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House Prices and Inflation

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The present paper examines the long-run impact of inflation on homeowner equity by investigating the relationship between house prices and the prices of nonhousing goods and services, rather than return series and inflation rates as in previous empirical studies on the inflation hedging ability of real estate. There are two reasons for this methodological departure from prior practice: (1) while the total return on housing cannot be accurately measured, the total return on housing is fully reflected in housing prices, and (2) given that using returns or differencing a time series leads to a loss of long-run information contained in the series, valuable long-run information can be captured by using prices. Also, unlike previous related studies, we exclude housing costs from goods and services prices to avoid potential bias in estimating how inflation affects housing prices. Monthly data series are collected for existing and for new house prices as well as the consumer price index excluding housing costs for the period 1968–2000. Based on both autoregressive distributed lag (ARDL) models and recursive regressions, the empirical results yield estimated Fisher coefficients that are consistently greater than one over the sample period. Thus, we infer that house prices are a stable inflation hedge in the long run.

In the United States, homeownership is enjoyed by two-thirds of the nation's households and homeowner equity constitutes about one-third of all household wealth. For most people, homeowner equity is their most important form of investment. Although corporate equity recently surpassed homeowner equity as the largest asset in the household sector, more than half of all households hold no corporate equity (Tracy, Schneider and Chan 1999). Because homeowner equity represents the largest portion of most households' investment portfolio, changes in the real value of homeowner equity have important implications for personal wealth as well as for the national economy. In this regard, the ability of homeowner equity to hedge against inflation compared to other forms of individual wealth, notably stocks and bonds, has been a subject of ongoing interest in the finance and economics literature.

The present paper examines the long-run impact of inflation on homeowner equity by investigating the relationship between house prices and prices of

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nonhousing goods and services, rather than return series and inflation rates as in previous empirical studies (to be discussed shortly) on the inflation hedging ability of residential real estate. There are two reasons for this methodological departure from prior practice: (1) while the total return on housing cannot be accurately measured, the total return on housing is fully reflected in housing prices, and (2) given that using returns or differencing a time series leads to a loss of long-run information contained in the series, valuable long-run information can be captured by using prices. Also, unlike previous studies that typically regress real estate returns on rates of change in the consumer price index (CPI) over time, we exclude housing costs from the CPI,[1] which historically ranges from 20% to 30% of the consumer price index. For example, as pointed out by an anonymous reviewer, Fama and Schwert (1977) regressed the housing component of CPI on CPI to examine whether real estate was a good inflation hedge. By excluding housing costs from our measure of goods and services prices, we avoid potential bias in estimating the so-called Fisher coefficient reflecting the relationship between inflation and asset values. Monthly data are collected for existing and new house prices as well as the consumer price index excluding housing costs for the 1968–2000 period. We employ autoregressive distributed lag (ARDL) models to obtain long-run Fisher coefficient estimates. To examine the variability of these estimates over the sample period, we utilize recursive regressions. The empirical results using the ARDL models generate estimated Fisher coefficients of 1.08 for existing home prices and 1.26 for new home prices, which are significantly greater than 1 as predicted by Darby's tax version of the Fisher equation. Recursive regressions indicate that time-varying Fisher coefficients are in the range of 1.19 to 1.42 during the period 1974–2000. These and other results lead to the conclusion that the estimated Fisher elasticities of house prices with respect to nonhousing prices of goods and services exceed 1 and are stable over the sample period. Thus, we infer that house prices are an inflation hedge in the long run.

The remaining sections of the paper overview related literature, provide a theoretical framework for the investigation of the relationship between house prices and inflation, discuss the econometric methods used for the estimation of the impact of inflation on house prices, report the empirical results, and provide the summary and conclusions.

Related Literature

Seminal work by Fama and Schwert (1977) compared U.S. government bonds and bills, private residential real estate, and common stocks in terms of their

[1] Our nonhousing cost index is CPI minus shelter costs as compiled by the U.S. Bureau of Labor Statistics.

ability to hedge against expected and unexpected movements in monthly interest rates in the 1953–1971 period. The home purchase price (HPP) component of the consumer price index (CPI), which is based on the prices of newly insured FHA home loans, was used as an index of house prices. Single equation time series analyses of the data for these three different asset categories revealed that expected changes in both government debt securities and real estate rates of return were close to unity with respect to a 1% change in the inflation rate. By contrast, and consistent with many other studies, common stock returns were negatively related to expected changes in inflation rates. Particularly relevant to the present paper, the regression coefficient measuring the relationship between housing rates of return and expected and unexpected movements in the inflation rate were 1.19 and 0.56, respectively, both significant at the 1% level. No other asset had a positive relationship to unexpected inflation. Thus, real estate was clearly the best hedge against expected and unexpected inflation. Fama and Schwert concluded that: (1) the expected responses of asset returns to inflation for government debt securities and real estate were consistent with the well-known Fisher (1930) hypothesis,[2] (2) real estate was the only "complete hedge" against expected and unexpected inflation in their sample period, and (3) results for common stocks were anomalous.

Some early studies on the inflation hedging ability of real estate focused on correlations between rates of return on real estate and inflation rates over time. For example, Reilly, Marquardt and Price (1977) compared the rates of return on U.S. land prices and common stock during the 1918–1974 period. While stock returns exceeded land returns over this long period, real estate returns exceeded rates of change in the CPI and exhibited a more stable relationship with CPI rates during periods of rapid inflation and deflation. Spellman (1981) also found that changes in housing prices grew more rapidly than both CPI and rents over the 1963–1978 period.[3] Moreover, Dougherty and Van Order (1982) reported evidence that, assuming that homeownership cost is overestimated in the CPI due to not taking into account tax and capital gains factors, about 15% to 25% of the increase in CPI in the 1968–1980 period could be spurious. As such, they concluded that the cost of housing was quite affordable in the United States relative to this lower index of consumer costs.

[2] Fisher (1930) and Fama and Schwert (1977) assumed that the real rate of return is constant, such that changes in nominal interest rates are a function of changes in inflation rates.

[3] Spellman (1981) proposed a model of home valuation based on the discounted value of expected future net rents (or revenues minus expenses). From this model the implied price–rent multiple can be determined from the capitialization rate for rent. He inferred the rapid rate of growth of housing prices from a 33% increase in the price–rent ratio over the sample period.

Other early studies of real estate as an inflation hedge conducted regression analyses similar in spirit to Fama and Schwert. For example, Rubens, Bond and Webb (1989) examined the inflation-hedging effectiveness of residential real estate, farmland, and business real estate in addition to corporate and government bonds and common stock over the 1960–1986 period. Based on regression analyses, they found that only residential real estate was a complete hedge against actual inflation shocks. Also, Treasury bills had some hedging ability, but other real and financial assets did not demonstrate any significant hedging effectiveness. The results changed to some degree using expected and unexpected changes in inflation, but residential real estate was the only consistent hedge against inflation over the sample period. Moreover, they found that by incorporating real estate in portfolios of assets, the risk per unit return was lowered and inflation hedging was improved. Another study by Bond and Seiler (1998) reported empirical evidence on the notion that residential real estate can decrease the variance of portfolio returns due to the fact that it is a significant hedge against expected and unexpected inflation whereas financial assets (*e.g.*, stocks and bonds) do not hedge against unexpected changes in inflation. Data for the 1969–1994 period on the percentage change in prices of existing homes per quarter, inflation rate (CPI), and a number of other explanatory variables that captured macroeconomic activity were incorporated in an added variable regression model. Strong goodness of fit of the regression models led the authors to conclude that residential real estate is an effective hedge against expected and unexpected inflation.

While not directly related to the present study, numerous studies have extended the analyses to the inflation hedging characteristics of commercial real estate. For example, Hartzell, Hekman and Miles (1989) found that commercial real estate returns were positively and significantly related to expected changes in inflation rates during the 1973–1983 period. Unexpected movements in inflation were not significantly related to real estate returns. The authors concluded that diversified commercial real estate portfolios can completely hedge against inflation risk.[4] More recently, Hamelink, Hoesli and MacGregor (1997) compared the hedging properties of bonds, commercial real estate, securitized real estate, and common stocks using U.S. and U.K. data for the 1975–1995 period. They distinguished between comovement of inflation and nominal returns on assets over time, or inflation hedging, and nominal returns that exceed the rate of inflation such that real returns are positive, or inflation protection. Rather then focusing on short-term comovement of asset returns and inflation, consistent with most investors' perspective, they sought to examine the

[4] For other early studies on real estate and its relationship to inflation, see Ibbotson and Fall (1979), Ibbotson and Siegel (1983), Brueggeman, Chen and Thibodeau (1984), Fogler, Granito and Smith (1985), and Sirmans and Sirmans (1987).

long-run hedging and protection performance of assets over time. In both the United States and United Kingdom, real estate and bonds were similar to one another in terms of providing long-run holding period protection against inflation, but were not superior to securitized real estate and stocks. Also, Chaudhry, Myer and Webb (1999) report cointegration tests using the Johansen method that reveal a long-run relationship between inflation and commercial real estate.[5]

The present paper argues that the long-run impact of inflation on homeowner equity can be better understood and examined by investigating the relationship between house prices and the prices of nonhousing goods and services instead of return series and inflation rates. There are two reasons for diverging from previously cited research. First, in order to estimate the return to owner-occupied housing, it is necessary to have the time series of actual cash flows for the implicit rent and for the repairs that the homeowner has incurred. However, because no actual time series of these cash flows are available, researchers have set cash flows for implicit rent and maintenance equal to what homeowners would receive or spend for equivalent rental properties. As noted by Shiller (1989, p. 319), "The owner-occupant of a home earns instead an implicit rent in the form of housing services, on which there is no market valuation. The best proxy for such implicit rents that we appear to have are rental indexes (computed from data on rental property)." Since the estimated time series of the returns to owner-occupied housing depends to a great extent on the underlying assumptions about the imputed values of rents and services performed by the owner, Crone (1995) has pointed out that it is difficult to estimate the long-run average rate of return on residential real estate. In the case of stocks and bonds, there is no potential estimation difficulty due to the fact that returns are calculated based on the actual dividend or interest payments, respectively. Given that the total return on housing cannot be accurately measured and that the total return on housing is fully reflected in housing prices, the use of housing prices avoids problems of estimating imputed rental income. Second, an extensive body of econometric literature has shown that differencing a time series leads to the loss of long-run information contained in the series (*e.g.*, Sargan 1964, Hendry and Mizon 1978, Granger and Joyeux 1980, and Juselius 1991). Since returns to owner-occupied housing are calculated by using the first difference of prices, it is likely that valuable long-run information contained in prices is lost. Finally, our paper differs from previously cited work on the impact of inflation on real estate values by excluding housing costs from the consumer price index

[5] For readers interested in related studies and further references on the subject of real estate (including both commercial and residential property) as an investment and inflation effects, see Webb (1990), Dokko *et al*. (1991), Wurtzebach, Mueller and Machi (1991), Froot (1995), Newell (1996), Liu, Hartzell and Hoesli (1997), Miles and Mahoney (1997), Chatrath and Liang (1998), and Blackley (1999).

employed in the regression analyses. In this way we mitigate potential bias in estimating the Fisher coefficient for real estate, which would otherwise tend to favor finding that real estate was an effective inflation hedge.

Theoretical Framework

An important feature of housing expenditure is that it constitutes both investment and consumption. In a survey of household motivations for purchasing houses, Case and Shiller (1988) found that the investment benefit was a major consideration for 44% to 64% of respondents in their decision to buy. Only 10% of respondents considered investment potential to be unimportant.

Because houses are both investment goods and consumption goods, there are two channels through which higher goods and services prices are transmitted to higher house prices. As a consumer good, inflation increases the construction costs of new houses through higher costs of building materials and construction wages. Higher construction costs of new houses result in higher new house prices. Since existing houses are close substitutes for new houses, higher new house prices increase the replacement costs of existing houses and, in turn, their prices.

Since a house is also an investment good, there is another channel of the relationship between house prices and goods and services prices. In a competitive market, the price of a house as an investment good is equal to the present value of future streams of actual or imputed net rents (*i.e.*, gross rental income minus maintenance and depreciation costs). In the absence of taxes on incomes and capital gains, the present value model for the stream of expected rents can be specified as

$$HP = PV = \sum_{k=1}^{n} \frac{E_t(R_{t+k})}{(1+r)^k}, \tag{1}$$

where PV denotes present value (equivalent to house price or HP), n is the life span of the house, $E_t(R_{t+k})$ is the net annual rent in period $t+k$ that is expected in period t, and r is the discount rate. Net annual rent is the gross rent minus depreciation, among other charges, and the accumulated depreciation charges at the end of the lifespan of the house are used to build another house on the land. Thus, the flow of net rent is permanent (*i.e.*, $n = \infty$). If rent and the discount rate are in real terms, then the present value or price is in real terms. Assuming that the annual rent is a constant, Equation (1) can be written as

$$HP = PV = \frac{R}{r}. \tag{2}$$

Using Fisher's (1930) proposition concerning the long-run relationship between real and nominal interest rates (*i.e.*, a 1% increase in expected inflation will increase nominal interest rates by 1% due to a constant real rate of interest), Equation (2) can be respecified in nominal terms to show the relationship between nominal house prices and goods and services prices adjusted for housing costs. Landlords, like lenders, seek to maintain the purchasing power of rental income in real terms and for this purpose incorporate expected inflation in rent agreements by relating the rent to a goods price index such as the consumer price index. Thus, Equation (2) can be written as

$$HP_t = PV_t = \frac{R\left(\frac{E_t(NH\_CPI_{t+1})}{NH\_CPI_b}\right)}{r}, \tag{3}$$

where $E_t(NH\_CPI_{t+1})$ is the expected nonhousing price index of goods and services for period $t+1$ based on all available information in period t, and NH_CPI_b is the nonhousing price index in the base period. Assuming that R and r are constants and that $NH\_CPI_b = 1$, and taking the log of both sides of Equation (3), we obtain

$$\ln HP_t = \alpha + \beta \ln E_t(NH\_CPI_t), \tag{4}$$

where the coefficient of the goods price index $\beta = 1$, and the constant term $\alpha = \ln R - \ln r$. Equation (4) suggests that in the absence of taxes, consistent with the so-called Fisher effect, we expect an inflation elasticity of unity for house prices with respect to goods and services prices adjusted for housing costs.

Tax considerations complicate the relationship between house prices and inflation. A landlord must pay income tax on rents as well as on capital gains when a property is sold, but deductions for depreciation and maintenance costs are allowed from rental incomes. By comparison, homeowners do not pay tax on imputed rent and normally are exempted from capital gains tax by living in the home for two of the previous five years. They are allowed to deduct mortgage interest payments from their income but are not allowed to deduct depreciation and maintenance expenses. Unfortunately, little data are available to examine the impact of taxes and exemptions on the return to housing or house prices. According to Darby (1975) and Carrington and Crouch (1987), the effects of all these taxes and exemptions should be reflected in the β coefficient. They show that if NIR_t, RIR_t, and INR_t are nominal interest rate, real interest rate, and inflation rate, respectively, and the tax rate is T, then the tax version of the Fisher relationship can be written as

$$NIR_t = (1 - T)^{-1} RIR_t^e + (1 - T)^{-1} INR_t^e. \tag{5}$$

Relevant to the present paper, Anari and Kolari (2001) have shown that the tax version of the Fisher relationship holds for the relationship between asset price and goods price indexes, such that the β coefficient in Equation (4) can be written as $\beta = (1-T)^{-1}$.

Econometric Methods

We use two econometric approaches for the investigation of the impact of nonhousing goods and services prices on house prices: (1) Pesaran and Shin's (1995) autoregressive distributed lag (ARDL) model and (2) recursive regressions. ARDL is employed to estimate the long-run relationship between house prices and nonhousing prices of goods and services, while recursive regressions are used to examine the variability of the relationship between house prices and nonhousing prices over time.

The ARDL model is a cointegration method for detecting the existence of long-run relationships between time series variables, as well as the subsequent estimation of their magnitude. The main advantage of the method is that it is applicable to time series data regardless of whether or not they are stationary. For this reason it is not necessary to pretest the series under investigation to determine the stationarity properties of the series.

The ARDL test for the existence of a long-term relationship between the variables under investigation is performed in two steps. First, a model in first differences of the variables is estimated; for instance, for house price and nonhousing consumer price indexes the underlying ARDL model is

$$\Delta HP_t = \mu + \sum_{k=1}^{n} B_k \Delta HP_{t-k} + \sum_{k=1}^{n} C_k \Delta NH\_CPI_{t-k}, \tag{6}$$

where B_k and C_k are the coefficients of lagged ΔHP and ΔNH_CPI to be estimated. Second, a variable selection test is conducted by computing an F-statistic to determine whether additional lags for HP and NH_CPI result in significant coefficients, as shown in the following error-correction model:

$$\begin{aligned} \Delta HP_t = \mu &+ \sum_{k=1}^{n} B_k \Delta HP_{t-k} + \sum_{k=1}^{n} C_k \Delta NH\_CPI_{t-k} \\ &+ \delta_1 HP_{t-1} + \delta_2 NH\_CPI_{t-1}. \end{aligned} \tag{7}$$

The null and alternate hypotheses are defined as $\delta_1 = \delta_2 = 0$ (no long-run relationship exists) and $\delta_1 \neq 0$ and $\delta_2 \neq 0$, respectively. Given the asymptotic distribution of this F-statistic is nonstandard, the critical values for the test statistic are given by Pesaran and Pesaran (1997).

Once the existence of a long-run relationship between a housing price index and the nonhousing price index is determined, the relationship is estimated using the error-correction model represented by Equation (7) based on different model selection criteria, including R-squared, Akaike (1973) information criterion (AIC), Schwarz (1978) Bayesian criterion (SBC), and Hannan-Quinn (1979) criterion (HQC).[6]

Like most regression methods, the ARDL procedure is a constant coefficient method and calculates a single coefficient for the impact of nonhousing goods and services prices on house prices, which is assumed to be stable throughout the sample period. However, as Lucas (1976) has argued, estimated econometric parameters are unlikely to be stable since market participants change the process of forming expectations in response to changes in policy regimes. It is possible that the impact of nonhousing goods and services prices on house prices has been changing over the sample period in response to changes in the financial and economic environment. To examine the stability of the impact of inflation on house prices over time, we apply recursive regressions to Equation (4). In this respect, Equation (4) is estimated repeatedly such that β_t and the variance of the disturbance term (Φ_t^2) are allowed to vary with t. Recursive regressions implicitly assume that market participants update their expected inflation rate (or expected goods and services prices) in each period t using an information set containing the history of goods price and house price indexes. In each round of estimation, or for each period t, the estimator uses a larger subset of the sample data and provides updated estimates of β_t and Φ_t^2. This process is repeated until all the n observations have been used, which provides $n - k$ estimates of β_t, where k is the number of variables ($i.e.$, $k = 1$). The recursive least squares can be solved using ordinary least squares (OLS) or instrumental-variable estimation methods. For a formal treatment of recursive regressions, see the Appendix.

Recursive regressions provide a test to determine whether or not the variation in a variable has been ordinary in the context of its past history. To investigate the variability of the impact of nonhousing prices of goods and services on

[6] The Akaike information criterion (AIC) is given as: $AIC = n \log rss + 2k$, where n is the number of observations, k is the number of regressors, and rss is the residual sum of squares. The Schwarz Bayesian criterion (SBC) is given as $SBC = n \log rss + k \log n$. The Hannan–Quinn criterion (HQC) is given as $HQC = n \log rss + 4k \log(\log n)$.

house prices over different parts of the sample period, the recursive residuals are examined. Recursive residuals outside the standard error bands suggest instability (*i.e.*, in the Fisher coefficient for house prices).

Empirical Results

The data set consists of monthly time series of home prices for new homes and existing homes as well as the consumer price index minus housing costs from January 1968 to June 2000.[7] All variables are transformed into logarithms and denoted as *PEH*, *PNH*, and *NH_CPI* for the price of existing homes, price of new homes, and the nonhousing price index of goods and services (as reported by the U.S. Bureau of Labor Statistics), respectively. Note that our *NH_CPI* index excludes housing costs to avoid the effect of house prices on *CPI* as discussed earlier.

To determine the lag length for the ARDL model in Equation (6), following Pesaran and Pesaran (1997), we specified a bivariate vector autoregressive model of house price and nonhousing price indexes and applied the Akaike information criterion (AIC) and the Schwarz Bayesian criterion (SBC) for lag order selection. Table 1 shows the results of AIC and SBC criteria for the determination of the lag order in the underlying test equations. For the pairs of nonhousing prices of goods and services and home prices, the tests suggest one monthly lag for both new and existing home prices.

Table 2 reports the results of the ARDL models. As shown in Panel A, all pairs of house price and nonhousing price indexes pass the test for the existence of a long-run relationship between the two variables.

Panel B of Table 2 reports the estimated long-run relationships between each pair of house price and nonhousing goods and services price indexes using the aforementioned criteria for model selection (*i.e.*, R-squared, Akaike information criterion, Schwarz Bayesian criterion, and Hannan–Quinn criterion). As shown there, the results appear to be robust to the choice of criteria for model selection. The estimated Fisher coefficients are 1.08 for existing homes and 1.26 for new homes. Because house price and nonhousing price indexes are in logarithms, the estimated Fisher coefficients are elasticities of house prices with respect to a 1% increase in nonhousing goods and services prices.

[7] The time series for existing home prices are from the National Association of Realtors, new home prices are from the Bureau of the Census, U.S. Department of Commerce, and the consumer price index minus housing costs is from the U.S. Bureau of Labor Statistics.

Table 1 ■ Results of Akaike and Schwarz tests for the determination of lag order in the autoregressive distributed lag (ARDL) models of home prices and nonhousing prices of goods and services.

Lag Selection Criterion and Bivariate Models	Lag Order			
	1	2	3	4
Akaike Information Criterion				
Model of existing home prices and nonhousing prices	2,717.3*	2,754.6	2,760.5	2,762.2
Model of new home prices and nonhousing prices	2,476.0*	2,582.2	2,593.7	2,594.0
Schwarz Bayesian Criterion				
Model of existing home prices and nonhousing prices	2,705.5*	2,734.9	2,732.9	2,726.8
Model of new home prices and nonhousing prices	2,464.1*	2,562.4	2,565.9	2,558.3

The Akaike information criterion is given as $AIC = n \log rss + 2k$, where k is the number of regressors, rss is the residual sum of squares, and n is the number of observations. For the pairs of home prices and nonhousing prices of goods and services, the *AIC* tests suggest one monthly lag. Test statistics for more than four lags were increasing and, therefore, are not reported. The Schwarz Bayesian criterion is given as $SBC = n \log rss + k \log n$, where all terms are defined above. For the pairs of home prices and nonhousing prices of goods and services, the tests suggest one monthly lag. Test statistics for more than four lags were increasing and, therefore, are not reported.

Statistical tests indicate that the estimated Fisher coefficients are significantly greater than one. As such, the results are consistent with Darby's tax version of the Fisher effect discussed previously. While tax differences between existing and new homes is one explanation for the differential between existing and new home coefficient estimates, another possible explanation is that these two types of property are different from one another. Evidence by Do and Grudnitski (1993) and others has revealed that the value of residential property declines with age. This age effect could also affect the rate of growth of existing home prices relative to new home prices over time.

Figure 1 shows the estimates of time-varying Fisher coefficients for existing and new home prices generated by recursive regressions for the period January 1973 to June 2000 (see Table 3 for estimated coefficients). We dropped coefficient estimates from January 1968 to December 1973 due to inadequate sample sizes for statistical purposes. After 1973 the Fisher coefficient estimates for both existing and new homes are fairly stable and lie in the range of 1.19 to 1.42. Like the results from the ARDL method, the estimated elasticities for new house prices tend to be somewhat larger than the elasticities for existing

Table 2 ■ Autoregressive distributed lag (ARDL) model analyses.[a]

Panel A: Results of ARDL tests for the existence of a long-run relationship between house prices and nonhousing prices of goods and services

	Existing House Prices and Nonhousing Prices	New House Prices and Nonhousing Prices
Test statistics	10.08†	3.64+

Panel B: Long-run elasticity of house prices in response to nonhousing prices of goods and services

Model Selection Criteria	Existing House Prices and Nonhousing Prices	New House Prices and Nonhousing Prices
R^2	1.08*(0.11)	1.26*(0.04)
Akaike	1.08*(0.11)	1.26*(0.04)
Schwarz	1.08*(0.11)	1.26*(0.04)
Hannan–Quinn	1.08*(0.11)	1.26*(0.04)

[a] In Panel A the symbol † means significantly different from zero at the 5% significance level and the symbol + corresponds to a 10% significance level. In Panel B, figures in parentheses are standard errors. An asterisk(*) denotes rejection of the null hypothesis that the elasticity of home prices with respect to nonhousing prices of goods and services is equal to one and acceptance of the alternative hypothesis that it is more than one.

houses. However, unlike the ARDL results, the difference in Fisher coefficient estimates for existing versus new home prices is relatively small.

To compare the results from the ARDL and recursive regressions, an anonymous referee suggested using a nonnested J test proposed by Davidson and MacKinnon (1981). For existing home prices, we test the null hypothesis of a Fisher coefficient of 1.08 from the ARDL model against an alternative hypothesis of a Fisher coefficient of 1.21 from the last estimate of the recursive regression model, that is, H_0: ARDL versus H_1: recursive. The ARDL model rejected the estimated coefficient from the recursive regression model which means that recursive regression does not provide additional information. We then tested H_0: recursive versus H_1: ARDL. In this test the recursive model did not reject the ARDL model, which means that ARDL provides additional information. We conclude that the ARDL method is superior to the recursive method for the estimation of Fisher coefficients for existing house prices. We repeated the same procedure for new home prices and found that each test accepted the estimated coefficients from the other test, which means that the test is inconclusive between the two methods.

To investigate whether the estimated elasticities of house prices in response to nonhousing prices of goods and services have been stable over the sample

Figure 1 ■ Time-varying Fisher coefficients for existing and new home prices: January 1973 to June 2000.

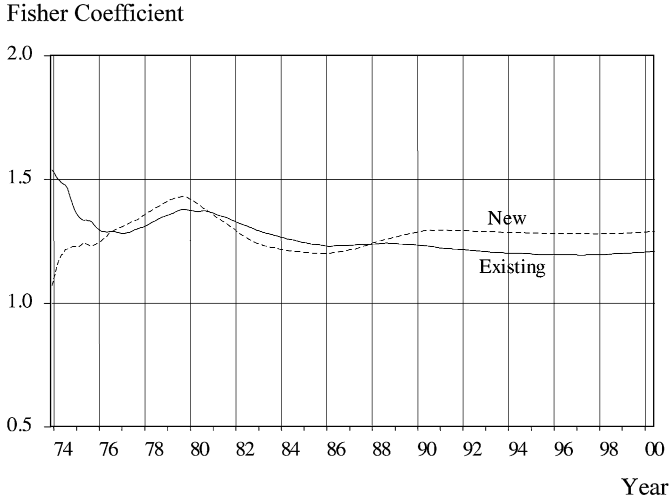
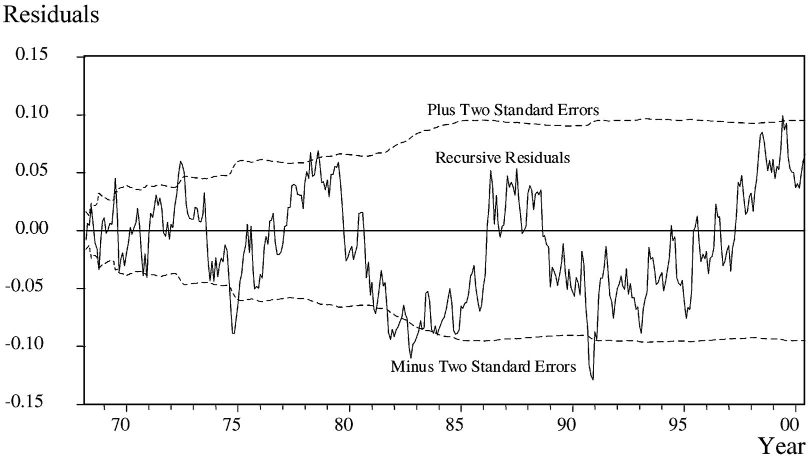


Table 3 ■ Time-varying elasticity of house prices with respect to nonhousing prices of goods and services using recursive regression analyses: 1973–2000.

Year/Month	Existing House Prices	New House Prices	Year/Month	Existing House Prices	New House Prices
1973:12	1.54*(0.04)	1.07*(0.12)	1987:12	1.24*(0.01)	1.24*(0.01)
1974:12	1.38*(0.03)	1.23*(0.07)	1988:12	1.24*(0.01)	1.26*(0.01)
1975:12	1.30*(0.02)	1.24*(0.05)	1989:12	1.24*(0.01)	1.29*(0.01)
1976:12	1.28*(0.02)	1.31*(0.04)	1990:12	1.22*(0.01)	1.29*(0.01)
1977:12	1.31*(0.01)	1.35*(0.03)	1991:12	1.22*(0.01)	1.29*(0.01)
1978:12	1.35*(0.01)	1.41*(0.03)	1992:12	1.21*(0.01)	1.29*(0.01)
1979:12	1.38*(0.01)	1.42*(0.02)	1993:12	1.20*(0.01)	1.29*(0.01)
1980:12	1.37*(0.01)	1.36*(0.02)	1994:12	1.20*(0.01)	1.28*(0.01)
1981:12	1.33*(0.01)	1.30*(0.02)	1995:12	1.20*(0.01)	1.28*(0.01)
1982:12	1.30*(0.01)	1.24*(0.02)	1996:12	1.19*(0.01)	1.28*(0.01)
1983:12	1.27*(0.01)	1.22*(0.02)	1997:12	1.20*(0.01)	1.28*(0.01)
1984:12	1.25*(0.01)	1.21*(0.02)	1998:12	1.20*(0.01)	1.28*(0.01)
1985:12	1.23*(0.01)	1.20*(0.01)	1999:12	1.21*(0.01)	1.29*(0.01)
1986:12	1.23*(0.01)	1.21*(0.01)	2000:06	1.21*(0.01)	1.29*(0.01)

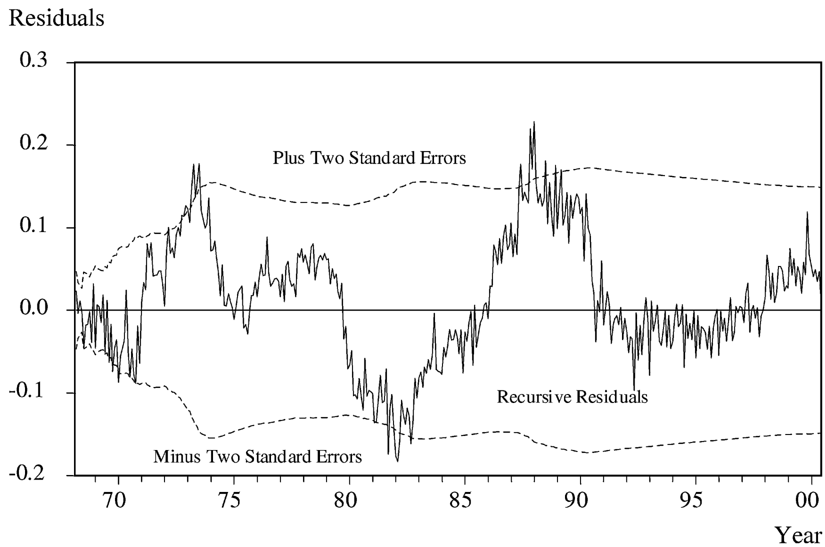
Figures in parentheses are standard errors. An asterisk (*) denotes rejection of the null hypothesis that the elasticity of home prices in response to nonhousing prices of goods and services is equal to one and acceptance of the alternative hypothesis that it is more than one.

Figure 2 ■ Recursive residuals for existing home prices: January 1968 to June 2000.



period, we plotted the recursive residuals in Figures 2 and 3 for existing and new house prices, respectively. In recursive least squares the Fisher equation is estimated recursively and at each step the last estimate of the Fisher coefficient is used to predict the next value of the nominal interest rate. Recursive residuals are the difference between the one-step-ahead forecast of nominal interest rate

Figure 3 ■ Recursive residuals for new home prices: January 1968 to June 2000.



and the actual interest rate. The forecast variances are estimated using these recursive residuals. Residuals outside plus and minus two standard error bands (5% significance level) suggest instability in the estimated Fisher's coefficients. As these figures show, the residuals for existing and new home prices stayed within this standard error band for the most part during the 1968–2000 sample period.

Summary and Conclusions

The present paper examined the long-run impact of inflation on homeowner equity by investigating the relationship between house prices and nonhousing goods and services prices, rather than return series and inflation rates as in previous empirical studies on the effects of inflation on real estate. This approach has the advantages that the total return on housing is fully reflected in housing prices and valuable long-run information can be captured by using prices. Also, unlike previous related studies, we excluded housing costs from the consumer price index to avoid potential bias in estimating the relationship between house prices and goods prices. Monthly data series were used for existing and new house prices as well as the consumer price index excluding housing costs for the 1968–2000 period. Both autoregressive distributed lag (ARDL) models and recursive regressions were employed to investigate the impact of nonhousing goods and services prices on house prices. The empirical results using the ARDL models indicated that house prices and nonhousing prices are cointegrated, such that a long-run relationship exists between these data series over the sample period.

The ARDL models generate Fisher coefficient estimates in the range of 1.08 for existing homes to 1.26 for new homes, both of which are significantly greater than 1 as predicted by Darby's tax version of the Fisher equation. Recursive regressions indicate that time-varying Fisher coefficients are in the range of 1.19 to 1.42 during the 1974–2000 period. These and other results from analyses of recursive residuals lead to the conclusion that the estimated Fisher elasticities of house prices with respect to nonhousing goods and services prices are stable over the sample period and exceed 1. An important implication of these findings is that house prices are a stable inflation hedge over time.

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Appendix: Recursive Regressions

The recursive least squares method provides the estimates of a linear regression model recursively, using ever larger subsets of the sample data. Denoting Y as the dependent variable, X as the independent variable, and u as the disturbance term, the underlying econometric model is given by

$$Y_t = \beta_t X_t + u_t \quad t = 1, 2, \ldots, n,$$

where the coefficient β_t and the variance of the disturbance term, Φ_t^2, are allowed to vary with time. This equation can be solved by the ordinary least squares method or by the instrumental variables method. The OLS option gives the following recursive coefficients:

$$\beta_t = (X_r' X_r)^{-1} X_r' Y_r, \quad r = k + 1, k + 2, \ldots, n,$$

where $X_r = (x_1, x_2, \ldots, x_r)'$, $Y_r = (y_1, y_2, \ldots, y_r)'$ and n is the number of observations (see Brown, Durbin and Evans 1975).

Recursive standard errors for the OLS method are given by

$$\Phi_r^2 = (Y_r - X_r \beta_t)'(Y_r - X_r \beta_t)/(r - k), \quad r = k + 1, k + 2, \ldots, n.$$

Recursive coefficients using the instrumental variables method are given by

$$\beta_t = (X_r' Z_r X_r)^{-1} X_r' Z_r Y_r, \quad r = k + 1, k + 2, \ldots, n$$

where $Z_r = S_r (S_r' S_r)^{-1} S_r N$, $S_r = (s_1, s_2, \ldots, s_r)'$, and s_t, $t = 1, 2, \ldots, n$ are the $sx1$ vector of observations on the s instrumental variables. Recursive standard errors for the instrumental variables method are estimated as shown above where β_t are recursive coefficients.