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Abstract

This paper analyses the contribution of interest rates to explain recent house price developments in Spain trying to reconcile different pieces of evidence. On the one hand, empirical evidence supports the view that interest rates are a key variable to explain house price developments. As a matter of fact, using simple asset pricing relations recent changes in house prices could be fully explained by movements in ex-post real interest rates. However, more refined asset pricing models show that the changes in the discount factor cannot fully explain the recent course of house prices in Spain. To resolve this puzzle we provide evidence that shows that the actual real cost of financing might have decreased significantly less than what the course of ex-post real rates would suggest.

JEL: E43, G12

Keywords: house prices, real interest rates, intertemporal marginal rate of substitution, stochastic discount factor

1 Introduction

One of the most relevant developments in the Spanish economy over the last 15 years has been the sharp reduction in nominal interest associated with the process of nominal convergence and EMU membership. Although inflation rates have also declined substantially over the same period, inflation-adjusted interest rates – often called ex-post real rates – have fallen by almost 10 percentage points since 1990 (see Graph 1). Clearly, using this variable as an indicator of the cost of capital for domestic agents, we can identify a huge reduction in financing costs and expect a substantial impact on agents' real and financial decisions and on asset prices.

Indeed, the Spanish economy has experienced significant transformations in the recent past which are all consistent with a substantial reduction in financial costs. In particular, in 2005 the household saving ratio was around four percentage points lower than the average over the first half of the previous decade. Net financial saving (the change in financial assets minus the change in financial liabilities) of households and firms has fallen dramatically. In the former case net financing saving has even changed its traditional positive sign to a negative figure. The debt of the private non-financial sector has risen to 160% of GDP, more than twice the 1995 ratio. In addition, the economy has recently witnessed a substantial real-estate boom which has led housing prices to increase by more than 100% in real terms since 1997. Finally, economic activity – heavily supported by domestic demand, especially consumption and investment in construction – has increased markedly in the last few years, with GDP growth averaging more than 3.5% since 1991.

Estimating the impact of lower interest rates on agents' balance sheets, asset prices and associated macroeconomic developments is not an easy task. For one thing, the economy has also faced other relevant structural changes. In particular, some labour market reforms and intensive immigration flows have reduced supply-side rigidities and contributed to substantial employment creation. These developments, together with the consolidation of an environment of macroeconomic stability within EMU, have prompted an upward revision of consumers' permanent income and reduced investors' uncertainty. Like low interest rates, these structural factors contribute to higher expenditure propensity and demand for financing.

In this paper we summarise and exploit the results in Ayuso and Restoy (2006a, 2006b) and Blanco and Restoy (2006) to analyse the contribution of interest rates to explain recent house price developments in Spain. In particular, the available empirical evidence supports the view that the interest rate is a key variable to explain recent house price developments in Spain. As a matter of fact, using simple valuation relations recent changes in house prices could be fully explained by movements in ex-post real interest rates. However, more refined asset pricing models show that changes in the discount factor alone do not explain the recent course of house prices in Spain. This puzzle is solved by providing evidence supporting the hypothesis that the observed decrease in ex-post real interest rates since mid 1990s is likely to overestimate the fall in the ex-ante real interest rates experienced by the Spanish economy.

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1. See Malo de Molina and Restoy (2005) for an analysis of the main financial developments affecting the Spanish economy.

The paper is structured as follows. In the second section we review the empirical results on the relationship between house prices and interest rates, including those which make it explicit the role of credit. In the third we exploit the long-term equilibrium condition between rents and house prices to assess the recent movements of the latter variable. In the fourth we provide some evidence on the recent course of ex-ante real interest rates in Spain by combining several macroeconomic and financial approaches. Section 5 concludes.

2 Some empirical evidence of the relationship between house prices and interest rates

From a theoretical perspective, the interest rates are a key variable to explain house prices through its effects on user cost but also on credit availability. Thus, the permanent income hypothesis states that spending decisions by individuals depend on permanent income and real interest rates. In this context, borrowing is the mechanism which allows agents to smooth their spending intertemporally and therefore adjusts passively to spending decisions. However, in the presence of liquidity constraints current income rather than permanent income and nominal rather than real interest rates have a role to play in explaining credit developments. Indeed, one of the key criteria banks apply in granting loans is the initial debt burden (ratio of interest payments plus repayment of principal over income), which depend on the level of the nominal interest rates. It is also worth noting that asymmetric information problems imply that borrowing capacity is affected by changes in house prices since they determine the collateral available for bank lending.

The available empirical evidence confirms that the interest rates are a key variable to explain credit developments in Spain. Nieto (2003) and del Río (2002) study the determinants of household borrowing estimating a single-equation error correction model. They find that the long-term equilibrium level of household borrowing (debt ratio in the second case) is negatively related to nominal interest rates (debt burden in the second case) and unemployment and positively related to spending (gross domestic income in the second case). As for the short-term dynamic of household borrowing, the specification include, among other variables, housing wealth (house prices in the second case) and the error correction term. The evidence reported in both papers suggests that one of the main drivers of household debt between mid 1990s and the beginning of 2000s is the interest rate (debt burden in the second case).

Also, Martínez-Carrascal and del Río (2004) estimate a Vector Error Correction Model (VECM) where consumption and household borrowing are modelled jointly. The results show that both consumption and borrowing are positively related to financial and/or housing wealth and labour income in the long term, and negatively related to nominal interest rates. They also find that disequilibria in the loan market are restored through adjustments in both consumption and borrowing. By contrast, when consumption is not in the long-term equilibrium level implied by fundamentals adjustment takes places only through movements in this variable (and income). Finally, the evidence they provide suggests that household borrowing developments between mid 1990s and mid 2000s can be largely explain by changes in the long-term fundamentals and, among them, the interest rates play a crucial role.

As to house prices, the expected return on the asset ”house” has to be equal to the return on an alternative investment with a comparable level of risk. As shown in Poterba (1984 and 1991), this condition, together with the equilibrium condition in the market for housing services, implies that real house prices depend on income, the housing stock and the user cost, which is the alternative return on investments with the same level of risk minus the expected increase in house prices net of depreciation. In empirical applications the latter variable is normally proxied by the risk-free interest rate due to difficulties in estimating both the expected increase in house prices and the return on alternative investments with the same risk.

Martínez-Pagés and Maza (2003) analyse the determinants of house prices in Spain following Poterba (1984, 1991) but using single-equation error correction models which permit them to distinguish between short-term and long-term equilibria. Interestingly, they find that when the parameters of the model are unrestricted house prices are only determined in the long run by income, with interest rates playing no role at all. However, they argue that the high correlation between income growth and nominal (and real) interest rates creates an identification problem that prevents them from properly estimating the effect of these variables on the long-term equilibrium house prices. When they estimate a model imposing that long-run income elasticity is equal to one, interest rates appear significant with a semi-elasticity of -4.5.

Both the restricted and the unrestricted models suggest that a large part of the rise in house prices in Spain between 1997 and 2002 can be explained by changes in the fundamentals but they also find that the expansionary behaviour of house prices also reflected an adjustment towards their long-term equilibrium level since in mid 1990s houses were somewhat undervalued. In the same way, by 2002 prices had increased beyond what was needed to restore their long-term equilibrium level leading to an overvaluation between 8% and 17%, compatible with the normal short-term dynamic adjustment of the Spanish housing market.

More recently, Gimeno and Martínez-Carrascal (2006) analyse the dynamic interaction between house prices and loans for house purchase using a VECM. Their results show that in the long run housing credit is positively related to house prices and income and negatively related to nominal interest rates and house prices are positively related to housing credit and income. They also find that when credit departs from the level implied by its long-term determinants the equilibrium is restored through movements not only in this variable but also in house prices. By contrast, house price disequilibria are fully corrected by changes in this variable. As for short-run dynamics, the evidence they report suggests that causality between house prices and loans goes in both directions, indicating the existence of mutually reinforcing cycles in both variables. Changes in house prices since mid 1990s can be largely explained by changes in fundamentals, among which nominal interest rates play a key role. However, in 2004 house prices were above their long-term equilibrium level, suggesting the existence of overvaluation.

3 Modelling house prices under a financial approach[2]

The relevance of interest rate changes to explain recent house price behaviour in Spain can be illustrated by analysing the latter from a purely financial approach which exploits the parallelism between a house which provides rents (or housing services) and a financial asset which provides different payoffs during a long period. The simplest case under this approach is the well-known Gordon dividend discount model. In this model the long-term equilibrium level of real house prices is derived as the present value of future real rents, discounted using a constant discount factor. Usually the discount factor is obtained adding a constant risk premium to the ex-post real interest rate:

$$\frac{D}{P} = r + RP - d \tag{1}$$

where *D* stands for real rents, *P* is the real house price, *r* is the risk-free real interest rate, *RP* is the risk-premium on the housing asset and *d* stands for the future growth rate of real rents. Expression (1) can be rewritten as:

$$P = \frac{D}{r+k} \tag{2}$$

where *k* stands for the spread $RP - d$. Given *D* and *k*, expression (2) can be used to estimate the impact on prices of changes in interest rates. The non-linear relationship between *P* and *r* implies that this impact is highly sensitive to the level of both *r* and *k*. The spread k can be estimated from (1) using information on the rent to price ratio and the real interest rates. Proxing the latter by the inflation-adjusted long-term bond yield[3] we obtain that in Q4 1997 *k* was approximately zero (both the inflation-adjusted long-term interest rates and the rent to price ratio were around 4%). Between Q4 1997 and Q4 2004 the inflation-adjusted interest rate fell by 3.5 pp. Using expression (2), as rents have not varied much in real terms, such a reduction in the real interest rates implies a 700% increase in real house prices for a nil value of *k*, much more than the 110% observed. Therefore, this simple valuation relation suggests that the movement in the real interest rates, proxied by the inflation-adjusted long-term bond yield, can more than explain the sharp rise of house prices in Spain between 1997 and 2004.

The static relation (1), however, is valid only under highly restrictive conditions including constant expected real rent growth and constant interest rates. Moreover, it does not consider any possible discrepancy between observed prices and fundamentals due to frictions that could prevent the immediate adjustment of supply and demand in this market. Finally, it also ignores the possible effects of taxes on the equilibrium relationship between house prices and rents.

As shown in Ayuso and Restoy (2006a, 2006b), these deficiencies can be partially solved by employing more general intertemporal asset pricing models to estimate the dynamic equilibrium relationship between house prices and rents. Possible transaction costs and regulations that hinder the immediate adjustment of house prices and rents to changing

2. This section heavily draws from Ayuso and Restoy (2006b).
3. The nominal bond yield is adjusted using the most recently published annual inflation rate.

conditions can also be taken into account by specifying a dynamic relationship that permits transitory but persistent deviations from equilibrium. The effects of taxes are more difficult to incorporate in the analysis, although it is possible to indirectly assess its likely role in the most recent developments in house prices in Spain.

3.1 The set-up

Following Campbell and Shiller (1988), we can approximate the (net) real return of any asset (say a house) between *t* and *t+1*, r_{t+1}, as

$$r_{t+1} \approx d_{t+1} + q_{t+1} - \frac{1}{\delta} q_t - k \tag{3}$$

where d_{t+1} is the real growth rate of dividends (rents) between *t* and *t+1*, q is the (log) asset price-dividend ratio and δ and k are linearisation constants, the former playing the role of a discount rate as will be explained later on.

We will now assume that the risk class of housing (understood as the equilibrium risk premium over any other asset) is constant.[4] For any portfolio *m,* let us then define π_m as

$$\pi_m \equiv E_t\ r_{t+1} - E_t\ r_{m,t+1}, \tag{4}$$

where r_m stands for the (log) real return on the portfolio *m*.

Solving (3) forward for q_t, assuming that bubbles are unfeasible, taking expectations on both sides of the resulting expression, and using (4) yields

$$q_t^* = -\left(k + \pi_m\right) \frac{\delta}{1-\delta} + E_t \sum_{s=1}^{\infty} \delta^s \left(d_{t+s} - r_{m,t+s}\right). \tag{5}$$

Therefore, we can approximate (up to a constant) the price-to-rent ratio as the present discounted sum of future expected rents minus the present discounted sum of future expected returns on any asset or portfolio of assets. The latter plays the role of the stochastic discount factor that should be applied to future payoffs.

As a particular case, we could consider as the reference portfolio a claim on future consumption growth, which in a standard representative agent economy will be equal in equilibrium to the economy's aggregate (wealth) portfolio. If intertemporal preferences are assumed to be of a generalised isoelastic form (see Epstein and Zin, 1989, and Weil, 1990), Restoy and Weil (1998) show that the equilibrium relationship between the return on the wealth portfolio (r_w) and consumption growth (x) can be approximated by the simple linear expression

$$E_t\left(r_{w,t+1}\right) = u + \rho\ E_t\ x_{t+1} ,$$

4. This is obviously the case in standard asset pricing models where returns are homoscedastic. Note that the expected capital gain on housing does not need to be constant as typically assumed in the standard empirical specifications of the relationship between house prices and rents.

where ρ is inverse of the elasticity of intertemporal substitution (which under GIP preferences may not to be equal to the relative risk aversion coefficient) and u is a constant. Ayuso and Restoy (2006a) show how this model could be extended to the case in which preferences are non-separable between good consumption and housing services and derive equilibrium house prices in that set-up. In particular, they provide the following equation for the equilibrium price-to-rent ratio in the housing market:

$$q_t^* = h + \tau\, E_t \sum_{s=1}^{\infty} \delta^s\, d_{t+s} - \rho\, E_t \sum_{s=1}^{\infty} \delta^s\, x_{t+s} \qquad (6)$$

where $h = \left(-k + \pi_w + u\right)\dfrac{\delta}{1-\delta}$ and τ is a constant that depends on both the elasticity of intertemporal substitution and the elasticity of substitution between housing and consumption.

Therefore, in this setting the price-rent ratio is 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. 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. It is worth noting that when $\rho = 0$, equation (6) collapses to the standard expression where prices are a function of future expected payoffs discounted at a constant rate δ.

Both expressions (5) and (6) are suitable for the empirical determination of the equilibrium house-to-price ratio. The latter permits a more genuine equilibrium analysis as housing prices are solely determined by rents and macroeconomic developments. It does however rely on a specific parameterisation and on relatively demanding intertemporal equilibrium conditions of a representative agent. The former only requires assuming that there are no arbitrage opportunities and that the risk premia of all assets are stable over time.

Therefore, it seems justified to explore the extent to which the estimation of the equilibrium price-to-rent ratio depends on the asset pricing model employed and, in particular, to investigate whether estimates vary with the stochastic discount factor applied to future expected rents. As this seems to be in essence an empirical question, we propose computing equilibrium prices according to both equations (5) and (6). The former will also admit several versions depending on the definition chosen for the reference portfolio *m*. We will then be able to assess the differences found and to identify those results which are robust to the choice of a particular discount factor. Also, before estimating long-run equilibria we have to enlarge our equations in order to capture the short-term dynamics arising from extant rigidities in this market.

3.2 The empirical approach

Let the *observed value* of the the (log) ratio of house prices to rents be denoted by q_t to distinguish it from the long-run *equilibrium* ratio q_t^*, which is the value consistent with expressions (5) and (6). Thus,

$$q_t = q_t^* + g_t \qquad (7)$$

where g_t is the gap, at *t*, between the observed ratio and its equilibrium value. The rationale for this gap can be found in both the rigidities that affect the adjustment of quantities in the property market as well as the slow adjustment of average prices in the rental markets.

As 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 and 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 behaviour. Ortalo-Magné and Rady (2006) 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.

By contrast computed rents tend to vary much less. 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 with some stickiness.

As obtaining an explicit theory-based expression for g_t is beyond the scope of this paper, we adopt a purely empirical approach. In particular, we characterise g_t as follows

$$\Phi(L)g_t = \beta\ w_t + \varepsilon_t, \tag{8}$$

where $\Phi(L)$ is a standard polynomial of order *p* in the lag operator *L* that meets the usual stationarity conditions, w_t is a zero-mean stationary variable capturing demand pressures (so that β is expected to be positive) and ε_t is standard iid white noise. Note that according to expression (8), prices can deviate from equilibrium only transitorily, although deviations may show some degree of persistence. In this respect, the slow adjustment of q_t towards its long-run equilibrium level after a shock resembles the adjustment pattern that characterises the behaviour of rigid variables in standard overshooting models.

Combining (6) and (8) we obtain:

$$\begin{aligned} q_t &= g(1-\sum_{i=1}^{p}\phi_i) + \sum_{i=1}^{p}\phi_i q_{t-1} + \tau\ E_t\sum_{s=1}^{\infty}\delta^s\ d_{t+s} - \rho\ E_t\sum_{s=1}^{\infty}\delta^s\ x_{t+s} \\ &-\tau\sum_{i=1}^{p}\phi_i\ E_{t-1}\sum_{s=1}^{\infty}\delta^s\ d_{t-1+s} + \rho\sum_{i=1}^{p}\phi_i E_{t-1}\sum_{s=1}^{\infty}\delta^s\ x_{t-1+s} + \beta\ w_t + \varepsilon_{1t} \end{aligned} \tag{9}$$

where $g = k + \pi_m$.

Likewise, combining (5) and (7) yields:

$$
\begin{aligned}
q_t = h(1-\sum_{i=1}^{p}\ \phi_i) + \sum_{i=1}^{p}\ \phi_i q_{t-1} + E_t\ \sum_{s=1}^{\infty}\delta^s\ d_{t+s} - \ E_t \quad \sum_{s=1}^{\infty}\delta^s\ r_{m,t+s} \\
-\sum_{i=1}^{p}\ \phi_i\ E_{t-1}\ \sum_{s=1}^{\infty}\delta^s\ d_{t-1+s} + \sum_{i=1}^{p}\ \phi_i E_{t-1} \quad \sum_{s=1}^{\infty}\delta^s\ r_{m,t-1+s}\ + \beta\ w_t + \varepsilon_{2t}
\end{aligned}
\qquad (10)
$$

In order to generate the regressors that incorporate consumers' expectations in equations (9) and (10) we make use of the VAR approach suggested by Campbell (1993) as adapted by Rodríguez et al. (2002). This VAR approach permits a very general interaction between house prices and rents (and the remaining variables) as we do not impose ex-ante any direction of causality between one and the other.

Thus, for each discount factor we define $y_t \equiv [y1_t \mid y2_t]$ as a k-vector where $y1_t$ is a vector of dimension 2 including d_t and the discount factor 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 a pp-dimensional variable z_t which includes y_t and *pp-1* lags of the variables in this vector. Denoting this transformed VAR(1) by

$$z_{t+1} = a + Az_t + \xi_{t+1} \qquad (11)$$

we can easily compute the expectations terms in (9)-(10) as

$$E_t\sum_{s=1}^{\infty}\delta^s d_{t+s} = i2'\left[\delta\,(I-A)^{-1}\left[\frac{1}{1-\delta}\ I - A\,(I-\delta\,A)^{-1}\right]a + \delta\,A\,(I-\delta\,A)^{-1} z_t\right] \qquad (12)$$

and

$$E_t\sum_{s=1}^{\infty}\delta^s j_{t+s} = i3'\left[\delta\,(I-A)^{-1}\left[\frac{1}{1-\delta}\ I - A\,(I-\delta\,A)^{-1}\right]a + \delta\,A\,(I-\delta\,A)^{-1} z_t\right] \qquad (13)$$

where *j* stands for the discount factor, *i2* (*i3*) is a *k* vector made up of zeros except for the component corresponding to the position of d_t (j_t) in z_t, which is equal to 1.

In the empirical application, we follow Campbell (1993) and replace δ by the sample average of the (log) consumption-wealth ratio.[5]

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 (9)-(10) by GMM and instrument these variables.

3.3 Results

We use quarterly data spanning the longest available period: 1987Q1-2004Q4. Although this is a relatively short period, it is worth noting that it covers the end of the mid-eighties boom in

5. Nothing substantial changes if we consider the alternative calibration suggested in Campbell (1993), based on the estimate of the time-preference parameter θ in $V_t = \sum_{i=0}^{\infty}\vartheta^i E_t\ U(C_{t+i})$.

the Spanish housing market, the subsequent correction and the new expansionary stage that started in the late nineties and is still going on. Thus, it seems to include a complete cycle in the market.

As usual in the literature, rents are computed from the corresponding component of the consumer price index.[6] As to house prices, we have used the average price per square metre of all dwellings released formerly by the Spanish *Ministerio de Fomento* and currently by the *Ministerio de la Vivienda*.

The return on the reference portfolio *m* in equation (10) has been proxied by three different empirical variables. As in most papers in the literature, we used first a broad stock index. In particular, we considered the return on the Ibex-35.[7] Then, we used a bond portfolio by considering the total return index released by the Banco de España, which measures the (monthly) total return on a theoretical portfolio made up of all outstanding bonds issued by the Spanish Treasury with a residual maturity longer than 1 year.[8] Finally, to include a proxy in which both stocks and bonds can play a role, we also considered the change in households' financial wealth as a proxy for the return on Spanish households' portfolio. In this regard, it is worth noting that, for quarterly data at least, it seems reasonable to expect price-change effects to dominate quantity movements in that portfolio.

As to the variables in $y2_t$ (i.e. the candidates to help predict rents and the different stochastic discount factors) we included (the adequate stationary transformations of) GDP, (net) financial and non-financial wealth, household consumption – except for the model where it is included in $y1_t$ – and the 10-year interest rate. Finally we proxy transitory demand pressure – i.e. the vector w_t in equation (8) – by the de-meaned and sign-changed changes in households' financial wealth. Thus, an increase in w_t captures the transitory positive effect on housing demand caused by a transitory fall in the price of alternative financial assets.

All variables used in the empirical estimates have been deflated by the Spanish CPI index excluding shelter.

Table 1 shows some descriptive statistics for the main variables used in the analysis, while Graph 2 depicts the path of (deflated) house prices, (deflated) rents and the (ln) price-dividend ratio corresponding to housing investment. Real house prices display an upward trend throughout the available sample, the level reached in 2004 being three times higher than in 1987. Real rents, which also display an upward trend, have increased less (38%) and, as a consequence, the price-dividend ratio has grown very sharply over the sample period. Also worth noting are the wide fluctuations in house prices and therefore in the ratio.

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6. See, for instance, Mankiw and Weil (1989) or Case and Shiller (1989 and 1990). For a critical view of this approach, see Clayton (1996).

7. We also tried the return on a world stock index (the MSCI World index) as well as a combination of returns on the domestic (until 1998Q4) and the world (since 1999Q1) stock exchanges. Both discount factors provided fairly similar results to those reported here.

8. For details on the index, see Banco de España (1991). It is also worth noting that non-government bonds have traditionally played a very minor role in Spanish households' portfolios.

The main results of the GMM estimates are shown in Table 2.[9] Here we present estimates for five specifications: i) the equilibrium ”consumption” model – as in equation (9); ii) three versions of equation (10) corresponding to the three different definitions of the reference portfolio *m*; and iv) a simple constant discount factor model. As can be seen, all but the constant discount-factor model fit the data well. Moreover, they show Sargan tests well above 10% and residual standard errors that vary between 6% and 12%. On the contrary, the results for the constant discount factor model are much poorer. The Sargan test rejects the instrument set and there is clear evidence of residual autocorrelation, maybe caused by an overestimate of the parameter ϕ_1 which violates stationarity conditions. As this does not change when more lags in $\Phi(L)$ or alternative instrument sets are considered, we interpret the results in column 5 as evidence in favour of time-varying discount factors.

Focusing therefore on the results for equations (9) and (10) we find clear evidence of strong price inertia. While one lag is enough to properly characterise the dynamics of g_t, ϕ_1 is estimated at around 0.9. Also, β is found to be significant in the four models considered. Finally, it is worth noting that the point estimate for ρ implies an elasticity of intertemporal substitution around 0.1, well in line with the results in the available literature.[10]

In any case, it is important to note that we are not interested in econometrically discriminating among the stochastic discount factors considered. Our objective is rather to analyse the implications of choosing different reasonable discount factors with a view to identifying those results which are robust across models.

To that end, we built both an estimated equilibrium price-dividend series and an estimated short-term adjustment path to equilibrium, where rigidities in the housing market are explicitly taken into account. 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 more quickly (if it is below equilibrium) or more slowly (if it is above equilibrium) than implied by the equation.

As our focus is on identifying the commonly shared implications of the four estimated model, we build ”average” long-run equilibrium and short-term adjustment paths as simple arithmetic averages of the results for each model and assess the uncertainty surrounding this ”average” behaviour by comparing them to the maximum and minimum model estimates in each year, which provide us with a sort of ”plausibility” range.

Graph 3 shows the average, minimum and maximum estimated equilibrium values for the price-to-rent ratio. As can be seen, the average equilibrium ratio displays some upward trend since the mid-1990s, which is later followed by a rapid increase in the observed ratio. The latter, moreover, tends at first to restore equilibrium in the market after the sharp decrease in the price-to-rent ratio that followed the end of the previous boom in the early nineties. The increase in the observed ratio, however, ends up going markedly beyond that required to restore equilibrium, thereby pointing to an overvaluation of house prices in relation to rents. It is worth remembering at this point that our approach does not allow us to discriminate between house price overvaluation and rent undervaluation.

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9. Regarding the estimation of the four different VARs, the number of lags has been chosen according to the Akaike criteria, after testing that no residual correlation was left. VARs tend to explain better the changes in real rents than in the discount factor, as is revealed by the different equation standard errors, which are in general reasonably small for the former (around 0.5%). For the latter, they range from 0.8% for consumption growth to 12% for the real return on the stock exchange index.
10. See, for instance, Campbell (1993).

The uncertainty in the estimated average behaviour is, as could be expected, quite significant. Notably, however, despite model diversity, the result that the increase in the ratio since 1998 contributed to restore equilibrium is fairly robust. In the same vein, all models coincide in pointing that house prices are above equilibrium at the end of the sample period.

Graph 4 replicates the exercise in Graph 3 for the estimated short-term adjustment path. As commented above, the differences are much less significant in this case. Thus all models allow it to be concluded that during the current boom in the housing market, the price-dividend ratio has moved well in line with its historical short-term adjustment pattern. At the very end of the sample period the observed ratio is only marginally above the short-term estimated path.

3.4 Taxes

Before concluding this section, it should be mentioned that both in the theoretical model and in its empirical application we have not considered taxes. The detailed analysis of the impact of the tax-subsidy system on rents and house prices is a complex task that falls outside the scope of the paper. In particular, it is important to note that the heterogeneous tax treatment of individual agents makes it very difficult to identify the "marginal" agent affecting asset pricing conditions.

Nevertheless, we can make use of the results reported in a recent paper by García-Vaquero and Martínez Pagés (2005) on taxes and the housing market in Spain to investigate the potential effects of ignoring taxes on our estimates.

Thus, note first that if we could identify the marginal agent and its tax treatment, we would be able to obtain a tax modified version of equation (5) for it. More concretely, we would have to take into account that the dividends obtained from the asset (house) in each period will include rents plus subsidies obtained minus taxes paid. The net flow of taxes and subsidies can always be expressed as a percentage of rents and the corresponding ratio could be seen as a sort of "net tax rate" on rents. Thus, if we still use D_t for real rents – d_t being its (log) growth rate – and represent that net tax rate as ψ_t it is easy to see that:

$$q_t \equiv \ln \frac{P_t}{D_t(1-\psi_t)} \approx q_t^b + \psi_t \tag{14}$$

where q_t^b is the price-to-rent ratio before taxes.

If we make the assumption that changes in the net tax rate are unpredictable for the marginal agent, it follows directly from equation (14) that nothing changes on the right-hand sides of equations (5) and (6), as $E_t \sum_{s=1}^{\infty} \delta^s \, \Delta\psi_{t+s} = 0$. Therefore, combining (7) and (14) to obtain an empirically workable equation for the observable before-tax price-to-rent ratio we have:

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

From the results in García-Vaquero and Martínez (2005) it is possible to obtain a proxy for ψ_t for different classes of households. More specifically, they obtain (annual) estimates of the wedge introduced by all relevant taxes and subsidies affecting housing demand (VAT, property tax, income tax relief, etc.) in the user costs of houses for different households according to their income, their age, and their property tenure (landlord, tenant, first-home owner occupier, second–home owner). Since user costs have to equal rents at (before tax) equilibrium, we can use the ratio of the estimated wedge to the user costs before taxes as a proxy for ψ_t. More precisely, if we allow for some measurement error we can write $\psi_t = \lambda_0 + \lambda\zeta_t + \upsilon_t$, where ζ_t is the estimated ratio of the tax wedge to the user costs before taxes and υ_t is iid white noise uncorrelated to the shock in equation (8).[11] After some algebra, we can obtain modified versions of (9) and (10) that explicitly include taxes:

$$q_t = n\ (1-\sum_{i=1}^{p}\ \phi_i) + \sum_{i=1}^{p}\ \phi_i q_{t-1} + \tau\ E_t \sum_{s=1}^{\infty}\delta^s\ d_{t+s} - \rho\ E_t \quad \sum_{s=1}^{\infty}\delta^s\ x_{t+s}$$

$$-\tau \sum_{i=1}^{p}\ \phi_i\ E_{t-1} \sum_{s=1}^{\infty}\delta^s\ d_{t-1+s} + \rho \sum_{i=1}^{p}\ \phi_i E_{t-1} \quad \sum_{s=1}^{\infty}\delta^s\ x_{t-1+s}\ + \beta\ w_t - \tag{16}$$

$$\lambda\,\zeta_t + \lambda \sum_{i=1}^{p}\ \phi_i \zeta_{t-i} + \varepsilon_{3t}$$

$$q_t = n'\ (1-\sum_{i=1}^{p}\ \phi_i) + \sum_{i=1}^{p}\ \phi_i q_{t-1} + E_t \sum_{s=1}^{\infty}\delta^s\ d_{t+s} -\ E_t \quad \sum_{s=1}^{\infty}\delta^s\ r_{w,t+s}$$

$$-\sum_{i=1}^{p}\ \phi_i\ E_{t-1} \sum_{s=1}^{\infty}\delta^s\ d_{t-1+s} + \sum_{i=1}^{p}\ \phi_i E_{t-1} \quad \sum_{s=1}^{\infty}\delta^s\ r_{w,t-1+s}\ + \beta\ w_t - \tag{17}$$

$$\lambda\,\zeta_t + \lambda \sum_{i=1}^{p}\ \phi_i \zeta_{t-i} + \varepsilon_{4t}$$

Graph 5 shows the behaviour of ζ_t for three representative household classes. Note that the net tax rate has always been positive during the sample period for landlords and negative, i.e. the net effect has been that of a subsidy, for owner-occupiers. For those who buy a house and decide to leave it unoccupied, the net tax rate is positive only since the late nineties. In all three cases, however, the estimated tax rates show an upward trend, although for landlords this trend does not start until the end of the nineties.

Table 3 shows the estimates of λ obtained from equation (16) and the three empirical versions of equation (17) for the representative landlord, where the quarterly net tax rate has been obtained by linearly interpolating the annual one. As can be seen, the coefficient estimates are small, non-significant and in one case even the sign is wrong.[12] If we made our computations including the non-significant point estimates for λ nothing substantial would change. These results support the view that taxes are unlikely to have been a key determinant of the sharp increase in the ratio observed in the last few years.

11. We thank the authors for providing us with these series. See the quoted reference for more details on how the representative agents are chosen and the tax wedges are computed.
12. These results do not change if the alternative net tax rates for representative households who are first-home owner occupiers or second-home owners are considered instead.

4 Estimating ex-ante real interest rates[13]

Although there are alternative ways of reconciling the results of the Gordon model and those found in section 3, in this section we provide evidence supporting the hypothesis that the observed decrease in ex-post real interest rates since mid 1990s is likely to overestimate the fall in the ex-ante real interest rates experienced by the Spanish economy.

According to the Fisher equation, ex-post real rates only provide an accurate proxy to the actual risk-free real interest rate if ex-post inflation does not differ much from expected inflation and the inflation risk premium is small. This means that, in stable economies where inflation does not show much volatility and remains close to a relatively low figure most of the time, average ex-post real rates over a certain period represent in that case a reasonably good approximation to the average actual real interest rate. Spain, however, cannot be presented as an economy with a stable macroeconomic regime during the 1990s. The economy underwent a very significant transformation, going from a period of exchange rate instability, large public deficits and high inflation at the beginning of the last decade of the 20th century to a new regime characterised by EMU membership, fiscal surpluses and moderate inflation. Moreover, the regime shift was not a gradual, predetermined process but a sinuous road whose end-point did not become certain until almost mid-1998. Therefore, it is very likely that the course of inflation expectations was substantially driven by the probability attached to a scenario of unsuccessful nominal convergence – which did not materialise – thereby creating a *peso problem*. At the same time, there are good reasons to believe that ex-post real rates during much of the previous decade incorporated a compensation for uncertain inflation. This means that the observed decline of ex-post real interest rates could be at least partially explained by overly pessimistic inflation expectations during the first half of the decade and by a decrease in the inflation risk premium as the economy approached EMU. This would mean that the low level of ex-post rates today reflects, at least to some extent, a higher predictability of inflation and lower inflation risk. That would, in turn, imply that the decrease in the actual real cost of capital could have been lower than suggested by the course of ex-post real rates.

There are, however, a number of difficulties in estimating directly the contribution of changes in the inflation regime to the observed course of ex-post real rates in Spain. In particular, inflation-indexed bonds have never been traded and there is no reliable series of inflation expectations at different horizons. We must therefore rely on economic and financial theory to derive implicit real interest rates.

4.1 The macroeconomic approach

As a starting point, it is useful to analyse international evidence on short-term interest rates. Assuming that capital markets are integrated, one should expect real short term rates not to diverge much across countries. It is therefore potentially helpful to use as a reference for Spanish real interest rates those of countries where this variable can be measured more accurately. This is the case of markets where there has long been an active market for inflation-indexed government bonds (as in the UK) and of countries where the relative stability of the inflation regime makes ex-post real rates a reasonable proxy for the actual riskless rate (as in Germany and, to a lesser extent, the United States).

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13. This section heavily draws from Blanco and Restoy (2006).

Table 4 presents average three-month inflation-adjusted interest rates for Germany, the UK and the United States, along with average 10-year indexed-bond yields for the UK. We present evidence for two periods: i) 1990-1998 and ii) 1999-2005. As can be seen, the actual level of average ex-post real rates differs somewhat across countries. However, the difference between periods is remarkably similar across countries, with the exception of Spain. For Germany, the United States and the UK, average ex-post real rates have declined somewhere between 1 1/2 and 1 3/4 percentage points. In Spain, however, the decrease is much sharper (more than 5 percentage points), thereby pointing either to a radical failure of the capital market integration hypothesis or to a missmeasurement of the actual decline in the risk-free real interest rate in the Spanish case.

Another possibility is to exploit intertemporal equilibrium relations for domestic producers and consumers. For example, one traditional rule of thumb is to set equilibrium real rates equal to potential output growth. Potential growth actually increased in Spain during the nineties, due essentially to high employment creation. According to various estimates, average potential GDP growth was about 0.5% higher in the period 1999-2005 than in the period 1990-1998.[14] A more refined measure could be a proxy for the marginal productivity of capital. According to Banco de España's internal estimates, the average ratio of Gross Value Added to the capital stock in the manufacturing sector actually went down from 1999, in comparison with the first period, by an amount close to 1.3%, a similar figure to that found for the decline in ex-post real rates in other countries.

Looking at the intertemporal marginal rate of substitution (IMRS) of a representative Spanish consumer, we could also derive a measure of equilibrium real interest rates. In Table 5 we provide the average implicit interest rate derived from the first order equilibrium conditions of a representative agent for several specifications of preferences. All data are drawn from Spain's Quarterly National Accounts.

Using first the standard isoelastic CRRA utility function, we find that average implicit real interest rates would have gone up and not down in the second sub-period for any reasonable value of the risk aversion parameter. This is not surprising as the IMRS is, in this case, a monotonic positive transformation of consumption growth and this has been, on average, almost 1% higher in the second sub-period. Using some form of multiplicative or additive external habits does not change the picture much. For plausible parameters, the average implicit real interest rate becomes either larger or, at most, slightly smaller in the second sub-period.

We have also checked whether non-separable preferences between consumption and leisure could provide somewhat different results. Indeed, as employment ratios have increased markedly in Spain in the recent past, one could conjecture that this might compensate the positive effect of consumption growth on the marginal rate of substitution. Indeed, using the KPR preferences[15] we find that the implicit risk-free rate falls in the second sub-period for high values of the elasticity of intertemporal substitution. But the maximum decrease we obtain is, for sensible parameter values, still less than 2%. This is a figure which lies well below the observed fall in ex-post real rates in Spain, although it is in line with that found in other countries.

14. See, for example, Denis et al (2006).
15. See King, Plosser and Rebelo (1988).

4.2 The financial approach

In Section 4.1 we have made use of equilibrium conditions of a representative agent. This analysis requires relatively strong assumptions on specific features of the economy, such as preferences, technology and the ability of agents to design intertemporal consumption and investment plans. A more robust approach is to exploit, as we did in section 3, pure non-arbitrage conditions in financial markets. As we saw, these conditions imply that all securities should be priced by applying to their future payoffs a common stochastic discount factor which, as shown in Huang and Litzenberger, 1988, is directly linked to the return on a risk-free security. This discount factor is also equivalent to the IMRS of the representative agent under the equilibrium conditions of section 4.1. As in section 3, we do not aim at identifying the precise stochastic discount factor that prevents arbitrage opportunities but consider instead some alternative methods useful to extract helpful information.

4.2.1 THE HANSEN-JAGANNATHAN FRONTIER

Hansen and Jagannathan (1991) derive regions for the admissible mean-standard deviation pairs for the discount factor with the sole assumption that markets are free of arbitrage opportunities. The expression for the standard deviation bound is given by:

$$\sigma(m) = [(E(p) - E(m)E(x))'\Sigma^{-1}(E(p) - E(m)E(x))]^{1/2} \tag{18}$$

where m is the discount factor, p is the vector of security prices, x is the vector of payoffs, Σ is the variance-covariance matrix of payoffs and E() is the unconditional expectation operator. It is apparent from expression (18) that to compute the HJ frontier we only need securities market data.

Note that, by restricting the standard deviation of the discount factor to a maximum level ($\bar{\sigma}$), we can obtain a lower (E_1) and an upper (E_2) bound for the average level of the discount factors (see Graph 6) and, implicitly, for the real interest rate (remember that $E(m) = 1/(1+r)$, where r stands for the real interest rate).

In this section we use this approach to find bounds for the average level of the ex-ante real interest rates. To do that we use monthly data for a sample of Spanish securities including 18 portfolios of stocks (10 size portfolios and 8 industry portfolios), 2 short-term securities with a time of 3 months and one year to maturity, respectively, and a portfolio of long-term debt.[16] Returns are computed in real terms (deflated by the Spanish CPI index) assuming a holding period of one month.

Graph 7 shows the HJ frontiers estimated for the periods 1990-96 and 1999-2005 using all securities in our dataset. We exclude from the analysis the years 1997 and 1998, which is an interim period where security prices are likely to incorporate already many of the relevant features of the monetary union regime. As can be seen in the graph, for reasonable values of the standard deviation, the ranges for the means of the discount factors are relatively narrow in both periods and they do not overlap. In particular, the means of the discount factors are higher in the second period, suggesting a fall in the average level of the real interest rate. Interestingly, the mid-point of the bound is similar to the level implied by the ex-post short-term real interest rates. However, as explained in the introduction, we suspect that this result might be contaminated by a *peso problem*. More specifically, if

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16. Annex 1 describes the composition of these portfolios and the computation of the monthly real returns.

inflation expectations during the first period were systematically higher than observed inflation, the average ex-post return and, therefore, the estimated mean of the discount factors would be overstated.

Therefore, we repeat the same exercise excluding short-term securities but retaining longer-term fixed-income instruments. Graph 8 shows the results. We can see that the size of the region of the admissible pairs of mean and standard deviation of discount factors increases dramatically for the two periods. Also, the two regions are now much closer compared with Graph 7. Thus, it is much harder to reject the hypothesis of equal average levels of real interest rates in the two periods.

Graph 9 shows the estimated HJ frontiers using only the 18 portfolios of stocks. In this case the HJ frontiers are even closer, making it harder to reject the hypothesis that the average level of real interest rates is the same in the two periods. However, the size of the range is very large. Therefore, once we exclude fixed-income securities the average level of the real interest rate is estimated with high uncertainty.

4.2.2 EXPLOITING IDIOSYNCRATIC RISK

Given the uncertainty of the previous approach in estimating the average level of real interest rates, in this section we rely on an alternative approach recently proposed by Flood and Rose (FR), which allows us to obtain point estimates for that variable as opposed to ranges.

FR consider the standard decomposition of the Euler equation:

$$p_t^j = E_t(m_{t+1}x_{t+1}^j) = COV(m_{t+1}, x_{t+1}^j) + E_t(m_{t+1})E_t(x_{t+1}^j) \tag{19}$$

where $COV_t()$ and $E_t()$ are, respectively, the covariance and expectations operators, both conditional on information available at t, m_{t+1} is the discount factor used to discount income accruing in period t+1, and p_t^j and x_{t+1}^j are, respectively, the price of asset j in period t and the payoff of that asset at time t+1. Equation (19) can be rewritten as

$$x_{t+1}^j = \delta_t(p_t^j - COV(m_{t+1}, x_{t+1}^j)) + \varepsilon_{t+1}^j \tag{20}$$

where $\varepsilon_{t+1}^j \equiv x_{t+1}^j - E_t(x_{t+1}^j)$ is a prediction error orthogonal to information at time t, and $\delta_t \equiv 1/E_t(m_{t+1})$.

The standard approach in finance to make equation (20) stationary is to normalise by p_t^j. FR propose normalising by the systemic component of this price ($\tilde{p}_t^j$), which is defined as the value of p_t^j conditional on idiosyncratic information available at t being set to zero.

$$x_{t+1}^j / \tilde{p}_t^j = \delta_t(p_t^j / \tilde{p}_t^j - COV(m_{t+1}, x_{t+1}^j / \tilde{p}_t^j)) + \varepsilon_{t+1}^j / \tilde{p}_t^j \tag{21}$$

FR rewrite equation (21) as

$$x_{t+1}^j / \tilde{p}_t^j = \delta_t(p_t^j / \tilde{p}_t^j) + u_{t+1}^j \tag{22}$$

where $u_t^j = \varepsilon_{t+1}^j / \tilde{p}_t^j - \delta_t COV(m_{t+1}, x_{t+1}^j / \tilde{p}_t^j)$. They note that assuming that $COV(m_{t+1}, x_{t+1}^j / \tilde{p}_t^j)$ moves only because of aggregate phenomena, δ_t in (22) can be consistently estimated using either OLS or GMM.

FR propose the following two-step strategy to estimate δ_t. In the first step they estimate the following J (the number of securities) time series regressions by OLS

$$\ln(p_t^j / p_{t-1}^j) = \alpha_0^j + \sum_{i=1}^{N} \alpha_i^j f_t^i + v_t^j \tag{23}$$

where f_t^i are a set of N aggregate factors and v_t^j is the residual, which captures the idiosyncratic part of asset price j return. Using estimated coefficients of regressions (23), the estimated systematic price is defined as

$$\hat{p}_t^j = p_{t-1}^j \exp\left(\hat{\alpha}_0^j + \sum_{i=1}^{N} \hat{\alpha}_i^j f_t^i \right) \tag{24}$$

In their empirical implementation, FR estimate regressions (23) using as factors the market-wide stock market return and the three Fama-French factors: the overall market return less the treasury-bill rate, the performance of small stocks relative to big stocks, and the performance of "value" stocks relative to "growth" stocks. In these time series regressions coefficients are estimated as fixed parameters using all the sample period.

In the second step they estimate cross-sectionally the following regressions for every period t

$$x_{t+1}^j / \hat{p}_t^j = \delta_t (p_t^j / \hat{p}_t^j) + u_{t+1}^j \tag{25}$$

FR note that using $\hat{p}_t^j$ in place of the unobservable $\tilde{p}_t^j$ might induce measurement error. Also the existence of a generated regressor in equation (25) might potentially understate the OLS standard errors. To handle both potential econometric problems, they estimate (25) using GMM. In these regressions variables are defined in nominal terms, whereby the parameter δ_t is interpreted as the inverse of the expected nominal discount factor in period t.

In this paper we employ the approach proposed by FR to test whether and by how much the average level of the real interest rate has fallen in the Spanish economy between the periods 1990-98 and 1999-2005. To do that we employ the 18 portfolios of stocks used to derive the HJ frontier. We estimate the time series regressions using only two factors: market-wide return and the performance of small stocks relative to big stocks. The former is the total return (including dividends) on the Madrid Stock Exchange General Index and the latter is the difference between the return on portfolios made up of securities in the decile of the smallest and largest stocks, respectively. Parameters are estimated using the last 60 monthly observations.

Unlike FR we are only interested in the average level of the real interest rates. In order to reduce noise we estimate the cross-section regression as a pool where the discount factor parameter is assumed to be fixed within the two periods of interest. More specifically, we estimate the following regression

$$x_{t+1}^j / \hat{p}_t^j = \delta_1 (p_t^j / \hat{p}_t^j) + \delta_2 (p_t^j / \hat{p}_t^j) D99_t + u_{t+1}^j \tag{26}$$

where $D99_t$ is a dummy variable which takes value 1 from January 1999. In regression (26) the payoffs x_{t+1}^j are deflated by the Spanish CPI. Therefore, parameters δ_1 and δ_2 should be interpreted in real terms. Note that δ_1 can be expressed as $\delta_1 = 1 + r_1$, where r_1 is the average real interest rate in the period 1990-96, and δ_2 as $\delta_2 = r_2 - r_1$, where r_2 is the average real interest rate in the period 1999-2005. Therefore, δ_2 measures the change in the average real interest rate level between the periods 1990-96 and 1999-2005.

Regression (26) is estimated by GMM using the first lag of the explanatory variables as instruments. Table 6 presents the estimated parameters together with their standard errors. Coefficient δ_2 is not significant at the standard levels, implying that the null hypothesis of equal real interest rates in the two periods cannot be rejected. The point estimate of coefficient δ_1 is 1.005, implying an annual real interest rate of around 6.2% (=1.00512-1), which seems very high, a result consistent with FR, who also obtained high average estimates for the implied (nominal) interest rates in their sample. However, the two-standard-error confidence interval band for the real interest rate is quite wide (0-16%), suggesting that this variable is estimated with much uncertainty.

All in all, results reported in this section based on pure arbitrage considerations show no evidence of a significant decrease in the implicit real risk-free rates since the late 1990s. Still, a natural follow-up would be to try to introduce greater structure into these models in order to increase the accuracy of the estimates.

5 Conclusions

Since the early nineties, real interest rates, proxied by the difference between the nominal ones and actual inflation, have decreased substantially in Spain. Given their role as a major component of the discount factor for any future payoff we should expect, other things equal, asset prices to have increased as a result. Against this background, some simple calculations allow the recent boom in house prices in this country to be fully explained as a simple consequence of the drastic reduction in (ex-post) real interest rates.

Our results, however, show that when the asset pricing exercise is done in a sufficiently refined way that takes into account that changes in the discount factor have to affect all assets, house prices are above their long-term equilibrium. Nevertheless, they are in line with the adjustment path we should expect when the long-term equilibrium increases as a consequence of movements in their fundamental determinants and prices overreact in the short term to these movements.

A first implication of this result is that contrary to a situation characterised by a bubble episode, the most likely scenario for future house prices in Spain is one where equilibrium is restored in an orderly way – which does not necessarily require an adjustment in nominal terms – and consequently, there is no need for policy action, beyond a prudent monitoring of the process aimed at promptly detecting any significant departure from the estimated short-term adjustment path.

And the second one is that, at least during the period considered, changes in ex-post real rates are a bad proxy for changes in ex-ante real rates, which are the relevant ones in standard asset pricing models. Underlying this result surely is the significant transformation underwent by the Spanish economy from a regime of exchange rate instability, large public deficits and high inflation to a new one characterised by EMU membership, fiscal surpluses and moderate inflation. In this process, which did not become certain until almost mid-1998, the course of inflation expectations is likely to have been substantially driven by the probability attached to a scenario of unsuccessful nominal convergence. Under these circumstances, we should expect nominal interest rates – and therefore ex-post real rates – to be affected by both a non-negligible inflation risk premium and a standard *peso problem* as, fortunately, the alternative scenario of no convergence never materialised.

ANNEX 1: SECURITIES MARKET DATA

In the empirical exercises we use monthly data for a sample of Spanish securities including 18 portfolios of stocks (10 size portfolios and 8 industry portfolios), 2 short-term securities and a portfolio of long-term debt. The sample period expands from January 1990 to December 2005.

The 10 size portfolios are made up from a dataset which includes all stocks traded on the electronic segment of the Spanish stock exchanges ("mercado continuo"). More specifically, at the end of each year stocks which have traded the following year are classified in 10 portfolios with the same number of stocks, according to the market value of the company on that date. Portfolio returns are computed as the equally weighted returns on individual stocks. Returns include dividends and are corrected by splits.

The industry portfolios are made up using the total return (including dividends) sectoral indices published by the Madrid Stock Exchange (MSE). Between 1940 and 2001 the MSE had been using 10 sectoral indices. Starting in 2002, these series were discontinued and new series were created. The new sectoral classification offers more detailed information. More specifically, there are 7 sectoral indices and 29 sub-sectoral indices. For 8 of the previous indices we were able to update the series using the new indices. These are the 8 industry portfolios we use in our empirical exercises. The sectors included are the following: banking, utilities, food, construction, investment companies, telecommunications, oil and basic materials.

The two short-term securities are notional bills issued with a time to maturity of 3 months and one year, respectively. Returns are computed using theoretical prices for these securities derived from the 3-month interest rates traded on the Madrid interbank market (EURIBOR rates since 1999) and one-year Treasury Bill yields, respectively.

Finally, the portfolio of long-term debt is the total return index of JP Morgan. This index is made up of bonds issued by the Spanish Treasury. The average duration of the portfolio over the sample period is 4.5 years. The index considers both changes in prices and coupon payments.

All returns are computed in real terms (deflated by the Spanish CPI index) assuming a holding period of one month.

REFERENCES

ABEL, A. B. (1990). "Asset prices under habit formation and catching up with the Jonese", *American Economic Review,* 80, 2 (May 1990), pp. 38-42.

AYUSO, J., and F. RESTOY (2006a). "House Prices and Rents. An Equilibrium Asset Pricing Approach", *Journal of Empirical Finance,* 13, pp. 371-378.

— (2006b). *House Prices and Rents in Spain: Does the Discount Factor Matter?,* Working Paper 0609, Banco de España.

BANCO DE ESPAÑA (1991). "Índice de rendimiento de una cartera de deuda del Estado", *Boletín Económico,* mayo.

BLANCO, R., and F. RESTOY (2006). *How much have real interest rates fallen in Spain,* Mimeo, Banco de España.

CAMPBELL, J. (1993). "Intertemporal asset pricing without consumption data", *American Economic Review,* 83, pp. 487-512.

CAMPBELL, J., and R. SHILLER (1988). "The dividend price ratio and expectations of future dividends and discount factors", *Review of Financial Studies,* 1, pp. 195-227.

CASE K., and R. SHILLER (1989). "The efficiency of the market for single-family homes", *American Economic Review,* 79, pp. 125-137.

— (1990). "Forecasting prices and excess returns in the housing market", *Journal of the American Real Estate and Urban Economics Association,* 18, pp. 253-273.

CLAYTON, J. (1996). "Rational expectations, market fundamentals and housing price volatility", *Real Estate Economics,* 24, pp. 441-470.

DEL RÍO, A. (2002). *El endeudamiento de los hogares españoles,* Working Paper 0228*,* Banco de España.

DENIS, D., D. GRENOUILLEAU, K. MC MORROW and W. RÖGER (2006). *Calculating potential growth rates and output gaps. A revised production function approach,* Economic Papers 247, European Commission.

DIPASQUALE, D., and W. C. WHEATON (1994). "Housing market dynamics and the future of housing prices", *Journal of Urban Economics,* 35, pp. 1-27.

EPSTEIN, L., and S. ZIN (1989). "Substitution, risk aversion and the temporal behaviour of consumption and asset returns: theoretical framework", *Econometrica,* 57, pp. 937-969.

FLOOD, R. P., and A. K. ROSE (2005). "Estimating the expected marginal rate of substitution: a systematic exploitation of idiosyncratic risk", *Journal of Monetary Economics,* 52, pp. 951-969.

GALI, J. (1994). "Keeping up with the Joneses: consumption externalities, portfolio choice, and asset prices", *Journal of Money, Credit and Banking,* 26 (1), pp. 1-8.

GARCÍA-VAQUERO, V., and J. MARTÍNEZ PAGÉS (2005). *La fiscalidad de la vivienda en España,* Occasional Paper 0506, Banco de España.

GENESOVE, D., and CH. MAYER (2001). *Loss aversion and seller behavior: evidence from the housing market,* NBER Working Paper 8143.

GIMENO, R., and C. MARTÍNEZ-CARRASCAL (2006). *The interaction between house prices and loans for house purchase. The Spanish case,* Working Paper 0605, Banco de España.

HANSEN, L. P., and R. JAGANNATHAN (1991). "Implications of security market data for models of dynamic economics", *The Journal of Political Economy,* 99, pp. 225-262.

HUANG, C. F., and R. H. LITZENBERGER (1988). *Foundations for financial economics,* Prentice Hall, Inc., Englewood Cliffs, New Jersey.

KENNY, G. (1999). *Asymmetric adjustment costs and the dynamics of housing supply,* Central Bank of Ireland Technical Paper 1999/0003.

KING, R. G., C. I. PLOSSER and S. REBELO (1988). "Production, Growth and Business Cycles: The Basic Neoclassical Model", *Journal Monetary Economics,* 21 (2/3), pp: 195-232.

MALO DE MOLINA, J. L., and F. RESTOY (2005). "Recent trends in corporate and household finances in Spain: Macroeconomic implications", *Moneda y Crédito,* 221, pp. 9-36.

MANKIW, N. G., and D. N. WEIL (1989). "The baby boom, the baby bust and the housing market", *Regional Science and Urban Economics,* 19, pp. 325-346.

MARTÍNEZ-CARRASCAL, C., and DEL RÍO (2004). *Household borrowing and consumption in Spain. A VECM approach,* Working Paper 0421*,* Banco de España.

MARTÍNEZ-PAGÉS, J., and L. A. MAZA (2003). *Analysis of house prices in Spain,* Working Paper 0307, Banco de España.

NIETO, F. (2003). "Determinantes del crecimiento del crédito a los hogares en España", *Boletín Económico*, Banco de España, April.

ORTALO-MAGNÉ, F., and S. RADY (2006). "Housing market dynamics: on the contribution of income shocks and credit constraints", *Review of Economic Studies,* 73 (2), pp. 459-485.

POTERBA, J. N. (1984). "Tax subsidies to owner-occupied housing: an asset market approach", *Quarterly Journal of Economics,* 99, pp. 729-752.

— 1991: "House price dynamics: the role of tax policy and demography", *Brookings Papers on Economic Activity,* 2, pp. 143-203.

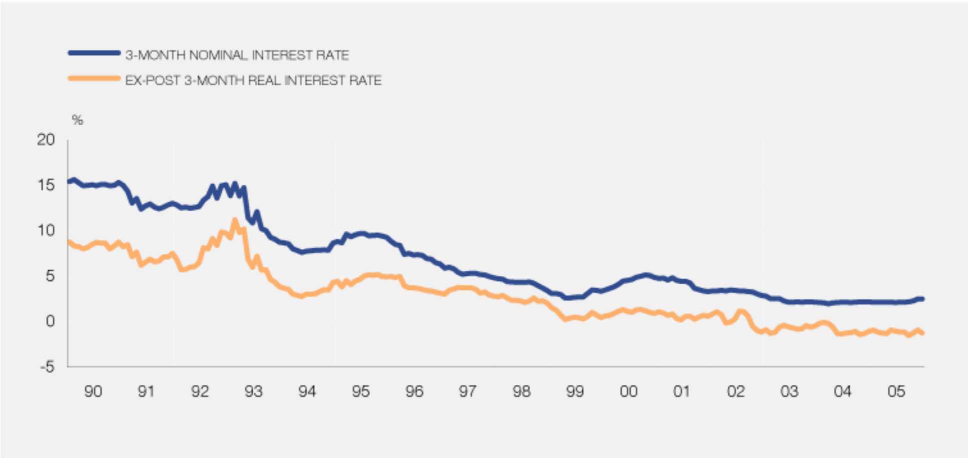
RESTOY F., and P. WEIL (1998). *Approximate equilibrium asset prices*, NBER Working Paper 6611.

RODRÍGUEZ, R., F. RESTOY and I. PEÑA (2002). "Can output explain the predictability and volatility of stock returns?", *Journal of International Money and Finance,* 21, pp. 163-182.

WEIL, P. (1990). "Non-expected utility in macroeconomics", *Quarterly Journal of Economics,* CV (1), pp 29-42.

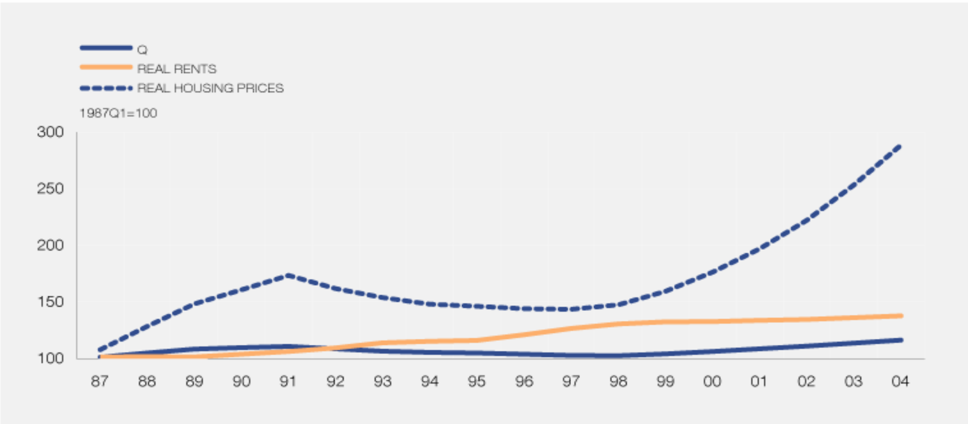
NOMINAL AND INFLATION ADJUSTED THREE-MONTH INTEREST RATES

GRAPH 1



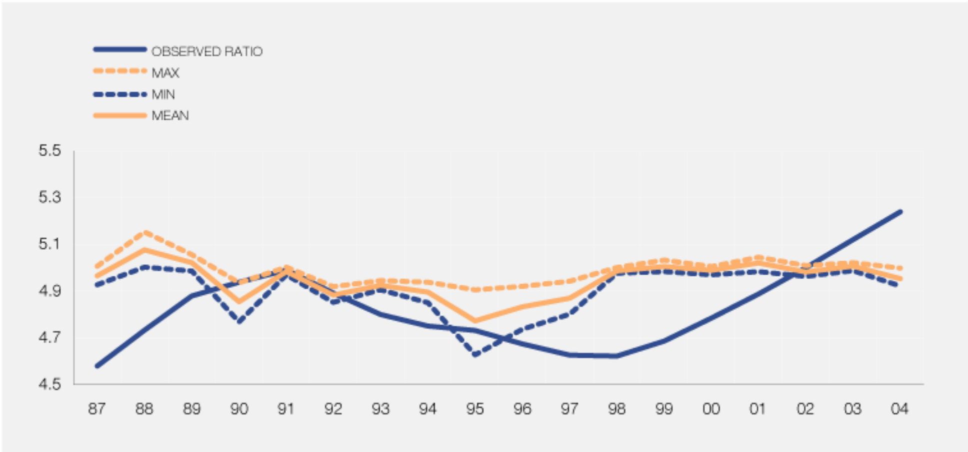
REAL HOUSE PRICES, REAL RENTS
AND (LOG) RATIO OF HOUSE PRICE TO RENT (Q)

GRAPH 2



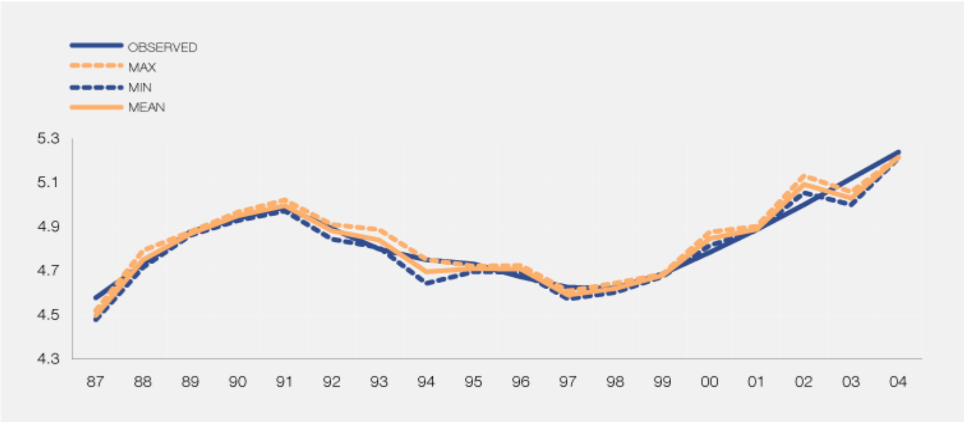
(LN) HOUSE PRICES / RENTS. ESTIMATED LONG-RUN EQUILIBRIUM

GRAPH 3



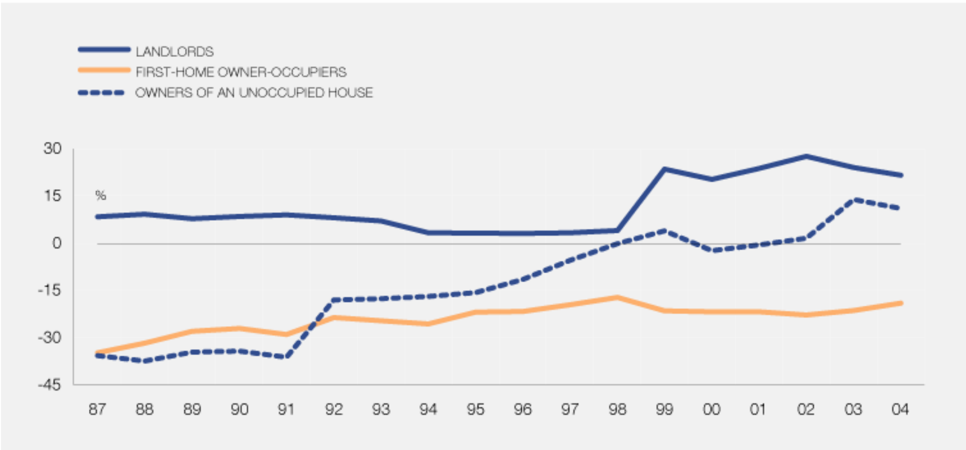
(LN) HOUSE PRICES / RENTS. ESTIMATED SHORT-TERM ADJUSTMENT PATH

GRAPH 4



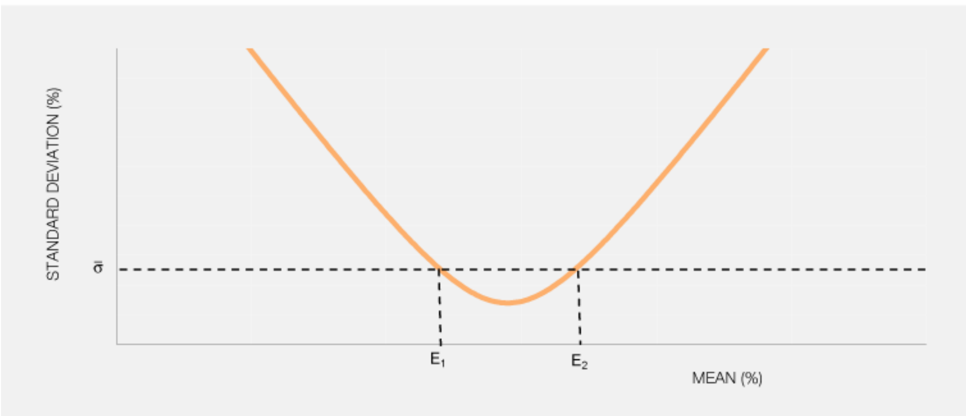
NET TAX RATE FOR DIFFERENT REPRESENTATIVE HOUSEHOLDS

GRAPH 5



HANSEN-JAGANNATHAN FRONTIER

GRAPH 6



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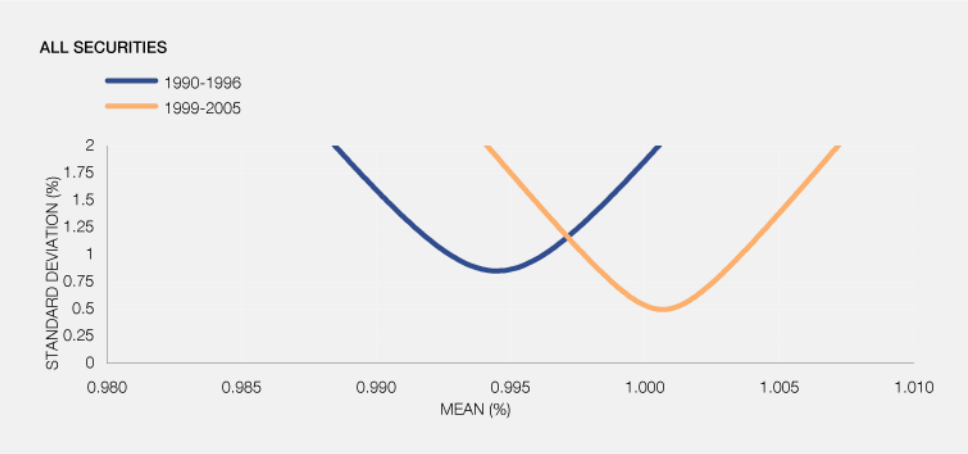
GRAPH 7



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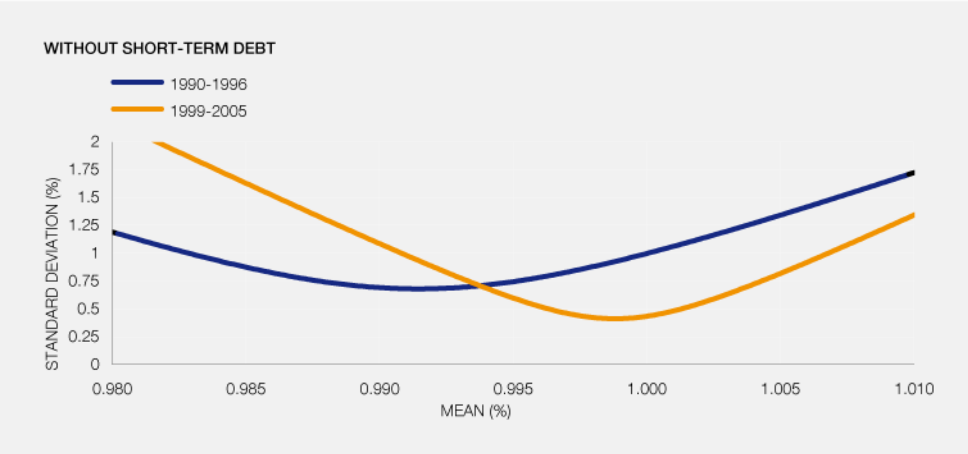
GRAPH 8



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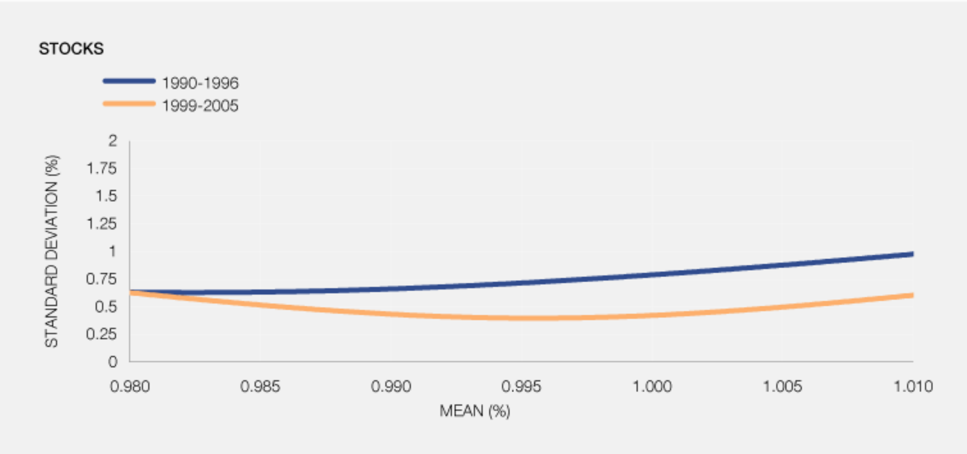
GRAPH 9



DESCRIPTIVE STATISTICS
SAMPLE PERIOD 1987Q2-2004Q4

TABLE 1

	Dlhp (a)	Dlr (a)	q (c)	Dlc (a)	Rm (b)	R10y (b)	Dly (a)	Dlw (a)	Rpd (b)
Mean	0.02	0.00	4.83	0.01	0.01	0.09	0.01	0.02	0.02
Median	0.02	0.00	4.80	0.01	0.02	0.09	0.01	0.02	0.02
Maximum	0.06	0.03	5.28	0.03	0.34	0.11	0.04	0.06	0.08
Minimum	-0.07	-0.02	4.50	-0.02	-0.35	0.07	-0.02	-0.05	-0.05
Std. deviation	0.02	0.01	0.18	0.01	0.14	0.01	0.01	0.02	0.02
Skewness	-0.59	0.26	0.60	-0.34	-0.46	0.01	-0.05	-0.28	0.03
Kurtosis	3.84	4.01	2.79	3.47	3.37	2.67	3.69	3.13	2.96
Jarque-Bera (d)	6.35	3.92	4.39	2.05	2.97	0.33	1.46	0.97	0.01
P-value	0.04	0.14	0.11	0.36	0.23	0.85	0.48	0.62	0.99

a. Dlhp, Dlr, Dlc, Dly and Dlw stand for the first difference of the logs of house prices, rents, household consumption, GDP and household tota net wealth, respectively. All them have been deflated by a CPI index excluding shelter.
b. Rm, R10y and Rpd stand for the return on the Ibex-35 stock exchange index, the return on 10-year public debt, and the bond total return index, respectively. Returns have been deflated by a CPI index excluding shelter. R10y has been de-trended to guarantee stationarity.
c. q stands for (log) ratio of house prices to rents.
d. Jarque-Bera stands for the Jarque-Bera normality test whose p-value is shown in the row below.

GMM ESTIMATES (a) (b) (c)
SAMPLE PERIOD 1987Q2-2004Q4

TABLE 2

	Discount factor				
	Consumption growth	Return on Ibex-35	Return on bond portfolio	Change in households' financial wealth	Constant
Constant	0.96 (.12)	0.58 (.09)	0.27 (.34)	0.25 (.28)	-0.06 (.62)
τ	-14.7 (.09)	--	--	--	--
ρ	10.5 (.09)	--	--	--	--
ϕ	0.88 (.00)	0.88 (.00)	0.95 (.00)	0.95 (.00)	1.02 (.00)
β	2.76 (.05)	3.14 (.00)	1.75 (.00)	1.74 (.00)	0.16 (.65)
Sargan test	0.28 (.96)	.42 (.81)	2.52 (.28)	4.30 (.12)	14.1 (.00)
σ	0.119	0.119	0.062	0.063	0.031
Q1 (d)	2.67 (.10)	0.89 (.34)	0.64 (.42)	0.06 (.81)	5.19 (.02)
Q4 (d)	9.70 (.05)	3.88 (.42)	0.86 (.93)	1.13 (.89)	29.1 (.00)
Q8 (d)	13.8 (.09)	6.31 (.61)	4.38 (.82)	4.72 (.79)	44.7 (.00)
ARCH1 (e)	0.73 (.39)	0.31 (.58)	0.33 (.56)	0.09 (.77)	0.09 (.36)
ARCH4 (e)	3.55 (.47)	4.83 (.30)	2.33 (.68)	4.31 (.37)	8.15 (.08)
ARCH8 (e)	9.92 (.27)	11.1 (.19)	6.43 (.60)	5.95 (.65)	10.9 (.21)
Bera-Jarque (f)	0.60 (.74)	0.10 (.95)	3.40 (.18)	0.65 (.72)	3.14 (.21)

a. Instruments: one lag of the variables involved. In the first column one lag of consumption growth, rent growth and 10-year de-trended interest rate has been added to overidentify parameters.
b. P-values in brackets.
c. w in columns 2 to 4 is the -(real log) rate of growth of households' financial wealth beyond its sample average.
d. Qi stands for the standard test on residual autocorrelation up to order i.
e. ARCHi stands for the standard test on residual ARCH-type heteroscedasticity up to order i.
f. Bera-Jarque stands for the Bera-Jarque test on residual normality.

GMM ESTIMATES OF TAX EFFECTS ON THE PRICE-TO-RENT RATIO (a)
SAMPLE PERIOD 1987Q2-2004Q4

TABLE 3

	Discount factor			
	Consumption growth	Return on Ibex-35	Return on bond porfolio	Change in households' financial wealth
λ	0.01	0.01	-0.00	0.00
(p-value)	(0.47)	(0.25)	(0.91)	(0.94)

a. Instruments: the same as in Table 3 plus the net tax rate at *t* and *t-1* .

REAL INTEREST RATES

TABLE 4

%	Ex-post 3-month real interest rate				10-year indexed bond yield. UK
	Spain	Germany	UK	USA	
1990-1998	5.31	3.17	4.18	2.21	3.76
1999-2005	-0.04	1.64	2.38	0.72	2.08
Change	-5.35	-1.53	-1.80	-1.50	-1.69

IMPLIED REAL INTEREST RATES DERIVED FOR ALTERNATIVE SPECIFICATIONS OF PREFERENCES TABLE 5

%

		1990-1998	1999-2005	Change
ISOELASTIC PREFERENCES	gamma=0.1	2.21	2.27	0.06
	gamma=1	3.93	4.52	0.59
	gamma=5	11.58	14.93	3.35
ABEL. PHI=-0.75	gamma=0.1	3.55	4.20	0.66
	gamma=1	5.28	6.49	1.21
	gamma=5	13.02	17.10	4.08
ABEL. PHI=-0.25	gamma=0.1	2.66	2.91	0.25
	gamma=1	4.38	5.17	0.80
	gamma=5	12.06	15.65	3.59
ABEL. PHI=0.25	gamma=0.1	1.77	1.63	-0.14
	gamma=1	3.47	3.87	0.39
	gamma=5	11.11	14.22	3.11
ABEL. PHI=0.75	gamma=0.1	0.88	0.37	-0.52
	gamma=1	2.57	2.57	0.00
	gamma=5	10.15	12.79	2.64
EXTERNAL ADDITIVE HABITS b=0.25	gamma=0.1	2.20	2.28	0.07
	gamma=1	3.80	4.56	0.76
	gamma=5	10.69	15.02	4.33
EXTERNAL ADDITIVE HABITS b=0.5	gamma=0.1	2.20	2.27	0.07
	gamma=1	3.78	4.47	0.70
	gamma=5	9.80	14.01	4.22
EXTERNAL ADDITIVE HABITS b=0.75	gamma=0.1	2.21	2.25	0.04
	gamma=1	3.59	4.14	0.55
	gamma=5	4.15	8.97	4.82
KPR PREFERENCES a=0.4	gamma=0.1	3.36	4.48	1.12
	gamma=1	3.93	4.52	0.59
	gamma=5	6.43	4.66	-1.77
KPR PREFERENCES a=0.5	gamma=0.1	3.17	4.11	0.94
	gamma=1	3.93	4.52	0.59
	gamma=5	7.29	6.31	-0.97
KPR PREFERENCES a=0.6	gamma=0.1	2.98	3.74	0.76
	gamma=1	3.93	4.52	0.59
	gamma=5	8.15	7.99	-0.16

Real interest rates in sub-period j (j=1, 2) are estimated using the expression $r_j = 1 - 1/(\sum_{i=1}^{N_j} m_{t_j+i} / N_j)$ where Nj is the number of quarters in sub-period j, mt is the discount factor in period t, which is is proxied using the several specifications of preferentes. For isolestastic preferences, $m_t = \beta(g_{t+1})^{-\gamma}$; where $g_{t+1} = c_{t+1}/c_t$ and ct is per capita seasonally-adjusted private non-durable consumption in real terms; for Abel's preferences $m_t = \beta(g_{t+1})^{-\gamma}(g_t)^{\Phi}$; for external additive preferences $m_t = \beta\left((c_{t+1} - bc_t)\ /(c_t - bc_{t-1})\right)^{-\gamma}$; and for KPR preferences $m_t = \beta(g_{t+1})^{-\gamma}(g_{t+1}n_{t+1})^{-(1-a)(1-\gamma)}$ where $n_{t+1} = (1 - N_t)\ (1 - N_{t+1})$ and Nt is the ratio of employment to the population aged over 16. We use quarterly data from Spain's National Quarterly Accounts and is set to 0.995.

ESTIMATION RESULTS FOR THE MEAN OF THE IMRS DERIVED FROM THE METHOD PROPOSED BY FLOOD AND ROSE (2005)

TABLE 6

	Coefficient	Std. error
δ_1	1.0053	0.0037
δ_2	0.0025	0.0050

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