HOUSE PRICE DYNAMICS: AN INTERNATIONAL EMPIRICAL PERSPECTIVE [1]

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Abstract

The paper compares the dynamics of housing prices in fifteen OECD countries. The data reveal a remarkable degree of similarity across countries and suggest rich dynamics for the first-differenced real annual house prices, with a significant structure of autocorrelation. We estimate a highly significant first-order autocorrelation coefficient at around .45 and obtain

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evidence of negative autocorrelation for lags up to five years. The contemporaneous GDP growth rate and the rate of change in real rate of interest are very significant along with the first-order lag, whose coefficient remains at .45. Lagged GDP growth and the real rate of interest exhibit significant predictive power. While house price dynamics across different open economies seem to be interdependent on descriptive grounds, our econometric results suggest weak evidence to that effect.

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1 Introduction

Almost all systematic evidence on house price dynamics comes from studies of individual countries or metropolitan housing markets. The comprehensive international asset price study by Cutler, Poterba and Summers (1991), which is based on returns for shares and other assets for a wide range of countries, only reports house prices for the United States. Poterba (1991) looks at house prices for four countries but not within a common statistical model. An exception is a recent BIS study by Kennedy and Andersen (1994) which analyzes house price data from 15 OECD countries, but focusses on the interaction between house prices and savings rather than on house price dynamics per se. In this paper we will draw on their data set covering the period 1970-1992. [2] Large swings in prices of owner-occupied homes are documented in Figure 1 for fifteen OECD countries. In nearly all these countries we see pronounced cycles with several consecutive years of rising prices followed by slumps. In fact it is in most cases quite easy to identify peaks and troughs. In particular the latter half of the 1980's witnessed housing price booms in almost all countries; see Table 1.

House price dynamics attract attention for several reasons. Owner-occupied homes comprise a major part of private-sector wealth. In Sweden, for example, housing wealth as a fraction of household net wealth has fluctuated between 50 and 75 per cent during the post-war period; see Berg (1988). Hence, house prices may have a major impact on the distribution of economic well-being and they should also be important in explaining household saving and consumption. From a theoretical point of view these relationships are not all that clear, since increased asset prices also imply increased rental cost of owner-occupied housing, which tends to offset the impact on real wealth using an appropriate cost-of-living index. Calculations by Miles (1993) in a life-cycle framework suggest small net effects from housing wealth on aggregate consumption. Nevertheless several empirical studies for various countries find large and statistically significant effects; Brodin and Nymoen (1992) for Norway, Berg and Bergström (1995) for Sweden, Bayoumi (1993) for U.K., Koskela, Loikkanen and Virén (1992) for Finland and others. One possible explanation is that large effects derive at least partly from the effects on household balance sheet composition via debt-equity ratios rather than from net wealth per se. With falling prices home equity is eroded and households find it more difficult to finance the purchase of a new and larger house, thereby depressing demand and exerting further downward pressure on house prices. Another mechanism may be related to redistribution between households at different stages of their housing career.

[2] The quality of the data, which derive from a variety of sources, differ widely from country to country. Quality adjusted nationwide price indices exist only for a few countries (Denmark, Finland, Sweden, United Kingdom). See our section on data sources below and Kennedy and Andersen (1994), Annex 1, pp. 52–56, for full details on the data.

Falling prices redistribute income away from households about to leave the owner-occupancy market towards new entrants.

The dynamics of house prices and the returns to holding housing are also interesting because they throw some light on the operation of the housing market. Several studies, mainly based on U.S. data, show that rates of change of house prices are quite strongly autocorrelated, more so than most other assets. This apparently stands in stark contrast with the notion of an efficient housing market; see e.g. Case and Shiller (1988, 1990) and Meese and Wallace (1994). The superficial impression one gets from the data presented in Figure 1 is also one of rather strong autocorrelation in most countries. This suggests either that housing markets are influenced by bubbles and fads or that the (no autocorrelation) asset market model used as an implicit yardstick is too simplistic.

It may be a useful background to briefly recapitulate what we take to be the standard model of house price dynamics based on the asset-market approach of Poterba (1984). In particular we may focus on the implications of this model for the time series patterns of house prices. This forms the basis for a brief review of some of the existing empirical literature. [3] The asset market approach is based on the notion of arbitrage in perfect capital markets. In such case the risk-adjusted rate of return on housing must equal the market interest rate. Housing return is the sum of the value of rental services and capital gains. The value of rental services is assumed to be determined in a perfect market for housing services, equating demand with supply which at each point in time is taken as predetermined. The dynamics of supply comes from a Tobin's-q type investment function. Given today's house prices this determines tomorrow's supply and hence tomorrow's rents, and thereby, via the arbitrage condition, the rate of capital gains. In the long run, absent any demand or supply shocks, the price level (or price trend) will be just sufficient to ensure that investment keeps pace with the trendwise rate of change of demand.

According to this model a sudden demand shock (e.g. a decrease following from a tax reform) causes market rents to fall to maintain equilibrium. This will lead investment to drop which will induce expectations of future rent increases from the new lower level as the housing stock adjusts downwards. The general implication is that any unexpected adverse shock will have an immediate negative impact on house prices. Further along the adjustment path back towards long run equilibrium there will be continuous price increases. Therefore, prices will have a mean-reverting tendency. Actual price movements will reflect a combination of shocks and adjustment mechanisms. Loosely speaking, if the sample path is dominated by

[3] Our analysis refrains from accounting for the theoretical and empirical possibility that expectations of capital gains may be so strong as to make housing demand slope upwards. On this, see Dusansky and Wilson (1993).

adjustment towards equilibrium rather than shocks away from it, then price changes will typically be positively autocorrelated, in particular at high frequencies. At lower frequencies autocorrelations should be close to zero if the adjustment process is reasonably fast. On the other hand, if the sample path is dominated by frequent white noise shocks, then price changes would also be approximately white noise.

Obviously one must conclude that this model is consistent with a wide variety of autocorrelation patterns in house prices, depending on the nature of the shocks hitting the housing market and the rate of speed of adjustment. On the other hand it is not consistent with arbitrary correlation patterns of rates of return. In fact, the standard version of the model which takes the rate of return requirement as constant immediately implies that realized returns (capital gains plus implicit rents) are constant or white noise. Of course, as is well known from tests of the efficient markets hypothesis in finance, it is in general not possible to distinguish between inefficiency and time-varying return requirements.

One implication of the standard model which may appear counterfactual is that it implies a smooth adjustment without oscillations towards equilibrium. Mechanisms that would give rise to cyclical adjustment include borrowing constraints [Stein (1995)] extrapolative expectations [Muellbauer and Murphy (1992)] and transactions costs [Ekman and Englund (1995)].

2 Selective Review of the Evidence

The most widely quoted studies on housing autocorrelation patterns are those by Case and Shiller (1988, 1990). Using pooled time-series cross-section data for four U.S. cities from the first quarter of 1970 up until the second quarter of 1986, Case and Shiller (1990) report strong evidence of positive autocorrelation at short lags and weaker evidence of negative autocorrelation at longer lags. Regressing the log price change between two successive years against four lags of price changes yields the following regression coefficients in lag order (absolute t-statistics in parenthesis): 0.38(3.3), -0.10(0.9), -1.12(1.1), -0.03(0.3). This conforms with what Cutler *et al.* (1991) find (using quarterly data) to be the general mean-reverting pattern of asset returns, holding for a wide range of assets. As pointed out above it is not necessarily inconsistent with an informationally efficient housing market with rational expectations. Similar findings are reported for Swedish metropolitan areas by Hort (1995).

The above results are cast in terms of house prices. Meese and Wallace (1994) in their study of the San Francisco area find similar patterns in annual housing returns, calculated as the sum of price increase and implicit rent. In separate regressions for 16 different municipalities they obtain consistently positive first-order coefficients ranging from .15 to .64. Based on a regression with three lags they can reject the hypothesis that all three lag coefficients are jointly zero at the five per cent level for all communities but one, and at the one per cent level for all communities but two.

Having found that house prices and housing returns are predictable from their own past values it is natural to ask what other factors might have predictive power. This is also considered by Case and Shiller (1990) who find that the level of construction costs and the percentage change of the adult population are the only two extra factors that perform significantly in a forecasting equation (both with positive sign). The importance of demographics is closely related to the controversial claim by Mankiw and Weil (1989) about the importance of demographic factors.[4] Similar results are obtained for the predictability of excess returns on housing over interest-bearing assets. Here income growth and a measure of mortgage costs are also significant. The studies quoted above throw some light on the "efficiency" of housing markets, but they are uninformative about what types of shocks appear to drive house prices and housing returns away from predicted values. These questions have been approached in various ways. Chinloy (1992) estimates a factor model to explain the housing returns. He finds changes in inflationary expectations to be the main macro factor to have an impact on housing returns. Hendershott and Abraham (1993; 1994) using pooled cross-section data of U.S. metropolitan areas find changes in the following factors to have an impact on house price changes: construction costs, employment growth, income growth, and the real after tax interest rate.

Hendershott and Abraham (1994) also find evidence of cyclical behavior. They estimate a model specification that is similar to an error correction model, which includes lagged house price changes among the explanatory variables. They find relatively slow adjustment towards equilibrium combined with a significantly positive impact from lagged price changes. These results imply a cyclical adjustment path, contradicting the simple asset-market model. This pattern is also in line with results in Muellbauer and Murphy (1992) where cycles are interpreted as coming from expectations formation. Further some studies, notably Koskela *et al.* (1992) and Muellbauer and Murphy (1992) highlight the empirical importance of household indebtedness and borrowing constraints.

[4]See *Regional Science and Urban Economics* (1991), vol. 31, No. 4 for a set of critical comments on the Mankiw and Weil study. Heiborn (1994) studies the impact of demographics on Swedish house prices.

3 International Patterns in House Prices

3.1 Are House Prices Predictable?

The first issue is whether the autoregression pattern found in U.S. and Swedish data also holds internationally. Table 2 reports results with regressions similar to those of Case and Shiller, *op. cit.*. It is based on straightforward pooling of the data and estimated with OLS. The results are quite similar to those of Case and Shiller. The first order coefficients are significantly higher than their value .312. The difference may partly reflect autocorrelated measurement errors, which they correct for. Further we also find signs of negative autocorrelation at lags up to the fourth order. Estimates of separate AR(1) equations for each country confirm that first- order autoregression is a general feature of house price changes. All estimates are positive ranging from .23 to .74, and half are significant at the one percent level. Estimates of separate AR(2) equations confirm the prevalence of positive autocorrelation at the first order and negative autocorrelation at the second order.

Next we ask whether individual country effects are significant with the first-differenced data. Allowing for fixed effects (FE) produces minuscule differences in the estimated coefficients. A variety of diagnostics reported in Table 2 suggest that, and fixed effects are not significant. The Lagrange multiplier test indicates that a random effect (RE) is significant (at 2%) for the specifications reported on Columns 1–3. However, a model with fixed effects is more appropriate for testing for individual effects here. Therefore, we conclude that after house prices have been differenced once house price dynamics display a remarkable degree of homogeneity. This implies that, at least as far as house price dynamics are concerned, housing markets in different countries are "all the same".

Next we examine whether other variables apart from lagged house prices have predictive power in explaining future house price changes. In addition to one-period lagged house price changes three variables are included all of which are also lagged one period: the rate of growth of real GDP, the rate of change of the real interest rate (ex post and pre-tax) and the rate of change of population in the house-buying ages, 20-30. The latter variable did not perform well in the regressions and is thus not reported.

The regression coefficients reported in column 1, Table 3 were obtained for the full sample by OLS. In addition to lagged price changes, lagged GDP growth has strong predictive power. One percentage faster GDP growth this year gives .77 percentage faster house price growth tomorrow. This may again be compared with Case and Shiller, who find that lagged income growth is insignificant but that demographics, measured as the rate of change of the population between 25 and 44, is significant with a positive sign. A number of diagnostics for regressions run with fixed effects and random effects again confirm that first-differencing removes virtually all differences in house price dynamics across countries in our sample.

It is interesting to contrast with tests based on the importance of individual effects by comparing regression estimates from regressions conducted separately for each of the fifteen countries in our sample. Degrees of freedom considerations suggest that regressions like those reported in Column 1, Table 3, of yearly house price changes against the lagged values of the GDP growth rate, the real rate of interest, and the own lagged value of the dependent variable may well serve as a benchmark case. The results we have obtained suggest the following. Belgium, Italy, the Netherlands and Sweden give overall significant regressions, with a highly significant coefficient for the lagged value of the house price change, with coefficients ranging from .879 for Belgium to .462 for Italy, and R^2 ranging from .266 to .674. For four other countries, that is, Finland, France, Germany and Japan, the estimated coefficients are in the range of .311 to .373. The coefficient of the GDP growth rate are significant for Finland, Ireland, Norway and the United Kingdom, ranging from 1.72 to 1.54. A χ^2 test rejects the hypothesis that the coefficients are equal: twice the difference in the llf's is 123.28, which exceeds the χ^2 value for 56 degrees of freedom at conventional levels of significance. While admittedly, a case for heterogeneity may be made, we remain quite impressed by the fit obtained with the pooled data, and continue to believe that the differences are fewer than the similarities across countries.

3.2 What Drives House Prices?

The results in the previous subsection confirm that house price changes appear to be quite predictable at a year's horizon, but they do not address the question of what types of shocks are important in explaining house price movements. In columns 2–5 of Table 3 are results from pooled regressions with explanatory variables the contemporaneous changes in the same set of factors as those in the regression report on column 1. All specifications except the first, reported in Column 2, include one-year lagged house price changes. The third, reported in Column 4, also includes fixed effects (country intercept dummies), which all turn out to be insignificant. The fourth, reported in Column 5, includes fixed effects and a linear time trend. When, year dummies are included only 1986 and 1988 are significant with coefficients of +5.3% and +5.8% respectively (1989 is the base year). Again the GDP growth variable turns out to be strongly significant, varying in magnitude between 1.15 and 1.22. The coefficient for the real interest rate has the expected negative sign, and is very significant in all specifications, that is regardless of whether or not autocorrelation is accounted for. Demographics do not appear to matter at all and are not reported here. We have also run separate regressions for each country using the Cochrane-Orcutt procedure to take account of autocorrelation. In those regressions the income coefficient is generally positive (except for Italy) and significantly so for eleven countries. The interest rate coefficient is negative and significant (at the ten per cent level) in six countries and positive and significant in one (Italy).

So far we have attempted to explain house prices by major macroeconomic variables. Common sense as well as some econometric studies suggest that policy changes may have had a major impact. The 1980's was a decade of major reforms in two areas: taxes and credit markets. The tax reforms generally reduced previously more favourable tax treatment of homeownership, which should be expected to have an adverse effect on house prices. Deregulating credit markets removed previous obstacles to borrowing for some households and should be expected to boost house prices. Of course, banking crises followed in several countries with opposite effects but that is another story. To investigate such policy effects in a simple manner we defined two sets of dummy variables indicating the years when major reforms became effective. Based on a brief survey by Christ (1994) the following tax reform years were identified: Denmark 1987, Norway 1988, Sweden 1983 and 1991, Holland 1990, Ireland 1990, Germany 1987, and the United States 1986. Obviously the exact dating may be somewhat controversial, in particular because announcement dates should be at least as important as when the reform went into effect. We tried to capture this in a very crude way by extending the dummy variable to include the year before the reform. Most of them had the expected negative sign but only the Danish reform of 1987 was significant at the five percent level, with a point estimate of -16%. The U.S. reform of 1986 was marginally significant with a t-value of 1.88 and a coefficient of - 13%. We also tried the same procedure for the Scandinavian and British deregulations of the credit market (Sweden 1985, Norway 1985, Finland 1986, U.K. 1981-83), but without any success. We do not report these regressions here.

3.3 Is There an International House Price Cycle?

In Englund and Ioannides (1993) we discuss evidence, which may be supported by the plots in Figure 1, that the intertemporal variation in house prices displays some degree of synchronization across countries whose financial markets are linked. We develop in that paper a model of an open economy, where foreign investment is allowed in all assets of the economy, that is, physical capital, land and housing. Our theoretical results do not yield unambiguous predictions about the effect of changes in the world interest rate upon domestic house prices and land prices. Whereas that particular model does not allow for separate domestic and world interest rates, one would expect the US interest rate to play the role of a world interest rate.

In an effort to examine possible interdependence of house price dynamics across the fifteen economies for which data are available in our sample, we performed a number of econometric experiments. We estimated the first- difference in log house price of each country as a function of its own lagged value, contemporaneous growth rate in GDP, the rate of change in the country's own and in the US real interest rate. We performed these estimations separately as well as jointly by means of generalized least squares (GLS).

Our GLS regressions gave rather mixed results. When judged in terms of R2adj. the fits are comparable across the two sets of regressions. When judged in terms of sums of squared residuals, the GLS regressions give better fits. For those countries for which the GLS regressions perform well, there is no apparent reason to prefer them over the separate OLS regressions with the same explanatory variables. We note that, in general, the coefficient for the rate of change in US real interest rate is negative, when it is significant, with numerical values being similar to those of the domestic interest rate. Inclusion of the US interest rate variable brings about little improvement in fit.

The GLS regressions produce an estimated variance-covariance matrix where the estimated covariance with the US equation is typically positive, whereas the covariances between other countries are typically small. Further analysis of correlations among the residuals of separate OLS regressions, obtained by regressing those residuals against one another in a variety of combinations, did not yield a general pattern of significant dependence among the majority of residuals. We obtain our best results when we work with residuals from each country's regressions against its own two lagged values for the first-differenced data. Regressions of residuals against those for the US are most significant for Denmark (2.14), the United Kingdom (1.83), Norway (1.75), Canada 1.72), Switzerland (1.47), and Germany (-1.41), where the $t-$ statistic is given in parentheses. However, only Denmark gives a regression coefficient which is significant at a conventional significance level (5%). When the own GNP growth rate and the change in the US interest rate are included in the regressions that produce residuals, the US residual remains significant only for Denmark (1.82), the United Kingdom (1.70), and Norway (1.59). In view of the paucity of the data, where only 18 years of data may be used for these regressions, we may not derive a firm conclusion on the existence of an international housing cycle.

4 Conclusion

The paper compares the dynamics of housing prices in fifteen OECD countries. The data reveal a remarkable degree of similarity across countries and suggest rich dynamics for the first-differenced real house prices, with a significant structure of autocorrelation. We estimate a highly significant first-order autocorrelation coefficient at around .45 and obtain signs of negative autocorrelation for lags up to up to the fifth order. These results imply oscillatory behavior for house prices around a trend.

When instead of own lagged values we include in the regressions the contemporaneous GDP growth rate and the rate of change in real rate of interest, they are are very significant along with the first-order lag, whose coefficient remains at .45. Lagged GDP growth and the real rate of interest exhibit significant predictive power. While house price dynamics across different countries seem, on descriptive grounds, to be interdependent, our econometric results suggest weak evidence to support an international house price cycle. Future research must restrict attention to groups of countries for which such evidence is strongest and more data are available.

5 Data Sources

Demographics: United Nations Population Yearbook. GDP: IMF Financial Statistics. CPI: IMF Financial Statistics.

The house price indices data come the Bank of International Settlements. Detailed documentation may be found in Kennedy and Andersen (1994), 54–56. We summarize here some highlights of the data as reported in *ibid.*.

`Australia:` weighted average index of prices for all capital cities and other areas; obtained from quarterly national census of home loan approvals, available annually. Commonwealth Bank of Australia, Housing Industry Association and Real Estate Institute of Australia.

`Belgium:` index based on annual transactions reports on small and medium sized dwellings from entire country, with outliers excluded, available annually. ANHYP S.A., Brussels.

`Canada:` average annual transactions prices reported by multiple listing services for entire country, covering 70% of all transactions. Bank of Canada, MLS Ottawa.

`Denmark:` average value of single-family houses, including only arms-length sales, available annually. Danmarks Statistik.

`Finland:` average price for apartments and *terraced* homes, obtained per square meter, as recorded by realtors (including 30% of all transactions), weighted by region, available quarterly. Bank of Finland.

`France:` index based on BIS's own estimate based on annual values for the Paris region, adjusted by four-year survey for entire country. INSEE and Chambre Interdépartementale des Notaires de Paris.

`Germany:` Transactions prices per square meter, obtained by realtors from the four largest cities, available annually. ring Deutcher Makler, Hamburg.

`Ireland:` average transactions price for existing homes, based on all loan approvals, available annually. Department of the Environment, Dublin, Housing Statistics Bulletin.

`Italy:` average price for new and completely refurbished dwellings in large and middle size cities and tourist areas, reported by realtors, available annually. CENSIS, Rome.

`Japan:` based on a survey of prices per square meter of land transactions in residentially zoned areas, appraised by realtors, conducted annually. National Land Agency.

`Netherlands:` weighted average sales price for existing single- and multi-family houses, reported by realtors (including 50-60% of transactions), available annually. Netherlands Association of Real Estate Agents.

`Norway:` average sales price of existing homes, weighted by type of dwelling, reported by Property Owners Association, covering about 50% of all transactions. Norges Ejendoms Forbund, Oslo.

`Sweden:` index based on owner-occupied one- and two-dwelling buildings, based on reports of title registrations for arm's-length transactions, weighted by type of dwelling, available annually. Central Statistical Office, Monthly Digest.

`United Kingdom:` index based on survey of all dwellings with building society mortgages, weighted by type of dwelling, available annually. Department of the Environment, London.

`United States:` index based on sales prices of existing single-family homes, based on realtor reports, adjusted by regional availability of single-family homes and homeowner mobility, available annually. National Association of Realtors.

TABLE 1

The House Price Boom of the 1980's

Real Price Increase (%)

Country	**Period**	**Period**	**Period**
Australia		1986-1989 37	
Belgium		1985-1992 38	
Canada		1984-1989 56	
Denmark	1980-1982 -24	1982-1986 49	1986-1992 -28
Finland	1980-1989 103	1985-1989 61	1989-1992 -41
France	1981-1986 -18	1986-1992 25	1985-1991 57
Germany	1982-1989 -23	1987-1992 36	1989-1992 38
Ireland	1980-1987 -21	1987-1990 23	
Italy		1987-1991 19	
Japan	1977-1990 117	1980-1990 88	1990-1992 -9
Netherlands	1980-1985 -33	1985-1989 21	
Norway	1981-1985 -25	1985-1987 +16	1987-1992 -36
Sweden	1980-1985 -30	1985-1990 +38	1990-1992 -13
UK	1980-1982 -12	1982-1988 +85	1988-1992 -26
USA	1980-1984 -6	1984-1988 +15	

TABLE 2

Autoregressions of yearly change in log real housing price $P_t - P_{t-1}$

t−statistics in parentheses

Constant	.0041 (.89)	.0035 (.77)	.0045 (.95)	.0075 (1.64)	.0073 (1.61)
$P_{t-1} - P_{t-2}$	.412 (7.91)	.472 (8.55)	.428 (7.28)	.426 (7.49)	.454 (7.55)
$P_{t-2} - P_{t-3}$		-.208 (3.74)	-.120 (1.93)	-.100 (1.65)	-.113 (1.86)
$P_{t-3} - P_{t-4}$			-.167 (2.87)	-.090 (1.49)	-.076 (1.26)
$P_{t-4} - P_{t-5}$				-.188 (3.35)	-.169 (2.86)
$P_{t-5} - P_{t-6}$					-.244 (.42)
Observations	314	299	284	269	254
R^2_{adj} – OLS	.164	.193	.202	.269	.277
F – OLS	62.59	36.67	24.94	25.59	20.41
LLF	343.7	339.1	324.2	325.4	316.99
R^2_{adj} – FE	.130	.160	.168	.241	.253
F – FE	4.11	4.55	4.36	5.72	5.51
LLF	344.5	340.3	325.4	327.7	320.2
R^2_{adj} – RE	.167	.196	.210	.279	.291
Lagrange Multiplier (prob)	6.28(.012)	5.51(.019)	5.48(.019)	3.84(.05)	2.78(.10)
Hausman FE vs. RE (prob)	.733(.39)	1.467(.48)	1.453(.69)	1.89(.76)	2.96(.70)

TABLE 3

Regression of yearly real house price changes against the rate of change of explanatory variables

Column 1: Lagged explanatory variables.

Column 2–5: Contemporaneous explanatory variables.

$t-$statistics in parentheses

	1	2	3	4	5
constant	-.018	-.022	-.0291		
	(3.01)	(3.49)	(5.24)		
GDP growth rate	.766	1.149	1.165	1.23	1.224
	(4.85)	(7.58)	(8.69)	(8.77)	(8.70)
real interest	-.012	-.016	-.011	-.012	-.012
	(4.07)	(5.32)	(4.21)	(4.41)	(4.33)
house prices	.299		.449	.448	.447
	(5.58)		(9.32)	(9.11)	(9.08)
Observations	300	314	314	314	314
R^2_{adj} – OLS	.220	.151	.334		.334
F – OLS		28.75	53.4		40.1
LLF	345.4	341.6	380.4		380.6
MSE	.077	.073	.073		.073
R^2_{adj} – FE	.192	.126		.315	.314
F – FE	5.17	3.83		9.48	8.96
LLF	347.4	344.4		383.3	383.3
MSE	.078	.083		.074	.073
R^2_{adj} – RE	.227	.156		.341	.342
Lagrange Multiplier (prob)	4.39(.04)	3.19(.07)		3.22(.07)	3.24(.07)
Hausman, FE vs. RE (prob)	1.50(.43)	1.59(.45)		3.33(.34)	3.45(.48)

Note: Column 4 reports a regression with country-specific fixed effects. Column 5 reports a regression with, in addition, a linear time trend.

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