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HAVE REAL INTEREST RATES REALLY FALLEN THAT MUCH IN SPAIN?*

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This paper analyses the behaviour of real interest rates in the Spanish economy between 1990 and 2005. Since inflation-indexed bonds are not available, changes in implicit real interest rates are estimated using several approaches suggested by macroeconomic and financial theory. In particular, we employ equilibrium conditions of a representative agent under several specifications of preferences. Moreover, we exploit no-arbitrage conditions in securities markets. The evidence we report indicates that inflation uncertainty could account for a notable part of the observed decrease in nominal rates. Consequently, the actual real cost of financing might have decreased significantly less than what the course of ex-post real rates would suggest.

Key words: real interest rates, intertemporal marginal rate of substitution.

JEL classification: E43, G12.

One of the most important developments in the Spanish economy between 1990 and 2005 was the sharp reduction in nominal interest associated with the process of nominal convergence and EMU membership. Although inflation rates also declined substantially over the same period, inflation-adjusted interest rates –often called ex-post real rates– fell by almost 10 percentage points over this period (see Figure 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.

Indeed, the Spanish economy 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

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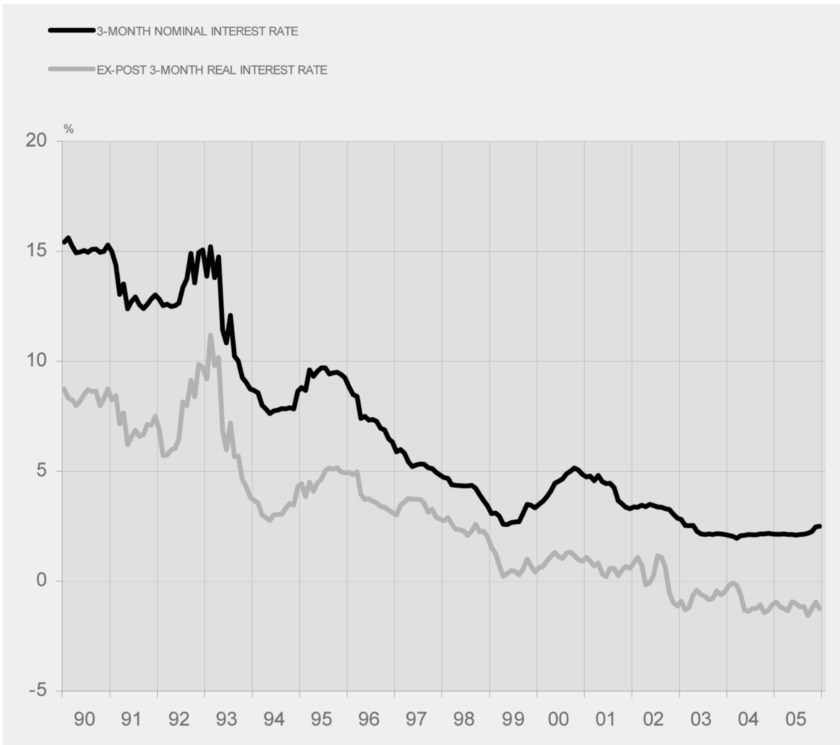
Figure 1: NOMINAL AND INFLATION ADJUSTED THREE-MONTH INTEREST RATES



Source: Banco de España.

points lower than the average over the first half of the previous decade. The debt of the private non-financial sector had risen to 160% of GDP, more than twice the 1995 ratio. In addition, the economy witnessed a substantial real-estate boom which led housing prices to increase by more than 100% in real terms between 1997 and 2005. Finally, economic activity –heavily supported by domestic demand– increased markedly until the recent crisis, with GDP growth averaging more than 3.5% between 1999 and 2005[1].

Still, estimating the impact of lower interest rates on agents' balance sheets and associated macroeconomic developments is not an easy task. For one thing,

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

the economy also faced other important structural changes. In particular, some labour market reforms and intensive immigration flows reduced supply-side rigidities and contributed to substantial employment creation. These developments, together with the consolidation of an environment of macroeconomic stability within EMU, 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.

Moreover, the measure of the actual cost of capital is not straightforward. Agents typically have access to different financing instruments whose relative value may not be stable over time. Yet conceivably, changes in the real return on a riskless asset are a good proxy for changes in the remuneration of capital (or the cost of debt), as that variation should also be reflected –in equilibrium- in the return on any other asset whose risk class remains unchanged. Inflation-indexed government bonds provide a good measure of these genuine riskless real interest rates but they are not available in many countries. Real interest rates are then often proxied by inflation-adjusted nominal interest rates. We know, however, that the ex-post real interest rate is only the real return on an asset –such as a non-indexed Treasury bond or bill– which is typically riskless in nominal terms but not in real terms. According to the Fisher equation, ex-post real rates only provide an accurate proxy to the actual real interest rate (i.e. the real return on a riskless security) 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 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 1990s 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 since 1999 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 directly estimating 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. One possibility is to exploit international data to conjecture about domestic real rates in a world of capital market integration. At the same time, we can make use of intertemporal equilibrium conditions of representative domestic consumers or producers to obtain interest rates implicit in estimates of marginal rates of substitution or transformation. The problem with these approaches is that we have to rely on relatively strong assumptions such as the absence of obstacles to capital mobility and liquidity constraints or a concrete specification of technology or preferences.

More hopeful, probably, is the use of financial market data to price –or to approximate the prices of– real riskless bonds using non-arbitrage conditions. For example, the approach suggested by Hansen and Jagannathan (1991) allows mean-variance frontiers to be derived for a common stochastic discount factor for future payoffs which is, of course, linked to the average implicit riskless rate. More promising, however, is the recent contribution by Flood and Rose (2005), which derives implicit riskless rates in non-arbitrage economies by exploiting the idiosyncratic risk of the securities traded in the financial markets.

In this paper we obtain some evidence on the course of real interest rates in Spain since the beginning of the 1990s by combining several macroeconomic and financial approaches. Our analysis is based on two sub-samples. The first subsample covers the pre-EMU period (1990 to 1998), whereas the second is the EMU period (1999 to 2005). We do not intend to provide point estimates of real interest rates and compute how much they have moved over this period. We analyse instead whether the fall in real interest rates between these two sub-periods suggested by inflation-adjusted interest rates can be reconciled with macroeconomic and financial theory. The evidence we provide does not support this hypothesis, suggesting therefore that real interest rates may have fallen much less than what the conventional estimates of this unobserved variable indicate.

The rest of the paper is organised as follows. In the first section we analyse foreign interest rate data and exploit several specifications of preferences and technologies to derive equilibrium conditions for domestic real interest rates. In the second section we analyse the extent to which Hansen-Jagannathan frontiers can help us to learn about the changes of real interest rates between the two subsamples considered. We then ask the same question by exploiting the Flood and Rose (2005) approach. Section 3 concludes.

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).

Table 1 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 1.5 and 1.8 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 mismeasurement of the actual decline in the riskless real interest rate in the Spanish case. The first hypothesis is, however, very unlikely. During the nineties there were not significant barriers to cross-trading within national debt markets in Europe. Indeed, non-residents held, on average, almost one fifth of the outstanding stock of the Spanish government debt market between 1990 and 1998.

Table 1: REAL INTEREST RATES

	Ex-post 3-month real interest rate				10-year idexed 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

Source: Banco de España.

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 higher employment and participation rates. According to various estimates, average potential GDP growth was about 0.5% higher in the period 1999-2005 than in the period 1990-1998[2]. 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 in the second period, 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.

(2) See, for example, Denis *et al.* (2006).

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. More specifically, we know from the first order equilibrium conditions of a representative agent that $E(m) = (1 + r)^{-1}$, where m is the IMRS and r is the actual real interest rate. In Table 2, we provide the average implicit interest rate derived from this expression for several specifications of preferences. All data are drawn from Spain's Quarterly National Accounts.

Table 2: Implied real interest rates derived for alternative specifications of preferences

		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
Constantinides b = -0.5	gamma = 0.1	2.21	2.27	0.06
	gamma = 1	3.90	4.54	0.64
	gamma = 5	11.51	15.09	3.57
Constantinides b = -0.25	gamma = 0.1	2.21	2.27	0.06
	gamma = 1	3.91	4.53	0.62
	gamma = 5	11.55	15.05	3.50
Constantinides b = 0.25	gamma = 0.1	2.22	2.27	0.05
	gamma = 1	3.96	4.47	0.51
	gamma = 5	11.37	14.38	3.02
Constantinides b = 0.50	gamma = 0.1	2.24	2.26	0.02
	gamma = 1	3.94	4.21	0.26
	gamma = 5	7.12	10.04	2.92

Table 2: IMPLIED REAL INTEREST RATES DERIVED FOR ALTERNATIVE SPECIFICATIONS OF PREFERENCES (continuation)

		1990-1998	1999-2005	Change
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 N_j is the number of quarters in sub-period j, and mt is the IMRS in period t, which is proxied using the several specifications of preferences. For isolelastic preferences, $m_t = \beta(g_{t+1})^{-\gamma}$; where $g_{t+1} = C_{t+1} / C_t$ and c_t 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 Constantinides preferences

$$m_t = \beta \frac{(g_{t+1}g_t - bg_t)^{-\gamma} - b\beta(g_{t+1}g_t)^{-\gamma}(g_{t+1} - b)^{-\gamma}}{(g_t - b)^{-\gamma} - b\beta g_t^{-\gamma}(g_{t+1} - b)^{-\gamma}}$$

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 N_t is the ratio of employment over the population aged over 16. We use quarterly data from Spain's National Quarterly Accounts and β is set to 0.995.

Source: Own elaboration.

We first use the standard isoelastic CRRA utility function,

$$U_t = \frac{C_t^{1-\gamma}}{1-\gamma}$$

where C_t is per capita consumption. The IMRS is given by $m_t = \beta(g_{t+1})^{-\gamma}$ where $g_{t+1} = C_{t+1} / C_t$ and β is the discount factor. In our empirical implementation, β is set to 0.995 and we use seasonally-adjusted private non-durable consumption in real terms. As is well known, in this case, preferences are not time-dependent and γ measures both the household degree of relative risk aversion and the inverse of the elasticity of intertemporal substitution between future and current consumption. Under this specification 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 was, on average, almost 1% higher in the second sub-period.

Using relative consumption effects and time dependence in preferences does not change the picture much. In particular, we use Abel's preferences [Abel (1990)], which assume the following utility function:

$$U_t = \frac{c_t^{1-\gamma}}{1-\gamma} C_{t-1}^{\Phi}$$

where c_t is individual consumption. In this case, household preferences not only depend on own consumption, but also on the aggregate level of consumption. The parameter Φ measures the dependence of the individual utility on the general level of consumption in the economy. Evaluated in equilibrium, where individual and per capita consumption coincide, the IMRS is given by $m_t = \beta(g_{t+1})^{-\gamma}(g_t)^{\Phi}$.

In addition, we use time dependence preferences that account for the existence of habits or durability in household decisions. In particular, we use the following utility function proposed by Constantinides (1990)

$$U_t = \frac{(c_t - bc_{t-1})^{1-\gamma}}{1-\gamma}$$

where b measures the degree of habits or durability. Note that b > 0 means the existence of habits (since current utility is a negative function of past consumption), whereas b < 0 implies the existence of durability. With these preferences, the IMRS is given by

$$m_t = \beta \frac{(g_{t+1}g_t - bg_t)^{-\gamma} - b\beta(g_{t+1}g_t)^{-\gamma}(g_{t+1} - b)^{-\gamma}}{(g_t - b)^{-\gamma} - b\beta g_t^{-\gamma}(g_{t+1} - b)^{-\gamma}}$$

For both specifications and using 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 increased markedly in Spain in the period under study, one could conjecture that this might compensate the positive effect of consumption growth on the marginal rate of substitution. In particular, we use the KPR preferences[3]:

$$U = \frac{1}{1-\sigma}\left[\left(C_t^i\right)^a \left(1 - N_t^i\right)^{1-a}\right]^{1-\sigma}$$

where N is the working hours. Parameter γ measures the intertemporal substitution attitudes of households and a is a parameter that allows us to pin down the steady state value of hours. In this model the IMRS is given by $m_t = \beta(g_{t+1})^{-\gamma}(g_{t+1}\, n_{t+1})^{-(1-a)(1-\gamma)}$ where $n_{t+1} = (1 - N_t)/(1 - N_{t+1})$. In our empirical application we proxied N_t as the ratio of employment over the population aged over 16. Using this model 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 2pp. 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

(3) See King, Plosser and Rebelo (1988).

in other countries. Therefore, these results suggest that most of the large fall in ex-post real interest rates in Spain cannot be explained by the main economic determinants of the actual real interest rate. This indicates that a significant part of the decline of the ex-post real interest rates could well be due to both expectational errors on inflation (i.e. realised inflation was lower than expected inflation) during the pre-EMU period and by the decrease in the inflation risk premium.

2. THE FINANCE APPROACH

In Section 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 pure non-arbitrage conditions in financial markets. These conditions imply that all securities should be priced by applying a positive stochastic discount factor to their future payoffs. The stochastic discount factor –which in equilibrium models would be equivalent to the IMRS of the representative agent– is directly linked to the return on a riskless security [see, for instance, Huang-Litzenberger (1988)]. Obviously, the exact identification of the stochastic discount factor that prevents arbitrage opportunities is not possible as markets are, in practice, incomplete, and also because econometricians can normally play only with a limited set of financial instruments. There are some methods, however, that can be used to extract some helpful information.

2.1. The Hansen-Jagannathan frontier

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

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

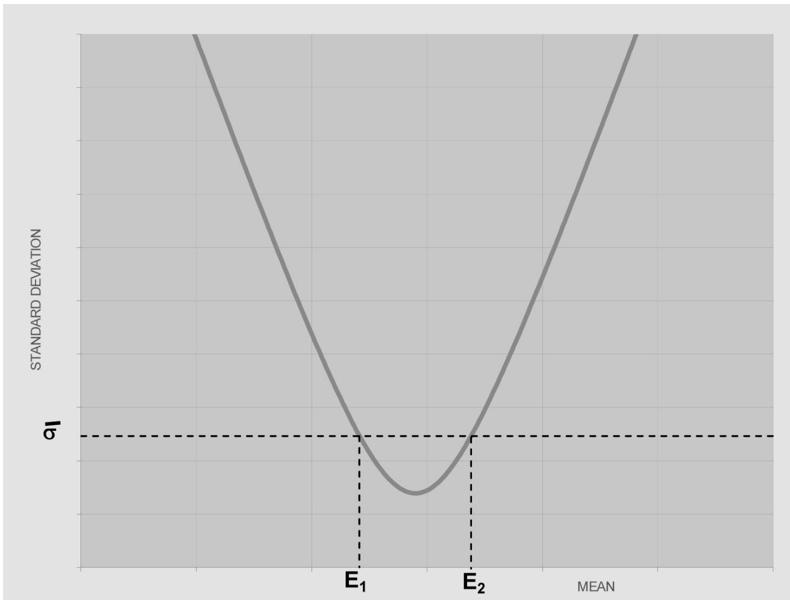
where p is the vector of security prices, x is the vector of payoffs, Σ is the variance-covariance matrix of payoffs and E() is the unconditional expectation operator. It is apparent from expression [1] that to compute the HJ frontier we only need securities market data.

Note that, by restricting the standard deviation of the IMRS to a maximum level (σ), we can obtain a lower (E_1) and an upper (E_2) bound for the average level of the IMRSs (see Figure 2) and, implicitly, for the real interest rate [remember that, $E(m) = (1 + r)^{-1}$].

In this section we use this approach to find bounds for the average level of the actual real interest rates. To do so, 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[4]. Returns are computed in real terms (deflated by the Spanish CPI index) assuming a holding period of one month.

(4) Annex 1 describes the composition of these portfolios and the computation of the monthly real returns.

Figure 2: HANSEN-JAGANNATHAN FRONTIER

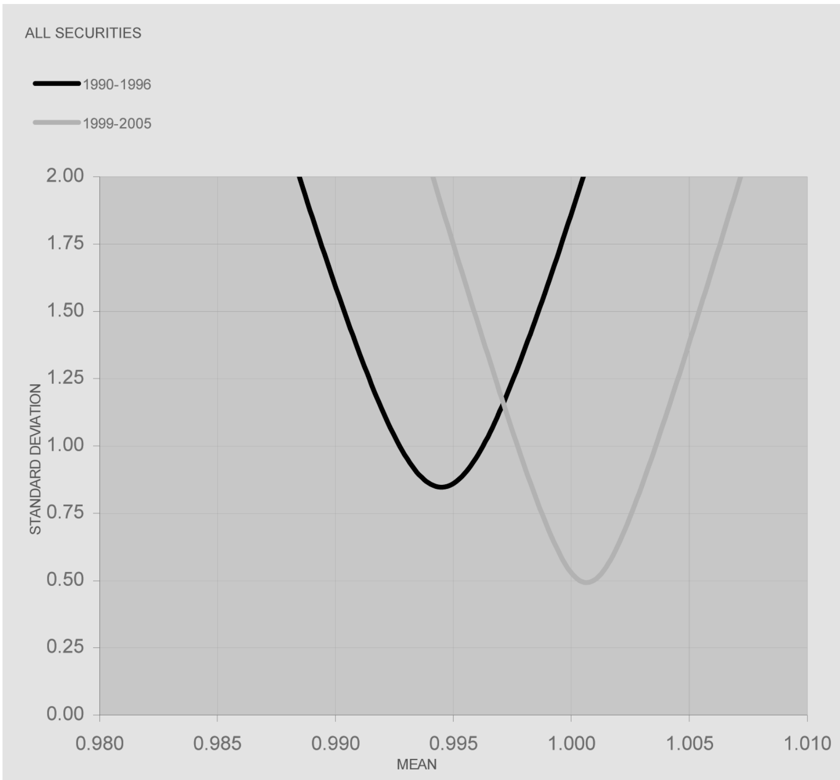


Source: Own elaboration.

Figure 3 shows the HJ frontiers estimated for the periods 1990-96 and 1999-2005 using all the securities in our dataset. We exclude the years 1997 and 1998 from the analysis because it is an interim period where security prices are likely to already incorporate 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 IMRSs are relatively narrow in both periods and they do not overlap. In particular, the means of the IMRSs 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 inflation expectations during the first period were systematically higher than observed inflation, the average ex-post return and, therefore, the inverse of the estimated mean of the IMRSs would be overstated[5].

(5) Interestingly, during the second period, the opposite could have happened. In particular, the mid-point of the bound suggests a negative real interest rate, consistent with the negative ex-post interest rate during most of this period (as can be seen in Figure 1). However, real rates could have been positive during this period if inflation expectations were lower than observed inflation.

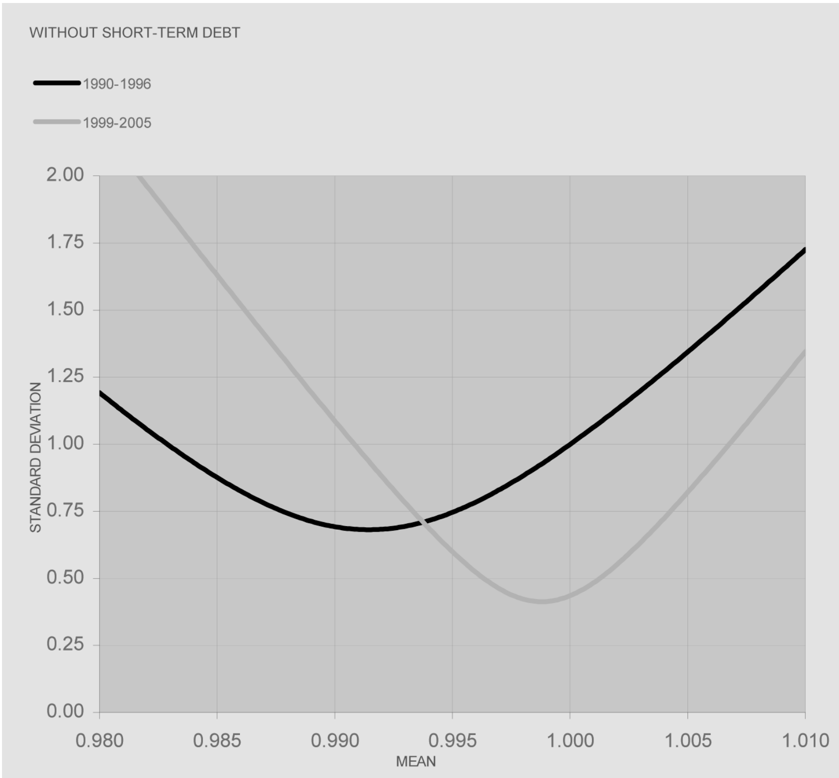
Figure 3: HANSEN-JAGANNATHAN FRONTIERS



Source: Own elaboration.

Therefore, we repeat the same exercise excluding the short-term securities but retaining the longer-term fixed-income instruments. Figure 4 shows the results. We can see that the size of the region of the admissible pairs of mean and standard deviation of IMRSs increases dramatically for the two periods. Also, the two regions are now much closer compared with Figure 3. Thus, it is much harder to reject the hypothesis of equal average levels of real interest rates in the two periods. However, even in this case, results can be contaminated by a peso problem since the cash flows associated with conventional bonds are fixed in nominal terms. Therefore, if inflation expectations were systematically higher than the realised inflation, the ex-post return would be overstated. This should not be the case, however, of stocks since the associated cash flows vary with realised inflation. Therefore, in this case the distribution of real returns should not be contaminated by the peso problem.

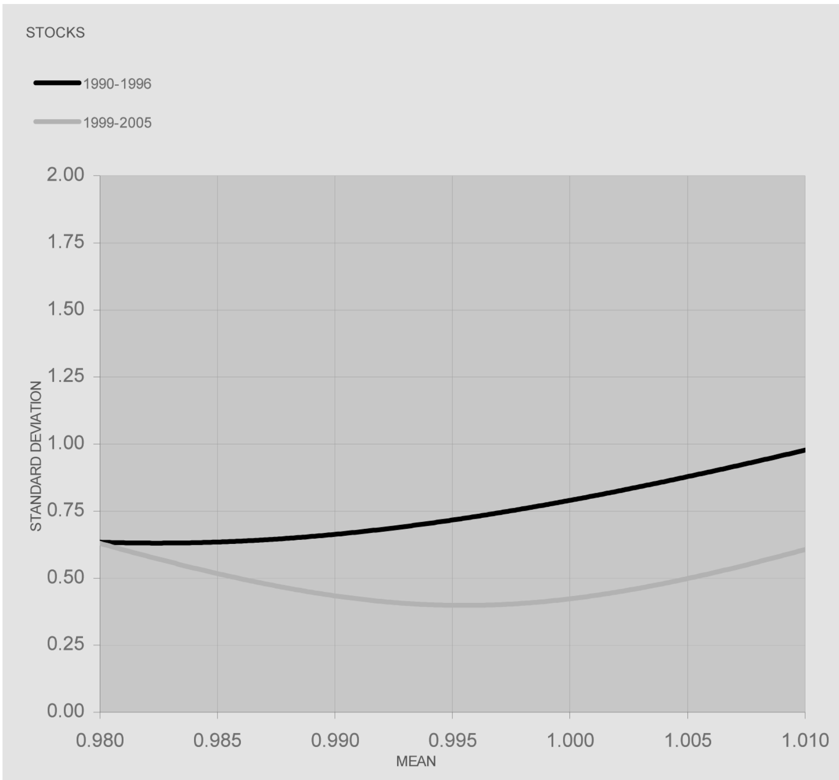
Figure 4: HANSEN-JAGANNATHAN FRONTIERS



Source: Own elaboration.

Figure 5 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.

Figure 5: HANSEN-JAGANNATHAN FRONTIERS



Source: Own elaboration.

2.2. Exploiting the 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) (2005), which allows us to obtain point estimates for this variable as opposed to ranges.

FR consider the standard decomposition of the Euler equation:

$$p_t^j = E_t(m_{t+1}x_{t+1}^j) = COV_t(m_{t+1}, x_{t+1}^j) + E_t(m_{t+1})E_t(x_{t+1}^j) \qquad [2]$$

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 IMRS 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 [2] can be rewritten as

$$x_{t+1}^j = \delta_t (p_t^j - COV(m_{t+1}, x_{t+1}^j)) + \varepsilon_{t+1}^j \qquad [3]$$

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 [3] stationary is to normalise by p_t^j. FR propose normalising by the systematic 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.

$$x_{t+1}^j / p_t^j = \delta_t (p_t^j / p_t^j - COV(m_{t+1}, x_{t+1}^j / p_t^j)) + \varepsilon_{t+1}^j / p_t^j \qquad [4]$$

FR rewrite equation [4] as

$$x_{t+1}^j / p_t^j = \delta_t (p_t^j / p_t^j) + u_{t+1}^j \qquad [5]$$

where $u_t^j = \varepsilon_{t+1}^j / p_t^j - \delta_t COV(m_{t+1}, x_{t+1}^j / p_t^j)$. They note that assuming that $COV(m_{t+1}, x_{t+1}^j / p_t^j)$ moves only because of aggregate phenomena, δ_t in [5] 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

$$\ln(p_t^j / p_{t-1}^j) = \alpha_0^j + \sum_{i=1}^{N} \alpha_i^j f_t^i + v_t^j \qquad [6]$$

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 [6], the estimated systematic price is defined as

$$\hat{p}_t^j = p_{t-1}^j \exp\left(\hat{\alpha}_0^j + \sum_{i=1}^{N} \hat{\alpha}_i^j f_t^i \right) \qquad [7]$$

In their empirical implementation, FR estimate regressions [6] using as factors the market-wide stock market return and the three Fama-French factors: the overall market return minus 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, the 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

$$x_{t+1}^j / \hat{p}_t^j = \delta_t (p_t^j / \hat{p}_t^j) + u_{t+1}^j \qquad [8]$$

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 [8] might potentially understate the OLS standard errors. To handle both these potential econometric problems, they estimate [8] 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 IMRS in period t.

Note that the basic idea of the estimation procedure suggested by FR is to use asset-idiosyncratic shocks (which are captured by regressor $p_t^j / \hat{p}_t^j$) to identify and measure the (inverse of the) expected marginal rate of substitution. While idiosyncratic shocks carry no information about individual asset risk premia, they are loaded with information relevant to market aggregates, as equation [8] shows, since these shocks earn the expected marginal rate of substitution. The main advantage of this method is that it does not rely on a specific asset pricing model since asset risk premia do not play any role in the pricing of idiosyncratic shocks.

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 IMRS parameter is assumed to be fixed within the two periods of interest. More specifically, we estimate the following regression

$$x_{t+1}^j / \hat{p}_t^j = \delta_1 (p_t^j / \hat{p}_t^j) + \delta_2 (p_t^j / \hat{p}_t^j) D99_t + u_{t+1}^j \qquad [9]$$

where $D99_t$ is a dummy variable which takes value 1 from January 1999 on. In regression [9] 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 [9] is estimated by GMM using the first lag of the explanatory variables as instruments. Table 3 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% (=1.005^{12}-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.

Table 3: ESTIMATION RESULTS FOR THE MEAN OF THE IMRS DERIVED FROM THE METHOD PROPOSED BY FLOOD AND ROSE (2005)

	Coeficient	Std. error
δ_1	1.0053	0.0037
δ_2	0.0025	0.0050

Source: Own elaboration.

All in all, results reported in this section, based essentially 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.

3. CONCLUDING COMMENTS

This paper has provided a number of arguments and evidence supporting the hypothesis that the observed decrease in ex-post real interest rates –of more than seven percentage points– between 1990 and 2005 is likely to overestimate the fall in the cost of capital –as measured by the actual riskless real interest rate– experienced by the Spanish economy.

Although our estimates are subject to much uncertainty, mostly as a consequence of the difficulty of measuring real interest rate levels with sufficient precision, we have seen that a decrease in real interest rates of a similar size to that in ex-post rates does not seem compatible either with the hypothesis of capital market integration or with rational optimising behaviour on the part of investors and consumers. Moreover, exploiting non-arbitrage conditions, we have shown that the behaviour of other security prices does not suggest such a large fall in the riskless real interest rate. Actually, our findings do seem compatible with the hypothesis that real interest rates when properly measured-might not have declined much more than in other more stable economies. If the ex-post real return on nominal bonds is lower now, this is partly the result of a change in the risk profile of these instruments as inflation uncertainty lessens. Specifically, the large fall in ex-post real interest rates would have to be explained, at least to some extent, by the impact of the new monetary regime on inflation expectations and the inflation risk premium and not only as a result of a genuine reduction in the cost of capital.

The implications of this hypothesis are potentially very relevant. The explanation of the significant expansion of the Spanish economy would probably have to rely somewhat less on low financial costs and more on employment creation, modernisation and competition in the financial sector and improved expectations as a consequence of the consolidation of an environment of macroeconomic stability. More reflection would, however, be needed to reassess the determinants of the marked expansion of private-sector –and, particularly, household– debt. It is a fact that such an increase has been larger than in other countries facing what now seems a less dissimilar reduction in actual real interest rates. One explanation could well be that the fall in nominal interest rates –even if it is not accompanied by a similar decrease in real rates– can actually relax credit constraints applied by banks. Indeed, there is some evidence that nominal rather than real rates explain developments in household credit in Spain[6]. But it might well be the case that the continuous expansion of demand for loans is partly due to a failure by borrowers to fully internalise the lower protection that they could expect from inflation in the new monetary union regime.

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 runs from January 1990 to December 2005.

The 10 size portfolios are constructed from a dataset which includes all stocks traded on the electronic segment of the Spanish stock exchanges ("mercado continuo")[7]. 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 constructed using the total return (including dividends) sectoral indices published by the Madrid Stock Exchange (MSE). Between 1940 and 2001 the MSE used 10 sectoral indices. Starting in 2002, these series were discontinued and new series were created. The new sectoral classification offers more detailed information. 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.

(6) See, for example, Nieto (2003) and Martínez-Carrascal and del Río (2004).

(7) These data were provided by Gonzalo Rubio for the period 1990 to 2003. We have updated the data until the end of 2005 using the same methodology.

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RESUMEN

En este trabajo se analiza la evolución de los tipos de interés reales en la economía española entre 1990 y 2005. Como no se dispone de bonos indiciados a la inflación, los cambios en los tipos de interés reales implícitos se estiman utilizando distintos procedimientos sugeridos por la teoría macroeconómica y financiera. En particular, se utilizan las condiciones de equilibrio de un agente representativo bajo varias especificaciones alternativas de las preferencias. Además, se explotan las condiciones de no arbitraje en los mercados financieros. La evidencia que se presenta indica que la incertidumbre ligada a la inflación podría explicar una parte importante de la caída observada en los tipos de interés nominales. En consecuencia, el coste real de financiación podría haberse reducido sustancialmente menos que lo que sugiere la evolución de los tipos de interés reales *ex post.*

Palabras clave: tipos de interés reales, tasa marginal de sustitución intertemporal.

Clasificación JEL: E43, G12.