	0.153	33.70	0.001
Dallas	-0.113	-0.047	0.165	0.095	0.067	0.188	-0.157	-0.112	11.92	0.452
San Fran.	0.216	0.171	0.198	0.139	0.153	0.018	0.040	0.038	12.09	0.438
PANEL B (1970:1-1979:3)										
Real Returns										
Atlanta	-0.484	0.302	-0.003	0.045	0.142	-0.023	0.024	-0.096	16.50	0.169
Chicago	0.208	0.048	0.153	0.143	-0.074	-0.300	-0.248	0.072	17.32	0.138
Dallas	-0.140	-0.069	0.151	0.059	0.031	0.212	-0.163	-0.165	8.40	0.753
San Fran.	0.046	-0.007	0.190	0.000	0.119	-0.071	-0.149	-0.053	6.17	0.907
PANEL C (1979:4-1986:1)										
Real Returns										
Atlanta	0.007	0.010	0.547	-0.114	0.083	0.286	-0.020	0.059	14.76	0.255
Chicago	0.003	-0.019	0.462	-0.067	0.038	0.270	-0.131	0.002	13.87	0.309
Dallas	-0.112	0.003	-0.117	0.195	-0.190	-0.137	-0.061	0.109	7.54	0.820
San Fran.	0.204	0.105	-0.073	0.080	-0.169	-0.213	-0.060	-0.055	6.50	0.889

TABLE 6-08 Autocorrelation Function of After-Tax Excess Quarterly Returns

Series	p_1	p_2	p_3	p_4	p_5	p_6	p_7	p_8	CHI-SQ (12)	PROB.
PANEL A (1970:1-1986:1)										
Exc. Returns										
Atlanta	-0.284	0.156	0.104	0.069	0.148	-0.073	-0.011	-0.050	12.95	0.373
Chicago	0.452	0.367	0.389	0.527	0.305	0.187	0.212	0.297	74.53	0.000
Dallas	0.048	0.048	0.267	0.200	0.160	0.244	0.185	-0.028	16.14	0.185
San Fran.	0.308	0.252	0.343	0.309	0.313	0.072	0.117	0.164	43.65	0.000
PANEL B (1970:1-1979:3)										
Exc. Returns										
Atlanta	-0.382	0.139	0.059	0.020	0.137	-0.077	0.015	-0.101	10.72	0.553
Chicago	0.111	0.032	0.015	0.323	-0.116	-0.320	-0.201	0.186	23.46	0.024
Dallas	-0.075	-0.124	0.204	0.067	0.078	0.203	-0.110	-0.088	8.59	0.738
San Fran.	0.030	0.023	0.120	0.039	0.122	-0.110	-0.105	0.054	3.56	0.990
PANEL C (1979:4-1986:1)										
Exc. Returns										
Atlanta	-0.064	-0.145	0.134	0.128	0.126	-0.273	-0.178	0.192	10.50	0.572
Chicago	0.150	-0.127	0.064	0.235	-0.020	-0.066	0.112	0.040	6.86	0.866
Dallas	0.079	0.186	-0.025	0.126	-0.228	-0.102	-0.263	0.020	10.45	0.576
San Fran.	0.208	-0.045	0.084	0.300	0.053	-0.323	-0.225	-0.005	15.91	0.196

the WRS price index accurately estimates movements in housing prices. The return model assumes, (2) that marginal income tax rates and property tax rates are constant, (3) inflation expectations were consistent with rational expectations during this period, and finally, (4) that the risk-premium associated with housing returns is not serially correlated.

While the WRS index methodology attempts to control for constant-quality housing, it has its shortcomings. When used over short time spans, the index may restrict observations to a disproportionate number of high turnover properties, potentially reflecting forced sales or renovation projects.[8] This biases the index and if the proportion of high turnover properties is autocorrelated, tests of autocorrelation in returns may be biased.

Despite these limitations, other price indices present other weaknesses. One-directional tests of efficiency can be conducted using the WRS index. If, for instance, autocorrelation is present in after-tax excess returns, it is not clear whether it is due to the model of housing returns or the actual nature of excess returns. However, if autocorrelation is rejected, the hypothesis of an efficient market cannot be rejected.

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[8]Case and Shiller indicate that transactions were selected such that properties which changed condition or quality between the time of the first sale and the second (repeat sale) were excluded from the data set.

The assumptions that marginal tax rates are constant, inflation accurately forecasted, and housing risk-premium constant are explicitly modeled in the following chapters and the robustness of the initial test findings, as reported in this chapter, examined.

CHAPTER 7
MARGINAL TAX RATES

Implied Bondholder Marginal Tax Rates

In this chapter the base model assumption that marginal tax rates (MTRs) are constant is removed. It is hypothesized that home buyer MTRs are serially dependent and negatively correlated with the after-tax excess return series previously estimated in Chapter 6. This suggests that any autocorrelation found in these after-tax excess returns may result from assuming a constant marginal tax rate.

Using yields on newly issued three-month prime-grade municipal securities and three-month U.S. Treasury securities, Figure 7-01 plots monthly estimates of the marginal bondholder's tax rate for the first quarter of 1970 through the second quarter of 1986.[1] This indicates that imputed bondholder marginal tax rates during this period

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[1]The marginal bondholder's tax rate, τ_b, is estimated as τ_b=1-(M/TB), where M denotes the prime-grade municipal yield, and TB the treasury yield. At this rate, after-tax treasury yields equal tax-exempt yields. It is recognized that exist complications with assuming this simple relationship. For instance, the risk premia on municipal bonds is assumed constant, while it is quite evident that during recessions (1974-75 and 1981-82) the risk premia on municipal bonds increases reducing the spread.

were not constant, but were in fact quite volatile. In addition, Panel A of Table 7-01 suggests that bondholder marginal tax rates exhibit significant first-order autoregressive, AR1, characteristics. Hence, they are not random, and would influence the patterns of any generated after-tax return series which neglected to take them into account.[2]

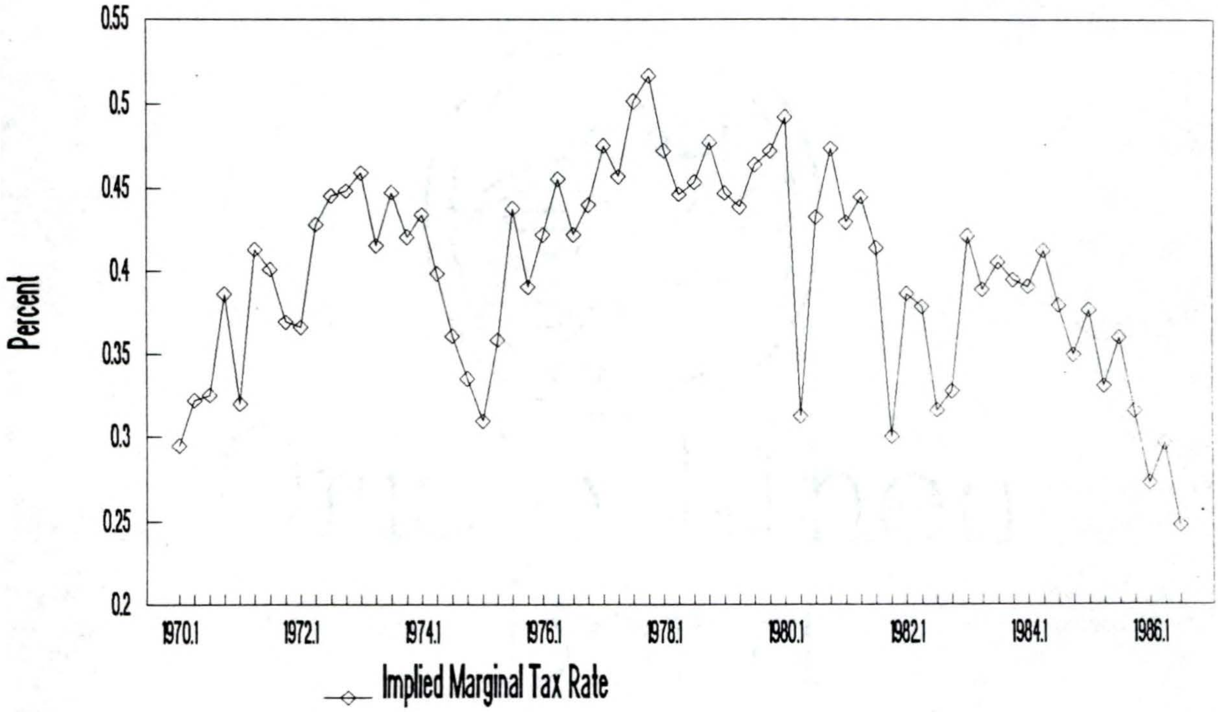

FIGURE 7-01 Implied Bondholder Marginal Tax Rates (1970:1 to 1986:2)

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[2]Marginal tax rates estimated from national tax data as the percentage tax on each additional dollar of income are published for selected years in the *U.S. Statistical Abstract*. These data suggest marginal tax rate trends similar to that indicated by the imputed bondholder's marginal tax rate. Marginal income tax rates generally increased during the 1970s and decreased during the early 1980s.

Unfortunately, implied bondholder marginal tax rates are not necessarily representative of the typical home buyer's marginal tax rate. They represent the marginal tax rate of the last (or marginal) bond investor for whom it is advantageous to invest in municipal securities rather than treasury bills.

TABLE 7-01

First-Order Autocorrelation of Marginal Tax Rates

$$MTR_{(t)} = \beta_0 + \beta_1 \, MTR_{(t-L)} + u_{(t)}$$

PANEL A

MTR Series	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
	1970:1-1986:2				
USA_MTR_b		65	0.145	0.641	.411
L=1		0.044	(3.685)	(6.635)	.402

PANEL B

MTR Series	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR2
	1970:1-1985:2				
ATL_MTR_h		61	0.106	0.537	.292
L=4		0.015	(2.338)	(2.716)	.280
CHI_MTR_h		61	0.076	0.661	.486
L=4		0.018	(2.130)	(4.105)	.477
DAL_MTR_h		61	0.064	0.738	.557
L=4		0.018	(1.806)	(4.736)	.549
SAN_MTR_h		61	0.054	0.812	.757
L=4		0.023	(1.907)	(7.454)	.753

Note: MTR_h and MTR_b denote the home owner marginal tax rates for each city and the implied U.S. bondholder marginal tax rate, respectively. L denotes number of lagged quarters. The coefficient standard errors in PANEL B are adjusted as indicated by Hansen and Hodrick (1980).

Home Buyer Marginal Tax Rates

The marginal tax rate of the home buyer is dependent not only on statutory changes in the tax structure but in changes in the 'out-of-pocket' cost of housing. For instance, if the current real price of housing or real mortgage rates increase, some home buyers are priced out of the market, unable to afford the 'out-of-pocket' housing costs. When this occurs, a tax bracket 'shift' takes place. The average buyer of the median-priced home has a different marginal tax rate than the average buyer of the median-priced home prior to price or mortgage increases. Average new buyers must have higher incomes in order to afford the increased out-of-pocket costs. Consequently, the MTRs of home buyers are not captured by simply examining statutory changes in the tax structure.

While it is impractical to model the entire distribution of MTRs for all home buyers, it is possible to estimate the average income of the buyer of the median-priced home in each city. Using marginal tax rate data, which are available by income bracket, the MTR of this home buyer can be estimated for each quarter.

Using the WRS housing price index, the median price of constant-quality housing is first estimated. Each city's respective WRS index is set to the median price of existing homes sold, as reported for that city in 1980 by the Office of Thrift Supervision (OTS, formerly the Federal Home Loan

Bank Board, FHLBB). This yields a series prices for constant-quality 1980 median priced housing.

Survey data reporting the national average loan-to-value ratios are available from both the OTS and the Chicago Title Insurance Company. Using the average contract mortgage rate in each quarter, a series of average mortgage payments is constructed for each city. The average income

TABLE 7-02
Sources of Data to Determine Home Buyer Marginal Tax Rates

Series	Dates Used	Freq	Source
Yield on newly-issued three-month prime-grade municipals	70:01-86:09	M	Salomon Brothers, *An Analytical Record of Yields and Spreads.*
Contract rate on 30-year fixed-rate conventional mortgages.	70:01-86:09	M	*Federal Reserve Bulletin.*
Median priced single-family housing unit (1980).	1980	A	FHLBB, *Savings and Home Financing Source Book.*
Marginal Federal Income Tax (Married).	70:1-86:3	Q	Ling, *The Price of Owner-Occupied Housing Services.*
Mortgage payment-to-Income Ratio; Loan-to-Value Ratio.	75:1-86:1	A	Chicago Title Insurance Co., *Chicago Guaranty.*
Loan-to-Value Ratio.	70:1-74:1	A	FHLBB, *Savings and Home Financing Source Book.*

Note: M, Q, and A denote monthly, quarterly, and annual data series, respectively.

of a representative home buyer of this representative housing is then determined from data indicating the average monthly payment as a percentage of income.[3] Finally, a series of median home buyer MTRs is estimated for each city using MTR data from Ling (1990).

EMH Test Results: Excess Returns Adjusted for Time-Varying Marginal Tax Rates

A series of home buyer marginal tax rates is estimated for each city using the methodology described in the previous section. These series are plotted in Figures 7-02 and 7-03, along with the previously estimated quarterly after-tax excess returns. These figures indicate that home buyer marginal tax rates were generally higher in the 1980s when after-tax excess returns, as previously estimated, declined. However, the inverted U shape, similar to that shown in Figure 7-01 remains prevalent. Panel B of Table 7-01 indicates that these series follow a first-order autoregressive (AR1) pattern, and quarterly after-tax excess returns to housing are shown to be negatively correlated with the constructed home buyer marginal tax rate series for each city in Panel C of Table 7-03. However, the correlation factors are relatively low.

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[3]Mortgage payment-to-income ratios were unavailable for the period 1970:1 to 1974:4. Interest rates during this period were relatively stable; hence an assumption that payment-to-income ratios were constant during this period was used.

Regressing after-tax excess returns constructed using the constructed series of time-varying marginal tax rates shows little difference from previous tests of autocorrelation reported in Chapter 6.

Tables 7-04 and 7-05 indicate that the earlier assumption, that marginal tax rates are constant, does not materially affect the results of autocorrelation tests. While this methodology lacks precision, it is preferable to assuming a strictly constant home buyer marginal tax rate and is applied in further tests of efficiency conducted in Chapters 8 and 9. In Table 7-04, after-tax annual excess returns are estimated as the product of four quarterly after-tax excess returns. This methodology is identical to simply using annual returns.[4]

In addition, Ling (1990) reports the tenure choice tax rate, using these data, results are reported in Tables B-02 and B-03 of Appendix B. While there is evidence that applying the tenure choice tax rate to the model reduces autocorrelation tendencies in the return series, Ling (1990) indicates that for the quantity of housing demanded decision (how much housing to invest in) the marginal tax rate is the applicable rate. Therefore, it is the rate applied to the models subsequently tested in Chapters 8 and 9.

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[4]Regression results applying this methodology to the base model appear in Appendix B, Table B-01. The results reported in Table B-01 are identical to those reported earlier in Table 6-04, indicating the test methodologies are equivalent.

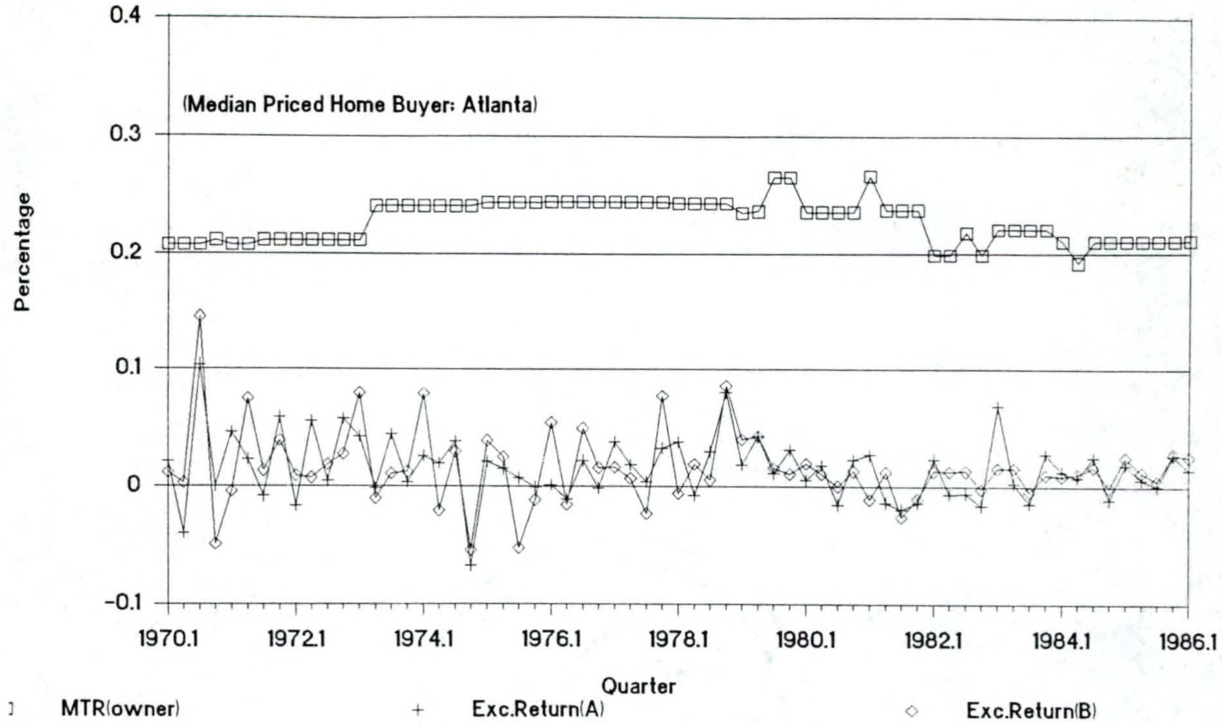

FIGURE 7-02 Homeowner Marginal Tax Rates and Quarterly Excess Returns: Atlanta

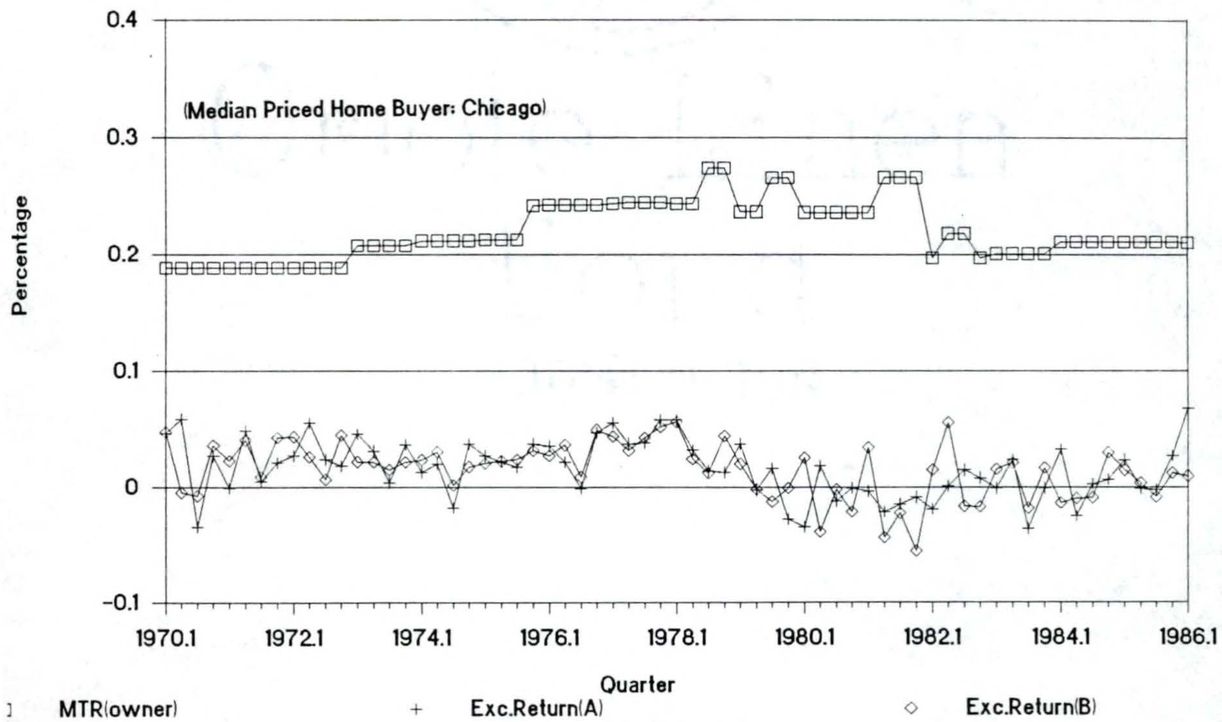

FIGURE 7-03 Homeowner Marginal Tax Rates and Quarterly Excess Returns: Chicago

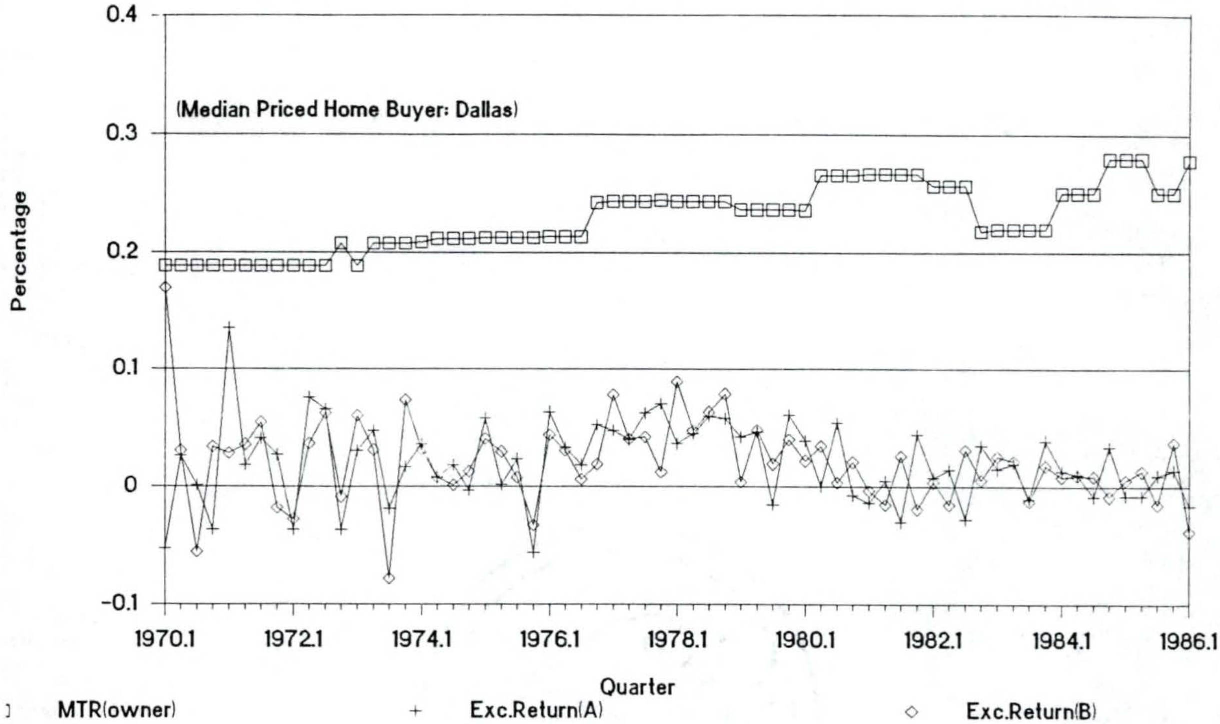

FIGURE 7-04 Homeowner Marginal Tax Rates and Quarterly Excess Returns: Dallas

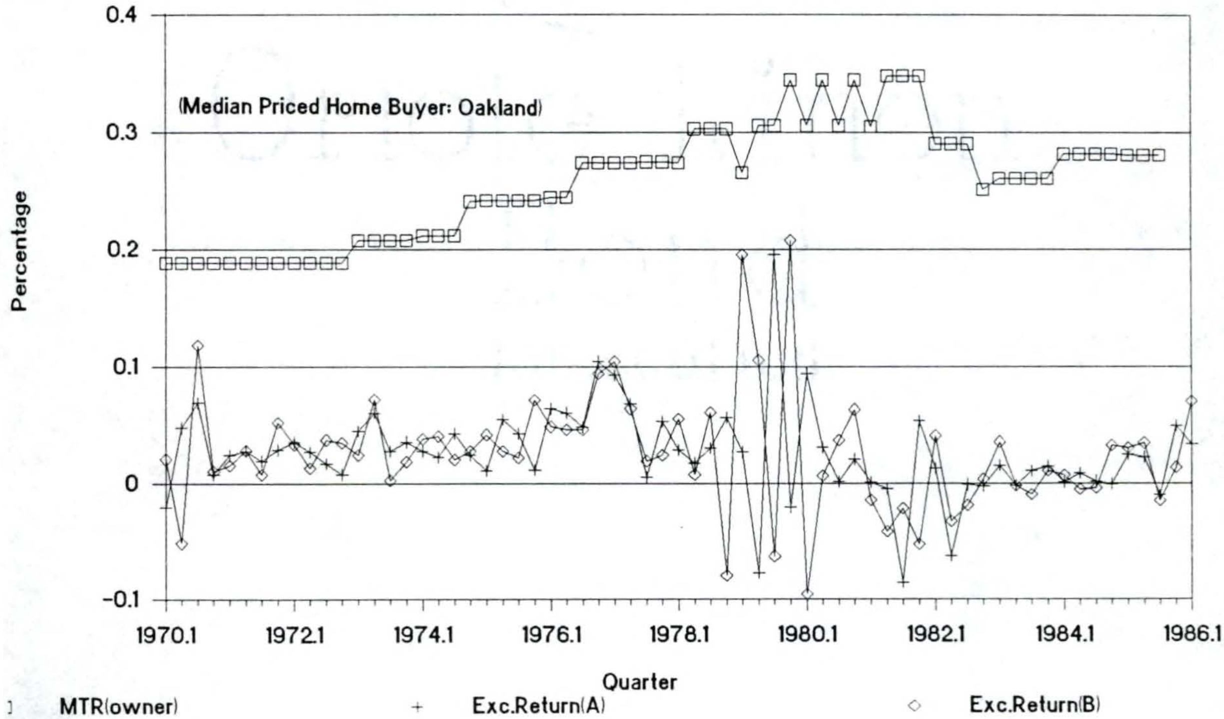

FIGURE 7-05 Homeowner Marginal Tax Rates and Quarterly Excess Returns: San Francisco

TABLE 7-03
Correlation Matrices of Quarterly Home Owner Excess Returns, Constructed Home Owner Marginal Tax Rates and Implied Bondholder Marginal Tax Rates

PANEL A	ATL_R_e	CHI_R_e	DAL_R_e	SAN_R_e
ATL_R_e	1.0000			
CHI_R_e	0.0675	1.0000		
DAL_R_e	0.1480	0.2939	1.0000	
SAN_R_e	0.3313	0.3225	0.1868	1.000

PANEL B	ATL_MTR_h	CHI_MTR_h	DAL_MTR_h	SAN_MTR_h
ATL_MTR_h	1.0000			
CHI_MTR_h	0.7277	1.0000		
DAL_MTR_h	0.1388	0.5612	1.0000	
SAN_MTR_h	0.3573	0.7691	0.8700	1.0000
USA_MTR_b	0.5033	0.4702	0.0233	0.1796

PANEL C	ATL_R_e	CHI_R_e	DAL_R_e	SAN_R_e
ATL_MTR_h	-0.0728			
CHI_MTR_h		-0.1538		
DAL_MTR_h			-0.1767	
SAN_MTR_h				-0.1914
USA_MTR_b	0.1464	0.1824	0.2866	0.2022

Note: Excess returns, denoted R_e, were computed using a constant 30 percent marginal tax rate for each city. Marginal home owner tax rates are denoted MTR_h for each city and the implied bondholder marginal tax rate is USA_MTR_b.

TABLE 7-04
Regression of After-Tax 4-Quarter Excess Returns
on After-Tax 4-Quarter Excess Returns Lagged
Adjusted for Nonconstant Home Owner Marginal Tax Rates

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 R_{e,m\_k(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A, k=B, L=0		62	0.010	0.830	.705
		0.047	(0.906)	(6.401)	.700
j=A, k=B, L=4		58	0.033	0.383	.167
		0.049	(1.895)	(1.866)	.152
j=B, k=A, L=4		58	0.034	0.348	.126
		0.047	(1.929)	(1.676)	.111
Chicago	1970:1-1985:2				
j=A, k=B, L=0		62	0.004	0.938	.867
		0.026	(0.406)	(9.843)	.865
j=A, k=B, L=4		58	0.015	0.706	.487
		0.053	(0.776)	(3.503)	.477
j=B, k=A, L=4		58	0.017	0.671	.457
		0.054	(0.885)	(3.344)	.447
Dallas	1970:1-1985:2				
j=A, k=B, L=0		62	0.023	0.814	.604
		0.048	(1.510)	(5.908)	.597
j=A, k=B, L=4		58	0.042	0.540	.272
		0.064	(1.735)	(2.540)	.259
j=B, k=A, L=4		58	0.029	0.522	.294
		0.061	(1.315)	(2.956)	.282
San Francisco	1970:1-1985:3				
j=A, k=B, L=0		63	0.040	0.614	.475
		0.069	(1.543)	(3.873)	.466
j=A, k=B, L=4		59	0.044	0.559	.395
		0.077	(1.642)	(3.541)	.384
j=B, k=A, L=4		59	0.053	0.515	.209
		0.098	(1.507)	(2.208)	.195

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE 7-05
Regression of After-Tax Quarterly Excess Returns
on After-Tax Quarterly Excess Returns Lagged
Adjusted for Nonconstant Home Owner Marginal Tax Rates

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 R_{e,m\_k(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1986:1				
j=A, k=B, L=0		65	0.006	0.521	.373
		0.022	(2.112)	(6.126)	.363
j=A, k=B, L=1		64	0.016	-0.129	.023
		0.028	(4.122)	(-1.207)	.007
j=B, k=A, L=1		64	0.018	-0.210	.032
		0.033	(3.848)	(-1.434)	.017
Chicago	1970:1-1986:1				
j=A, k=B, L=0		65	0.008	0.535	.269
		0.021	(2.423)	(4.820)	.258
j=A, k=B, L=1		64	0.006	0.564	.306
		0.021	(2.176)	(5.231)	.295
j=B, k=A, L=1		64	0.009	0.267	.073
		0.024	(2.757)	(2.207)	.058
Dallas	1970:1-1986:1				
j=A, k=B, L=0		65	0.014	0.253	.071
		0.034	(3.034)	(2.191)	.056
j=A, k=B, L=1		64	0.016	0.224	.058
		0.033	(3.353)	(1.947)	.042
j=B, k=A, L=1		64	0.016	0.091	.010
		0.032	(3.396)	(0.793)	-.006
San Francisco	1970:1-1986:2				
j=A, k=B, L=0		66	0.026	-0.037	.002
		0.041	(4.601)	(-0.380)	-.013
j=A, k=B, L=1		65	0.021	0.163	.044
		0.040	(3.836)	(1.695)	.028
j=B, k=A, L=1		65	0.004	0.842	.422
		0.040	(0.734)	(6.787)	.413

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters.

CHAPTER 8
INFLATION EXPECTATIONS

Expected and Unexpected Inflation

Tests of efficiency using returns generated from the models described in Chapters 6 and 7 assume that unexpected inflation (differences between actual and expected inflation) is stochastic with mean zero. Under these conditions, market participants are said to hold "rational" expectations with respect to inflation.

In defining after-tax excess returns as the nominal return to housing in excess of the after-tax risk-free interest rate, $R_e = R_n - (1-\tau)TB$, inflation expectations (embedded in the interest rate structure) are subtracted from the nominal return to housing.[1] If inflation expectations are rational during the sample period and returns to housing are consistently correlated with unexpected inflation, then returns due to unexpected inflation should provide be random and provide no new information in explaining *ex post* after-tax excess returns.

Under this condition, the assumption,

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[1]Consistent with the Fisher hypothesis, nominal risk-free interest rates are set equivalent to the required real risk-free interest rate plus the expected rate of inflation.

$$E(R_{ui,t}) = 0, \tag{8-01}$$

holds and Equation (4-04), the formal market efficiency model rewritten below for convenience, is unaffected,

$$E[(R_{n,t}-(1-\tau)RF_t-R_{ui,t})-E((R_{n,t}-(1-\tau)RF_t)|\Phi_{t-1},\Phi_{\pi,t-1})] = 0.^{2} \tag{8-02}$$

However, if unexpected inflation is systematically biased, any subsequent tests of the EMH will be affected.

While the assumption that Equation (8-01) holds is consistent with the rational expectations hypothesis and tests of market efficiency, the position that market expectations of inflation are rational is replete with controversy. When considering the volatile inflationary period of the 1970s and early 1980s, substantial evidence exists suggesting that inflation expectations during this period are were not rational. This evidence suggests that both the rapid increases and decreases in the rates of inflation were systematically underestimated. This has led some to reject the entire notion of the rational expectations hypothesis.

However, this position appears extreme and myopic. It is quite reasonable to expect that rational markets may not meet full rational expectations conditions (in a statistical sense) when examined over relatively short periods of time.

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[2]The variables are defined earlier in Equation (4-04).

In addition, single factors that highly influence a market, such as inflation expectations relative to housing returns or bond returns, may not appear rational for short time intervals. This, however, should not necessarily lead to a full rejection of the EMH for these markets. Studies need to include sufficient time periods to fully reject the efficient markets hypothesis.

It is possible, for example, to identify short periods, (days and weeks) where *ex post* price changes in the S&P 500 and *ex post* returns to bonds are serially dependent. However, these dependencies are not consistent when longer time spans are considered and should not lead to a rejection of the entire EMH hypothesis. Unfortunately, adequate housing market data for long periods do not currently exist. However, it is important to look at the sensitivity of the previous test results to the assumption that inflation expectations over this period were rational.

This chapter first presents evidence suggesting that inflation surprises over the period studied were autocorrelated. Using a model to estimate unexpected inflation, after-tax excess returns to housing are adjusted for any returns due to unexpected inflation. Finally, the random-walk characteristics of the abnormal returns to housing, residuals from this model, are examined and the EMH test results reported.

Past Inflation Expectation Models and Tests of Rationality

Since inflation expectations are not observable, alternative methods have been used to estimate both expected and unexpected inflation. Although there has been much work done evaluating these methodologies, there is no general consensus as to the 'best' or 'dominant' method.

In estimating expected inflation, one of three general methodologies is typically used. The first extracts inflation expectations from the market risk-free interest rate. For example, an interest rate model developed by Fama (1975) is used to predict the rate of inflation one-month ahead using the CPI. This model requires the assumption that real risk-free returns are constant. Later, Fama and Gibbons (1982, 1984) modify the interest rate model to account for nonconstant real interest rates. While they suggest this model dominates other methodologies, this finding is not consistent with work by others, such as Hafer and Hein (1990).

Univariate time-series forecasting models represent the second method. A number of alternative model specifications have been suggested. Adaptations of autoregressive distributed lag (DL) models or autoregressive integrated moving average (ARIMA) models are most commonly used. All these models use some form of past inflation information, available to market participants, to estimate future inflation. The shortcoming of the DL model is that it will lead to systematic forecast errors during periods of rising or falling inflation. However, the ARIMA model yields residuals which are consistent with rational expectations, making that class of models highly attractive and widely used.

Consensus survey data of market participants or experts, represents the third method. In this case survey data on expected inflation is used as a proxy for market inflation expectations. While this method is also appealing, widely used, and considered by many as the most viable method, the data are often found to reject the rational expectations hypothesis. This has led researchers to question both the survey procedures and whether survey respondents typify the sentiment of the aggregate market.

A number of studies have examined observable inflation expectations using the Livingston survey data.[3] Work by Pesando (1975) found that the survey data reject the rational expectations hypothesis.[4] Carlson (1977),

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[3]These data come from a survey of professional economists conducted by Joseph Livingston. Taken semi-annually (June and December) since 1947, the survey asks economists to estimate the consumer price index (CPI) that will exist six and twelve months from the survey date, as well as forecasts of other key economic variables. The forecasted levels are then typically converted to six- and twelve-month inflation forecasts by using the actual CPI level at the time the survey was taken. In addition, these data have been widely used to empirically test the effect of anticipated inflation on interest rates.

[4]Pesando could not reject a weaker hypothesis of rationality using the six-month-ahead survey data and indicates that they are more consistent with the notion of

examining the same data, indicates that respondents are likely to know only the April CPI figures when surveyed in June and the October CPI when surveyed in December. He suggests that the data reflect eight- and fourteen-month forecasts, rather than six- and twelve-month. Using these modified assumptions, Carlson rejects the rational expectations hypothesis. Mullineax (1978), later suggests that, in fact, Carlson's revised data are consistent with rationality under alternative tests.

While the above tests suggest that the data from 1959 to 1969 may be consistent with rationality, none of the tests includes data from the period of the 1970s, characterized by rapid inflation. Using data from 1959 to 1975, Pearce (1979) finds that the Livingston survey data are not rational. In addition, Hafer and Hein (1985) examine a survey conducted jointly by the American Statistical Association and the National Bureau of Economic Research (ASA-NBER).[5] They find that using data from the first quarter of 1970 through the first quarter of 1984 the ASA-NBER data display systematic bias in their forecast errors. The Livingston data, shown in Figure 8-01, appear to follow very similar trends during this period.

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rationality than the twelve-month-ahead data.

[5]The ASA-NBER surveys economists quarterly, asking for expected levels of the GNP one- to six- quarters in the future. From this data an expected implicit GNP deflator is then derived which some studies have employed as a proxy for expected inflation.

These studies suggest that survey respondents were unable to forecast the historically high rates of inflation experienced during the 1970s and early 1980s. An exception to this result appears in a study by Pearce (1987). He examined a monthly survey of money market participants conducted by Money Market Services International and found that these data, including data from the 1970s and 1980s, are consistent with rational expectations. However, this survey typically takes place within two weeks of the actual CPI announcement date. This, in addition to the findings of Pesando (1977), suggests that the accuracy of survey respondents inflation forecasts deteriorate as the term of the forecast lengthens.

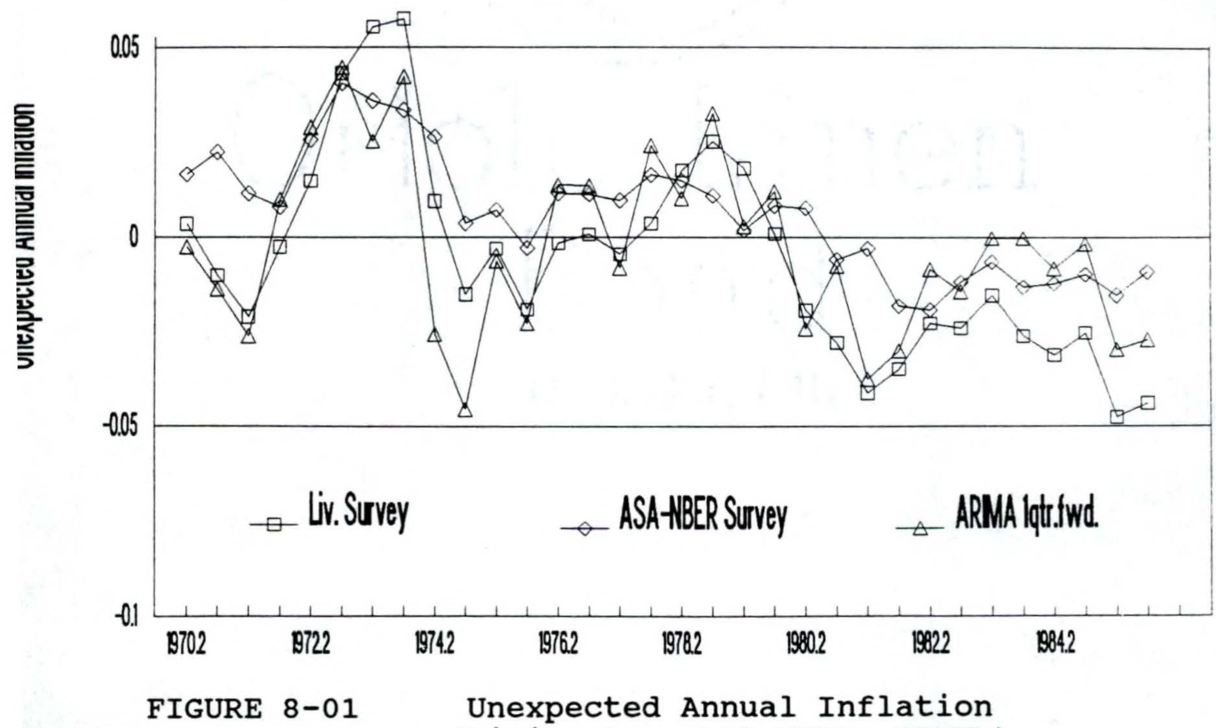

FIGURE 8-01 Unexpected Annual Inflation (Livingston, ASA-NBER, ARIMA)

In addition, changes in money and capital market expectations appear to lag movements of inflation rates during this period. Interest rates lag both increases and decreases in expected inflation rates.[6] This results in a autocorrelated series of real interest rates, typified by periods of prolonged negative real rates as in Figure 8-02.

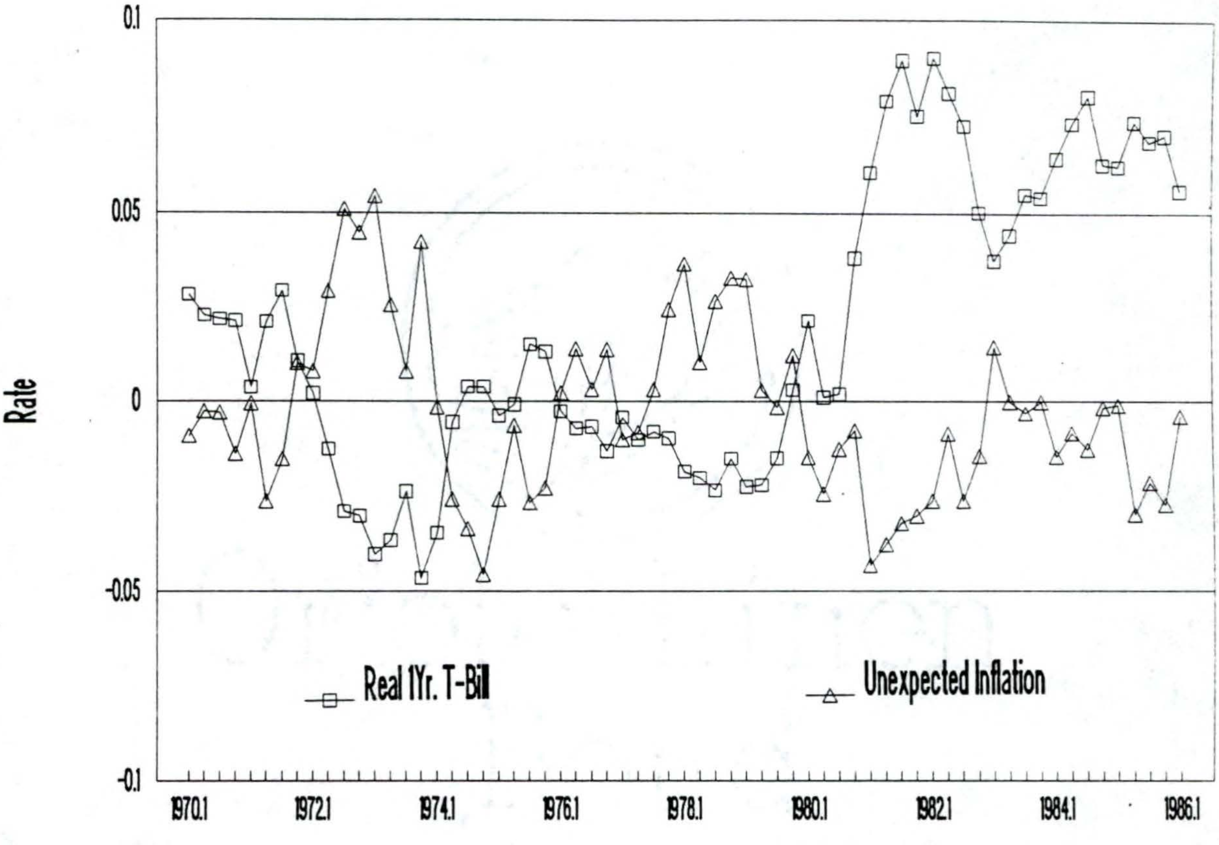

FIGURE 8-02 Real 1-Year Treasury Bill Yields and Unexpected Annual Inflation

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[6]Typically, real interest rates are low when inflation expectations are high; and, conversely, real interest rates are relatively high when expected inflation rates are declining. See Taylor (1982) and Dokko and Edelstein (1987). Mundell (1963) and Tobin (1965) hypothesize that this results from people shifting to interest-bearing assets during times of high interest and expected inflation rates, reducing the equilibrium real rate of return on such assets.

Finally, the study period is marked by changes in monetary policy, which have been characterized as a 'regime shift'. Blanchard (1984) indicates that the impact (or event) of changes in the monetary policy in October 1979 were not immediately believed. "Direct, informal evidence suggests that it took more than a year to fully change the beliefs of financial markets, perhaps more than that to change those of labor markets." (Blanchard, p212.) Regressing after-tax excess returns prior to and subsequent to the October 1979 policy shift indicates that the hypothesis of an efficient market can no longer be rejected using the base model data. Results of this regression appear in Table 8-01.

While there exists little consensus to the 'dominant' methodology for estimating expected inflation, both the survey data and the interest rate structure suggest that inflation shocks during the 1970s and early 1980s were not fully anticipated. Systematic forecast errors appear evident during this period. Throop (1988) finds that adaptive expectation models, not rational expectation, best explain both investment and labor market inflationary expectations during this period. Given this possibility, inflation expectations are modeled in the following sections, and after-tax returns to housing are adjusted for autocorrelated inflation shocks.

TABLE 8-01 Regression of Annual Excess Returns on Annual Excess Returns Lagged
(Prior and Subsequent to Monetary Policy Shift)

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters		Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2		Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	PANEL1970:1 -1979:3					PANEL 1979:4 -1985:2				
j=A, k=B, L=0	[A]	39	0.020	0.772	.658	[B]	23	0.007	0.763	.356
		0.032	(1.187)	(4.638)	.649		0.024	(0.426)	(1.872)	.325
j=A, k=B, L=4		35	0.058	0.296	.091		19	0.023	0.266	.042
		0.052	(2.124)	(1.190)	.063		0.029	(1.086)	(0.486)	-.012
j=B, k=A, L=4		35	0.057	0.302	.082		19	0.033	0.002	.000
		0.056	(2.017)	(1.112)	.055		0.025	(2.226)	(0.007)	-.056
Chicago										
j=A, k=B, L=0		39	-0.007	1.049	.800		23	0.003	0.610	.360
		0.024	(-0.446)	(6.697)	.795		0.029	(0.310)	(1.889)	.330
j=A, k=B, L=4		35	0.067	0.338	.051		19	0.007	0.245	.057
		0.022	(1.078)	(0.648)	.022		0.034	(0.504)	(0.576)	.005
j=B, k=A, L=4		35	0.097	0.096	.008		19	-0.001	0.001	.000
		0.047	(1.927)	(0.231)	-.022		0.037	(-0.045)	(0.001)	-.056
Dallas										
j=A, k=B, L=0		39	0.041	0.712	.472		23	0.018	0.763	.366
		0.055	(1.332)	(3.161)	.458		0.032	(0.969)	(1.913)	.336
j=A, k=B, L=4		35	0.084	0.397	.172		19	0.035	-0.000	.000
		0.067	(2.207)	(1.545)	.147		0.028	(1.962)	(-0.000)	-.056
j=B, k=A, L=4		35	0.070	0.390	.180		19	0.024	0.041	.005
		0.067	(2.102)	(1.796)	.155		0.024	(1.448)	(0.164)	-.050
Oakland										
j=A, k=B, L=0		39	0.117	0.273	.109		24	0.012	0.604	.438
		0.073	(2.626)	(1.173)	.085		0.045	(0.647)	(2.279)	.413
j=A, k=B, L=4		35	0.101	0.410	.149		20	0.014	0.255	.282
		0.073	(1.899)	(1.417)	.123		0.051	(0.702)	(1.615)	.249
j=B, k=A, L=4		35	0.156	0.127	.010		20	0.016	0.172	.072
		0.094	(2.371)	(0.335)	-.021		0.064	(0.591)	(0.591)	.030

A and B denote half sample indices for each city and L denotes number of lagged quarters.
Coefficient standard errors adjusted for MA3 process (Hansen and Hodrick, 1980)

Inflation Expectations Model

This study models inflation expectations using one-step-ahead forecasts from an ARIMA (0,2,1) model and employing a similar methodology to that used by Pearce (1979).[7] Following Schwert (1987) the autocorrelation function of the second differences of the natural log of quarterly CPI-U, net-of-shelter, index is estimated for the period 1960:1 to 1989:1. The first twelve autocorrelations of the second difference of the variable, which represent changes in the inflation rate, are close to zero with the exception of the first. This is typical of a first-order moving average process,

$$(1-L)X_t = (1-\theta_1 L)a_t, \qquad (8\text{-}01)$$

were X_t is the inflation rate; a_t, the independent errors; θ the moving average parameter; and L the lag operator.[8] The ARIMA model specified by Equation (8-01), using the time series techniques of Box and Jenkins (1970), is estimated for the quarterly CPI-U series from, 1962:3 to 1972:2.[9] Using the coefficient estimates of the period selected, the

[7]The autoregressive integrated moving average model is usually denoted by the notation ARIMA(p,d,q), where p is the degree of the autoregressive process, d is the degree of differencing, and q is the degree of the moving-average process.

[8]Box and Jenkins (1970) describe this process in detail.

[9]The use of future inflation information to estimate the model coefficients is consistent with the notion that market participants anticipate future events.

CPI-U price level in quarter one of 1970 is forecasted using a 'forecast back' procedure of the SAS step-wise model. The model is reestimated for the period 1962:4 to 1972:3 and then used to forecast the CPI-U for 1970:2. This procedure is repeated until the last forecast date, 1986:3, is reached.[10]

The estimates of the CPI-U, net-of-shelter, index are used to derive estimates of the expected quarterly inflation. Unexpected quarterly inflation (ARIMA 1qtr.fwd.) is then defined as the difference between actual inflation and the estimated expected inflation.

Two methods were considered to convert the unexpected quarterly inflation rate to unexpected annual inflation rates on a quarterly basis. The first method used the following formula,

$$UI_{a(t)}=1(1+UI_{q(t)})(1+UI_{q(t+1)})(1+UI_{q(t+2)})(1+UI_{q(t+3)})-1 \tag{8-02}$$

where $UI_{a(t)}$ is the unexpected annual inflation (ARIMA 4qtrs.cum.) in period t and $UI_{q(t)}$ is the unexpected quarterly inflation in period t from the ARIMA model. Thus, the annual inflation is simply the product of the unexpected inflation in each subsequent quarter. However, this should underestimate annual inflation, since inflation expectations

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[10]The model residuals are evaluated by examining the Q-statistics for each quarter forecast. Residuals for all periods were determined to be "white-noise".

are generally more accurate for shorter terms. Considering a second methodology, the quarterly inflation series already includes information on future inflation in the estimation of the coefficients, thus quarterly unexpected inflation rates (ARIMA annualized) were annualized by multiplying each value by four. The unexpected inflation values are plotted in Figure 8-03.

Comparing the annualized estimate with the Livingston survey data on a semiannual basis, Table 8-02 indicates that the ARIMA (0,2,1) model yields a lower value of mean absolute unexpected inflation. While the data do not follow a random-walk, (see Figure 8-01) they have characteristics nearly identical to the Livingson survey data (the correlation coefficient is equal to .81). In addition, quarterly inflation expectations using the model, plotted in Figure 8-01, appear consistent with rational expectations.

TABLE 8-02

Comparative Statistics:
Livingston Survey vs ARIMA-1 Quarter Forward Model
of Unexpected Inflation

$$I_{u,LIV\_(t)} = \beta_0 + \beta_1 \ I_{u,n\_(t)} + u_{(t)}$$

PANEL A	Period	Obs. Root MSE	ß0 (t-stat)	ß1 (t-stat)	R2 adjR2
	1970:2 - 1985:4	32	-0.003	0.913	.635
		0.016	-1.222	7.221	.622

PANEL B	Absolute Unexpected Inflation Mean	Std.Dev.
Livingston Survey	0.0216	0.0160
ARIMA-1 Qtr.Frwd.	0.0189	0.0132

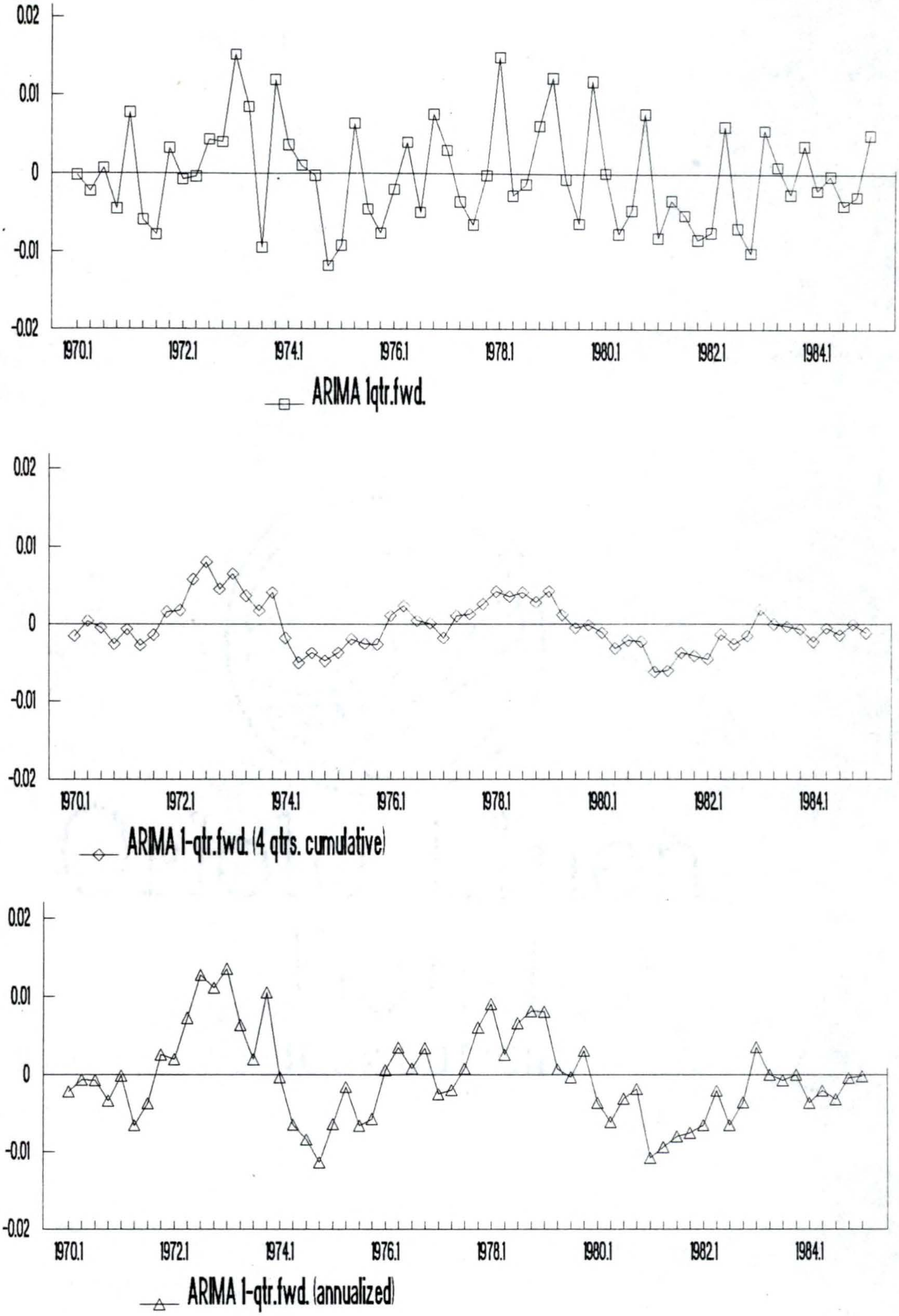

FIGURE 8-03 Unexpected Inflation
ARIMA Model: Quarterly, 4-Quarter, Annual

Unexpected Inflation and Housing Returns

Studies by Fama and Schwert (1977) and Hartzell, Hekman, and Miles (1987) indicate that both expected and unexpected inflation are highly correlated with housing returns (estimated coefficients are close to 1.0). If unexpected inflation is autocorrelated, it is necessary to adjust *ex post* housing returns due to unexpected inflation when evaluating *ex ante* after-tax excess returns. To remove the component of housing returns associated with unexpected inflation, after-tax excess returns, $R_{e,m}$, are regressed on unexpected inflation using the linear model

$$R_{e,m\_(t)} = \beta_0 + \beta_1 \, UI_{(t)} + e_{(t)}. \qquad (8\text{-}03)$$

The intercept, β_0, is interpreted as the expected after-tax excess returns uncorrelated with unexpected inflation, while the coefficient, β_1, reflects unexpected after-tax excess returns associated with unexpected inflation. The model residual, $e_{(t)}$, is then interpreted as the abnormal return.

It is anticipated that, consistent with past work, unexpected inflation is positively correlated with returns to housing. If inflation is systematically underestimated or overestimated by the market, treasury bill yields, as well as nominal housing returns, are expected to be affected. If inflation rates were one percent higher than expected this would result in, given previous findings, *ex post* nominal returns to housing being overstated by one percent. In addition, after-tax treasury bill yields may be understated by about three-quarters percent, assuming a marginal tax rate of about twenty-five percent. Hence, the estimated β_1 coefficient in Equation (8-01) is anticipated to be approximately 1.75.

The results of regressing after-tax excess returns for each city on unexpected inflation are listed in Tables 8-03, 8-05, 8-07, and 8-09.[11] These results suggest that after-tax excess returns, as modeled, are significantly correlated with unexpected inflation, which is serially dependent. In addition, the estimated β_0 and β_1 coefficients are close to their anticipated values.

EMH Test Results: Residual Returns Adjusted for Unexpected Inflation

Further tests of first-order autocorrelation are then conducted on the abnormal return series (the series of model residuals for Eq. 8-03) using a linear model such that,

$$Res_{n\_j(t)} = \beta_0 + \beta_1 \, Res_{n\_k(t-L)} + e_{(t),} \qquad (8\text{-}04)$$

where $Res_{n\_j(t)}$ is the abnormal return in period t constructed from sample j, $Res_{n\_k(t-L)}$ is the abnormal return in period t

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[11]The results of a strong positive relationship between unexpected inflation and excess return are quite robust. Unexpected inflation was modeled in a number of different ways on both an annual and quarterly basis. The results of these regressions are found in Tables B-04, B-06, B-08, and B-10 in Appendix B. In every case, results indicate that housing returns provide a significant hedge to unexpected inflation.

constructed from sample k lagged L periods; and e an error term. This model represents a minor modification of the previous first-order tests of autocorrelation.

The results of the regressions for each city are indicated in Tables 8-04, 8-06, 8-08, and 8-10. These results indicate that after the effect of unexpected inflation is removed from the after-tax excess return series, the hypothesis $\beta_1 = 0$ is rejected at the 5 percent confidence level consistently in only one city, Chicago. Both the β_0 and the β_1 coefficients are close to zero for Atlanta, suggesting little possibility for investors to identify trading rules which yield abnormal returns. In Dallas and in San Francisco findings suggest that abnormal returns exhibit some inertia. However, these series are not significantly different from zero at a five percent confidence level, and little confidence is placed in the San Francisco series because of the previously demonstrated measurement error.[12] While there exist autocorrelated returns in Chicago, they are substantially less significant than those previously reported by Case and Shiller (1989).

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[12]In addition, Proposition 13 was enacted in California during the sample period. This resulted in a substantial decrease in property taxes and a two percent ceiling was placed on annual property reassessments.

TABLE 8-03

Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Unexpected Inflation: Atlanta

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \, UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1986:1				
j=A, n=1		65	0.015	1.822	.142
		0.026	(4.685)	(3.226)	.128
j=B, n=1		65	0.016	1.839	.105
		0.031	(4.067)	(2.717)	.091
j=T, n=1		65	0.015	1.827	.153
		0.025	(4.943)	(3.371)	.139
j=T residuals:	DW=2.925	Q(24)=27.12			

Note: A and B denote half sample indices and T denotes a full sample index while L indicates the number of lagged quarters. ARIMA 1-qtr. forward (annualized) model of inflation indicated by n=1.

TABLE 8-04

Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged (Adjusted for Inflation Shocks): Atlanta

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 \, R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A;k=B;n=1;L=0		62	0.001	0.612	.461
		0.025	(0.153)	(4.526)	.452
j=A;k=B;n=1;L=4		58	-0.005	0.122	.026
		0.028	(-0.820)	(1.010)	.008
j=B;k=A;n=1;L=4		58	-0.007	0.000	.000
		0.034	(-1.033)	(0.002)	-.018

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE 8-05

Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Unexpected Inflation: Chicago

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \ UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Chicago	1970:1-1986:1				
j=A, n=1		65	0.016	1.484	.007
		0.024	(5.277)	(2.838)	-.009
j=B, n=1		65	0.015	1.647	.103
		0.023	(5.180)	(3.309)	.089
j=T, n=1		65	0.015	1.403	.144
		0.020	(5.884)	(3.283)	.131

j=T residuals: DW=1.309, Q(24)=51.65

Note: A and B denote half sample indices and T denotes a full sample index while L indicates the number of lagged quarters. ARIMA 1-qtr forward (annualized) model of inflation indicated by n=1.

TABLE 8-06

Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged (Adjusted for Inflation Shocks): Chicago

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 \ R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Chicago	1970:1-1985:2				
j=A;k=B;n=4;L=0		62	-0.004	0.897	.843
		0.025	(-0.697)	(8.062)	.841
j=A;k=B;n=4;L=4		58	-0.006	0.593	.359
		0.047	(-0.469)	(2.911)	.348
j=B;k=A;n=4;L=4		58	0.001	0.549	.304
		0.049	(0.066)	(2.567)	.292

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE 8-07

Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Unexpected Inflation: Dallas

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \, UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Dallas	1970:1-1986:1				
j=A, n=1		65	0.021	2.097	.120
		0.033	(5.094)	(2.937)	.106
j=B, n=1		65	0.021	1.754	.076
		0.035	(4.738)	(2.282)	.062
j=T, n=1		65	0.017	1.622	.127
		0.025	(5.341)	(3.052)	.113

j=T residuals: DW=2.878 Q(24)=25.63

Note: A and B denote half sample indices and T denotes a full sample index while L indicates the number of lagged quarters. ARIMA 1-qtr. forward (annualized) model of inflation indicated by n=1.

TABLE 8-08

Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged (Adjusted for Inflation Shocks): Dallas

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 \, R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Dallas	1970:1-1985:2				
j=A;k=B;n=1;L=0		62	0.007	0.730	.458
		0.045	(0.757)	(4.522)	.449
j=A;k=B;n=1;L=4		58	0.005	0.410	.159
		0.054	(0.376)	(1.949)	.144
j=B;k=A;n=1;L=4		58	-0.008	0.260	.085
		0.053	(-0.605)	(1.409)	.068

Note: Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980). See notes above.

TABLE 8-09

Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Unexpected Inflation: San Francisco

$$R_{e,m\_j(t)} = \beta_0 + \beta_1\, UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
San Francisco	1970:1-1986:2				
j=A, n=1		66	0.025	0.825	.015
		0.040	(5.059)	(0.969)	-.001
j=B, n=1		66	0.027	2.061	.055
		0.050	(4.306)	(1.933)	.040
j=T, n=1		66	0.025	1.290	.060
		0.030	(6.841)	(2.035)	.046
j=T residuals:	DW=1.500 Q(24)=46.32				

Note: A and B denote half sample indices and T denotes a full sample index while L indicates the number of lagged quarters. ARIMA 1-qtr. forward model of inflation indicated by n=1.

TABLE 8-10

Regression of After-Tax 4-Quarter Residual Returns After-Tax 4-Quarter Residual Returns Lagged (Adjusted for Inflation Shocks): San Francisco

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1\, R_{res,n\_k,(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
San Francisco	1970:1-1985:3				
j=A;k=B;n=1;L=0		63	0.002	0.598	.410
		0.065	(0.145)	(3.510)	.401
j=A;k=B;n=1;L=4		59	0.001	0.513	.301
		0.073	(0.072)	(2.882)	.290
j=B;k=A;n=1;L=4		59	0.001	0.422	.155
		0.086	(0.037)	(1.831)	.141

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

In addition, these results suggest that the degree of efficiency in the single-family housing market is related to city size. It is suspected that search costs are positively correlated to city size, limiting the ability for market participants to correct pricing errors.

Finally, the Chicago market was less active over the period studied. Atlanta, Dallas, and San Francisco all experienced substantial growth during this time period. These findings preliminarily suggest that active markets are relatively more efficient than inactive markets. This suggests abnormal returns may be easier to identify in slower markets periods; however, this issue requires more research.

CHAPTER 9
TIME-VARYING RISK PREMIA

Evidence of Time-Varying Volatility

Given the "momentum" in the after-tax excess returns adjusted for unexpected inflation, this chapter explores the possible intertemporal relationship between expected risk and expected housing returns. The chapter concludes by examining a trading strategy to determine if the resultant after-tax excess return, the study's adjusted return structure, is exploitable.

Returns to real estate, including returns to housing, have often been described as dominating other asset classes both before and after adjusting returns for risk. However, these returns are often generated from appraisal based indices, which are argued to smooth the volatility of the return series.[1] This data set provides possibly a unique opportunity to analyze the relationship between expected risk, as proxied by expected return volatility, and expected returns to housing.

In particular, expected after-tax excess returns to housing (the expected risk premia) are examined and

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[1]See Geltner (1989) and Giliberto (1988).

hypothesized to be positively correlated to expected housing return volatility.[2] In addition, if housing expected return volatility is serially dependent and correlated with the expected risk premium to housing, it will bias subsequent tests of autocorrelation of the after-tax excess return series. This condition is commonly referred to as autoregressive conditional heteroskedasticity (ARCH). To date, no real estate studies have been published which address this issue.

Casual observation of the *ex post* quarterly after-tax excess returns (Figures 7-02, 7-03, 7-04, and 7-05) suggest that volatility may be time-varying and correlated with returns. It appears that *ex post* return volatility in each market is greater in the periods prior to 1979. This would relate to the periods marked by high *ex post* after-tax excess annual returns shown previously in Figure 6-01.

Using the model described by Equation 8-03, a series of *ex ante* after-tax excess returns can be constructed, excluding the return associated with unexpected inflation. A moving average variance using subsequent four periods is calculated from the previously constructed *ex ante* after-tax excess return series. This series of variances is defined

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[2]French, Schwert, and Stambaugh (1987) report that the expected stock market risk premium is positively correlated to expected return volatility. In addition, they find that unexpected return volatility is negatively correlated with the expected market risk premium, providing further indirect evidence.

as expected volatility. Using 16 data observations, every fourth period, Table 9-01 reports the correlation coefficients of expected after-tax excess housing returns in each city with expected volatility. While the correlations are found to be positive, as hypothesized, the number of observations is so small that there is little reliability in any drawing conclusions from these data.

TABLE 9-01
Correlation of Expected Quarterly After-Tax Excess Returns and Expected Volatility

City Parameters	Atlanta E[Re]	Chicago E[Re]	Dallas E[Re]	San Fran E[Re]
E[Atl_vol]	0.65			
E[Chi_vol]		0.01		
E[Dal_vol]			0.27	
E[San_vol]				0.13

Autoregressive Conditional Heteroskedasticity

The autoregressive conditional heteroskedastic (ARCH) model, initially developed by Engle (1982), allows the full data set of 62 quarterly observations to be used. In examining the a simple autoregressive model,

$$y_t = \Gamma y_{t-1} + e_t \qquad (9\text{-}01)$$

where e_t is assumed independent and normally distributed with mean, 0, and variance, σ^2. Maddala (1988) indicates, that the conditional mean, $E(y_t|y_{t-1}) = \Gamma\ y_{t-1}$, is a function of t while the conditional variance, $var(y_t|y_{t-1})$ is constant. While the simple autoregressive model assumes that the variance is constant, the ARCH model allows the variance to be a function of the past. Thus, the variance is autocorrelated. Engle (1982), generalizing Equation (9-01) specifies the conditional mean as a linear function such that,

$$y_t = \Gamma\ y_{t-1} + e_t,$$

with e_t ~$IN(0,h_t)$, where

$$h_t = var\ e_t = \alpha_0 + \alpha_1 e^2_{(t-1)} + \ldots + \alpha_n e^2_{(t-n)}. \qquad (9\text{-}02)$$

Under this condition the trend in the variance is conveyed into the mean, resulting in a trend in the mean. Therefore, while diagnostic statistics may suggest that an autocorrelated error structure of the OLS regression exists, it is simply the result of serially dependent variance.

The ARCH model has been recently used effectively by Engle, Lilien, and Robins (1985), and French, Schwert, and Stambaugh (1987), and others, to model speculative security prices. They indicate that the variance structure of the series analyzed is predictable and this is translated to the characteristics of the expected mean.

Modeling Risk-Premia

This study uses a modification of the ARCH model developed by Engle, Lilien, and Robins (1982) to adjust expected after-tax excess returns for systematic variance in the return series. It is commonly referred to as the ARCH-M model (ARCH in mean). The ARCH-M model allows the conditional variance to be a determinant in modeling the expected mean. Hence, expected after-tax excess returns to housing are modeled as,

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \, UI_{n\_(t)} + \beta_1 \, \sigma_{(t)} + e_{(t)}$$

where,
$$\sigma^2_{(t)} = \alpha_0 + \alpha_1 \, e^2_{(t-1)} + \alpha_2 \, e^2_{(t-2)} \qquad (9\text{-}03)$$

This suggests that the variance of the return series follows an autoregressive process. The expected variance is modeled as a function of the squared residuals of the previous two quarters. Equation 9-03 is estimated using the quarterly returns generated from the after-tax excess returns constructed in Chapter 7, for each city.

The nonlinear model is estimated using a maximium likelihood function. Employing the Regression Analysis of Time Series (RATS) software available from VAR Econometrics, the ARCH model coefficients specified by Eq. 9-03 are estimated.

ARCH-M Test Results

The test results indicate a wide dispersion of model estimates across the four cities. Two model specifications were estimated, the first does not control for unexpected inflation, while the second model simultaneously estimates a β_1 coefficient.

The first model suggests that with the exception of Dallas, which is negatively related, there is little significance in any of the estimated α_1 or α_2 coefficients. This indicates that there exists little predictability in the variance structure of quarterly after-tax excess returns to housing. San Francisco indicates signs of positive momentum in the variance structure, however the α_1 is not significant at the 5 percent level.

Estimating the second model yields results similar to those of the first. Again, in this model the variance structure does not appear to be consistently predictable. Only in San Francisco do the coefficient estimates follow the hypothesized results. In the other cities, contrary to prior expectations, the expected variance appears to be negatively correlated with the expected risk-premium to housing. This condition, while not robust, is speculated to be due to the measurement error in the return series itself and the limited number of observations available. Despite using quarterly return data as opposed to the annual data used previously, the β_1 coefficient remains quite significant and the coefficient estimate near 1.75 as anticipated.

Tests of the model residuals are determined to be invalid due to the unpredictability of the expected variance term. Thus, returns to housing are not consistent with the ARCH-M model as specified. While the estimated model is quite sensitive to the specification of the expected variance function, other model specifications were examined with little change in the results.

TABLE 9-02
Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Unexpected Inflation and Volatility

$$R_{e,m\_j(t)} = \beta_0 + \beta_1\ UI_{n\_(t)} + \beta_2\ \sigma_{(t)}$$

City Obs.	β_0 (t-stat)	β_1 (t-stat)	β_2 (t-stat)	α_0 (t-stat)	α_1 (t-stat)	α_2 (t-stat)
Atlanta	0.023		-0.385	0.001	0.019	-0.428
61	(0.352)		(-0.151)	(4.532)	(0.060)	(-0.818)
	0.041	1.897	-1.344	0.001	-0.181	-0.549
	(5.885)	(5.095)	(-2.990)	(5.277)	(-2.840)	(-4.049)
Chicago	0.021		-0.308	0.000	0.255	0.131
61	(0.691)		(-0.211)	(2.739)	(0.623)	(0.213)
	-0.283	1.642	15.656	0.000	0.006	0.017
	(-0.028)	(3.281)	(0.029)	(2.898)	(0.030)	(0.028)
Dallas	0.033		-0.780	0.001	-0.151	-0.345
61	(10.714)		(-3.919)	(6.114)	(-2.334)	(-0.816)
	0.042	1.900	-1.242	0.001	-0.169	-0.503
	(7.307)	(5.063)	(-3.419)	(4.337)	(-1.195)	(-0.271)
San Fran.	0.020		0.746	0.001	0.836	-0.457
61	(2.235)		(0.235)	(3.037)	(1.693)	(-0.702)
	0.017	1.504	0.252	0.000	1.010	-0.533
	(1.346)	(4.147)	(0.527)	(2.647)	(2.237)	(-3.090)

Note: Unexpected inflation model with ARIMA 1-qtr. forward model.

Trading Strategies

As described previously in Chapter 2, formal testing of the random-walk model is limited in it's ability to describe fully the necessary conditions of the EMH. In order to reject the EMH, random-walk inconsistencies must be exploitable. Autocorrelation in the return structure alone, does not imply a rejection of the EMH.

The results reported in Chapter 8 indicate that after adjusting for unexpected inflation, abnormal returns are significantly correlated in Chicago. In addition, although not statistically significant, Dallas and San Francisco exhibit substantial inertia in their abnormal return structure. However, the returns to housing in Atlanta appear consistent with that of the random-walk and unexploitable.

Exploiting any known serial dependency in the housing market appears quite difficult. Their currently exist no futures contracts or short selling opportunities for the housing asset. Thus, any expected decrease in the price of housing is very difficult to exploit. In addition, transaction costs, which are assumed by Linneman (1986) to be as high as 12 percent, make it largely prohibitive for the typical home buyer to exploit any expected price increases.[3]

Case and Shiller (1989) indicate, home buyers who already intend to purchase a home, but are indifferent to the time of purchase, may be able to exploit the market. In this case, all transaction costs are irrelevant, since they will be incurred anyway.

The following trading rule is suggested by Case and Shiller: buy now if the after-tax excess return modeled by the regression in Table 6-04 is greater than the after-tax excess full sample mean, otherwise delay the purchase for one year. Using this methodology they indicate the average trading profits for the years invested earn range about 1 to 3 percent of the value of the house.[4]

Looking closer at this strategy, it implies that there are no costs to waiting a year. However, under such a strategy any no excess housing returns can be earned in the year waited. Hence, it is important to include the "opportunity cost" of waiting a year.

Under this scenerio, which includes the "cost" of delaying the purchase, one buys if the expected after-tax

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[3]Transaction costs often include such items as brokers fees, moving expenses, and search costs. In addition, other types of costs may be incurred which are difficult to quantify, such as new schools or neighborhood amenities.

[4]Case and Shiller define the average trading profit as the difference between the buy signal average returns and the total sample average returns times the proportion of quarters indicating a buy signal.

excess return is greater than the sample mean, otherwise the home buyer delays purchase for one year and invests in one-year treasury bills. The average return from this strategy is listed below in column B of Table 9-03. Column C indicates a strategy where the buyer buys if the expected after-tax excess return is greater than the treasury bill, otherwise dely the purchase and invest in the security. Only in San Francisco does this trading rule earn excessive returns; however, the rule is not consistent across cities.

TABLE 9-03

Average Quarterly Returns
Using Alternative Trading Strategies

City Regression	A Ave. Sample R_e	B Mean Strat. $E[R_e]$	C TBill Strat. $E[R_e]$	D Ave. Sample R_e,m	E Mean Strat. $E[R_e,m]$	F TBill Strat. $E[R_e,m]$
Atlanta						
AB	.0615	.0376	.0615	.0536	.0213	.0009
BA	.0622	.0398	.0622	.0541	.0290	.0000
Chicago						
AB	.0646	.0605	.0644	.0569	.0549	.0528
BA	.0636	.0550	.0646	.0559	.0519	.0509
Dallas						
AB	.0949	.0551	.0949	.0870	.0477	.0553
BA	.0846	.0477	.0827	.0766	.0423	.0414
San Fran.						
AB	.1105	.0923	.1128	.1059	.0685	.0685
BA	.1141	.1018	.1152	.1094	.0801	.0842

Note: Column A, B, and C denotes the average quarterly return for each sample using unadjusted after-tax excess return data. Columns D, denotes the sample average quarterly return using return data adjusted for home buyer marginal tax rates. Columns E and F applies trading rules on the basis of the expected after-tax excess returns adjusted for unexpected inflation.

Column D reports the average annual after-tax excess returns using time-varying home buyer marginal tax rates. Using expected after-tax returns calculated from the regression in Tables 8-03, 8-05, 8-07, 8-09 in Chapter 8 to signal the buy or delay purchase signal. The actual returns that would have resulted from such a signal are listed in columns E and F of Table 9-01. Column E gave a buy signal for values above the mean of the expected after-tax returns adjusted for unexpected inflation, while column F gave a buy signal for expected returns greater than the after-tax one-year treasury bill. These strategies do not consistently outperform the buy-and-hold strategy, hence the EMH cannot be rejected.

CHAPTER 10
SUMMARY AND CONCLUSIONS

Previous results examining the efficiency of the single-family housing market have been mixed. This study extends past work by incorporating time-varying home buyer marginal tax rates, unexpected inflation and volatility into the analysis of after-tax excess returns to housing.

Findings indicate that excess housing returns are negatively correlated with home buyer marginal tax rates, which are determined to be autocorrelated. However, when time-varying marginal tax rates are introduced into the housing return model and the resulting return series tested for autocorrelation, there is not significant change from previous estimates. This suggests that findings of autocorrelated after-tax excess returns to housing are robust to alternative marginal tax rate specifications.

More importantly, excess housing returns are found to be positively correlated with unexpected inflation. Evidence is presented that suggests inflation shocks were autocorrelated during the late 1970s and early 1980s, and not fully anticipated by either the bond or housing markets. When after-tax excess returns are adjusted for both time-varying marginal tax rates and autocorrelated inflation shocks, findings indicate that there remains persistent return inertia in three of the four markets studied, however, trading rules cannot exploit this autocorrelation. Hence the hypothesis of an efficient market cannot be rejected.

In addition, preliminary evidence suggests that autocorrelated return patterns are correlated with city size. It is speculated that inactive markets may induce inefficient results and that substantial search costs in larger cities may inhibit competing properties to be identified allowing greater price inertia to occur. This would appear to explain the conditions of predictable returns consistently found in Chicago.

Finally, the volatility of the return structure is explored. While expected return volatility is found to be positively correlated to expected after-tax quarterly excess returns, employing an ARCH-M model using quarterly returns yields little predictability in the volatility of city return series. Hence, there appears to be no ARCH impact on the expected after-tax excess returns. Although the data do not confirm the ARCH model, it seems likely that home buyers would require greater expected returns during times of volatile inflation (and volatile housing returns), however, this remains an area to be examined in the future.

APPENDIX A
WEIGHTED-REPEAT-SALES (WRS) PRICE INDICES

The median sales price of single-family housing for different markets is widely reported in the popular literature. Such price statistics have been highly criticized as measures of housing appreciation primarily because price changes are contaminated by changes in the age, characteristics and size of the units sold over the period reported. Several methods have been suggested in the literature that attempt to identify the changes in price of "equivalent quality" housing.[1]

Case and Shiller (1987) propose a weighted-repeat-sales (WRS) methodology for constructing existing home price indices. They argue that the WRS methodology is superior to other return indices because it most accurately controls for size and quality changes in house transaction data.

The WRS method is a modification of the work of Bailey, Muth, and Nourse (1963) (BMN), controlling for heteroskedasticity in the regression error term. The BMN method can be described by first assuming:

$$P_s = P_i(1+r_1)^{D1}(1+r_2)^{D2}(1+r_3)^{D3}\ldots\ldots(1+r_n)^{Dn} \qquad \text{(A-01)}$$

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[1]Nourse (1982) and Greenlees (1984) provide useful critics of alternative housing price indices.

where P_i is the initial sales price; P_s is the second sales price; r_t is the rate of appreciation in period t; Dt is the exponent equal to -1 if period t is the initial sale, +1 if it is the repeat sale, and 0 for all other t. Taking the natural log of (6.01) yields,

$$P_s{}'=P_i{}'+D_1(1+r_1)'+D_2(1+r_2)'+D_3(1+r_3)'+\ldots+D_n(1+r_n)'; \qquad \text{(A-02)}$$

thus,

$$(P_s/P_i)'=D_1(1+r_1)'+D_2(1+r_2)'+D_3(1+r_3)'+\ldots+D_n(1+r_n)'; \qquad \text{(A-03)}$$

where ( )' denote the log function.

Using data which identify a large number of repeat sales, (B.03) can be estimated by linear regression. The independent variables, D_t, represent a series of dummy variables for each period t in the sample, except for the initial period 0. The dummy variables take the value -1 for the year in which the initial sales tales place, 1 for the year in which the second sale takes place, otherwise 0. The estimated coefficients represent the natural logs of the constructed price index.

Case and Shiller indicate that the variance of the regression error term from (A-03) is likely to be related to the time interval between repeat sales. Further, they argue that there is to be a random drift component due to the differences in the amount of house depreciation and upkeep among houses. A modification is made to the BMN method by assuming that the log price of the jth house at time t is given by:

$$P_{jt} = C_t + H_{jt} + U_{jt} \qquad \text{(A-04)}$$

where C_t is the log of the citywide level of housing prices at time t; H_{jt} is a random walk drift component of zero mean and variance $\sigma^2{}_h$; and U_{jt} is a house specific random error term having zero mean and variance $\sigma^2{}_n$.

The WRS method consists of three stages. First, the BMN procedure is followed by estimating (B.03), and a vector of regression residuals is calculated. A weighted regression of the squared residuals in the first stage is then regression on the time interval between the sales using a constant term. The constant term of the second-stage regression is an estimate of the house specific random error ($2\sigma^2{}_n$), and the slope term is the estimate of drift ($\sigma^2{}_h$). Finally, a generalized least squares regression (weighted) is run by first dividing each observation in the first-stage regression by the square root of the fitted value in the stage-two regression and then running the stage-one regression again.

Sample bias can enter the WRS methodology in applying the data to the regressions. When used to estimate indices for repeat-sales over short time periods, the index restricts observations to a disproportionate number of high turnover properties, potentially reflecting the pricing of forced sales or renovation properties. This effect will influence the index, particularly if the proportion of high turnover properties are not constant over time. In this case any subsequent tests of autocorrelation will be biased.

Despite these limitations, it is unlikely that spurious correlation of this type, introduced by the methodology, could result in the degree of autocorrelation reported by Case and Shiller. It appears reasonable to assume that one directional tests of efficiency can be conducted using the methodology. While the WRS methodology may induce some limited time pattern to the return series, real estate transactions costs are generally sufficient to suppress the likelihood of establishing any profitable investment rules under such conditions. Therefore, if the return series are found to follow random-walk properties, it is reasonable not to reject the hypothesis of an efficient market.

APPENDIX B
SUPPLEMENTARY TABLES AND STATISTICS

TABLE B-01

Regression of After-Tax 4-Quarter Excess Returns on After-Tax 4-Quarter Excess Returns Lagged

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A, k=B, L=0		62	0.011	0.820	.695
		0.029	(1.106)	(6.428)	.690
j=A, k=B, L=4		58	0.038	0.331	.114
		0.049	(2.085)	(1.478)	.098
j=B, k=A, L=4		58	0.037	0.363	.152
		0.046	(2.147)	(1.741)	.137
Chicago	1970:1-1985:2				
j=A, k=B, L=0		62	0.004	0.935	.861
		0.026	(0.504)	(10.616)	.859
j=A, k=B, L=4		58	0.019	0.663	.447
		0.053	(1.132)	(3.697)	.437
j=B, k=A, L=4		58	0.017	0.698	.476
		0.053	(1.008)	(3.921)	.466
Dallas	1970:1-1985:2				
j=A, k=B, L=0		62	0.024	0.810	.600
		0.048	(1.440)	(5.213)	.593
j=A, k=B, L=4		58	0.045	0.540	.271
		0.064	(1.845)	(2.508)	.258
j=B, k=A, L=4		58	0.031	0.520	.290
		0.061	(1.292)	(2.629)	.277
San Francisco	1970:1-1985:3				
j=A, k=B, L=0		63	0.040	0.619	.483
		0.069	(1.736)	(4.150)	.474
j=A, k=B, L=4		59	0.053	0.523	.216
		0.099	(1.498)	(2.177)	.202
j=B, k=A, L=4		59	0.044	0.562	.398
		0.077	(1.676)	(3.374)	.387

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE B-02
Regression of After-Tax Quarterly Excess Returns on
After-Tax Quarterly Excess Returns Lagged
Using Nonconstant Tenure Choice Marginal Tax Rates

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 $adjR^2$
Atlanta	1970:1-1986:1				
j=A, k=B, L=0		65	0.007	0.516	.368
		0.022	(2.212)	(6.058)	.358
j=A, k=B, L=1		64	0.016	-0.139	.027
		0.028	(4.326)	(-1.302)	.011
j=B, k=A, L=1		64	0.018	-0.225	.037
		0.032	(4.026)	(-1.534)	.021
Chicago	1970:1-1986:1				
j=A, k=B, L=0		65	0.008	0.534	.268
		0.022	(2.449)	(4.804)	.266
j=A, k=B, L=1		64	0.006	0.562	.304
		0.021	(2.205)	(5.200)	.292
j=B, k=A, L=1		64	0.010	0.263	.071
		0.024	(2.800)	(2.172)	.056
Dallas	1970:1-1986:1				
j=A, k=B, L=0		65	0.015	0.244	.066
		0.034	(3.166)	(2.103)	.051
j=A, k=B, L=1		64	0.017	0.215	.053
		0.033	(3.493)	(1.863)	.038
j=B, k=A, L=1		64	0.017	0.078	.007
		0.032	(3.583)	(0.673)	-.009
San Francisco	1970:1-1986:2				
j=A, k=B, L=0		66	0.026	-0.050	.004
		0.040	(4.833)	(-0.508)	-.012
j=A, k=B, L=1		65	0.022	0.155	.040
		0.039	(4.006)	(1.613)	.024
j=B, k=A, L=1		65	0.005	0.837	.416
		0.040	(0.775)	(6.701)	.407

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters.

TABLE B-03

Regression of After-Tax 4-Quarter Excess Returns on After-Tax 4-Quarter Excess Returns Lagged Using Nonconstant Tenure Choice Tax Rates

$$R_{e\_j(t)} = \beta_0 + \beta_1 R_{e\_k(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A, k=B, L=0		62	0.011	0.818	.690
		0.029	(1.092)	(6.353)	.685
j=A, k=B, L=4		58	0.036	0.363	.115
		0.048	(2.031)	(1.485)	.099
j=B, k=A, L=4		58	0.035	0.333	.151
		0.046	(2.096)	(1.737)	.136
Chicago	1970:1-1985:2				
j=A, k=B, L=0		62	0.004	0.937	.866
		0.026	(0.468)	(10.822)	.864
j=A, k=B, L=4		58	0.017	0.669	.455
		0.054	(1.040)	(3.758)	.445
j=B, k=A, L=4		58	0.015	0.705	.486
		0.053	(0.921)	(4.166)	.477
Dallas	1970:1-1985:2				
j=A, k=B, L=0		62	0.024	0.801	.583
		0.048	(1.464)	(5.042)	.577
j=A, k=B, L=4		58	0.045	0.528	.260
		0.062	(1.903)	(2.440)	.247
j=B, k=A, L=4		58	0.032	0.507	.278
		0.060	(1.341)	(2.558)	.266
San Francisco	1970:1-1985:3				
j=A, k=B, L=0		63	0.043	0.597	.452
		0.069	(1.832)	(3.903)	.443
j=A, k=B, L=4		59	0.058	0.496	.193
		0.097	(1.606)	(2.030)	.179
j=B, k=A, L=4		59	0.047	0.550	.383
		0.076	(1.774)	(3.269)	.371

Note: A and B denote half sample indices for each city respectively and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE B-04
Regression of After-Tax Quarterly Excess Returns
Adjusted for Home Owner Marginal Tax Rates
on Alternative Unexpected Inflation Models: Atlanta

$$R_{e,m\_j(t)} = \beta_0 + \beta_1\ UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1986:1				
j=A, n=2		65	0.012	1.242	.095
		0.027	(3.597)	(2.577)	.081
j=B, n=2		65	0.013	1.000	.045
		0.032	(3.216)	(1.719)	.030
j=A, n=3		65	0.015	1.487	.123
		0.026	(4.526)	(2.970)	.109
j=B, n=3		65	0.015	1.839	.060
		0.032	(3.840)	(2.717)	.045
j=A, n=4		65	0.013	0.886	.016
		0.028	(3.563)	(1.015)	.001
j=B, n=4		65	0.015	-0.103	.000
		0.033	(3.436)	(-0.100)	-.016
j=A, n=5		65	0.012	1.983	.062
		0.027	(2.743)	(2.041)	.047
j=B, n=5		65	0.011	1.946	.082
		0.032	(3.101)	(2.377)	.068
j=A, n=6		65	0.014	0.899	.053
		0.027	(4.195)	(1.882)	.038
j=B, n=6		65	0.015	0.911	.040
		0.032	(1.882)	(1.613)	.024

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. n denotes the model used for inflation expectations

n=2: ARIMA model (in sample)
n=3: ASA-NBER
n=4: Polynomial Distributed Lag (6,3)
n=5: Harris PDL
n=6: ARIMA model (1-step forward, using city specific inflation data.

TABLE B-05
Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged: Atlanta
(Using Alternative Models)

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Atlanta	1970:1-1985:2				
j=A;k=B;n=2;L=0		62	0.002	0.770	.647
		0.027	(0.472)	(5.767)	.641
j=A;k=B;n=2;L=4		58	-0.003	0.380	.183
		0.039	(-0.632)	(1.948)	.168
J=B;k=A;n=2;L=4		58	-0.006	0.346	.120
		0.043	(-1.113)	(1.522)	.105
j=A;k=B;n=3;L=0		62	0.001	0.689	.561
		0.026	(0.346)	(4.816)	.554
j=A;k=B;n=3;L=4		58	-0.004	0.355	.180
		0.033	(-0.813)	(1.925)	.165
j=B;k=B;n=3;L=4		58	-0.006	0.317	.094
		0.039	(-1.270)	(1.326)	.078
j=A;k=B;n=4;L=0		62	0.001	0.766	.690
		0.026	(0.245)	(6.357)	.685
j=A;k=B;n=4;L=4		58	-0.002	0.321	.133
		0.044	(-0.207)	(1.610)	.117
j=B;k=A;n=4;L=4		58	-0.005	0.355	.156
		0.048	(-0.463)	(1.488)	.100
j=A;k=B;n=5;L=0		62	0.002	0.699	.581
		0.026	(0.276)	(5.017)	.574
j=A;k=B;n=5;L=4		58	-0.001	0.233	.068
		0.038	(-0.076)	(1.115)	.051
j=B;k=B;n=5;L=4		58	-0.003	0.205	.038
		0.042	(-0.343)	(0.813)	.020
j=A;k=B;n=6;L=0		62	0.002	0.738	.563
		0.026	(0.262)	(4.837)	.556
j=A;k=B;n=6;L=4		58	-0.004	0.243	.079
		0.035	(-0.513)	(1.205)	.063
j=B;k=A;n=6;L=4		58	-0.007	0.140	.023
		0.038	(-0.806)	(0.627)	.005

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE B-06
Regression of After-Tax Quarterly Excess Returns
Adjusted for Home Owner Marginal Tax Rates
on Alternative Unexpected Inflation Models: Chicago

$$R_{e,m\_j(t)} = \beta_0 + \beta_1\ UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Chicago	1970:1-1986:1				
j=A, n=2		65	0.015	0.048	.000
		0.026	(4.561)	(0.106)	-.016
j=B, n=2		65	0.012	0.832	.055
		0.024	(4.084)	(1.908)	.040
j=A, n=3		65	0.015	0.320	.007
		0.025	(4.739)	(0.659)	-.009
j=B, n=3		65	0.014	1.205	.103
		0.023	(4.914)	(2.693)	.089
j=A, n=4		65	0.013	1.187	.035
		0.025	(4.108)	(1.505)	.019
j=B, n=4		65	0.012	1.208	.038
		0.024	(3.865)	(1.582)	.023
j=A, n=5		65	0.011	2.263	.134
		0.024	(3.634)	(3.119)	.120
j=B, n=5		65	0.010	2.432	.164
		0.023	(3.340)	(3.517)	.151
j=A, n=7		65	0.016	1.066	.104
		0.024	(5.320)	(2.698)	.089
j=B, n=7		65	0.015	1.117	.121
		0.023	(5.160)	(2.941)	.107

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. n denotes the model used for inflation expectations

n=2: ARIMA model (in sample)
n=3: ASA-NBER
n=4: Polynomial Distributed Lag (6,3)
n=5: Harris PDL
n=7: ARIMA model (1-step forward, using city specific inflation data.

TABLE B-07

Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged: Chicago (Using Alternative Models)

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 $adjR^2$
Chicago	1970:1-1985:2				
j=A;k=B;n=2;L=0		62	-0.003	0.955	.862
		0.025	(-0.675)	(10.628)	.859
j=A;k=B;n=2;L=4		58	-0.005	0.744	.521
		0.049	(-0.461)	(4.289)	.512
j=B;k=A;n=2;L=4		58	-0.001	0.607	.394
		0.053	(-0.046)	(3.317)	.383
j=A;k=B;n=3;L=0		62	-0.004	0.961	.838
		0.027	(-0.616)	(9.681)	.835
j=A;k=B;n=3;L=4		58	-0.005	0.760	.522
		0.048	(-0.474)	(4.303)	.514
j=B;k=B;n=3;L=4		58	0.000	0.548	.334
		0.054	(0.021)	(2.916)	.322
j=A;k=B;n=4;L=0		62	-0.005	0.932	.690
		0.025	(-0.790)	(9.886)	.685
j=A;k=B;n=4;L=4		58	-0.005	0.641	.133
		0.050	(-0.374)	(3.372)	.117
j=B;k=A;n=4;L=4		58	-0.003	0.592	.156
		0.052	(-0.223)	(3.079)	.100
j=A;k=B;n=5;L=0		62	-0.005	0.905	.792
		0.025	(-0.794)	(8.303)	.788
j=A;k=B;n=5;L=4		58	-0.002	0.514	.259
		0.048	(-0.210)	(2.435)	.246
j=B;k=B;n=5;L=4		58	-0.004	0.436	.195
		0.050	(-0.332)	(2.025)	.181
j=A;k=B;n=7;L=0		62	-0.005	0.873	.790
		0.025	(-0.820)	(8.270)	.787
j=A;k=B;n=7;L=0		58	-0.006	0.489	.250
		0.048	(-0.555)	(2.375)	.236
j=A;k=B;n=7;L=0		58	-0.002	0.469	.218
		0.049	(-0.021)	(2.173)	.204

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE B-08

Regression of After-Tax Quarterly Excess Returns Adjusted for Home Owner Marginal Tax Rates on Alternative Unexpected Inflation Models: Dallas

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \, UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Dallas	1970:1-1986:1				
j=A, n=2		65	0.017	1.613	.103
		0.033	(4.075)	(2.689)	.089
j=B, n=2		65	0.018	1.388	.069
		0.036	(3.939)	(2.162)	.054
j=A, n=3		65	0.020	1.629	.009
		0.033	(4.914)	(2.566)	.008
j=B, n=3		65	0.020	1.765	.101
		0.035	(4.771)	(2.653)	.086
j=A, n=4		65	0.017	1.975	.051
		0.034	(3.848)	(1.845)	.036
j=B, n=4		65	0.019	0.493	.003
		0.037	(4.008)	(0.428)	-.013
j=A, n=5		65	0.017	1.965	.054
		0.034	(3.661)	(1.892)	.039
j=B, n=5		65	0.017	2.093	.055
		0.036	(3.503)	(1.919)	.040
j=A, n=8		65	0.021	1.578	.130
		0.033	(5.152)	(3.061)	.116
j=B, n=8		65	0.021	0.958	.043
		0.036	(4.598)	(1.687)	.028

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. n denotes the model used for inflation expectations

n=2: ARIMA model (in sample)
n=3: ASA-NBER
n=4: Polynomial Distributed Lag (6,3)
n=5: Harris PDL
n=8: ARIMA model (1-step forward, using city specific inflation data.

TABLE B-09

Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Excess Returns Lagged: Dallas (Using Alternative Models)

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 R_{res,n\_k,(t-L)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
Dallas	1970:1-1985:2				
j=A;k=B;n=2;L=0		62	0.008	0.771	.535
		0.046	(0.773)	(4.569)	.527
j=A;k=B;n=2;L=4		58	0.005	0.514	.252
		0.045	(0.377)	(2.391)	.239
J=B;k=A;n=2;L=4		58	-0.006	0.500	.290
		0.056	(-0.837)	(2.627)	.277
j=A;k=B;n=3;L=0		62	0.008	0.777	.511
		0.045	(0.791)	(4.357)	.503
j=A;k=B;n=3;L=4		58	0.005	0.530	.251
		0.054	(0.407)	(2.382)	.238
j=B;k=B;n=3;L=4		58	-0.010	0.443	.241
		0.050	(-0.852)	(2.320)	.228
j=A;k=B;n=4;L=0		62	0.008	0.754	.561
		0.025	(0.746)	(4.814)	.554
j=A;k=B;n=4;L=4		58	0.007	0.530	.284
		0.050	(0.520)	(2.595)	.272
j=B;k=A;n=4;L=4		58	-0.009	0.523	.284
		0.052	(-0.637)	(2.594)	.272
j=A;k=B;n=5;L=0		62	0.008	0.786	.537
		0.025	(0.831)	(4.584)	.529
j=A;k=B;n=5;L=4		58	0.009	0.514	.254
		0.048	(0.672)	(2.404)	.241
j=B;k=B;n=5;L=4		58	-0.007	0.431	.224
		0.050	(-0.519)	(2.212)	.210
j=A;k=B;n=8;L=0		62	0.007	0.672	.418
		0.046	(1.148)	(6.561)	.408
j=A;k=B;n=8;L=0		58	0.002	0.401	.161
		0.053	(0.232)	(1.802)	.146
j=A;k=B;n=8;L=0		58	-0.010	0.289	.098
		0.053	(-1.415)	(1.357)	.082

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

TABLE B-10
Regression of After-Tax Quarterly Excess Returns
Adjusted for Home Owner Marginal Tax Rates
on Alternative Unexpected Inflation Models: San Francisco

$$R_{e,m\_j(t)} = \beta_0 + \beta_1 \, UI_{n\_(t)} + u_{(t)}$$

City Parameters	Sample Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
San Francisco	1970:1-1986:2				
j=A, n=2		66	0.025	-0.059	.000
		0.041	(4.840)	(2.689)	-.016
j=B, n=2		66	0.022	2.138	.085
		0.050	(3.554)	(2.436)	.071
j=A, n=3		66	0.025	0.193	.001
		0.040	(4.964)	(0.252)	-.015
j=B, n=3		66	0.027	2.450	.098
		0.049	(4.420)	(2.639)	.084
j=A, n=4		66	0.024	0.788	.006
		0.040	(4.525)	(0.631)	-.009
j=B, n=4		66	0.024	1.129	.008
		0.052	(3.625)	(0.708)	-.008
j=A, n=5		66	0.021	2.507	.064
		0.039	(4.057)	(2.097)	.050
j=B, n=5		66	0.022	2.336	.034
		0.051	(3.312)	(1.504)	.019
j=A, n=9		66	0.026	0.797	.037
		0.039	(2.884)	(0.854)	.022
j=B, n=9		66	0.027	1.594	.089
		0.050	(2.389)	(1.361)	.074

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. n denotes the model used for inflation expectations

n=2: ARIMA model (in sample)
n=3: ASA-NBER
n=4: Polynomial Distributed Lag (6,3)
n=5: Harris PDL
n=9: ARIMA model (1-step forward, using city specific inflation data.

TABLE B-11
Regression of After-Tax 4-Quarter Residual Returns on After-Tax 4-Quarter Residual Returns Lagged: San Francisco
(Using Alternative Models)

$$R_{res,n\_j,(t)} = \beta_0 + \beta_1 R_{res,n\_k,(t)} + u_{(t)}$$

City Parameters	Period	Obs. MSE	β_0 (t-stat)	β_1 (t-stat)	R^2 adjR^2
San Francisco	1970:1-1985:3				
j=A;k=B;n=2;L=0		63	0.002	0.643	.458
		0.065	(0.153)	(3.952)	.450
j=A;k=B;n=2;L=4		59	0.002	0.596	.393
		0.071	(0.089)	(3.340)	.382
J=B;k=A;n=2;L=4		59	-0.001	0.457	.187
		0.087	(-0.058)	(1.993)	.173
j=A;k=B;n=3;L=0		63	0.003	0.628	.419
		0.067	(0.171)	(3.650)	.410
j=A;k=B;n=3;L=4		59	0.103	0.603	.385
		0.071	(-0.441)	(3.283)	.374
j=B;k=B;n=3;L=4		59	-0.001	0.451	.191
		0.083	(-0.073)	(2.015)	.176
j=A;k=B;n=4;L=0		63	0.002	0.591	.442
		0.064	(0.155)	(3.826)	.433
j=A;k=B;n=4;L=4		59	0.002	0.530	.356
		0.071	(0.114)	(3.086)	.345
j=B;k=A;n=4;L=4		59	0.001	0.462	.168
		0.091	(0.028)	(1.866)	.153
j=A;k=B;n=5;L=0		63	0.002	0.526	.367
		0.061	(0.166)	(3.272)	.357
j=A;k=B;n=5;L=4		59	0.004	0.431	.259
		0.068	(0.256)	(2.454)	.246
j=B;k=B;n=5;L=4		59	0.003	0.368	.099
		0.087	(0.127)	(1.380)	.084
j=A;k=B;n=9;L=0		63	0.001	0.598	.358
		0.065	(0.053)	(3.209)	.348
j=A;k=B;n=9;L=4		59	-0.001	0.579	.338
		0.064	(-0.037)	(2.969)	.327
j=B;k=A;n=9;L=4		59	-0.004	0.366	.135
		0.067	(-0.211)	(1.638)	.120

Note: A and B denote half sample indices for each city respectively, and L denotes number of lagged quarters. Coefficient standard errors have been adjusted to yield estimates of the corrected standard errors for a MA3 process as described by Hansen and Hodrick (1980).

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I certify that I have read this study and that in my opinion it conforms to acceptable standards of scholarly presentation and is fully adequate, in scope and quality, as a dissertation for the degree of Doctor of Philosophy.

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