HOUSE PRICES AND RENTS
AN EQUILIBRIUM ASSET PRICING APPROACH

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Abstract: We use a relatively general intertemporal asset pricing model where housing services and consumption are non-separable to measure overvaluation of housing in relation to rents in Spain, the UK and the US. Part of the increase in real house prices during the late nineties can be seen as a return to equilibrium following some undershooting after previous price peaks. However marked increases in house prices have led price-to-rent ratios above equilibrium by mid 2003 (around 30% above equilibrium in the UK, 20% in Spain and 10% in the US). Part of that overvaluation – particularly in Spain and the UK – may be attributable to the sluggishness of supply in the presence of large demand shocks in this market and/or the slow adjustment of observed rents.

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1. INTRODUCTION

Since the late 1990s, many industrialised economies have experienced a protracted period of significant increases in property prices. In some of these economies the process has become particularly acute after the heavy correction of stock market prices set in train at the end of 2000. It is now becoming conventional wisdom that such a surge in real estate prices has substantially contributed to the resilience of private consumption thereby moderating the depth and length of the cyclical downturn that both the US and the European economy experienced in that period. Actually, a number of studies (see e.g. Case, Quigley and Shiller, 2001 and BIS, 2002) have documented the fact that the wealth effects associated with real estate are more significant than those linked to financial asset holdings in most economies.

Given the important role that non-financial wealth seems to play as a determinant of aggregate demand, it is worth exploring the determinants of house prices in order to better assess the prospects of the economy. In particular, a key issue in this regard is to assess whether a particular level of prices is sustainable in the sense of being consistent with economic fundamentals.

Most of the empirical literature on this matter has made use of simple asset pricing models where the expected return on housing is equal (up to a constant) to the return on alternative investments. In most cases (Poterba, 1984, 1991; Topel and Rosen, 1988; Mankiw and Weil, 1989; Muellbauer and Murphy, 1997) the asset return condition is used together with a housing services supply rule to express house prices in terms of a set of macroeconomic variables. Others, however (Case and Shiller, 1988; Clayton, 1996), have directly tested an implicit relation between house prices and rents, using present value models that resemble those used to explain the price-to-dividend ratios in the financial literature. This approach has the advantage of not requiring the specification of the user cost of housing and/or of the market price of housing services. One drawback, however, is that rents are normally measured with noise. In addition, the relationship between prices and rents is normally affected by supply restrictions, regulations and contractual practices in the rental market that are not easily captured by standard asset financial models. Moreover, the available empirical studies typically exploit theoretical relationships that imply either a constant or a relatively restrictive specification of the discount factor for future asset payoffs which is not consistent with a meaningful intertemporal asset pricing model.

Indeed, as for any other asset, equilibrium house prices should be related to future rents through a stochastic discount factor, determined by agents' intertemporal marginal rate of substitution which, as usual, would depend on their aversion to risk and to fluctuations in their consumption bundle. Moreover, insofar as houses yield utility by providing services that are not likely to be separable from private consumption, the relevant discount factor would itself be affected by housing market conditions. This suggests that an adequate modelling of house prices would ideally require a general intertemporal model with more than one good. Recent work by Aoki, Proudman and Vlieghe (2002), who use an equilibrium model to explain the impact on house prices of some institutional features of the UK real estate market, confirms this intuition.

In this paper we use a relatively general intertemporal asset pricing model with non-separability between housing services and consumption in the relationship between house prices and rents. The model allows us to derive a benchmark that can be used to obtain a measure of under- or over-valuation of house prices. In order to estimate the model we make use of the approximation technique developed by Restoy and Weil (1998) to obtain explicit equations for equilibrium returns and asset prices. We also take into account in our estimates possible transaction costs and regulations that hinder the immediate adjustment of house prices and rents to changing conditions by specifying a dynamic relationship that permits transitory but persistent deviations from equilibrium. In order to control for measurement errors and the potential endogeneity of explanatory variables, the model is estimated by GMM.

The dynamic model is estimated for Spain, the United Kingdom and the United States, which have historically experienced large fluctuations in real-estate returns and sharp increases in house prices in the most recent past, both in absolute terms and in relation to the trend of rents. According to our results, part of the surge in real house prices in the late nineties can be seen as broadly compatible with the historical adjustment pattern of prices and rents to equilibrium conditions. However, current levels are, in the three countries considered, clearly above what their long-term fundamentals would suggest.

The paper is structured as follows. In Section 2 we introduce the intertemporal asset pricing model. In Section 3, we explain our empirical strategy, comment on the database we use, and present the main empirical results. Lastly, in Section 4 we outline the main conclusions of the paper.

2. THE MODEL

2.1. Set-up

We assume an economy where agents' preferences are defined over an aggregator, C, of consumption goods, c, and housing services, h, which has the following CES form

$$C_t = \left[\alpha^{\frac{1}{\eta}} c_t^{\frac{\eta-1}{\eta}} + (1-\alpha)^{\frac{1}{\eta}} h_t^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}} . \tag{1}$$

Agents' intertemporal preferences are of a Generalised Isoelastic (GIP) form (see Epstein and Zin, 1989, and Weil, 1990)

$$V_t = \left[(1-\theta) C_t^{1-\rho} + \theta \left(E_t V_{t+1} \right)^{\frac{1-\gamma}{1-\rho}} \right]^{\frac{1-\rho}{1-\gamma}} \tag{2}$$

where ρ is the inverse of the elasticity of intertemporal substitution and γ is the coefficient of relative risk aversion.

The budget constraint that agents face in this economy may be written as

$$W_{t+1} = \widetilde{R}_{w,t+1} \left(W_t - \frac{P_{c,t}}{P_t} c_t - \frac{P_{h,t}}{P_t} h_t \right), \tag{3}$$

where W_t is their real wealth, expressed in terms of a composite price index (P_t), $P_{h,t}$ and $P_{c,t}$ are the prices of consumption and housing services, respectively, and $\widetilde{R}_w$ is the real return (in terms of *P*) on the equilibrium portfolio of (real, human and financial) assets.

2.2 Intra-period decisions

According to the consumption aggregator function (1), for a given C, the optimal distribution of real expenditure between consumption and housing services is given by

$$c_t = \alpha \left(\frac{P_{c,t}}{P_t} \right)^{-\eta} C_t \tag{4}$$

and

$$h_t = (1-\alpha) \left(\frac{P_{h,t}}{P_t} \right)^{-\eta} C_t \, . \tag{5}$$

where the composite price-index is

$$P_t = \left[\alpha \, P_{c,t}^{1-\eta} + (1-\alpha) \, P_{h,t}^{1-\eta} \right]^{\frac{1}{1-\eta}} . \tag{6}$$

2.3 Intertemporal decisions and asset pricing

From (4), (5) and (6), the budget constraint (3) can be written after intra-period maximisation as

$$W_{t+1} = (W_t - C_t) \, \widetilde{R}_{w,t+1} \tag{7}$$

Therefore, equilibrium intertemporal consumption and portfolio decisions can be obtained by simply solving a standard single consumption good maximisation problem where preferences are of a GIP type. Since our goal is to exploit explicit equilibrium relationships between an asset price and its fundamentals, it is useful to adopt the "approximate" equilibrium approach, outlined in Restoy and Weil (1998).

More specifically, given an asset i (say a house), we can approximate its (net) return between *t* and *t+1*, $\widetilde{r}_{i,t+1}$, as

$$\tilde{r}_{i,t+1} \approx \tilde{d}_{i,t+1} + q_{i,t+1} - \frac{1}{\delta_i} q_{i,t} - k_i , \tag{8}$$

where $\tilde{d}_{i,t+1}$ is the real growth rate of dividends (rents) expressed in terms of P_t, between *t* and *t+1*, q_i is the (log) asset price-dividend ratio and δ_i and k_i are linearisation constants, the former being equal to $1 - e^{E(-q_i)}$.

Assuming that consumption growth and asset returns are all homokedastic, the difference between the expected return on any two assets is constant.[1] This, of course, applies to the return on the equilibrium portfolio. Let us then define π_{iw} as

$$\pi_{iw} \equiv E_t \ \tilde{r}_{i,t+1} - E_t \ \tilde{r}_{w,t+1} , \tag{9}$$

where $\tilde{r}_w = \ln \tilde{R}_w$.

Solving (8) forward for $q_{i,t}$, assuming that bubbles are unfeasible, taking expectations on both sides of the resulting expression, and using (9) yields

$$q_{i,t} = -\left(k_i + \pi_{iw}\right) \frac{\delta_i}{1-\delta_i} + E_t \sum_{s=1}^{\infty} \delta_i^s \left[\tilde{d}_{i,t+s} - \tilde{r}_{w,t+s}\right]. \tag{10}$$

Restoy and Weil (1998) show that the equilibrium relationship between the expected return on the wealth portfolio and the expected consumption growth (consumption function) can be approximated by the simple linear expression

$$E_t \left[\tilde{r}_{w,t+1}\right] = u + \rho \ E_t \ X_{t+1} ,$$

[1] Note that in this homokedastic framework both covariances of individual asset returns with aggregate consumption and with market returns are constant. For GIP preferences, this implies that the risk premia of all assets do not vary over time. See Epstein and Zin (1989), Weil (1990) and Giovannini and Weil (1989).

where $X_t \ (\equiv \ln C_t - \ln C_{t-1})$ is the growth rate of the consumption aggregator, and u is a constant. This implies that (10) can be written as

$$q_{it} = m + E_t \sum_{s=1}^{\infty} \delta_i^s \ \tilde{d}_{i,t+s} - \rho \ E_t \quad \sum_{s=1}^{\infty} \delta_i^s \ X_{t+s}, \tag{11}$$

where $m = \left(-k_i + \pi_{iw} + u\right) \dfrac{\delta_i}{1-\delta_i}$

Therefore, the price-to-rent ratio can be approximated by a linear function of the discounted sum of future expected growth rates of rents and the discounted sum of future expected growth of the consumption aggregator. Significantly, the coefficient of relative risk aversion γ does not explicitly appear in equation (11). Note, however, that it affects the equilibrium ratio through its effects on both π_{iw} (and therefore, on *m*) and the discount term δ_i.

Not surprisingly, the higher the future expected rents the higher the current price of houses. Also, the higher the future expected consumption aggregator, and the higher the inverse of the elasticity of intertemporal substitution, then the higher the equilibrium discount rate of future rents and, therefore, the lower the current asset price. If $\rho = 0$, expression (11) collapses to the standard expression of prices as a function of future expected payoffs discounted at a constant rate δ.

Expression (11) is not however suitable for empirical exercises since housing services and, therefore, the consumption aggregator *C*, are not observable. Moreover, the variable $\tilde{d}_i$ is expressed in terms of a composite price index *P* that incorporates the shadow price of housing services, which is normally not available either. In order to circumvent those difficulties, we could first make use of the intraperiod equilibrium condition (4) in order to express *C* only in terms of non-housing consumption, *c*, and the relative price of consumption (P_c / P). We could then calculate the relative price of consumption by using the approximation[2]

$$\Delta \ln \tfrac{P_c}{P} \approx a - b\, \Delta \ln \tfrac{P_h}{P_c}\,, \qquad \text{with } b > 0 \tag{12}$$

Finally, assuming that the housing market is in equilibrium implies that rental prices should be equal to the shadow price of housing services. This means that

$$\Delta \ln \tfrac{P_h}{P} = \tilde{d} \tag{13}$$

Therefore, putting together (4), (11), (12) and (13) we can express the equilibrium price of houses in the economy as

$$q_t^* = m' + \tau\, E_t \sum_{s=1}^{\infty} \delta^s\, d_{t+s} - \rho\, E_t \sum_{s=1}^{\infty} \delta^s\, x_{t+s} \tag{14}$$

where $x_t = \Delta \ln c_t$, m' is a constant, $\tau = 1 - b + b\rho\eta$ and $d_t \left(\equiv \tilde{d}_t - \Delta \ln(\tfrac{P_{ct}}{P_t})\right)$ is the real growth rate of rents, expressed now in terms of the price of the consumption good, and the subscript *i* has been dropped from *q* and *d* to denote aggregate variables.

Expression (14) has of course the same rationale as expression (12). It is worth noting, however, that τ, the coefficient of the expected sum of future growth of rents when they are deflated by the price of the consumption good, is in general different from 1. This is so because in our two-good equilibrium setting deflated rents are not only dividends of an asset but also a relative price of housing services.

Thus, the first component in the equation for τ – i.e. the "1" – reflects that, as before, the higher (lower) the growth rate of rents, the higher (lower) the payoff of housing investment.

[2] Under a Cobb-Douglas price aggregator this approximation would be exact with *a* = 0 and *b* equal to the share of housing expenses in total household expenditure. Under a CES aggregator like (6) it can be seen as a linear approximation around η = 1, where there is no structural interpretation of the coefficient *b*.

But a higher (lower) relative rental price also increases (reduces) the overall price level and therefore reduces (increases) the purchasing power of a given growth rate of dividends (expressed in terms of the consumption good). This makes real assets less (more) attractive. This additional effect is captured by *b*, which, according to expression (12) determines how the change in the relative rental price affects the overall price level under our linear approach.[3]

Finally, an increase in the relative price of rentals also affects the optimal distribution between consumption goods and housing services. When this relative price increases (decreases), consumption goods increase (decrease) their share, meaning an increase (decrease) in x_t. As this is a change in the distribution of the components of the CES aggregator (it is not a change in the aggregator itself) it should not affect *q.* Thus, the third element in τ corrects the change in *q* caused by this change in x_t.

3. THE EMPIRICAL ANALYSIS

Empirical strategy

Equation (14) provides an equilibrium relationship for the (ln) ratio of house prices to rents. Let the corresponding *observed value* of the ratio be denoted by q_t to distinguish it from the *equilibrium* ratio q_t^*. Thus,

$$q_t = q_t^* + g_t \tag{15}$$

where g_t is the gap, at *t*, between the observed ratio and its equilibrium value. To empirically test the model, we make a number of assumptions regarding g_t. First, we allow g_t to depend on a state variable capturing transitory demand pressures in the housing market. It is worth noting that because of the specific nature of this market – building a new house takes a long time – the responses by supply to unexpected demand shocks are very likely to show a high degree of sluggishness. Therefore, prices would tend to overreact in the short run to such shocks. DiPasquale and Wheaton (1994), Kenny (1999) or Genesove and Mayer (2001) have documented the relevance of supply adjustment costs to explain house price

[3] Note that the other parameter in equation (12), *a*, is embedded in *m'.*

behaviour. Ortalo-Magné and Rady (2001) also show how house prices may overreact in the short run to income shocks because of the interaction between young credit-constrained households and old non-constrained ones.

Additionally, we also allow g_t to show some degree of persistence. In this case, the rationale is related also to the information available on rents. As will be explained later, we compute rental prices from the corresponding domestic CPI shelter components. However, as the average maturity of rental contracts is typically well above one quarter, it will take some time for changes in the equilibrium value of rents to be fully incorporated into the corresponding CPI component. This lag will be all the greater the longer the average maturity of rental contracts. Thus, even if there are no unexpected demand shocks, the price-dividend ratio will converge to its equilibrium – when needed – with some stickiness.

Therefore, we characterise g_t behaviour as follows

$$g_t = \phi \, g_{t-1} + \beta \, w_t + \varepsilon_t \, , \tag{16}$$

where $-1 < \phi < 1$, w_t is a zero-mean stationary variable capturing demand pressures – so that β is expected to be positive – and ε_t is a standard iid white noise. Note that according to expression (16), prices can deviate from equilibrium only transitorily, although deviations may show some degree of persistence.

Combining (15) and (16) and using (14) to eliminate the unobservable equilibrium ratio we obtain

$$\begin{aligned} q_t = m' \, (1-\phi) + \phi \, q_{t-1} + \tau \, E_t \sum_{s=1}^{\infty} \delta^s \, d_{t+s} - \rho \, E_t \sum_{s=1}^{\infty} \delta^s \, x_{t+s} \\ - \tau \, \phi \, E_{t-1} \sum_{s=1}^{\infty} \delta^s \, d_{t-1+s} + \rho \, \phi \, E_{t-1} \sum_{s=1}^{\infty} \delta^s \, x_{t-1+s} + \beta \, w_t + \varepsilon_t \end{aligned} \tag{17}$$

In order to generate the regressors that incorporate consumers' expectations in equation (17) we make use of the VAR approach suggested by Campbell (1994)[4]. This VAR approach permits a very general interaction between house prices and rents – and the remaining included variables – as we do not impose ex-ante any direction of causality between one and the other.

[4] See also Dumas, Harvey and Ruiz (2003) and Rodríguez, Restoy and Peña (2002)

Thus, we define $y_t \equiv [y1_t \mid y2_t]$ as a k-vector where $y1_t = [d_t, x_t]$ and $y2_t$ is a (k-2) vector incorporating other variables which help predict $y1_t$. As shown in Campbell and Shiller (1988), we can re-write any VAR for y_t as a VAR (1) model for z_t, where z_t includes y_t and *p-1* lags of the variables in this vector, *p* being the number of lags in the original VAR. Denoting this transformed VAR(1) by

$$z_{t+1} = a + Az_t + \xi_{t+1} \tag{18}$$

we can easily compute the expectations terms in (17) as

$$E_t \sum_{s=1}^{\infty} \delta^s d_{t+s} = i2' \left[\delta \ (I-A)^{-1} \left[\frac{1}{1-\delta} \ \ I - A \ (I - \delta \ A)^{-1} \right] a + \delta \ A \ (I - \delta \ A)^{-1} z_t \right] \tag{19}$$

and

$$E_t \sum_{s=1}^{\infty} \delta^s x_{t+s} = i3' \left[\delta \ (I-A)^{-1} \left[\frac{1}{1-\delta} \ \ I - A \ (I - \delta \ A)^{-1} \right] a + \delta \ A \ (I - \delta \ A)^{-1} z_t \right] \tag{20}$$

where *i2* (*i3*) is a *k* vector made up of zeros but the component corresponding to the position of d_t (x_t) in z_t, which is equal to 1. In the empirical application, δ is replaced by its sample value – i.e. $\delta = 1 - e^{-\bar{q}}$.

As to the estimation methodology, it is worth noting that both the expectations variables at *t* and w_t might be correlated with the error term. Therefore, we estimate (17) by GMM and use lags of all regressors as instruments.

Data

We use quarterly data covering the longest available common period for the three countries considered: 1987Q1-2003Q2. As usual in the literature rents are computed from the corresponding component of the consumer price index.[5] As to house prices, we have used

[5] See, among other, Mankiw and Weil (1989) or Case and Shiller (1989, 1990). For a critical view on this approach, see Clayton (1996).

the average price per square metre of all dwellings released by the Spanish *Ministerio de Fomento*, the average house sale price released by the Office of the Deputy Prime Minister, and the median house price index released by the OFHEO, for Spain, the United Kingdom and the United States, respectively. Both rents and house prices have been deflated by the corresponding domestic CPI indices excluding shelter. It is worth noting that because of the heterogeneity of data sources, variables in levels (including *q*) are not directly comparable.

As to the variables in $y2_t$ (i.e. the candidates to help predict rents and consumption growth) we have included (the adequate stationary transformations of) GDP, (net) financial and non-financial wealth and a 10-year interest rate.[6] They are standard arguments in the demand functions for both non-durable consumption and housing. Data have been taken from the websites of the domestic statistics offices – and deflated by the domestic CPI index excluding shelter.[7] Finally, the set of candidate variables to proxy transitory demand pressures – i.e. the vector w_t in equation (16) – includes (deviations from the mean of the stationary transformation of) GDP and the (sing-changed-deviation-from-its-mean) return on two alternative investments: a standard broad stock index[8] and changes in financial wealth, used as a proxy for a broader return index also including corporate bonds and public debt.

Table 1 shows some descriptive statistics for the main variables used in the analysis, while Charts 1 to 3 depict the path of (deflated) house prices, (deflated) rents and the (ln) price-dividend ratio corresponding to housing investment in the three countries considered. As expected, real house prices display an upward trend throughout the available samples.[9] This trend is more marked in Spain and the UK – where prices have multiplied respectively by 2.5 and 2.4 since the beginning of the sample – than in the United States – where they have multiplied by 1.3. Prices have fluctuated more, too, in the former two countries. Real rents, which also display an upward trend, have increased more markedly in the United Kingdom and Spain too, although the rises have been lower than those in house prices – 49%, 36% and 12%, respectively, in the United Kingdom, Spain and the United States. As a consequence, price-dividend ratios have grown in all three countries although to a higher

[6] Replacing the latter by a 3-month interbank interest rate does not significantly affect the main results of the analysis.
[7] Except for Spanish non-financial wealth, which has been taken from internal estimates at the Banco de España. See Estrada and Buisán (1999).
[8] Namely, *Índice General de la Bolsa de Madrid* (Spain), *FTSE-All* (United Kingdom) and *S&P 500* (United States).
[9] See, for instance, Englund and Ionnides (1997), who find that trend in their analysis of 15 developed countries.

extent in Spain and in the United Kingdom than in the United States. Finally, it is worth noting that the fluctuations in this ratio have also been important in Spain and the United Kingdom and less acute in the United States.

Results

Regarding the first step of our empirical approach, Table 2 shows the main results of the estimation of the three different VARs focusing on the two equations of interest (namely, those for x_t and d_t). The number of lags has been chosen according to the Akaike criteria, after testing that no residual correlation was left.

Our VAR models explain better changes in rents than in real consumption, as revealed by the corresponding residual standard deviations, which are reasonably low – below 1% in all cases. Standard goodness-of-fit tests, on the other hand, are comfortably passed, as shown in the table.

As to the second step, the main results of the GMM estimates are shown in Table 3, where, first of all, it is worth noting that the different models satisfactorily pass the usual goodness-of-fit tests. Thus, there are no signs of residual autocorrelation and the results of the Sargan test do not reject the overidentifying restrictions imposed on any of the three models.

There are some differences between the three equations worth noting. Thus, in the United States the gap between the observed price-dividend ratio and its equilibrium is not persistent. In the other two countries, however, the gap exhibits important persistence – ϕ is estimated around 0.9. This might be related to a higher flexibility in the American market as suggested by Meen (2002). In the three countries, however, the gap responds to (sign-changed) changes in the real return on alternative financial investments, the coefficient being, as predicted, positive – that is, an increase in the real return on alternative assets tends to reduce housing demand and, therefore, the gap[10].

As was explained in section 2, τ is a non-structural parameter which depends on η, *b* and ρ. While the latter is directly estimated, we cannot retrieve from our results an estimate of

[10] As far as expectations in the empirical counterpart of equation (17) are generated regressors, the p-values of the different coefficients in table 3 could be underestimated depending on the correlation between the coefficient estimators in equations (17) and (18). It is also worth noting, however, that both equations are not estimated jointly but in two steps, which reduces efficiency and therefore pushes up standard errors.

η. Yet for both Spain and the UK, we can use the estimates of τ and ρ to support the use of a CES aggregator instead of the simpler Cobb-Douglas approach because for $\eta = 1$ our estimates of ρ would imply a strictly positive τ. As a matter of fact, non-positive values of τ point towards $\eta < 1$, that is, towards some degree of complementarity between consumption goods and housing services.

ρ varies between .53 in the United Kingdom and 7.19 in USA. In Spain, it is estimated at 1.73. The corresponding elasticities of intertemporal substitution (1.9, 0.58 and 0.14) are somewhat higher than those estimated in the literature from aggregated consumption data for the UK and Spain, which oscillate around 0.1-0.2.[11] For the latter, it is however still lower than the estimates based on micro studies, which tend to be around 1.[12]

To assess the consistency between current house prices and their fundamentals, it is worth starting by noting that apart from the observed and the equilibrium price-dividend ratios we may build a third price-dividend series: the *adjusted* price-dividend, defined as the observed one minus the estimated error in equation (17). Thus, while an observed ratio above the estimated equilibrium value would imply some overvaluation, an observed ratio above the adjusted one would mean that it is reverting to its equilibrium quicker (if it is below equilibrium) or more slowly (if it is above equilibrium) than implied by the equation.[13] Charts 4.1 to 4.3 show the different price-dividend series (observed, adjusted and the estimated equilibrium) for each country.

In all three countries, the increases in the last years have brought house price-to-rent ratios above their equilibrium values. In Spain, the equilibrium ratio has kept relatively stable, although it exhibits a slight upward trend since the mid nineties. In addition, part of the increase in real house prices since 1998 may be seen as a return to equilibrium values. As a matter of fact, the fall after the peak in 1991 resulted in relatively severe overreaction. Nevertheless, in 2001 real house prices already exceeded the estimated equilibrium and in 2003 they show – in relation to rents – an overvaluation close to 20%, the highest in the available sample.[14] In addition, while in 2001 and 2002 price-to-rent ratios grew on average

[11] See, for instance, Hall (1988), Campbell (1999) and, for Spain, López-Salido (1995).
[12] For the United Kingdom and the United States, see Attanasio (1999) and the references therein. For Spain, see López-Salido (1995).
[13] In the United States, however, the adjusted ratio and the equilibrium ratio coincide, as we have not been able to model the gap.
[14] Martínez-Pagés and Maza (2003) estimate a simple reduced-form equation for house prices in Spain and finds a similar overvaluation.

along the adjusted path, in 2003 – during its first two quarters, at least – they grew above such a path.

In the United Kingdom the picture is somehow similar. After some overreaction following the previous peak, the price-to-rent ratio has been above the equilibrium path during the 2000s. Estimated overvaluation is around 30%, on average, in 2003 and there are no signs of a reversion towards equilibrium in this year.[15] It is worth noting, too, that the equilibrium ratio is currently rather similar to that in 1987.

Finally, house prices in relation to rents also seem to be overvalued in USA, albeit to a lesser extent: around 10%. This overvaluation, however, also reaches its peak at the end of the sample. Contrary to Spain and the United Kingdom, overreactions are less marked[16] in this case. The equilibrium ratio exhibit a slight upwards trend since the late 1990s as in Spain. As a matter of fact, the increase in the observed ratio between 1999 and 2002 may be explained in terms of a similar increase in its equilibrium value.

A potentially relevant caveat must however be made. Both in the theoretical model and in its empirical application we have not considered taxes. In principle, the equilibrium relationship (14) should only hold for the after-tax values of prices and rents. The analysis of the impact of tax-subsidy national systems on the transaction and rental value of prices for a representative agent is a complex task that falls outside the scope of the paper. But it is worth noting that, assuming tax changes are essentially unpredictable by economic agents, that feature would only affect the value of the intercept term *m'* in equation (14). More importantly, making the plausible assumption that discretionary variations in taxes or subsidies are unrelated to the right-hand side variables in equation (17) would imply that our estimates of the slope parameters would be robust to those changes in taxation. Finally, even if we could estimate the impact of tax changes on prices and rents, it is not clear how taxation could be fully incorporated into our intertemporal framework. Additional non-neutral assumptions would have to be made on other variables for both private agents' and government budget constraints to hold at all times.

4. CONCLUSIONS

[15] Both the magnitude and the timing of the overvaluation are very similar to those reported by some private analysts. See, for instance, Goldman Sachs (2002).
[16] Meen (2002) finds that the price elasticity of house supply is higher in the United States than in the United Kingdom. This might explain why overreactions are less important in the former.

House prices have increased markedly in many industrialised countries in recent years, apparently backing the strength of households' consumption in a context of otherwise weakening activity. Against this background, this paper has sought to assess whether current price levels are broadly consistent with their economic fundamentals.

Thus, we have used a relatively general intertemporal asset pricing model where housing services and consumption are non-separable and, therefore, they both affect the equilibrium stochastic discount factor of future payoffs. In order to take into account slow adjustment dynamics of housing prices to changing conditions, we have adopted an approach that allows for systematic transitory deviations of price-to-rent ratios from equilibrium conditions. We have then made use of quarterly data spanning the period 1987-2003 on three different countries – Spain, the United Kingdom and the United States – where house price increases have been particularly sizable in recent years to estimate the potential overvaluation of housing in relation to rents in each case.

According to our results, which given the approximations implicit in the model and the difficulties of taking into account all possible frictions present in this market have to be taken with due caution, part of the increase in real house price-to-rent ratios during the late nineties can be seen as a return to equilibrium following some overreaction of house prices after previous peaks. This overreaction was particularly marked in Spain and the United Kingdom. However, more recently, marked increases in house prices led price-to-rent ratios to well above equilibrium in all three countries by 2003. More specifically, the price-to-rent ratios were around 30% above equilibrium in the United Kingdom, around 20% in Spain and around 10% in the United States. It is worth noting, however, that part of that overvaluation – particularly in Spain and the United Kingdom – may be attributable to the sluggishness of supply in the presence of large demand shocks in this market and/or the slow adjustment of observed rents to the conditions prevailing in the housing market.

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Table 1. Descriptive statistics
Sample period 1987Q2-2003Q2

	SPAIN			UK			USA		
	Dlhp	Dlr	q	Dlhp	Dlr	q	Dlhp	Dlr	q
Mean	0.0149	0.0044	4.7882	0.0123	0.0065	4.2625	0.0036	0.0018	4.1581
Median	0.0135	0.0027	4.7679	0.0130	0.0065	4.2496	0.0038	0.0019	4.1448
Maximum	0.0672	0.0272	5.0940	0.1466	0.0275	4.5913	0.0197	0.0160	4.2824
Minimum	-0.0563	-0.0196	4.4746	-0.1280	-0.0057	4.0358	-0.0252	-0.0165	4.0887
Std. Deviation	0.0242	0.0082	0.1423	0.0479	0.0070	0.1525	0.0090	0.0060	0.0514
Skewness	-0.0658	0.2892	0.0714	-0.1763	0.5820	0.4022	-0.4808	-0.0801	0.7271
Kurtosis	3.1418	3.9536	2.2378	3.5576	3.5029	2.0584	3.4438	3.2583	2.6621
Jarque-Bera	0.1030	3.4204	1.6537	1.1969	4.421	4.2182	3.0843	0.2540	6.1299
Probability	0.9498	0.1808	0.4374	0.5497	0.1096	0.1213	0.2139	0.8807	0.0467

Notes:
- Dlhp stands for the first difference of (ln) house prices deflated by a CPI index excluding shelter
- Dlr stands for the first difference of (ln) rents deflated by a CPI index excluding shelter
- q stands for (ln) ratio of house prices to rents

Table 2. VAR estimates
Endogenous variables: rents, consumption, GDP, (net) financial and non-financial wealth and a 10-year interest rate. All are in real terms. All but the last are in first (ln) differences.
Sample period: 1987Q1:2003Q2

	SPAIN		UK		USA	
	Rents	Consumption	Rents	Consumption	Rents	Consumption
R^2	0.66	0.58	0.70	0.56	0.48	0.41
σ	0.005	0.006	0.004	0.007	0.005	0.006
Q1	1.10 (0.29)	0.17 (0.68)	1.89 (0.17)	0.03 (0.88)	0.03 (0.87)	0.00 (0.98)
Q4	3.33 (0.50)	3.90 (0.42)	11.39 (0.03)	2.67 (0.62)	2.31 (0.68)	0.13 (1.00)
Q8	4.54 (0.81)	7.66 (0.27)	16.35 (0.04)	4.58 (0.80)	9.99 (0.27)	4.13 (0.85)

Notes:
- All VARs include four lags of the endogenous variables.
- Qi stands for the standard test on residual autocorrelation up to order i. P-values in brackets.

Table 3. GMM estimates

$$q_t = m'(1-\phi) + \phi\, q_{t-1} + \tau\, E_t \sum_{s=1}^{\infty} \delta^s d_{t+s} - \rho\, E_t \sum_{s=1}^{\infty} \delta^s x_{t+s}$$

$$- \tau\, \phi\, E_{t-1} \sum_{s=1}^{\infty} \delta^s d_{t-1+s} + \rho\, \phi\, E_{t-1} \sum_{s=1}^{\infty} \delta^s x_{t-1+s} + \beta\, w_t + \varepsilon_t$$

Sample period: 1987Q3:2003Q2

	SPAIN	UK	USA
m'	0.70 (.36)	0.34 (.20)	6.27 (1.44)
τ	-2.03 (.78)	--	7.16 (4.61)
ρ	1.73 (.69)	0.53 (.35)	7.19 (4.70)
ϕ	0.91 (.04)	0.93 (.04)	--
β	1.09 (.34)	0.66 (.28)	1.10 (.52)
Sargan test	7.05 (.13)	1.93 (.86)	12.20 (.03)
σ	0.043	0.055	0.088
Q1	0.20 (0.65)	0.30 (0.58)	0.55 (0.45)
Q4	2.18 (0.70)	1.97 (0.74)	4.57 (0.33)
Q8	2.51 (0.96)	6.45 (0.60)	6.70 (0.57)

Notes:
- Instruments: Lags 1 and 2 of all variables in the LHS
- Standard errors (first block) and p-values (second block) in brackets. Both robust to heteroscedasticity
- w = -(real log) rate of growth of households' financial wealth beyond its sample average
- Qi stands for the standard test on residual autocorrelation up to order i

Chart 1

Real house price indexes. Annual averages (1987Q1=100)

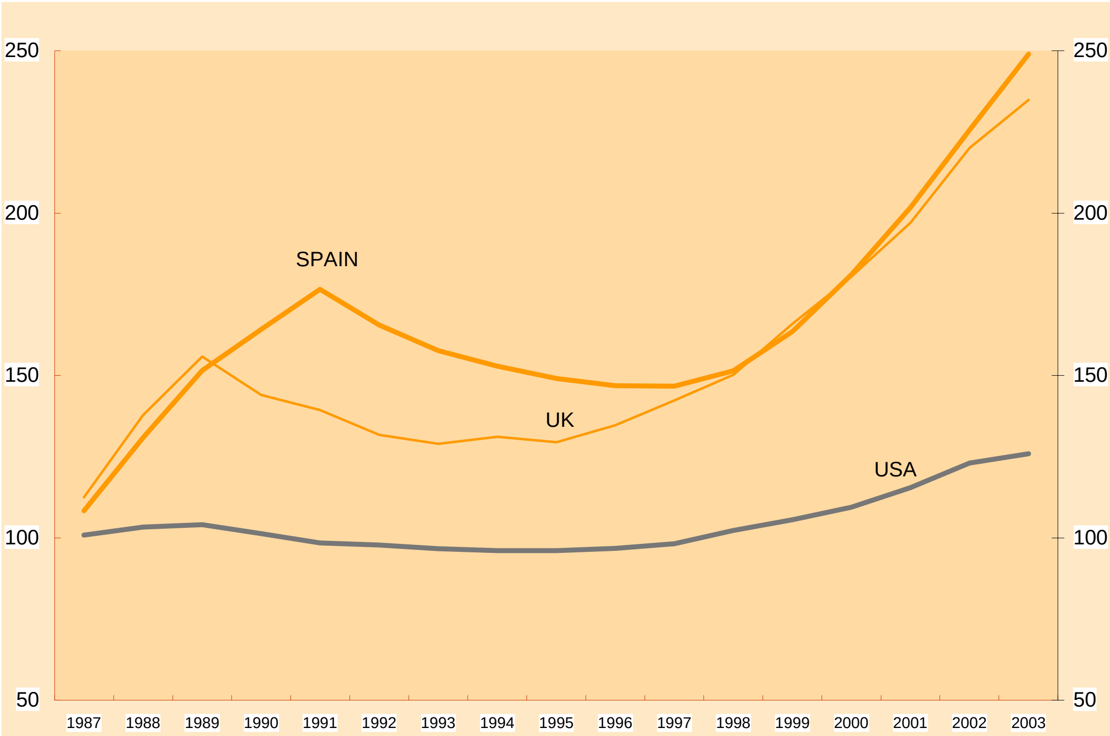

Chart 2

Real rent indexes. Annual averages (1987Q1=100)

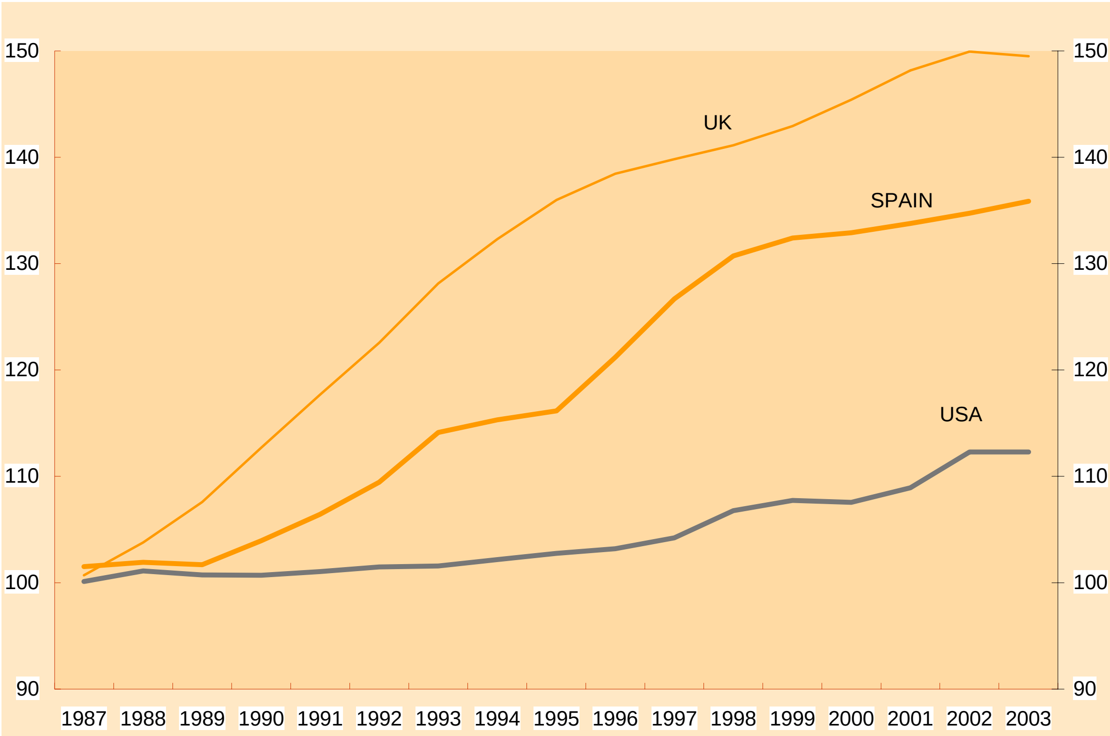

Chart 3

(Ln) House price-to-rent ratio. Annual averages (1987Q1=100)

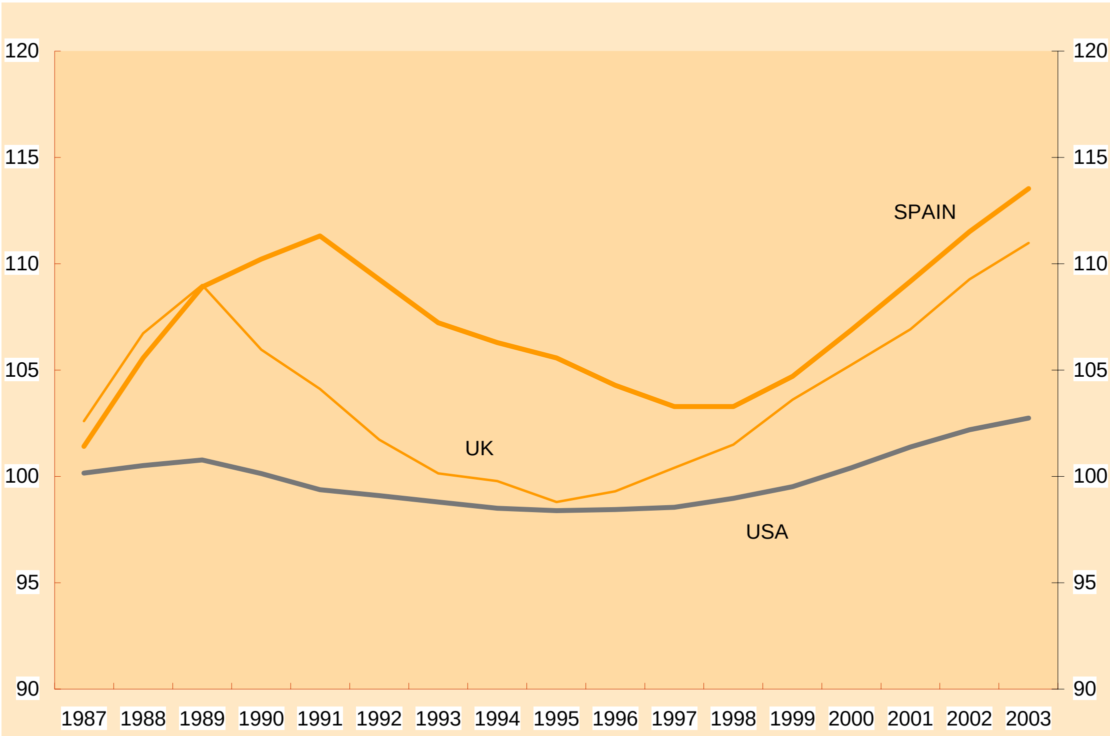

Chart 4.1

(Ln) House prices / rents. Spain

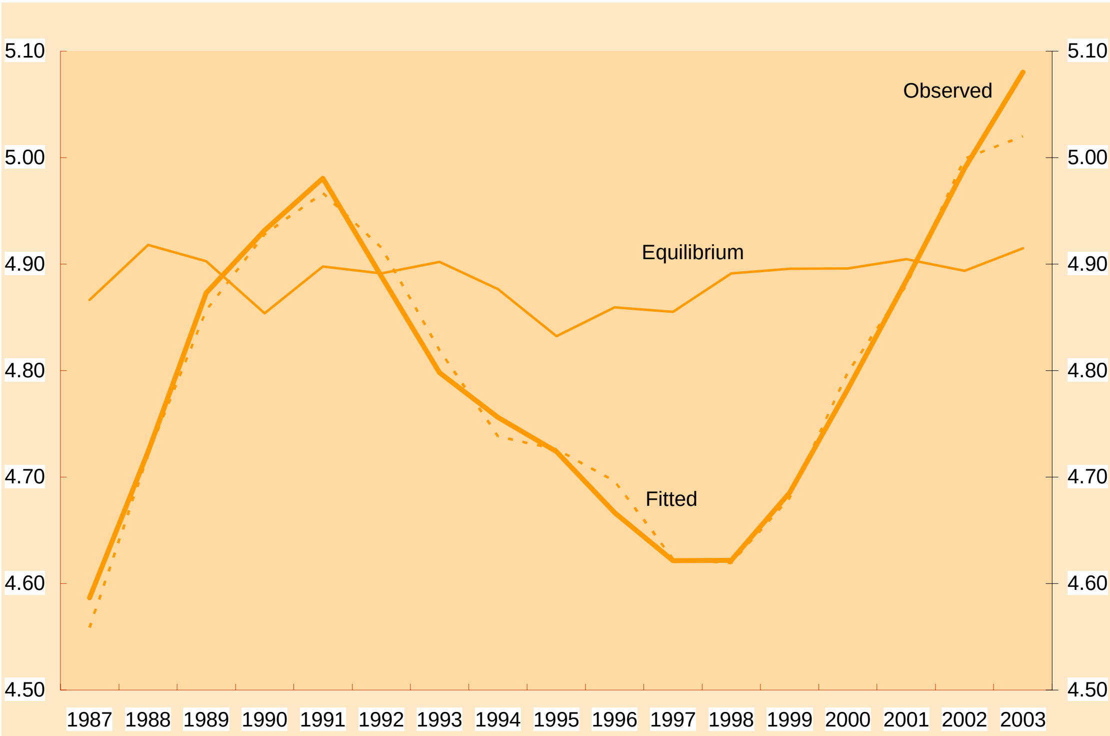

Chart 4.2

(Ln) House prices / rents. UK

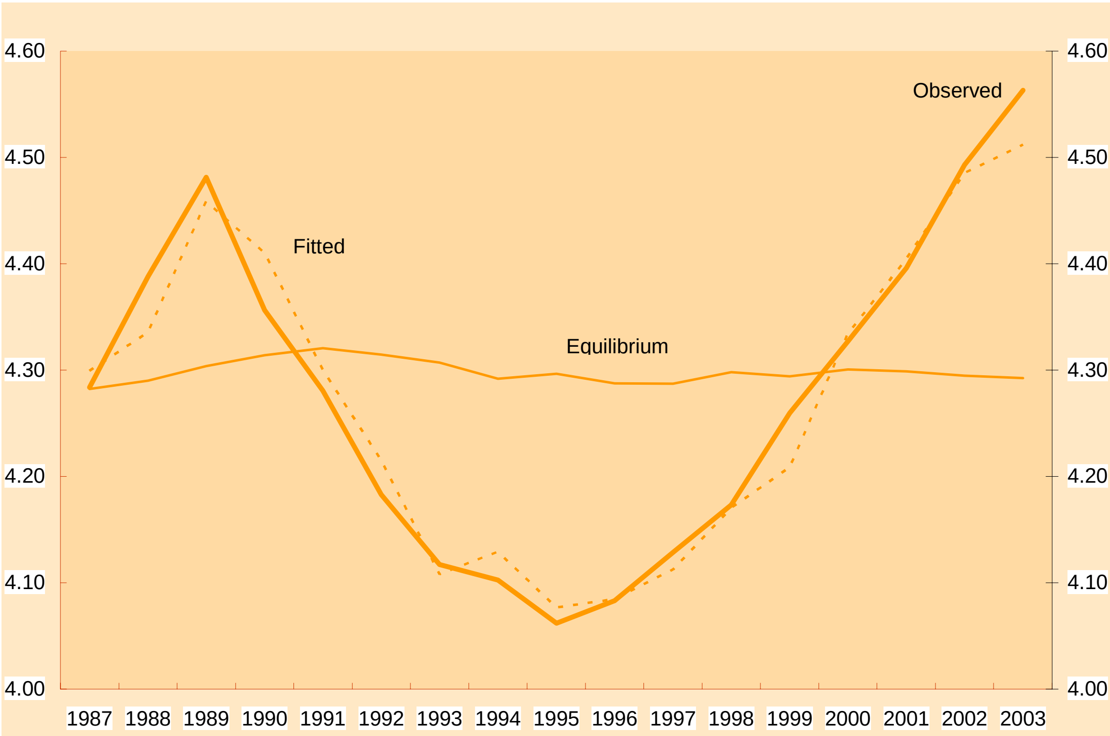

Chart 4.3

(Ln) House prices / rents. USA