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(*) We thank for helpful comments José Luis Escrivá and Soledad Núñez.

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ABSTRACT

We use an integrated framework based on the CCAPM to jointly estimate *ex-ante* real interest rates, (bounds on) inflation risk premia and (bounds on) agents' inflation expectation errors in four countries -France, Spain, UK and US- under three different preference specifications. According to our results, 1-year *ex-ante* real interest rates exhibit a rather low correlation with the 1-year *ex-post* rate. For those *ex-ante* real rates showing the highest correlation with the *ex-post* real rate, the difference between both real rates seems to be mainly explained in terms of agents' inflation expectation errors, while the inflation premia play a minor role.

1. INTRODUCTION

Real interest rates are a critical variable in economics. They measure the real cost of capital, thus playing a crucial role in determining long-run output growth. Moreover, real as opposed to nominal interest rates determine agents' consumption and investment decisions. Thus, when central banks steer (short-term) nominal interest rates in order to attain monetary policy targets, disentangling the relationships between nominal interest rates, real interest rates and inflation expectations is crucial to understand the relative importance of different channels in the monetary policy transmission mechanism.

Unfortunately, real interest rates are non-observable and are usually proxied by the so-called *ex-post* real interest rates, i.e., the difference between the nominal interest rate and the *ex-post* observed inflation rate. As is well known, however, *ex-post* rates include two disturbing components which can render them a misleading proxy for non-observable *ex-ante* real interest rates: inflation risk premia and agents' inflation expectation errors. Like *ex-ante* real interest rates, these two variables are also non-observable.

There are reasons to think that both disturbances are probably negligible. First, inflation premia can hardly be relevant if the inflation rate is not very volatile. And second, if agents are rational when forming their inflation expectations, the expectation error should be zero on average. Although it is still an open question, this view has been recently challenged in the literature, particularly in relation to the magnitude of the inflation expectation error. Thus, a series of papers have found that, due to informational or to (monetary policy) credibility problems, inflation rates can be successfully characterised by switching-regime models à la Hamilton, not only in high-inflation countries like Argentina, Israel or Mexico (see Kaminsky and Leiderman, 1996) but also in countries whose inflation rates are lower and more stable like the US (Evans and Lewis, 1995) or Canada (Bank of Canada, 1996). These switching-regime models produce inflation expectation errors which have zero-mean *ex-ante* but, *ex-post*, can show a non-zero mean. Similarly, according to King (1996) if agents do not immediately learn about central bank behaviour, disinflationary processes will probably be characterised by inflation targets (and, therefore, by actual inflation) below agents' inflation expectations. Lasting inflation expectation errors are also predicted by models à la Backus-Driffill (1985) where central bankers face credibility problems and need time to build their anti-inflationary reputation.

In this paper we use an integrated framework that provides a joint estimate of the *ex-ante* real interest rate, (bounds on) inflation premia and (bounds on) inflation expectation errors, on the basis of the Consumption Capital Asset Pricing Model (CCAPM). This framework allows, first, for a comparison between *ex-ante* and *ex-post* real interest rates to analyse to what extent the latter are a good proxy for the former. Second, relative to other alternative methods that estimate *ex-ante* real interest rates, this approach also offers a simultaneous estimate of inflation premia and inflation expectation errors, thus providing an explanation of the discrepancies between both interest rates. This ability to analyse the three variables jointly is one of the advantages of this method. An additional advantage of our approach lies in its data requirements. All the results are obtained using information on non-durable consumption alone which, thus, acts as a sufficient statistic for the three non-observable variables.

Most of the existing empirical evidence is not generally favourable to the CCAPM (see, for instance, Sentana 1993). Yet such evidence, is usually based on a joint test of the CCAPM and the isoelastic preference assumption for data on the US economy. When other preference specifications are considered, however, the evidence is more favourable to the CCAPM (see Cochrane and Hansen, 1992; and Campbell and Cochrane, 1995). The same applies when other economies are analysed (see Ayuso, 1996 for Spanish evidence, or Bakshi and Naka, 1997 for Japanese data). In this paper, we address this issue through the analysis of four countries and three alternative preference specifications. Moreover, we do not confine ourselves to any particular estimate of the parameters that characterise each utility function but consider, for each preference specification, a (sensible) set of values of the relevant parameters. Thus, we generate *ex-ante* real interest rates, inflation premia and inflation expectations under different assumptions, which allows for an analysis of the sensitivity of those variables to changes in important features in agents' utility functions: time preference, risk aversion, intertemporal substitution, habits and keeping-up-with-the-Joneses effects.

As usual in any calibration analysis, we look at the variance and correlation of the *ex-post* real rates and the (model-) generated *ex-ante* real rates under alternative utility functions and parameterizations. In doing so, we can select for each country in our sample the most suitable preference specification in terms of those statistics. Once we select the specification, we duly construct bounds on inflation expectations and risk premia. This allows us to discuss the sources of discrepancy between the *ex-post* real rates and the (model-)generated *ex-ante* rates.

The paper is organised as follows. Section 2 briefly states the main theoretical relationships between *ex-ante* real interest rates, *ex-post* real interest rates, inflation risk premia and inflation expectation errors, within the CCAPM framework. Section 3 presents the empirical results, based on data from Spain, France, UK and US. These results show that, in general, *ex-post* real interest rates are a poor proxy for *ex-ante* real rates, the correlation between them being rather low. For those *ex-ante* real rates showing the highest correlation with the *ex-post* real rate, the different behaviour of both series seems to be mainly due to the existence of inflation expectation errors, while the inflation premia seem to play a minor role. Finally, Section 4 draws the main conclusions of the paper.

2. NOMINAL INTEREST RATES, REAL INTEREST RATES, INFLATION PREMIA AND INFLATION EXPECTATIONS IN A CCAPM FRAMEWORK.

As is well-known, the CCAPM states that, in equilibrium, the expected returns on all financial assets must be the same, provided they are properly discounted. The model also gives the proper discount factor: the marginal rate of substitution between future and current consumption. Formally,

$$E_t\,[MRS_{t+1}\; \Pi^{-1}_{t+1}\; R^{j}_{t+1}] = 1, \quad \forall t, j, \qquad [1]$$

where MRS_{t+1} is the marginal rate of substitution between tomorrow's and today's consumption[1], Π_{t+1} is the ratio between tomorrow's and today's price levels, R^{j}_{t+1} is today total 1-period-ahead return (including the principal returned) on asset j, and E_t represents the expectation operator conditional on information available at t.

For the purpose of this paper, we are interested in two particular financial assets. First, we consider a 1-period default-free zero-coupon bond. Its riskless *nominal* interest rate -i.e. known at t- is denoted by i_t. Next, we also consider a perfectly indexed 1-period default-free zero-coupon bond. Its riskless *real* interest rate, known at t, is denoted by r_t. The latter can be understood as a "real bond" that returns units of consumption at its maturity date. It is seen that in those two cases equation [1] takes the following form:

$$E_t[MRS_{t+1}]\,(1+r_t) = 1, \quad \forall\, t. \qquad [1.a]$$

$$E_t[MRS_{t+1}\; \Pi^{-1}_{t+1}]\,(1+i_t) = 1, \quad \forall\, t. \qquad [1.b]$$

Equations [1.a] and [1.b] impose statistical restrictions on the comovements between interest rates (nominal and real), expected inflation and consumption:

$$\frac{1}{(1+i_t)} = \frac{1}{(1+r_t)} E_t[\Pi^{-1}_{t+1}] + COV_t[RMS_{t+1}, \Pi^{-1}_{t+1}], \ \forall t. \qquad [2]$$

Thus, this model provides a general framework to relate nominal and real interest rates. This relationship is, moreover, broadly consistent with the well-known Fisher equation. To illustrate that point in a simplified way, we can use the lognormal version of equation [2], i.e., the relationship among the relevant variables when MRS_{t+1}, Π^{-1}_{t+1} and R^{j}_{t+1} follow

[1] That is, the time preference parameter β times the ratio of the marginal utilities of tomorrow's and today's consumption.

lognormal distributions. In that case, it can be shown (see, for example, Evans and Watchel, 1992) that equation [2] can be rewritten as:

$$i_t = r_t + E_t[\pi_{t+1}] + Cov_t[rms_{t+1}, \pi_{t+1}] - \frac{1}{2} V_t[\pi_{t+1}] \quad [2']$$

where small letters denote logs of capital ones. Equation [2'] states that the nominal interest rate is positively related to the real interest rate and the expected inflation as in the conventional Fisher equation. But the conditional covariance between the ratio of marginal utilities to inflation also enters into this relationship. In fact, this term can be interpreted as an inflation risk premium(2): if the covariance term is positive (negative), the nominal bonds provide a poor (good) hedge against unanticipated consumption changes and households require higher (lower) nominal interest rates.

From expression [2'], and the usual definition of *ex-post* real rates:

$$r_t^{ex\text{-}post} = i_t - \pi_{t+1}$$

it is clear that inflation premium acts as a first potential wedge between the *ex-post* real interest rate $r_t^{ex\text{-}post}$ and the *ex-ante* real interest rate r_t. But even if the inflation premium is nil, there might still be another wedge if agents do not perfectly forecast future inflation. Note, then, that if there is no inflation premium (and we obviate the Jensen inequality term) then

$$i_t = r_t + E_t[\pi_{t+1}] = r_t + \pi_{t+1} + (E_t[\pi_{t+1} - \pi_{t+1}]) ==>$$

$$r_t^{ex\text{-}post} = r_t + (E_t[\pi_{t+1}] - \pi_{t+1})$$

The difference between the expected inflation rate and the observed one, i.e. the expectation error, will also contaminate the *ex-post* real interest rate.

(2) The conditional variance is devoid of economic content, and appears as a consequence of the log-approximation due to the Jensen inequality.

As commented in the Introduction, *ex-ante* real interest rates, inflation premia and inflation expectation errors are non-observable. Equations [1.a], [1.b] and [2] can be used to estimate them.

Real interest rates

Real interest rates can be readily estimated under equation [1.a]:

$$1 + r_t = \frac{1}{E_t(MRS_{t+1})}, \qquad \forall t. \qquad [3]$$

Observe that, in order to estimate *ex-ante* real interest rates, we only need information on the stochastic discount factor MRS_{t+1} , i.e. information on preferences and per-capita consumption.

Bounds on inflation expectations and on inflation premia

Estimating inflation expectations and inflation premia is not so straightforward. Nevertheless, it is still possible to derive bounds for inflation expectations through the implied bounds for the risk premium term in equation [2], as in Ireland (1996).

Since the risk premium is a covariance, the following relationship holds:

$$\rho_t[MRS_{t+1}, \Pi_{t+1}^{-1}] = \frac{Cov_t[MRS_{t+1}, \Pi_{t+1}^{-1}]}{\sigma_t[MRS_{t+1}]\, \sigma_t[\Pi_{t+1}^{-1}]}$$

where ρ_t and σ_t denote conditional correlation and conditional standard deviation, respectively. But the correlation coefficient must lie inside the interval [-1, 1]. Thus:

$$|Cov_t(MRS_{t+1}, \Pi_{t+1}^{-1})| \leq \sigma_t(MRS_{t+1})\, \sigma_t(\Pi_{t+1}^{-1}) \qquad [4]$$

Under an additional assumption on inflation volatility, namely:[3]

[3] According to the available empirical evidence (see Ireland, 1996 and Ayuso and López-Salido, 1996) this does not seem to be too restrictive an assumption .

$$\sigma_t[\Pi_{t+1}^{-1}] \leq E_t[\Pi_{t+1}^{-1}] \qquad [5]$$

we can eliminate inflation from equation [4], thereby reducing our data requirements to consumption information. Thus, equations [1.a], [2], [4] and [5] yield

$$(1+i_t)\ \{E_t[MRS_{t+1}] - \sigma_t[MRS_{t+1}]\} \leq \{E_t[\Pi_{t+1}^{-1}]\}^{-1} \leq \{E_t[MRS_{t+1}] + \sigma_t[MRS_{t+1}]\}\ (1+i_t) \qquad [6]$$

Since $\{E_t(\Pi_{t+1}^{-1})\}^{-1} \approx E_t(\Pi_{t+1})$, equation [6] defines bounds on expected inflation that can be compared with observed inflation, thus providing bounds on the expectation errors. Observe that, as in the case of real interest rates, all we need to estimate those bounds is data on consumption and information on agents' preferences, in addition to data on nominal interest rates. Moreover, the width of the band for expected inflation also provides information on the magnitude of the inflation premium. Observe that by using equation [3], equation [6] can be rewritten as

$$\frac{(1+i_t)}{(1+r_t)} - (1+i_t)\ \sigma_t[MRS_{t+1}] \leq \{E_t[\Pi_{t+1}^{-1}]\}^{-1} \leq \frac{(1+i_t)}{(1+r_t)} + (1+i_t)\ \sigma_t[MRS_{t+1}]$$

and $(1+i_t)\,/\,(1+r_t)$ is the inflation expectation estimate provided the inflation premium is zero. Therefore, equation [6] yields a band for inflation expectations centred around the inflation expectation under risk neutrality, (half) the width of that band being the (absolute) maximum value for the inflation premium.

As stated in the Introduction, the empirical performance of the CCAPM seems to depend on the chosen agents' utility functions. We close this section with a brief overview of three different utility functions -commonly used in the asset-pricing literature- that will be used later in the empirical part of the paper.[4] The first one is the well-known isoelastic utility function, and the other two encompass the isoelastic as a particular case.

Constant relative risk aversion

The most usual utility function in the financial literature is the isoelastic utility function:

$$U_t = \frac{C_t^{1-\gamma}-1}{1-\gamma}, \qquad \gamma \neq 1.$$

where $\gamma = 1$ implies the log utility function. Notice that preferences are not time-dependent and γ measures both the household degree of (constant) relative risk aversion and the inverse of the (constant) elasticity of intertemporal substitution between future and current consumption.[5] It is easy to prove that, in this case

$$MRS_{t+1} = \beta \, (g_{t+1})^{-\gamma} \qquad [7]$$

where $g_{t+1} = C_{t+1}/C_t$

Habit formation or durability

A simple way of considering time dependence in preferences is assuming habits in household decisions. Following Constantinides (1990) and Heaton (1993), a parsimonious way of capturing these effects is to specify the following utility function:

(4) See, for instance, Cochrane and Hansen (1992) and Kocherlakota (1996).

(5) In order to circumvent this restriction, we tried using the generalised isoelastic preferences proposed by Epstein and Zin (1989) and Weil (1990). The empirical implementation of this preference specification poses some problems (see Ayuso and López-Salido, 1996) which can not be successfully addressed for all the countries concerned. These results are available from the authors upon request.

$$U_t = \frac{(C_t - \lambda C_{t-1})^{1-\gamma}}{1-\gamma}, \qquad \gamma \neq 1.$$

where λ measures the degree of habits (durability when negative) and, for U_t to be well defined, $\lambda < Min\{g_t\}$ is assumed.[6] Notice that $\lambda > 0$ implies that it takes more consumption today to make an investor happier if he consumed more yesterday -i.e. he has habits. The degree of relative risk aversion is a function of γ, λ and g_t, and the corresponding marginal rate of substitution is:

$$MRS_{t+1} = \beta \frac{(g_{t+1}g_t - \lambda g_t)^{-\gamma} - \lambda\beta(g_{t+1}g_t)^{-\gamma}(g_{t+1}-\lambda)^{-\gamma}}{(g_t-\lambda)^{-\gamma} - \lambda\beta g_t^{-\gamma}(g_{t+1}-\lambda)^{-\gamma}} \qquad [8]$$

Relative consumption

The last extension of household preferences we consider is the presence of externalities in the utility functions. In particular, following Abel (1990) or Galí (1994), household preferences may not only depend on own consumption but also on the aggregate (economy-wide) level of consumption. A simple way of capturing this is assuming the following utility function:

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

where c_t is individual consumption at t, as opposed to per-capita consumption in the economy C_t. A positive (negative) value of ϕ implies that the individual is altruistic (invidious) in that the more the society consumes the better off (worse) he is.[7] In equilibrium, where individual and per-capita consumption coincide, the marginal rate of substitution takes the following form:

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[6] By setting $\lambda = 0$, these preferences reduce to the isoelastic utility function.

[7] This utility function reduces to the isoelastic case when $\phi = 0$.

$$MRS_{t+1} = \beta\,[g_{t+1}]^{-\gamma}\,[g_t]^{\phi} \qquad [9]$$

3. EMPIRICAL EVIDENCE

In this section we empirically address the question posed in the title of the paper. We consider 1 year as the empirical definition of the theoretical concept of "1 period" and therefore, we aim at estimating 1-year *ex-ante* real interest rates and at bounding inflation expectation errors and inflation premia to this horizon. However, so as to increase the number of observations we use quarterly per-capita private non-durable consumption and CPI data on France, Spain, UK and US from 1970:I to 1995:IV.(8) Moreover, in order to bound inflation expectations we also use the corresponding 1-year nominal interest on public debt, except for Spain, for which the longest available series is the 1-year domestic interbank market. Even so, the periods covered are 1970:I to 1995:IV for UK and US, 1978:I to 1995:IV for France, but only 1979:III to 1995:IV for Spain.(9)

The strategy we propose for analysing the potential error when *ex-post* real interest rates are used instead of *ex-ante* real interest rates is the following. First, for each of the three considered preference specifications, we choose a sensible range for the parameters involved. For each parameter set, we compute the corresponding *ex-post* marginal rate of substitution. Then, the expected MRS_{t+1} is obtained as the forecast from an ARMA model for the *ex-post* marginal rate of substitution so

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(8) Our data source are the QNA of the OECD, but for Spanish data that have been provided by the INE. Notice that using quarterly data induces a data overlapping problem that is taken into account when obtaining the 1-year *ex-ante* real interest rates. Figure 1 shows consumption growth patterns in each country.

(9) We also tried using German data but we did not find data on non-durable consumption and available data on total consumption showed strange patterns around German reunification difficult to address.

obtained.[10] Given that expectation, equation [3] immediately provides the corresponding model-generated 1-year *ex-ante* real interest rate.

Next, we compare the different model-generated *ex-ante* real interest rates to the *ex-post* rate according to the following criterium. Unless there are implausible negative correlations between *ex-ante* real rates, inflation premia and inflation expectation errors, it is clear from Section 2 that the variance of the *ex-ante* real interest rate must be below that of the *ex-post* rate. Thus, we focus on those parameters yielding *ex-ante* real interest rates which fulfil that requirement. For those model-generated real rates, their correlation with the *ex-post* rate is then calculated.

Finally, we investigate whether differences between *ex-ante* and *ex-post* real interest rates can be mainly explained in terms of inflation premia or inflation expectation errors. In doing that, we use the variance of the 4-quarter-ahead forecasting error of the corresponding ARMA model for the marginal rate of substitution and the 1-year nominal interest rate to build bounds on inflation expectations (equation [6]). These bounds are then compared to *ex-post* observed 1-year inflation rate. Thus, if the observed inflation rate is outside the bounds we can infer that inflation expectation errors play an important role. On the contrary, inflation premia will be the main explanatory factor if the observed inflation is inside the bounds and these make up a narrow band. However, if the band is too wide and contains the observed inflation inside, nothing can be said about the relative importance of each wedge between *ex-ante* and *ex-post* real interest rates.

3.1. Real Interest Rates

As previously commented, we do not estimate agents' preferences but consider a reasonable range for the parameters involved. According to equations [7], [8] and [9], there are 4 parameters in our analysis. The first

[10] Although data overlapping implies a MA(3) component, a long AR(8) proved to be enough to capture the correlation in the marginal rate of substitution series.

two are common to all utilities functions: β and γ. The remaining two, λ and ϕ, are related to time non-separabilities in preferences.

Regarding β, it is important to note that it acts as a simple scale factor when *ex-ante* real interest rates are obtained. Thus, choosing a value for β does not affect the correlation between *ex-ante* and *ex-post* real rates, although it affects the variance comparisons. Accordingly, we first, choose a value for β of .9588 in annual terms, which corresponds to a value of .9965 in monthly terms. Empirical support for the monthly figure can be found in Canova and Marrinan (1996) or Ayuso (1996). And second, the variance of the *ex-post* real rate is divided by $\beta = .9588$ in order to obtain an upper bound for the variances of the model-generated *ex-ante* real rates. This is equivalent to assuming β is equal to 1 when calculating the variance of the *ex-ante* real interest rate and therefore, to minimising that variance (i.e. maximising the probability of fulfilling the second moment requirement). Thereafter, changing β simply changes the mean level of the model-generated real interest rate, and therefore, we do not undertake any sensitivity analysis, nor comment on estimated real interest rate levels.

As for γ, five values are regarded: .1, .5, 1.5, 5 and 10. These values cover the most commonly considered range in the asset-pricing literature, varying from figures close to risk neutrality to high risk aversion cases. Turning now to λ, its maximum value is bounded so as to guarantee a well-defined utility function. We consider a case of weak habits $(\lambda = .2)$ as well as a case of strong habits $(\lambda = .8)$.[11] Although we use non-durable consumption, we also look at negative values for λ (-.2 and -.8) to keep some symmetry in our analysis. Finally, four ϕ values are considered: .75, .25, -.25 and -.75. As before, we look at negative values (envy) as well as at positive ones (altruism). The maximum absolute value for ϕ has been chosen in such a way that, except for very high risk aversion, own consumption and social consumption have sensible weights in agents' utility function.

(11) Notice that the possibility of negative consumption growth implied that the bound for λ had to be below 1.

Tables 1 to 4 summarise the results of our empirical comparisons in each country. In particular, they show the standard deviations of the different model-generated 1-year *ex-ante* real interest rates, as well as their correlations with the corresponding 1-year *ex-post* real interest rates.

As can be seen, the four tables offer, in general, low correlation degrees.[12] This is particularly true for France and US for which the maximum correlations among those preference specifications satisfying the variance criterium -shaded squares in the Tables- are .199 and .221, respectively. In Spain and UK maximum correlations are higher, but still low: .537 and .621. In all countries but France, the highest correlations are obtained when agents have habits in their preferences. Those habits are stronger for the US $(\lambda = .8)$ than for Spain or UK $(\lambda = .2)$.[13] For France, Abel's preferences offer the best results. In any event, the isoelastic utility functions provide correlations not too distant from those obtained when more general preferences are considered.

The previous view of markedly different patterns of *ex-ante* and *ex-post* real interest rates is reinforced by Figure 2, which shows, for each country, the 1-year *ex-post* real interest rate and the (three) closest model-generated *ex-ante* real rates for each utility function family, i.e. those showing the maximum correlation among those fulfilling the variance restriction.

A clear message seems to emerge from tables 1 to 4 and from Figure 2, . Even allowing for a wide range of preference specification families and of utility function parameters, it is virtually impossible to generate an *ex-ante* real interest rate close enough to the *ex-post* one in any of the four countries considered. Accordingly, and always conditional on our general

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(12) In order to test whether this result is a consequence of the (possibly poor) predictive power of our AR(8) models, we also used *ex-post* observed marginal rates of substitution in equation [3]. The so obtained *ex-ante* real rates exhibited lower correlations for all considered countries and preferences.

(13) Choosing an intermediate value for λ (.4) does not provide higher correlations.

framework choice, it can be said that 1-year *ex-post* real interest rates seem to be a poor proxy for *ex-ante* real rates. It must also be pointed that this result, which is common to the four countries, is obtained despite national consumption data displaying rather different patterns, as Figure 1 shows.

Given that *ex-ante* and *ex-post* real rates behave so differently, it is worth analysing now whether inflation premia or inflation expectation errors are the main explanatory factor behind those differences. This is the aim of next the section.

3.2. Inflation Expectations and Risk Premia

As shown in our theoretical section, the choice of parameters determining the stochastic discount factor (i.e. MRS_{t+1}) crucially affects the joint evolution of those two unknown variables. Thus, the main aim of this section is to shed light on the implications for expectation errors and risk premia of different preference assumptions. This is not an easy task since our approach does not allow us to know which is the right utility function. The strategy we have adopted is the following. First, we look at those utility functions yielding the *ex-ante* real rates which exhibit the highest correlation with the *ex-post* rate, i.e., we consider those cases in which the *ex-post* real rate would be a better proxy for the unobservable *ex-ante* real rate. And second, we analyse how the results change when other preferences providing *ex-ante* rates less volatile than the *ex-post* rate are allowed for.

Thus, Figures 3 to 6 show for each country in our sample bounds on inflation expectations corresponding to that preference specification providing the model-generated *ex-ante* real interest closest to the *ex-post* rate -i.e. that with the highest correlation among those with a variance lower than that of the *ex-post* rate.

Figures 4 and 6 present a similar picture in terms of the band-width of the inflation expectations bounds. In those cases (Spain and the US), we find that inflation expectations bounds are very closely related and move together. This clearly indicates that risk premium terms are quantitatively small and rather stable for the sample periods at hand. In those cases, what mainly seems to determine the differences between *ex-post* and *ex-ante* real interest rates are inflation expectation errors.

The evidence for France and UK is somewhat different. As it is apparent from Figures 3 and 5, the band-width for inflation expectations has increased relative to the evidence presented for the US and Spain. Nevertheless, in both cases the inflation bounds also reflect relatively long periods of under- and overpredictions, less clear-cut for the UK.

The foregoing comments are based on the preference specifications providing the *ex-ante* real rates the closest to the *ex-post* real rate. How robust are the results to changes in the utility functions? The effects of considering alternative preferences can be briefly summarised as follows: the higher the values of λ, $|\phi|$ and, particularly γ, the wider the bands for inflation expectations. Comparing Figure 5 to Figures 4 and 6 gives a good example of how an increase in γ from .1 to 1.5 -higher values tend to produce too volatile *ex-ante* real rates- affect the width of the band. Similarly, Figure 6 in comparison to Figure 4 shows that increasing λ from .2 to .8 has a much lower effect. Finally, increases in $|\phi|$ have been found to produce intermediate effects on the band width.

As to the implications for the analysis of wider expectation bands, as commented at the beginning of this section, we are bounding inflation expectations. We are not estimating them. Thus, it should be clear that the wider the bands, the lower their informational content. Therefore, in cases in which bands are relatively wide, it is not possible to conclude whether inflation premia or expectation error are behind the differences between *ex-post* and *ex-ante* real rates.[14]

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[14] At this point, the main message emerging from Figures 3 to 6 can be reinforced by the available direct estimates of inflation premia. Thus, Alonso and Ayuso (1996) estimated inflation premia below 40 basis points for Spain; Söderlind (1995) found inflation premia below 30 basis points for the American economy and, finally, Levind and Copeland (1993) estimated the sum of the inflation premium and the (always negative) Jensen inequality term to be around -.16 basis points in UK.

Before concluding this section, one additional comment is in order, related to our choice of a single value for β . As noted before, such a choice affects the mean value of the model-generated real interest rates and, thus, the mean level of the inflation expectation bands presented. For that reason, it is important to highlight that, by allowing β to vary, the bands can be moved up or down but without any change in their patterns. By doing this, overprediction (underprediction) episodes would be removed only by enlarging the existing underprediction (overprediction) ones. This is true even for quite wide bands, as in Figure 5.

A reliable β estimate would allow for a deeper analysis of the rationale behind the over- or underprediction episodes. This has been done for the Spanish case in Ayuso and López-Salido (1996). They found a clear overprediction episode following the relatively brisk disinflationary process in the early 80s, which could be justified if agents had doubts about the permanent or transitory nature of the process.

By way of a summary, after having showed significant differences in *ex-ante* and *ex-post* real interest rates in the previous section, the empirical evidence in this section points to inflation expectation errors as the main factor behind such discrepancies in those cases in which the *ex-post* real interest rate could be a better proxy for the unobservable *ex-ante* rates. In these cases, the other potential wedge between both interest rates, i.e. inflation premia, seems to play a minor role. For other preferences for which *ex-post* rates are a worse proxy, high values of $|\phi|$ and, mainly, of γ , render our approach less conclusive on the relative importance of expectation errors and risk premia.

To conclude this section, it is worth noting that the evidence in Figures 3 to 6 is in line with several different recent works that have, on the one hand, detected relatively long periods in which agents fail to correctly forecast inflation and, moreover, offered a theoretical framework in which those expectation errors do not imply irrational agents. Thus, for the UK, King (1996) also shows relatively long inflation overprediction periods since 1982 which are justified in terms of a disinflationary process in which both, agents and the central bank do not immediately learn about the other's reactions to their decisions. For the US economy, the existence of lasting overprediction episodes has been recently stressed from a different perspective by Ceccheti (1996) and Watson (1996). Finally, overpredicting periods during disinflationary processes have also been detected in Canada (Ragan, 1995), New-Zealand (Walsh, 1996) or Denmark (Christensen, 1996), among others.

4. CONCLUSIONS

Non-observable *ex-ante* real interest rates are commonly proxied by the so-called *ex-post* real interest rates, i.e., the difference between the nominal interest rate and the *ex-post* observed inflation rate. *Ex-post* rates, however, can be a misleading proxy if nominal interest rates include a non-negligible inflation risk premium or agents' inflation expectation errors are, *ex-post*, far from the usual zero-mean white-noise case.

In this paper we have shown how the CCAPM provides a simple and integrated framework to jointly estimate the *ex-ante* real interest rate, (bounds on) inflation premia and (bounds on) expectation errors. Thus, this framework allows not only for a comparison of *ex-ante* and *ex-post* real rates, but also provides an explanation of their discrepancies.

Using data on four different economies (France, Spain, UK and US) and considering three different utility function families and a wide range for their corresponding parameters, we have illustrated that 1-year *ex-post* real interest rates are very different from the model-generated *ex-ante* rates. Irrespective of the country and the preference specification, both interest rates display a markedly different pattern, the maximum correlation between them oscillating between .20 (France) and .62 (UK).

For the *ex-ante* real rates showing the highest correlation with the *ex-post* real rate, the discrepancy between both series seems to be mainly explained in terms of agents' inflation expectation errors, the inflation premia playing a minor role. For other preference specifications, high values of γ (and, to a lesser extent, of $|\phi|$), render our approach less conclusive on the relative importance of each wedge between *ex-ante* and *ex-post* real interest rates. Although this is still an open question in the literature, the existence of relatively lasting inflation expectation errors is in line with several different recent works that have detected relatively long periods in which agents fail to correctly forecast inflation and, moreover, offered a theoretical framework in which those expectation errors do not imply irrational agents. These works emphasise the role of information problems and learning processes in the search for an optimal monetary policy.

Table 1. Correlation between *ex-ante* and *ex-post* 1-year real interest rates and their standard deviations: FRANCE 1978.I-1995.III.

PREFERENCES	γ = .1	.5	1.5	5	10
ISOELASTIC	.04	.21	.63	2.34	4.90
	.146	.144	.139	.126	.100
HABITS/DURAB. λ = -.8	.02	.10	.30	1.02	2.15
	(-)	(-)	(-)	(-)	(-)
-.2	.02	.08	.23	.75	1.53
	.069	.070	.071	.074	.078
.2	.07	.034	1.03	3.48	12.8
	.098	.097	.093	.098	.059
.8	2.11	15.0	(*)	(*)	(*)
	.025	(-)	(-)	(-)	(-)
ABEL ϕ = -.75	.27	.23	.37	2.15	4.71
	.195	.158	.132	.154	.115
-.25	.08	.12	.36	2.21	4.85
	.199	.134	.136	.135	.105
.25	.14	.27	.53	2.34	5.02
	(-)	.030	.068	.113	.095
.75	.34	.51	.80	2.53	5.22
	(-)	(-)	.028	.090	.084

Notes: 1. In each square, the top figure is the standard deviation of the *ex-ante* 1-year real interest rate and the bottom figure the correlation between that rate and the *ex-post* one.

2. The standard deviation of *ex-post* 1-year real interest rate is 2.92, which implies an upper bound for that of *ex-ante* rate of 3.05. Shaded squares contain standard deviations lower than the bound.

3. (-) indicates a negative correlation and (*) a st.dev. higher than 100.

Table 2. Correlation between *ex-ante* and *ex-post* 1-year real interest rates and their standard deviations: SPAIN 1979.III-1995.IV.

PREFERENCES	γ = .1	.5	1.5	5	10
ISOELASTIC	.12	.60	1.85	9.26	19.5
	.476	.475	.473	.396	.385
HABITS/DURAB.	.13	.63	1.95	7.00	15.49
λ = -.8	.527	.527	.526	.524	.522
-.2	.12	.62	1.91	6.87	15.16
	.532	.532	.532	.530	.527
.2	.12	.59	1.79	6.42	14.04
	.5372	.5371	.5369	.5366	.5370
.8	.45	2.20	6.16	23.02	286
	(-)	(-)	(-)	(-)	(-)
ABEL	1.05	1.57	2.89	10.5	20.6
ϕ = -.75	.529	.510	.491	.386	.380
-.25	.43	.93	2.21	9.43	19.3
	.519	.494	.484	.395	.386
.25	.17	.28	1.52	8.26	18.1
	(-)	.475	.476	.405	.392
.75	.76	.24	.84	7.13	16.8
	(-)	(-)	.474	.416	.397

Notes: 1. In each square, the top figure is the standard deviation of the *ex-ante* 1-year real interest rate and the bottom figure the correlation between that rate and the *ex-post* one.

2. The standard deviation of *ex-post* 1-year real interest rate is 2.56, which implies an upper bound for that of *ex-ante* rate of 2.67. Shaded squares contain standard deviations lower than the bound.

3. (-) indicates a negative correlation.

Table 3. Correlation between *ex-ante* and *ex-post* 1-year real interest rates and their standard deviations: UK 1973.I-1995.III.

PREFERENCES	γ = .1	.5	1.5	5	10
ISOELASTIC	5.18	5.22	5.37	7.76	9.86
	.464	.478	.508	.541	.491
HABITS/DURAB. λ = -.8	5.26	5.23	5.17	5.16	6.07
	(-)	(-)	(-)	.027	.104
-.2	3.48	3.51	3.64	4.59	6.84
	.544	.555	.574	.570	.481
.2	4.96	4.99	5.17	6.77	10.6
	.608	.617	.621	.512	.318
.8	4.90	21.5	40.3	13.5	47.1
	(-)	(-)	(-)	.127	(-)
ABEL ϕ = -.75	4.97	3.98	3.91	7.89	10.3
	.564	.399	.513	.587	.510
-.25	4.07	3.66	4.62	7.86	10.1
	.497	.461	.575	.595	.516
.25	4.95	4.18	4.99	7.74	9.27
	.509	.459	.564	.598	.521
.75	5.33	3.94	4.25	7.52	9.66
	.533	.403	.483	.595	.524

Notes: 1. In each square, the top figure is the standard deviation of the *ex-ante* 1-year real interest rate and the bottom figure the correlation between that rate and the *ex-post* one.
2. The standard deviation of *ex-post* 1-year real interest rate is 5.82, which implies an upper bound for that of *ex-ante* rate of 6.07. Shaded squares contain standard deviations lower than the bound.
3. (-) indicates a negative correlation.

Table 4. Correlation between *ex-ante* and *ex-post* 1-year real interest rates and their standard deviations: US 1973.I-1995.III.

PREFERENCES	γ = .1	.5	1.5	5	10
ISOELASTIC	5.05	5.09	5.27	12.2	18.2
	.332	.342	.361	.369	.329
HABITS/DURAB.	5.82	5.88	6.12	7.85	12.3
λ = -.8	.229	.235	.248	.259	.230
-.2	5.89	5.93	6.12	7.64	11.8
	.255	.261	.271	.275	.237
.2	6.06	6.12	6.33	7.69	11.4
	.329	.332	.335	.320	.259
.8	2.44	4.87	(*)	(*)	(*)
	.221	(-)	.047	.268	(-)
ABEL	5.05	4.86	5.46	11.4	17.5
ϕ = -.75	.326	.280	.311	.348	.297
-.25	4.80	5.11	5.90	11.0	16.8
	.278	.292	.321	.363	.308
.25	6.29	6.18	6.49	10.6	16.1
	.374	.377	.348	.377	.320
.75	5.41	7.20	6.26	10.3	15.4
	.344	.373	.385	.390	.331

Notes: 1. In each square, the top figure is the standard deviation of the *ex-ante* 1-year real interest rate and the bottom figure the correlation between that rate and the *ex-post* one.
2. The standard deviation of *ex-post* 1-year real interest rate is 3.70, which implies an upper bound for that of *ex-ante* rate of 3.86. Shaded squares contain standard deviations lower than the bound.
3. (-) indicates a negative correlation and (*) a st.dev. higher than 100.

FIGURE 1. Annual consumption growth (%)

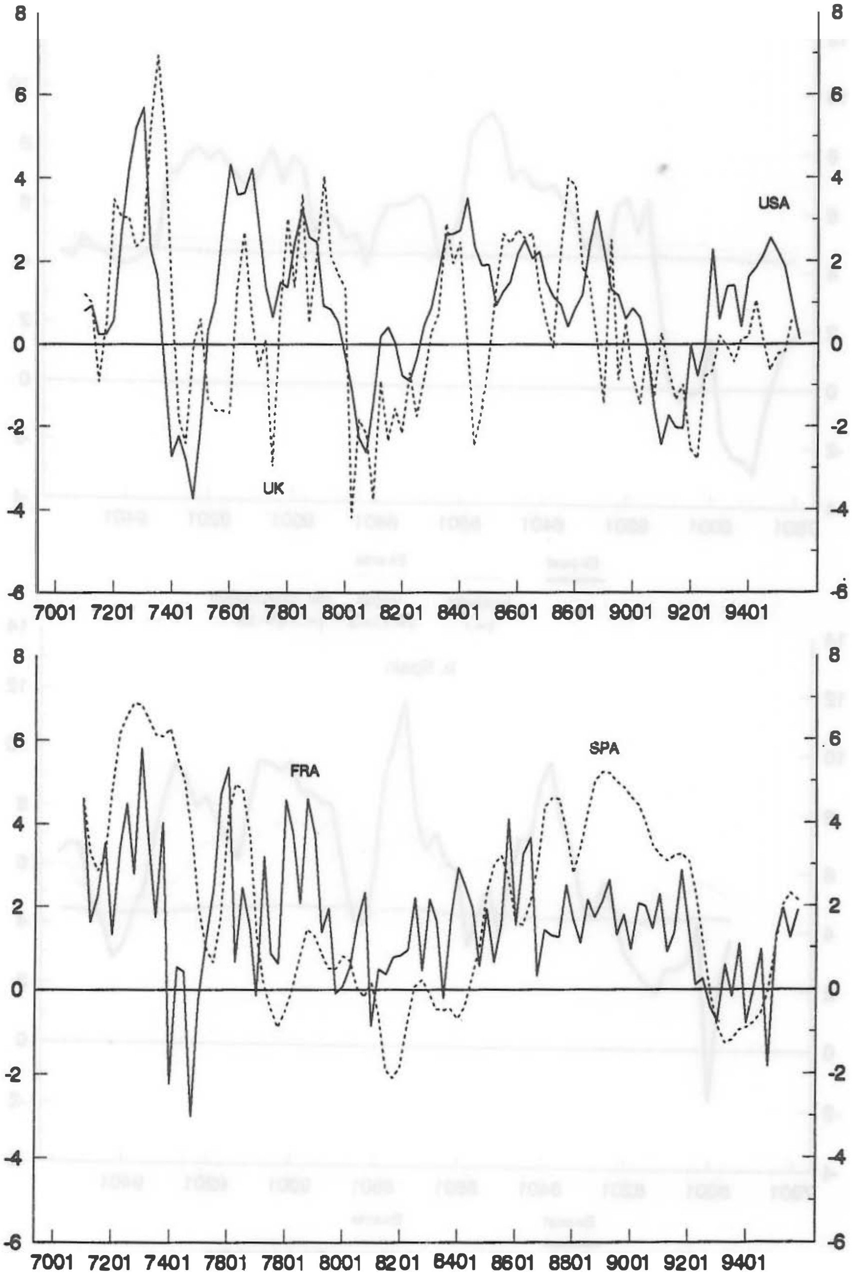
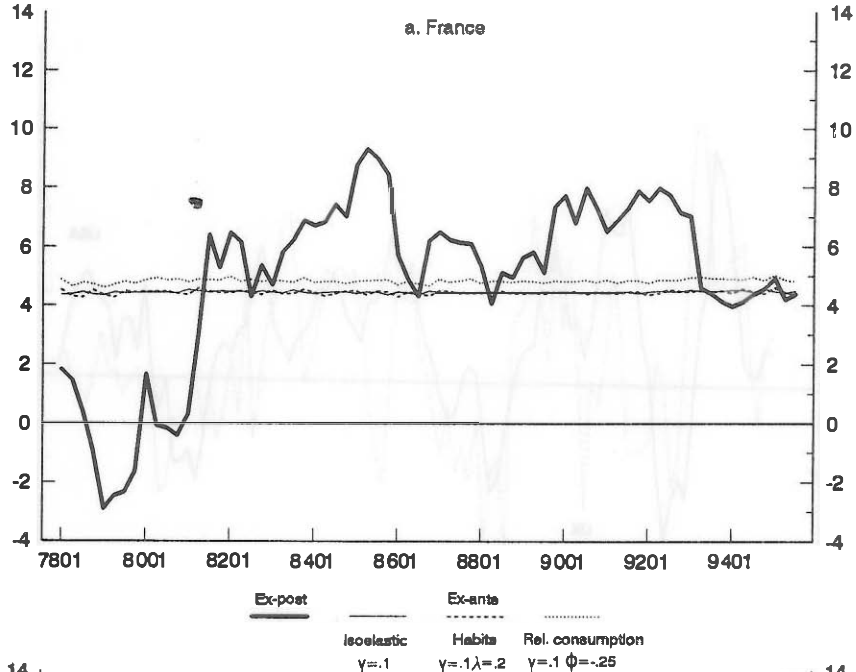


Figure 2. 1-year real interest rates (%)



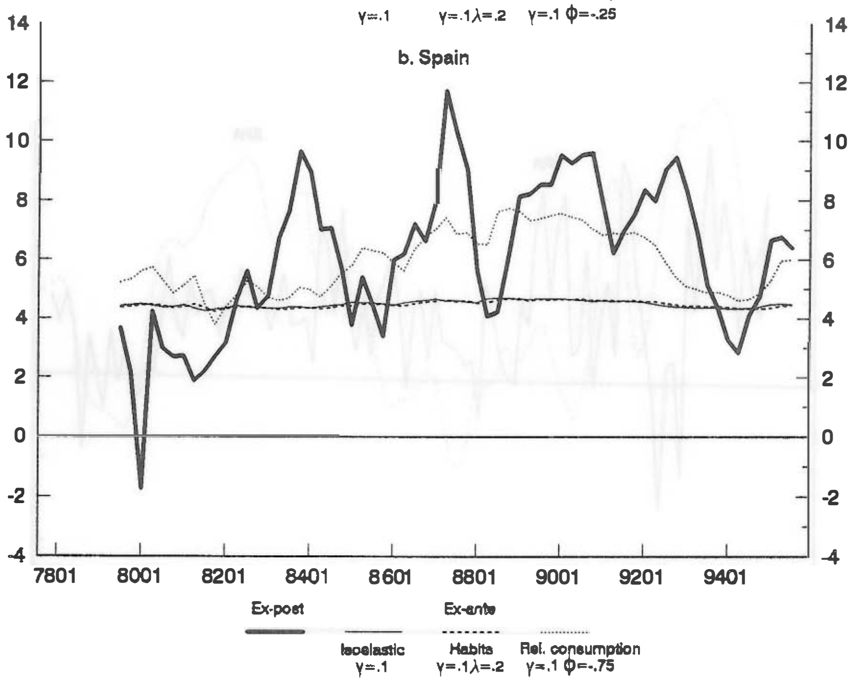
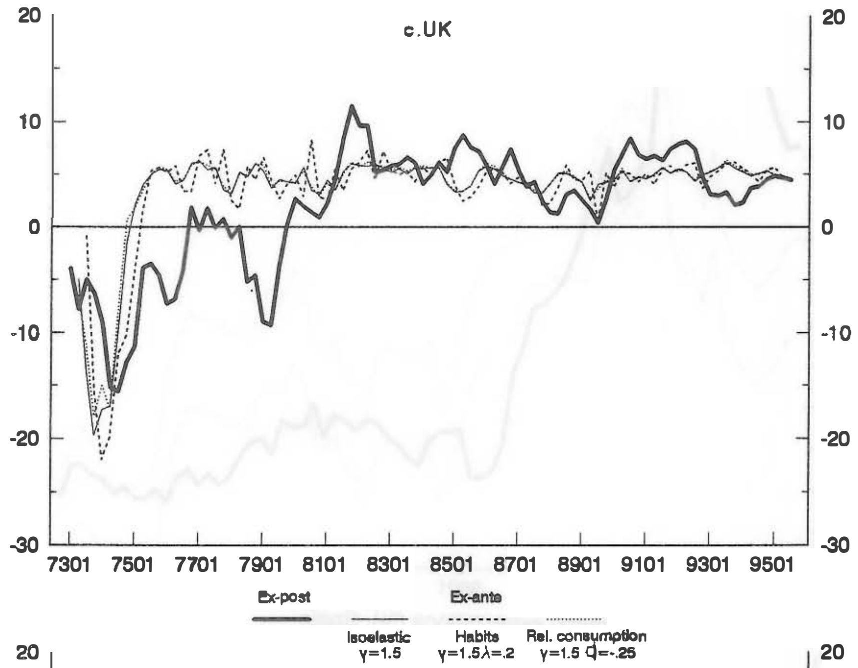


Figure 2 (cont.). 1-year real interest rates (%)



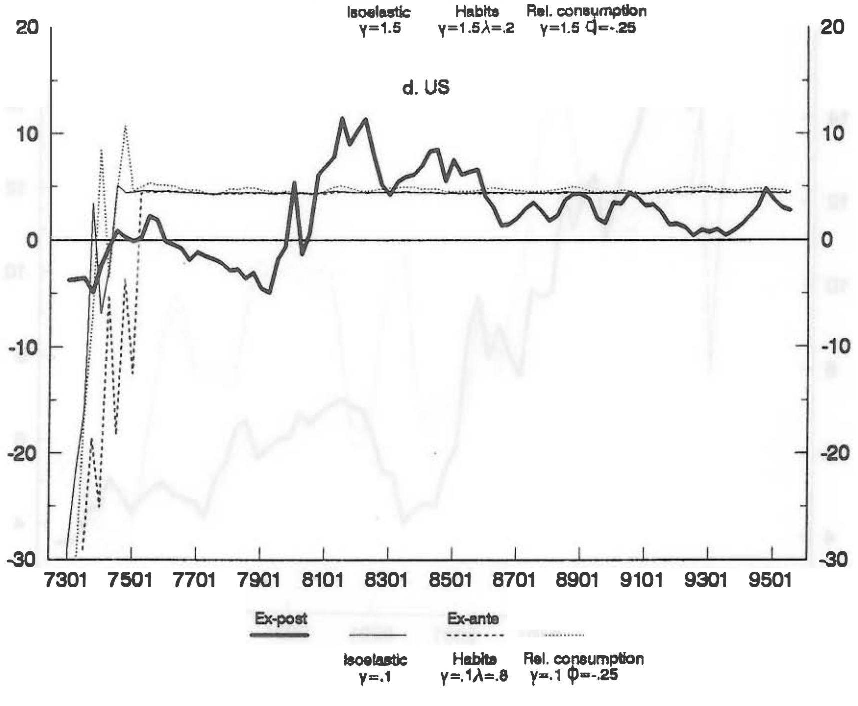


Figure 3. 1-year-ahead inflation expectations. France

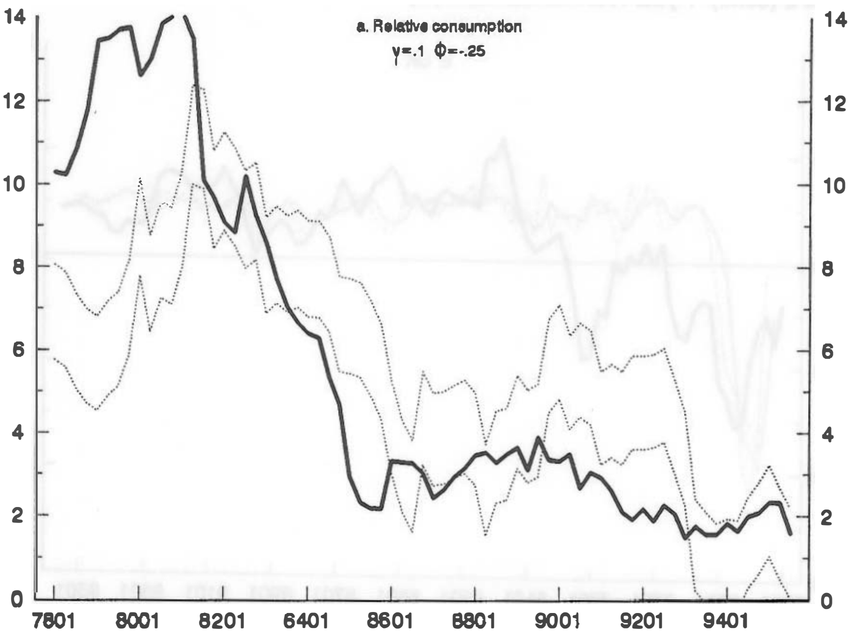


Figure 4. 1-year-ahead inflation expectations (%). Spain

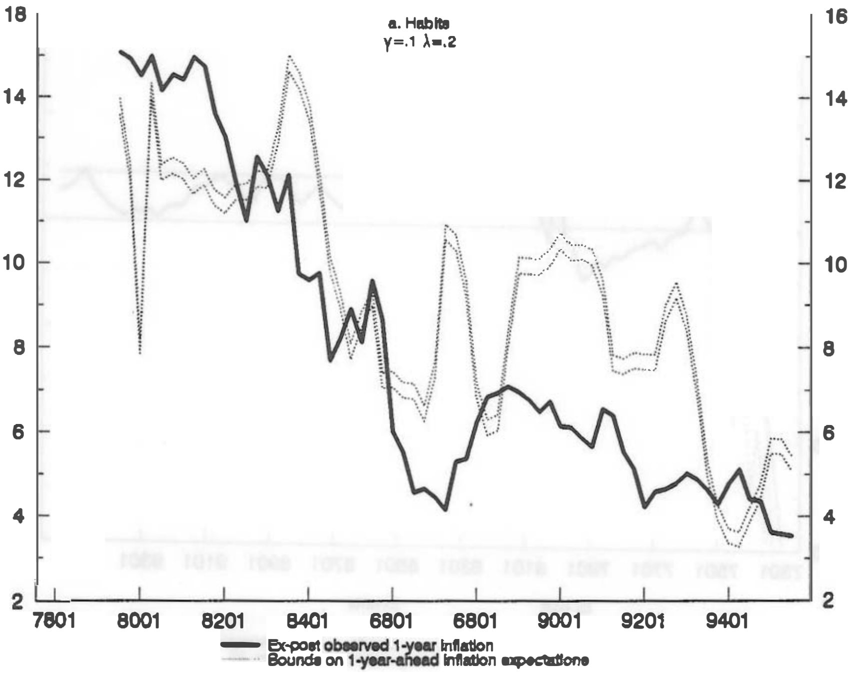


Figure 5. 1-year-ahead inflation expectations. UK

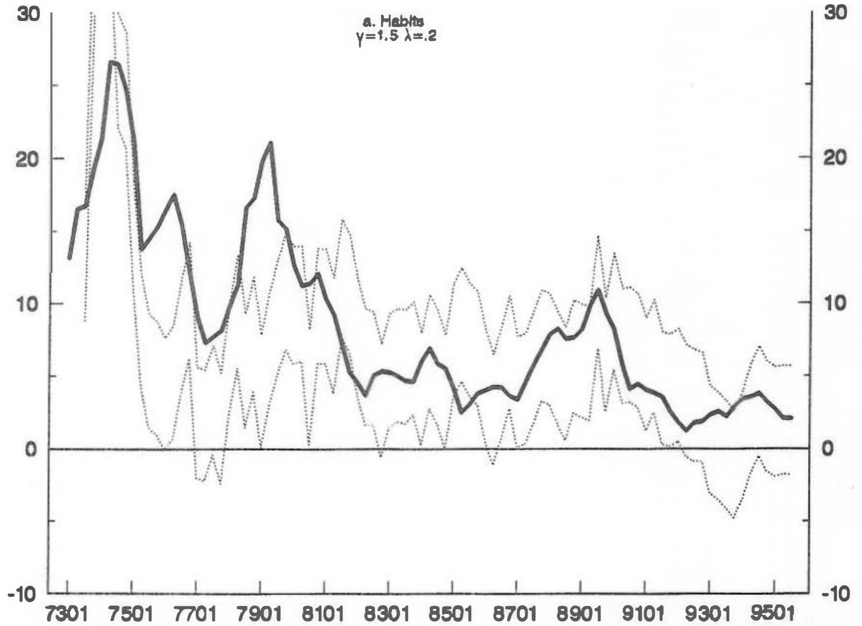
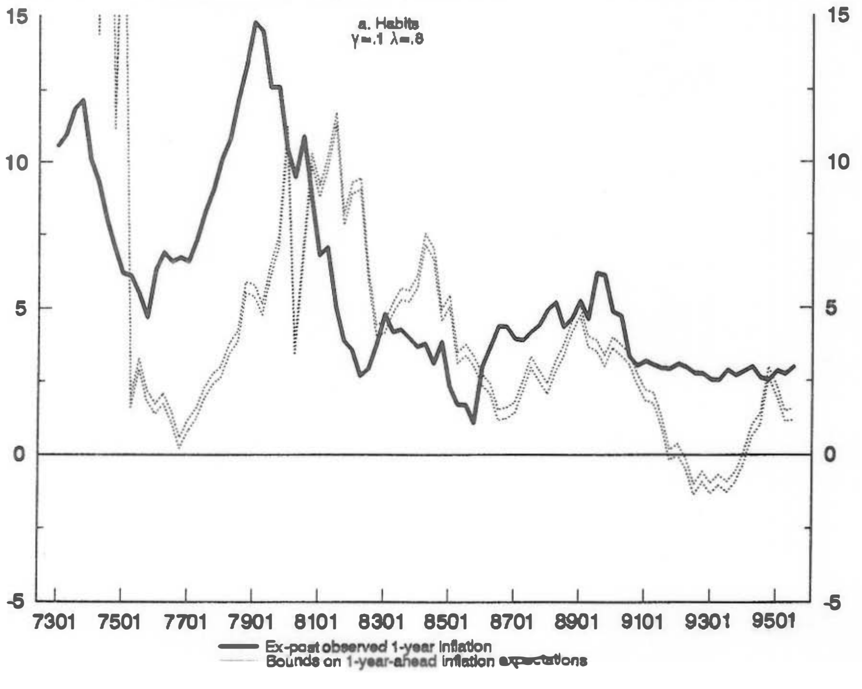


Figure 6. 1-year-ahead inflation expectations. USA



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