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A Model of Cross-Country House Prices

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Abstract

The widespread nature of the recent international house price boom suggests that the underlying forces behind this sustained price increase may be common across countries. Many OECD countries have, over the past decade, witnessed sustained increases in living standards while housing affordability has further improved in recent years with the low interest rate environment experienced by many of these countries. In this paper we propose a theoretical model of house price determination that is driven by changes in income and interest rates. In particular, the current level of income and interest rates determine how much an individual can borrow from financial institutions to purchase housing and ultimately this is a key driver of house prices. The model is applied to a panel of 16 OECD countries from 1980 to 2005 using both single country-by-country and panel econometric approaches. Our results support the existence of a long-run relationship between actual house prices and the amount individuals can borrow and we find plausible and statistically significant adjustment, across countries, to this long run equilibrium.

1 Introduction

Over the five year period 2000-2005, estimates by *The Economist*[1] reveal that the value of residential property in developed countries rose by over 30 trillion dollars - an increase equivalent to 100 per cent of those countries combined GDPs. In North America and across Europe, countries have experienced record highs in terms of house price to income ratios. Inevitably, the concern amongst policy-makers is the inherent stability and sustainability of this asset price increase - are property markets overvalued and if so, by how much? As noted by Case and Shiller (2003), the international media has, of late, been saturated with stories/analyses documenting the imminent “collapse” of property bubbles.

During this period, most countries within the OECD experienced economic conditions highly conducive to house price growth. Macro-economic growth has been strong and relatively stable, while international monetary conditions have also been benign. For example, many European countries have enjoyed a low interest rate environment associated with membership of the European Monetary Union (EMU). Compared with the relatively turbulent nature of interest rates in the 1980’s and early 1990’s, countries such as the United States, the United Kingdom, Canada, Australia and Sweden have also experienced more favourable interest rate conditions. This combination of low interest rates and continued improvements in living standards has increased the ability of households to finance higher mortgage levels with resulting upward pressure on house prices. The role of interest rates and income levels in determining house prices across countries has been commented upon in a variety of studies such as Ahearne et al. (2005), OECD (2005), Tsatsaronis and Zhu (2004) and ECB (2003).

Less agreement, however, is forthcoming on the theoretical and empirical approaches used to model these potential determinants of house prices. For example, it is not uncommon for price levels in the same property market when analysed with two different (and popular) approaches to be deemed either “determined by fundamentals” and consequently, sound or, conversely, “dangerously overvalued”.

It is possible to separate much of the existing literature into two broad approaches. The first we call the “econometric” approach whereby a reduced form

[1] Volume 375, Number 8431, 2005.

price equation is estimated based on some underlying notion of the determinants of supply and demand. Typically, house prices are regressed on a set of potential determinants. The fitted values from the regression are then interpreted as the price level justified by fundamentals within the economy and the potential stability of the asset price increase is gauged by comparing this fundamental price with the actual price level.[2] One of the problems with this approach is that variables which are believed, a priori, to be important in house price determination such as interest rates often appear with the wrong sign or are found to be insignificant. For example, in models estimated for eight different US States, Case and Shiller (2003) acknowledge that the mortgage rate had an insignificant coefficient in all but one of the regression models. Mayer (2003) also notes that the results from such regression models suggest that, historically, house purchase behaviour and housing values may not have been very responsive to changes in interest rates.

An alternative, more finance-based, approach taken in the literature can be characterised by an underlying notion of arbitrage where the returns to investing in housing relative to some other asset are evaluated or the costs and benefits of renting relative to buying are compared. One standard metric used in this context is the ratio of rental income to house prices. Deviations of the current rental price ratio from its long-run average are frequently taken to be an indication of over or undervaluation.[3] A more sophisticated implementation of this approach, based on the methodology of Campbell and Shiller (1988a,b) has been recently applied to the US housing market by Campbell, Davis, Gallin and Martin (2006). In this type of model, a tight relationship is imposed between house prices and interest rates. This contrasts with the former, econometric approach where the interest rate variable enters in freely into the regression specification and can often be "swamped" in the estimation yielding a very small and minor semi-elasticity effect.

However, one of the potential drawbacks of many finance based approaches is that underlying supply and demand factors such as income or demographics are not modelled. Rather, these factors enter indirectly by affecting either the growth rate of rental income or in terms of a changing discount factor. Moreover, this approach

[2]Examples of this type of approach can be observed in Poterba (1991), Mankiw and Weil (1989), Muellbauer and Murphy (1997), Roche (2001) and Fitzpatrick and McQuinn (2007) amongst others.

[3]The Economist magazine regularly posts a survey based on house price developments in a number of country capitals based on rental price ratios.

has little to say regarding any adjustment path for house prices if house prices are away from their fundamental level. In recent times many of these finance-based indicators such as the rental price ratio have deviated substantially from their long-run average for a number of different housing markets. OECD (2005) illustrate this fact for 14 out of the 17 international housing markets examined.[4] However, the implied overvaluation from such measures is, at times, at variance with the results from reduced form econometric models, which tend to suggest far less evidence of overvaluation.

In this paper, we propose a simple intuitive theoretical model of the housing market which captures the important role of credit, income and interest rates as drivers of housing demand but also resolves some of the difficulties of previous approaches already highlighted. More specifically, we model the demand-side determinants of house prices as a function of the average amount borrowed by households given current disposable income levels and interest rates. In reality, the amount lent by a mortgage institution to an individual is critically dependent on current disposable income and interest rates. Based on this observation, we back out how much a financial institution would lend an individual given plausible assumptions regarding the fraction of income that goes to mortgage repayments and the duration of the mortgage using a standard annuity formula. Ultimately, this value should be an important determinant of housing demand. We believe this model captures the fact that most house purchases are mortgage-financed and the amount that mortgage providers are willing to lend is ultimately a function of income and interest rates.

In contrast to the finance approach, however, we do not derive a "fundamental" price level and then compare it with the actual level. Instead, we estimate both a long-run relationship between house prices and the amount that can be borrowed and a short-run model that examines the speed of adjustment when there is a deviation from the long run equilibrium. We apply the model to 16 OECD country housing markets for the period 1980:Q1 to 2005:Q4. Using data for house prices, income levels and relevant mortgage interest rates in each country enables us to estimate the model both on a single country-by-country and panel data basis. In both cases, we find plausible and robust results in terms of the long-run relationship

[4] Campbell, Davis, Gallin and Martin (2006) find similar results for the four census regions of the US.

between actual prices and the level suggested by the amount that households can borrow. We also find substantial evidence of error correction to the long run price implied by our theory irrespective of whether the model is estimated in a panel or in a country-by-country manner.

We believe our model draws upon the advantages of both the econometric and finance based models while avoiding some of their pitfalls. First, the model is intuitively appealing, familar as it is to most people who have taken out a mortgage. In addition, it models, in a plausible fashion, how mortgage institutions decide how much to lend.

Secondly, since we impose a realistic theoretical relationship between interest rates, income and how much one can borrow, we avoid the shortcomings of having an insignificant or incorrectly signed interest rate response - something that is characteristic of much of the previous literature. Moreover, we highlight one possible reason for the failure of standard regression specifications to find a significant response of house prices to interest rates. Our theoretical model suggests there is a nonlinear relationship between house prices and interest rates while standard approaches only permits interest rates to enter linearly. In support of this hypothesis we report evidence that if one includes higher order powers of the interest rate, the coefficient on the interest rate term switches from being insignficant in the linear specification to being significantly negative in the more general specification. The inclusion of higher powers of the interest rate is entirely consistent with our model.

Finally, in estimating our long and short-run models, we achieve plausible and robust results in terms of the relationship between the actual and predicted price levels. This contrasts with issues of fit which can arise with the more finance-based models where the price suggested by, say, rental price ratios, are often quite out of kilter with the actual observed price.

The rest of the paper is structured as follows. In the next section we outline our proposed model which aims to resolve some of these difficulties. The following section describes the data used and the empirical approach adopted when our model is taken to the data. Results are then discussed in both a single-country and panel context. A final section offers some conclusions.

2 A Theoretical Model of Cross-Country House Prices

The model uses the following variables

$$
\begin{aligned}
P_t &= \text{actual house prices.} \\
B_t &= \text{amount that can be borrowed.} \\
S_t &= \text{supply of housing.} \\
Y_t &= \text{GDP per capita.} \\
R_t &= \text{mortgage interest rate.} \\
\tau &= \text{duration of mortgage.} \\
\kappa &= \text{proportion of household income going on mortgage repayments} \\
\omega &= \text{mulitple of GDP per capita as a proxy for household income}
\end{aligned}
$$

The model focuses on the role played by the demand-side factors income and interest rates. The demand for housing is taken to be a function of the amount that can be borrowed from a financial institution based on current disposable income and the existing mortgage interest rate. In particular, the amount lent out by financial institutions to their customers is based on the present value of an annuity, where the annuity is some fraction of current disposable income discounted at the current mortgage interest rate for an horizon equal to the term of the mortgage. This amount which can be borrowed is given by the following formula

$$B_t = \omega\kappa Y_t \left(\frac{1 - (1 + R_t)^{-\tau}}{R_t} \right). \tag{1}$$

Clearly, an upward shift in income or downward movements in the interest rate yields an increase in the amount which can be borrowed prompting additional demand for housing. Our approach is closely related to the notion of a housing affordability index frequently used in assessments of the housing market.[5]

[5]This concept measures the ratio of an average monthly mortgage payment based on cur-

We now seek to nest this expression for income and interest rates within a general model of the housing market. Firstly, we decompose B_t into its time-varying, (X_t), and constant components, (η), i.e.

$$\begin{aligned} X_t &= Y_t \left(\frac{1-(1+R_t)^{-\tau}}{R_t} \right) \\ \eta &= \omega\kappa. \end{aligned} \tag{2}$$

Both expressions are then incorporated within the following inverted demand function:

$$P_t^D = \eta X_t S^{-\mu}. \tag{3}$$

The supply variable S enters negatively in this function through the own price elasticity of demand μ. An inverted housing supply equation is given by the following

$$P_t^S = \delta S^{\phi}. \tag{4}$$

where δ, the intercept in the supply function, can be regarded as a standard supply side shifter.

In the short-run, supply is assumed to be inelastic, i.e. $S = \overline{S}$. Therefore, the short-run price of housing depends on the amount that can be borrowed. In order to derive the long-run equilibrium price level, we set $P_t^D = P_t^S$ and solve, yielding the following equilibrium expression for S^{LR}

$$S^{LR} = \left(\frac{\eta X_t}{\delta} \right)^{\frac{1}{(\phi+\mu)}}. \tag{5}$$

The corresponding expression for the long-run price is given as

rent interest rates to average family monthly income. The National Realtors Association in the United States publishes a monthly Housing Affordability Index (HAI), which is quoted frequently by the Wall Street Journal in its commentaries on the US market. See, for example, http://www.realestatejournal.com/buysell/markettrends/20051223-simon.html

$$P^{LR} = \eta^{\frac{\phi}{(\phi+\mu)}} \delta^{\frac{\mu}{(\phi+\mu)}} X_t^{\frac{\phi}{(\phi+\mu)}}. \tag{6}$$

Taking logs of equation (6) yields the following, where lower case denotes a variable is in logs

$$p^{LR} = \left(\frac{\phi}{\phi+\mu}\right) log(\eta) + \left(\frac{\mu}{\phi+\mu}\right) log(\delta) + \left(\frac{\phi}{\phi+\mu}\right) x_t. \tag{7}$$

Grouping the constants together, we simplify this expression to

$$p_t = \alpha + \psi x_t. \tag{8}$$

From the long-run model, we can retrieve an estimate of $[\frac{\phi}{\mu+\phi}]$ from the coefficient ψ. House prices are a function of how much can be borrowed and the own price elasticities of the demand and supply. The intercept α is a composite of the supply shifter δ and the parameters ϕ, μ and η. It is evident, therefore, from an estimation perspective, we do not actually have to make assumptions about either the proportion of disposable income going on mortgage repayments or the multiple of GDP per capita required to arrive at household income.

3 Data

All data used in the paper are quarterly and cover the period 1980:Q1 to 2005:Q4 for 16 OECD countries.[6] The data comes from three main sources. Real quarterly house price data are taken from a Bank of International Settlements (B.I.S.) dataset and is an increasingly popular source for studies on international house price movements. Examples of such studies include Ceron and Suarez (2006), Ahearne et al. (2005), OECD (2005) and Tsatsaronis and Zhu (2004).[7] Prices are available in index form and have been rebased such that 1980:Q1 = 100.

[6]The countries are respectively Australia, Canada, Denmark, Finland France, Italy, Ireland, Japan, Spain, Sweden, Switzerland, the Netherlands, Norway, the US and the UK.

[7]Information on the country-level sources of this data can be obtained from Table III.4 in OECD (2005).

Quarterly GDP, interest rates and the GDP deflator data are taken from the IMF's International Financial Statistics (IFS) database. Typically, most country level mortgage markets are characterised, in the aggregate, by a preference for variable or fixed rate mortgages. A recent survey paper, ECB (2003), based on questionnaires conducted by national Central Banks (NCBs), provides some information on the nature of mortgage contracts in individual EU countries. The interest rate adjustment in each country, is characterised as being fixed (F) or variable (V). For an interest rate to be classified as fixed, it must be fixed for more than five years, or, until final maturity, whereas in the case of the variable rate, it is either negotiable after one year, or, is tied to market rates, or, is adjustable at the discretion of the lender.[8] Based on these observations, each country in our sample is classified into a variable or fixed rate category where the variable (fixed) rate mortgage rates are proxied by country specific short-term money market rates (long-term Government bond rates). Annual population data is taken from either a country's national statistical agency or EuroStat's NewCronos. These series are then interpolated and along with the GDP data are combined to arrive at a quarterly GDP per capita series for each country.

Table 1 provides a summary of the core data used. Of the 16 countries, 6 are assumed to have fixed rate mortgages with the remainder having variable rate mortgages. Over the period, the countries registering the greatest increase in prices are Italy, Spain, the UK and Ireland.

3.1 Empirical Approach

The relatively long nature of the time-series dimension of the dataset enables both 'panel' and individual 'country-by-country' empirical approaches to estimation. While this is beneficial from a robustness point of view, it can yield a relatively large set of results. In the interests of clarity, therefore, we outline our estimation strategy as follows; in the next section we discuss and present results for panel unit root and cointegration tests. Panel and country-by-country approaches to the estimation of the parameters within the long-run relationship are then outlined, while, finally, both panel and country-by-country error correction models are presented and estimated. The results from the different models are also discussed in the context of

[8] See Table 5.1 of ECB (2003) for more details.

known characteristics of the mortgage markets in the different countries.

3.2 Panel Unit Root Tests

Here, we discuss three unit root tests along with results from their application to the data in this study. Consider the following model where the variable of interest is observed for N cross sectional units and T time periods,

$$\Delta y_{i,t} = \alpha_i + \rho y_{i,t-1} + \sum_{t=1}^{p_i} \Delta y_{i,t-i} + \epsilon_t \quad i = 1, .N; t = 1, ..T. \tag{9}$$

Levin and Lin (1992) consider a model in which the coefficient on the lagged dependent variable, (ρ), is restricted to be homogenous across all units (countries) of the panel. The null hypothesis is $H_0 : \rho = 0$ while the alternative hypothesis is $H_1 : \rho < 0$. The alternative hypothesis is restrictive since it implies that the autoregressive parameter is identical across the panel.

In contrast to the assumption of a constant ρ across all countries, Im, Pesaran and Shin (2003) suggest implementing a separate ADF test for each of the separate countries. The test statistic is then calculated as the average of the individual Augmented Dickey Fuller (ADF) statistics. Assuming no cross-country correlation among the errors and the same time dimension for all countries, the normalised statistic converges to a standard normal distribution.

Rather than basing the test statistic on the average of the ADF tests for each unit in the panel, Maddala and Wu (1999) and Choi (2001) suggest collecting the *p values* associated with the ADF statistic and calculating the following test statistic

$$P_{NW} = -2 \sum_{i=1}^{N} \log(p_i).$$

The resulting test statistic is asymptotically χ^2 distributed with *2N* degrees of freedom. There is some evidence that this test has better size properties than that of Im, Pesaran and Shin.

Table 2 reports the results of these three panel unit root tests for both the actual nominal house prices and the house price based on equation (1). All unit root tests suggest the series are non-stationary.[9]

3.3 Panel Cointegration

In testing for cointegration within a panel-data context we adopt two approaches. The first is the Pedroni (1999) single equation framework, which we complement with the systems approach of Larsson et al. (2001). Pedroni's cointegration tests are all single-equation methods based on estimating the static cointegrating regression given by

$$y_{it} = \alpha_i + \delta_u t + \beta_i x_{it} + \epsilon_{it} \quad i = 1, 2.., N; \quad t = 1, 2.., T \tag{10}$$

where x is a vector of regressors and β consists of its associated parameters. The tests are constructed by using the residuals $\hat{\epsilon}_{it}$ from the above cointegrating regression.

The cointegration tests proposed by Pedroni are sufficiently flexible so as to enable the investigation of heterogeneous panels, in which heterogeneous slope coefficients, fixed effects and individual specific deterministic trends are permitted. In total, Pedroni proposes the use of seven panel cointegration statistics. Four of these statistics, called *panel* cointegration statistics, are based on within-country based statistics. The other three statistics, called *group mean* panel cointegration statistics, are between-country panel statistics. Within the first category, three of the four tests are non-parametric corrections, the fourth is a parametric ADF test. In the second category, the first two use non-parametric corrections, the third is again an ADF test. Denoting the autoregressive coefficient of the residuals in the ith unit by γ_i, the within-country tests impose a common coefficient under the alternative hypothesis

$$H_0 : \gamma = 1, \quad H_1 : \gamma = \gamma < 1 \tag{11}$$

while the between dimension tests allow for heterogenous coefficients under the alternative hypothesis

[9]We have also checked whether the series are I(2), however, we find no evidence to support this.

$$H_0 : \gamma = 1, \quad H_1 : \gamma_i = \gamma < 1. \tag{12}$$

The standardized statistics tend in distribution to the normal density under the null hypothesis. Pedroni (1999) tabulates the required moments for the standardization by simulation, for different specifications of deterministics included in the models.

Larsson et al. (2001) develop a likelihood ratio panel test of cointegrating rank based on the average of the individual rank trace statistics developed by Johansen (1995). Given N countries with time dimension T, and a set of p I(1) variables, the heterogeneous vector error-correction model is given by

$$\Delta y_{i,t} = \pi_i y_{it-1} + \sum_{i=1}^{p} \Delta\Gamma y_{i,t-1} + \epsilon_{it} \tag{13}$$

where y consists of the vector of possible cointegrated variables. This equation is estimated for each country, and the average of the individual trace statistics are then calculated. The panel cointegration rank trace test statistic, is the standardised mean of the average of the N individual trace statistics and is distributed N(0,1).[10]

In Table 3, we report the results of Pedroni's cointegration tests between the log of actual nominal house prices and the log of the amount that can be borrowed. In all cases, with the exception of the *panel v test*, we can reject the null hypothesis of no cointegration at or beyond the 5 percent signficance level. We also find evidence against the null of no cointegration when we apply Larsson et al's (2001) test.

[10]The test statistic is given by

$$P_{TR} = \frac{\sqrt{N}(TR_W - EZ_W)}{VAR(Z_W)}, \tag{14}$$

where N is the number of countries in the panel, T is the time dimension (t=1, ,T); TR_W is the average of the individual trace statistics. $E(Z_W)$ and $VAR(Z_W)$ are the expected mean and variance of the asymptotic trace statistics which are tabulated by the authors through stochastic simulations.

3.4 Estimating the Cointegration Relationship

3.5 DOLS and FMOLS

In terms of estimating the long run relationship between the variables, we employ two different cointegration estimators: the first type, in the panel data case, is based on fully modified OLS (FM-OLS) while the second type, for the country-by-country models, is dynamic OLS (DOLS). Among these we can distinguish between three different forms of estimator depending on the way they pool the data: estimators that pool information along the between dimension Pedroni (2004), estimators that pool information along the within dimension weighting all the variables by their long run covariances Pedroni (2004) and Kao and Chiang (2000) and within estimators that do not scale the variables due to Mark and Sul (2003) and Pedroni (2004). The between or group mean estimators can be obtained as

$$\hat{\psi}_{GFM} = \frac{\sum_{i=1}^{N} \hat{\psi}_{FM,i}}{N}$$

$$\hat{\psi}_{GD} = \frac{\sum_{i=1}^{N} \hat{\psi}_{D,i}}{N}$$

where $\hat{\psi}_{GFM}$ and $\hat{\psi}_{GD}$ are the group mean FM-OLS and group mean DOLS estimates, $\hat{\psi}_{FM,i}$ is the FM-OLS estimator applied to the ith member of the panel $\hat{\psi}_{D,i}$ is the DOLS estimator applied to the ith member. An important advantage of the between dimension estimator is that the form in which the data is pooled allows for greater flexibility in the presence of heterogenity of the cointegrating vector. Point estimates for the between dimension estimator can be interpreted as the mean values for the cointegrating vectors.

3.6 FM-OLS and DOLS Results

We estimate the following long run relationship identified in (8) between the log of actual house prices and the log of X - the average amount which can be borrowed calculated from (2):

$$p_t = \alpha + \psi x_t.$$

Based on equation (7), ψ is a convolution of the own price elasticities of supply and demand. The panel FM-OLS results for this long run relationship are reported in Table 4. All the long-run parameters are statistically significant from zero. The largest parameter is 1.15 for the Netherlands suggesting a highly elastic response by actual prices to any change in x. Japan, with a coefficient value of 0.17, has the lowest response to a change in the price based on the average amount borrowed. The overall group estimate for the panel is 0.59.

Also reported in Table 4 are the results of t-tests for the null hypothesis: $H_0 : \psi_i = 0.59$, i.e., can all the long-run parameters for each country be restricted to equal 0.59? As can be seen, the null cannot be rejected in the case of seven countries: Canada, Denmark, Finland, France, Norway, Sweden and the US at the 5% level of signficance. Consequently, for the resulting panel data error correction model, we estimate two variants: one where the long-run parameter ψ is homogenous across all countries (set at the value 0.59) and the second, a more heterogenous approach, where we allow each country to have the long-run parameter set at its FM-OLS estimate.

The results for the long-run FM-OLS estimates can be compared with those from the individual country-by-country dynamic OLS (DOLS) estimates. The Stock and Watson (1993) DOLS estimator falls under the single equation Engle and Granger (1987) approach to cointegration, while explicitly allowing for potential correlation between the explanatory variable and the error process. This is done by adding both leads and lags of the differenced regressors to the long-run specification. Asymptotically, the FM-OLS estimator approximates DOLS, therefore, a reasonable robustness check on our results is to compare both sets of results. The DOLS estimates are presented in Table 5. The results are almost identical to those of FM-OLS - the correlation coefficient between both sets is 0.99 and the ranking of the countries is also almost identical.

The relative size of the different long-run parameters, ψ, i.e. the long-run response of each country's actual price to changes in the average amount borrowed can be potentially rationalised by cross country differences in individual country housing markets. For example, the relative size of the long-run parameters can be a function of the stickiness of supply in a particular country. Recall, that the more elastic supply is, i.e. the greater the size of $[\frac{\phi}{\mu+\phi}]$ in (7), the smaller will be the long-run relationship between the actual price and the amount borrowed. ECB (2003) contends, that while information on the supply response in different EU countries may be sketchy, what information is available suggests that the supply of new housing is more responsive to house prices in Germany than in the UK, the Netherlands or the Nordic countries. Our estimate of the long-run parameter for Germany is the second lowest in the sample.

Alternatively, cross country differences in the coefficients in the long-run relationship may reflect heterogenities across countries on the demand side. One potential difference is the degree of flexiblity of credit markets in a particular country. Two recent survey papers - OECD (2005) and Giuliodori (2004) examine mortgage markets in a number of countries. For example, Giuliodori (2004), quoting EMF (1998) and the ECB (2003) amongst others, suggests that the UK, which has one of the largest ψ's, has a very high loan to value ratio by international standards. Similarly, the OECD (2005), quoting Scanlon and Whitehead (2004) and the Canada Mortgage and Housing Corporation (2005), suggests that innovation in mortgage products tends to be highest in countries such as the UK, Australia and the Netherlands. All three of these countries have long-run parameters that are amongst the highest in our sample.

3.7 Nonlinear Effect of Interest Rates

An alternative way of thinking of the approach adopted here is that interest rates have a nonlinear effect on housing demand which isn't captured by a standard regression specification where interest rates enter linearly. This nonlinear effect is illustrated in Figure 1. The value of an annuity is plotted as the interest rate varies for three different annuity maturities, i.e., the value of an annuity that pays out one euro each year for 10 years, 20 years and 30 years respectively. The annuity value is clearly a nonlinear function of the interest rate and regression specifications where the interest rate enters linearly will not capture this phenonemon.

To further explore this issue, we estimate for each country two specifications where interest rate enter in a standard linear or nonlinear fashion. Hence, we estimate the following two variants of the standard reduced form house price regression

$$p_t = \alpha + \beta y_t + \sum_{i=1}^{2} \omega_i R_t^i. \qquad (15)$$

where i = 1 and i = 2. In the first regression, the interest rate variable enters in a standard linear fashion along with disposable income, while in the second specification, both the level and the square of the interest rate variable are included.

In the linear specification, the interest rate variable enters the regression with either a positive and/or an insignificant coefficient in all but one of the 16 countries. This result highlights the issue outlined earlier concerning the problematic nature of the interest rate response in reduced form estimates of house prices. However, the introduction of the square of the interest rate results in a significant and negative coefficient on the level interest rate variable in 9 countries. The impact of the level interest rate variable is, also, considerably larger under the augmented model. In Table 6, we report the "linear" and "nonlinear" estimates of equation (15) for these nine countries. Apart from Denmark and the United States, the coefficient on the level interest rate in the linear model is positive.[11] Including the square of the interest rate as a regressor brings about a significant change in the overall interest rate effect on house prices. In the case of some countries, this change is quite substantial. For the Netherlands, the coefficient on the interest rate variable goes from 0.025 to -0.308.

In the next section, we turn our attention to the short-run models based on the long-run estimates. Error correction models are presented both on a panel data and on a country-by-country basis.

3.8 Error Correction Models

Using a panel data approach, the estimated error correction model is specified as follows

$$\triangle p_{i,t} = \lambda\,(p_{i,t-1} - \alpha - \psi\, x_{i,t-1}) + \sum_{i=1}^{4} \theta_i \triangle p_{i,t-i} + \sum_{i=0}^{4} \theta_{i+5} \triangle x_{i,t-i} + u_{i,t}. \qquad (16)$$

[11] In the US case, however, the coefficient is insignificant.

In the first estimated panel model, henceforth labelled Model 1, ψ in the long-run relationship is set equal to 0.590 for all countries based on the FM-OLS group estimate. In the second specification, henceforth labelled Model 2, the long-run coefficient is allowed to vary across countries and the values assumed are based on the country-specific results in Table 4. We adopt a fixed effects estimator. The results along with p-values for the inclusion of the country-specific dummies are reported in Table 7.

In both cases, it is evident that error correction actually takes place - λ the coefficient on the error correction term, (ECT), is negative and significant. Unsurprisingly, the degree of error correction is greater for Model 2, i.e., where we allow country-specific long-run parameters. For Model 1, the degree of correction is just over one per cent per quarter while in model 2, the degree of correction is two per cent per quarter.[12]

The estimation of dynamic panel data models has, of late, attracted considerable interest. In the presence of dynamics, Bond (2002), amongst others, note certain biases, which can affect the estimated coefficients. In particular, OLS and fixed effects estimators can be shown to exert biases in opposite directions on the lagged dependent variable in such regressions. The latter fixed effects bias has been documented analytically by Nickell (1981). This upward bias tends to zero as the $T \rightarrow \infty$. Given the time dimension in this study is 104 quarters, the effect of this "Nickell" bias is likely to be very small. Nonetheless, Models 1 and 2 are also estimated via OLS and the coefficients on the lagged dependent variable were compared with those in Table 7. The coefficients in all four cases are almost identical - at 0.09,[13] therefore, we believe that our results are relatively free of these biases.

We next turn to the results from the country-by-country regressions. Two models are, again, estimated. In the first, which we label Model 3, the long-run parameter ψ is estimated simultaneously with the rate of error correction. Thus, the following model is estimated for each of the 16 countries

[12] We also run the model in an unconstrained fashion i.e. where both the degree of error correction λ and the long-run parameter ψ are simultaneously determined. We achieve statistically significant estimates with a value of -0.012 for λ and 0.77 for ψ.

[13] Full regression results are available, upon request, from the authors.

$$\triangle p_t = \lambda \left(p_{t-1} - \alpha - \psi\, x_{t-1}\right) + \sum_{i=1}^{4} \theta_i \triangle p_{t-i} + \sum_{i=0}^{4} \theta_{i+5} \triangle x_{t-i} + u_t. \qquad (17)$$

enabling the generation of 16 country-specific rate of corrections and long-run coefficients.

The next specification, labelled Model 4, also estimates short-run parameters. However, in this instance they are conditional on the DOLS long-run results. Hence, the following specification is also estimated

$$\triangle p_t = \lambda \left(p_{t-1} - \alpha^{DOLS} - \psi^{DOLS} x_{t-1}\right) + \sum_{i=1}^{4} \theta_i \triangle p_{t-i} + \sum_{i=0}^{4} \theta_{i+5} \triangle x_{t-i} + u_t. \quad (18)$$

where λ is again the speed of error correction and γ^{DOLS} and α^{DOLS}, are the previous estimates of the long run parameters from Table 5 based on DOLS. A summary of the estimation results for all countries are presented in Table 8 - with the respective error correction coefficients λ and $\overline{R}^2$ are presented for both models and long-run coefficients, ψ, are presented for Model 3.

In terms of the ψ coefficient in Model 3, 12 of the 16 individual country ψ coefficients are significant at the 5 per cent level. Again, these long-run results are very similiar to those of the FM-OLS and the DOLS estimates. A large number of individual country error-correction terms are also significant for both models - 12 in the case of Model 3 and 13 in the case of Model 4. Of the countries, Ireland has the fastest rate of convergence to its long-run level, with a speed of four per cent per quarter in terms of error correction.

Four countries report insignificant coefficient estimates across both models for both the error correction term and long-run terms - Denmark, Switzerland, Japan and the US. The result for the US is not altogether surprising - studies, including that of Gallin (2006) and Himmelberg, Mayer and Sinai (2005), highlight the regional diversity in US house price movements, with national-level data often obscuring important economic differences between major US cities. Moreover, Gallin finds an absence of a cointegrating relationship between house prices, income and other variables both at a national and regional level for the US.

In the case of Japan, the period 1990-2004 was a period of considerable asset price instability, with real house prices, in particular, falling by over 38 per cent between 1991 and 2004. Furthermore, since the mid 1990's, interest rates have been exceptionally low, which, given the nonlinear nature of the present value formula as a function of interest rates, results in a large effect on the predicted price.

From an international perspective, the results from the short-run models are somewhat reassuring as they suggest that most countries prices do "error correct" over the longer term. By this we mean, that if individual countries experience differences between the actual price level and that suggested in the long-run by income and interest rates, then, any subsequent correction which occurs can be affected through changes in the growth rate of house prices rather than solely through changes in the level. However, the policy conclusions based on these empirical results are clearly different from those suggested by models underpinned by assumptions of instantaneous adjustment.[14]

4 Concluding Comments

Capturing the dynamics of cross-country house prices would appear to be a formidable task. Many country-specific factors can impact on the performance of individual property markets. However, the strong co-movement across countries in house prices is matched by similar patterns in underlying macro-economic indicators such as interest rates and income, which are considered central to any model of the property sector. In this paper we propose that a cross-country house price demand schedule can be adequately represented by a price suggested by the average amount that can be borrowed in each country with the latter being determined by current disposable income levels and interest rates.

This approach has a number of advantages. Chief amongst these is the theoretical rigour imposed on the relationship between house prices, interest rates and income levels. It implies a specific role for interest rates, which is particularly im-

[14] We assess the robustness of our results by varying the value assumed for the parameters assumed in equation (1). Initially a mortgage term of 20 years is assumed. Results are also generated for when mortgage terms of 15 and 25 years are assumed for the same econometric specifications as shown in Table 8. These results, which are available, upon request, from the authors, are broadly similar for the different mortgage maturities.

portant for scenario analysis in the light of recent monetary tightening witnessed in many OECD countries. The formula is also well known to housing market participants - either prospective buyers or credit institution lenders as it corresponds with popular notions of what determines mortgage lending. Additionally, the formula is straightforward and easy to compute given that it requires information on just income and interest rates.

Our results reveal cointegration between actual prices and the predicted price based on (1) across the sample. This finding is robust across seven out of the eight cointegration tests applied. Results for individual countries both within the panel context and on a country-by-country basis tend to correlate with *a priori* expectations given recent survey information concerning individual mortgage markets. Furthermore, we find error correction to the long-run price across the panel and for all but four of the 16 countries when short-run models are estimated individually.

Understanding the role played by fundamental variables in determining house price movements is important and advantageous on a number of fronts. Using the approach presented here, for example, future research could examine the extent to which recent increases in OECD house prices have been generated by movements in market fundamentals, or whether the increase is built on altogether less secure foundations. If they are not driven by fundamentals, are periods of overvaluations correlated across countries? Additionally, it may be possible to identify common patterns vis-à-vis the relationship between actual prices and fundamental prices, during periods of significant price changes.

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Table 1: Summary of Data: 1980:1 - 2004:4

	Sample Means			R_t
Country	P_t	Y_t	R_t	Classification
Australia	320.2	25686	8.9	V
Canada	226.8	26303	7.6	V
Denmark	190.1	179365	9.0	F
Finland	291.4	18059	6.2	V
France	213.2	18941	8.3	F
Germany	128.3	18854	6.3	F
Ireland	368.4	16058	8.9	V
Italy	384.5	14132	10.3	V
Japan	158.3	3408818	3.4	V
The Netherlands	163.3	19340	6.9	F
Norway	295.5	214478	8.8	V
Spain	508.5	10194	9.8	V
Sweden	195.6	178767	8.6	V
Switzerland	158.0	47887	4.3	F
UK	326.8	11340	8.2	V
USA	190.7	26000	7.7	F

Note: All monetary variables are in nominal terms. House prices, (P_t), are in index form with 1980 quarter 1 = 100, GDP per capita, (Y_t), is in the national currency and interest rates, (R_t), are in percentages. F = fixed and V = variable. Fixed interest rates are fixed for more than five years, or, until final maturity, variable interest rates are renegotiable after one year, or, are tied to market rates, or, are adjustable at the discretion of the lender.

Table 2: Panel Unit Root Test Results

Test	*Variable* p_t	x_t
Levin & Lin ADF Test	0.448	-1.454
Im, Pesaran Shin ADF Test	0.697	-1.630
Maddala & Wu	37.572	9.617

Note: Maddala & Wu unit root test is χ^2 distributed with the null hypothesis of a unit root. All tests include a constant and trend.

Table 3: Panel Cointegration Test Results

Tests	
Pedroni	
panel v-stat	-2.366
panel rho-stat	2.977
panel pp-stat	2.297
panel adf-stat	2.555
group rho-stat	4.408
group pp-stat	4.232
group adf-stat	3.728
Larsson et al	
Panel Trace $H(r = 0 \mid r = 1)$	7.810

Note: Panel-v is a non parametric variance ratio statistic; Panel-rho and panel-pp are analogous to the nonparametric Phillips-Perron rho- and t-statistics. Panel-adf is a parametric statistic based on the augmented Dickey-Fuller ADF-statistic. Group rho, is analogous to the Phillips-Perron rho-statistic, while Group-pp and group-adf are analogous to the Phillips-Perron t-statistic and the adf-statistic. The Larsson et al. panel trace test is distributed N(0,1).

Table 4: Panel FM-OLS Long-Run Estimates

		$H_0 : \psi_i = 0.0$	$H_0 : \psi_i = 0.59$
Country	Coefficient	t-stat	t-stat
Australia	0.74	-13.35	-2.72
Canada	0.54	-10.99	-0.91
Denmark	0.53	-13.34	-1.56
Finland	0.56	-9.08	-0.44
France	0.53	-12.00	-1.27
Germany	0.25	-8.80	-11.88
Ireland	0.73	-21.61	-4.19
Italy	0.46	-11.80	-3.30
Japan	0.17	-2.42	-5.80
The Netherlands	1.15	-14.75	-7.2
Norway	0.72	-9.48	-1.71
Spain	0.75	-14.23	-3.01
Sweden	0.52	-11.26	-1.52
Switzerland	0.34	-3.28	-2.41
UK	0.76	-12.01	-2.68
USA	0.59	-19.37	-0.12
Panel Group Estimate	0.59	45.91	

Note: Number of observations = 16 (N) $\times$ 104 (T) = 1664.

Table 5: Country-by-Country DOLS Long-Run Estimates

Country	Coefficient	T-Statistic
Australia	0.75	4.27
Canada	0.55	3.58
Denmark	0.52	1.48
Finland	0.54	5.54
France	0.52	0.893
Germany	0.28	2.94
Ireland	0.73	6.49
Italy	0.45	3.68
Japan	0.17	0.23
The Netherlands	1.20	15.41
Norway	0.73	2.98
Spain	0.75	3.56
Sweden	0.53	3.69
Switzerland	0.32	1.18
UK	0.77	3.26
USA	0.59	1.18
Correlation with FM-OLS Estimates = 0.998		

Note: In our application the error process in the DOLS regression is assumed to follow an AR(2) process, while k - the number of leads and lags is set equal to 2. This results in 86 degrees of freedom for the German house price regression and 88 degrees of freedom for the remaining countries.

Table 6: Select Country-by-Country Reduced-Form Estimates

	Linear		*Nonlinear*		
Country	β	ω_1	β	ω_1	ω_2
Denmark	0.639	-0.026	0.353	-0.113	0.003
	(5.204)	(-2.761)	(3.036)	(-6.786)	(5.985)
France	1.066	0.012	1.023	-0.056	0.003
	(10.886)	(1.264)	(11.258)	(-3.130)	(4.329)
Ireland	1.036	0.005	0.870	-0.182	0.009
	(18.128)	(0.481)	(23.411)	(-11.611)	(13.050)
Italy	1.003	0.020	1.122	-0.019	0.002
	(16.033)	(3.181)	(19.109)	(-2.158)	(5.634)
The Netherlands	1.815	0.025	1.572	-0.308	0.021
	(22.387)	(1.998)	(24.362)	(-8.448)	(9.448)
Norway	1.114	0.032	1.059	-0.112	0.008
	(37.495)	(6.135)	(42.843)	(-5.817)	(7.662)
Sweden	0.882	0.000	0.808	-0.020	0.001
	(22.071)	(-0.069)	(16.139)	(-2.197)	(2.360)
UK	1.348	0.028	1.289	-0.051	0.004
	(30.583)	(4.654)	(28.583)	(-2.185)	(3.496)
USA	0.932	-0.001	0.784	-0.108	0.005
	(16.357)	(-0.166)	(14.436)	(-5.868)	(6.147)
Regression:	$p_t = \alpha + \beta y_t + \sum_{i=1}^{2} \omega_i R_t^i$				

Note: T-stats are in parentheses.

Table 7: Panel Data Error Correction Models

		Model 1		Model 2	
Parameter	Variable	Coeff.	t-stat	Coeff.	t-stat
α	Constant	-2.747	-14.683	-4.540	-48.038
λ	ECT_{t-1}	-0.011	-4.918	-0.021	-7.329
θ_1	$\triangle p_{i,t-1}$	0.376	15.02	0.366	14.72
θ_2	$\triangle p_{i,t-2}$	0.232	8.78	0.230	8.78
θ_3	$\triangle p_{i,t-3}$	0.161	6.12	0.163	6.27
θ_4	$\triangle p_{i,t-4}$	-0.054	-2.16	-0.047	1.93
θ_5	$\triangle b_{i,t}$	0.017	2.65	0.018	2.79
θ_6	$\triangle b_{i,t-1}$	0.015	2.28	0.010	1.54
θ_7	$\triangle b_{i,t-2}$	0.017	2.70	0.013	1.98
θ_8	$\triangle b_{i,t-3}$	0.013	2.05	0.008	1.34
θ_9	$\triangle b_{i,t-4}$	0.005	0.79	0.002	0.24
H_O : No Country Dummies		0.000		0.000	
$\overline{R}^2$		0.383		0.396	

Note: As a test for autocorrelation in the residuals of both regressions we calculate the Baltagi-Wu LBI statistic (Baltagi and Wu (1999)). A score well below 2 suggests the presence of positive serial correlation. We get scores of 1.868 and 1.874 respectively for Models 1 and 2.

Table 8: Country-by-Country Error Correction Models

	Model 3			Model 4	
Country	ψ	λ	$\overline{R}^2$	λ	$\overline{R}^2$
Australia	0.795	-0.025	0.410	-0.022	0.430
	(7.690)	(-2.619)		(-2.320)	
Canada	0.657	-0.034	0.143	-0.038	0.130
	(4.972)	(-2.063)		(-2.277)	
Denmark	2.960	-0.002	0.470	-0.006	0.440
	(0.211)	(-0.175)		(-0.329)	
Finland	0.552	-0.030	0.740	-0.030	0.740
	(5.774)	(-3.101)		(-3.363)	
France	0.741	-0.016	0.730	-0.015	0.710
	(5.876)	(-2.253)		(-2.378)	
Germany	0.274	-0.009	0.920	-0.009	0.920
	(4.978)	(-2.272)		(-2.289)	
Ireland	0.923	-0.031	0.130	-0.044	0.130
	(5.758)	(-1.642)		(-2.267)	
Italy	0.555	-0.021	0.590	-0.024	0.580
	(6.270)	(-2.378)		(-2.739)	
Japan	-3.206	-0.001	0.91	-0.002	0.910
	(-0.248)	(-0.283)		(-0.775)	
The Netherlands	1.469	-0.026	0.600	-0.019	0.59
	(9.836)	(-2.822)		(-2.245)	
Norway	0.770	-0.023	0.410	-0.024	0.400
	(5.141)	(-2.477)		(-2.612)	
Spain	0.735	-0.019	0.550	-0.018	0.520
	(7.598)	(-2.574)		(-2.517)	
Sweden	0.629	-0.027	0.580	-0.028	0.570
	(8.235)	(-3.446)		(-3.518)	
Switzerland	0.169	-0.031	0.350	-0.032	0.36
	(0.888)	(-2.869)		(-3.010)	
UK	0.866	-0.019	0.530	-0.019	0.510
	(6.847)	(-2.431)		(-2.454)	
USA	-1.264	0.001	0.640	-0.006	0.436
	(0.076)	(0.110)		(-0.782)	

Note: T-stats are in parentheses.

Figure 1: Value of Annuity for Differing Interest Rates and Maturities