Credit, liquidity and market risk in the term structure of swaps in pesetas*

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Abstract: Some characteristics of the term structure in interest rate swap (*IRS*) markets are influenced by the idiosyncrasy of this financial instrument itself, which might explain the rejection of the Expectations Hypothesis in the formation of interest rates. After testing and rejecting the Expectations Hypothesis, we present evidence supporting the existence of significant time-varying *premia*. We then focus on characterizing some properties of realized *ex-post premia*, and provide explanatory variables for these *premia*. We pay particular attention to the extent to which the levels of *market risk, default risk* and *liquidity risk* explain the time evolution of *risk premia* at different maturities.

Keywords: Term structure, interest rate swaps, expectations theory, forward rate, premium.

1. INTRODUCTION

Investors in financial markets use the term structure of interest rates (*TSIR)* to estimate a correct price for fixed income assets, as well as to design their investment and hedging strategies. The *TSIR* in fixed income markets can also be used to obtain information on market consensus on the future evolution of interest rates. The huge increase in liquidity in *interest rate swap* (*IRS*) markets, the heterogeneity in public debt issuing among *EMU* countries, and the fact that *IRS* can be homogeneously traded across Europe, have converted the *IRS* structure into the reference curve for capital markets in the *EMU*.

Characterizing the main properties of the *TSIR* for the *IRS* market is therefore central for risk management in fixed income portfolios. In particular, the market for *IRS* in pesetas presents some specific characteristics that make it somewhat different from the analysis of the *TSIR* in other fixed income markets. Following a standard practice in fixed income markets, we use estimates of the relationship between *forward* rates implicit in the current *TSIR* and future spot rates to test the *Expectations Hypothesis* (*EH*) in the formation of interest rates. Given the overwhelming evidence in favour of the non-stationary nature of spot and forward rates, we explore the possibility

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that current forward and future spot rates are cointegrated, with the coefficients imposed by the *EH*. Since we show empirical evidence clearly rejecting both the weak and the strong versions of the *EH* as a representation of the *TSIR* in the *swap* market in pesetas, we explore the possibility that term- or risk-premia may explain the observed deviations from the *EH*.

After providing evidence on the existence of premia, the paper focuses on characterizing their behaviour over time as well as on finding some explanatory factor for them. Relative to the latter question, there is some consensus in fixed income markets that term-premia may arise due to interest rate risk. Nevertheless, since *IRS* are exposed to different types of risk (interest rate or market risk, credit risk and liquidity risk), we use proxies for them in an attempt to evaluate their relative importance. We approximate market risk by a measure of interest rate volatility, while credit and liquidity risk are jointly approximated by the spread between zero coupon rates from the secondary market for Spanish public debt and the market for *IRS* in pesetas. We find statistically significant evidence that both indicators contain explanatory ability for realized *ex-post premia*.

In Section 2 we briefly review the Expectations Theory on the formation of the term structure of interest rates as well as that of premia, along with the main results in the empirical literature. Section 3 contains a description of the data. The *EH* is tested in Section 4, while Section 5 contains evidence on the existence of ex-post premia together with an analysis of their main characteristics. In Section 6 we analyse the role of the level of risk as an explanatory factor of realized *risk-premia*. The paper closes with some conclusions.

2. THE EXPECTATIONS HYPOTHESIS AND RISK-PREMIA

Several alternative explanations on the relationship between interest rates across the term structure have been advanced in the financial literature. According to the *EH*, the shape of the *TSIR* at each point in time is the result of an equilibrium in which, given current expectations of future interest rates, the investor is indifferent between short- and long-term positions. In such a case, *forward-premia is zero*. As defined by Hicks (1946), a *term-premium* is the difference between the returns of two investment strategies with the same maturity. Specifically, the time t *term-premium* ($P_{t,n,m}$) compares the strategy consisting in investing at time $t+n$ over m periods, whose return $r_{t+n,m}$ is unknown as of time t, with the *forward* rate determined at time t for an investment that will take place at time $t+n$ over m periods, with $m<n$, ($f_{t,t+n,m}$):

$$P_{t,n,m} = m\left[f_{t,t+n,m} - E\left(r_{t+n,m}\right)\right] \tag{1}$$

where E_t denotes the conditional expectation operator, based on the information available to market participants at time t.

The weak form of the *EH* allows for the returns on alternative investment strategies to differ by a constant, which may depend on the investment horizon, but not on time. Rewriting the definition of the *forward* premium under the assumption that agents form their expectations rationally[1]:

1. Under rational expectations: $E_t(r_{t+i,n}) = r_{t+i,n} - u_{t+i}$ where u_{t+i} is the forecast error, which is unpredictable from the information available at time t.

$$r_{t+n,m} = -\frac{P_{t,n,m}}{m} + f_{t,t+n,m} + u_{t+n} \quad (2)$$

which, under the assumption of a constant premium, suggests estimating the model:

$$r_{t+n,m} = a + bf_{t,t+n,m} + u_{t+n} \quad (3)$$

The strong version of the Expectations Theory implies: a=0, b=1 and u_{t+n} uncorrelated with any variable known[2] as of time t. It is clear that u_{t+n} must satisfy the described lack of correlation, since otherwise there would be some relevant information on the future evolution of spot rates, available at time t and not incorporated in *forward* rates. The test of the joint hypothesis above is known as testing for the *forward as an unbiased predictor of future spot rates*.

Rejection of the *EH* under the assumption of rational expectations is usually taken as evidence on the existence of time-varying premia. Hence, characterizing the determinants of the sign and level of premia becomes a crucial issue for interest rate forecasting and risk management. Seminal work on characterizing the sign of term-premia under rational expectations in fixed income markets is that of Fama (1976, 1984a, 1984b). Fama finds positive premia, increasing with maturity, in line with the findings reported in McCulloch (1987). But these results do not seem very robust over time: Fama and Bliss (1987) find that premia for maturities between 1 and 5 years change sign relatively often. Working with data between 1964 and 1988, Evans and Lewis (1994) show premia at the longer maturities in Treasury bills to be non-stationary.

Pioneer work on the determinants of premia in fixed income markets was that of Kessel (1965), who works under the assumption that the relationship between premia and their determinants is linear. Empirical results in this line of research have been rather controversial: using USA data, Kessel (1965) and Nelson (1976) use regression methods to show that observed spot rates are a determinant of premia, but with coefficients of the opposite sign to those imposed by the Expectations Theory. Shiller (1979) runs a similar regression with USA and UK data for longer maturities, and interprets the resulting coefficients as an indication of excess volatility in interest rates. In a similar regression with maturities around 5 years, Campbell and Shiller (1987) find a negative coefficient for interest rates, which they interpret as an insufficient reaction of longer-term interest rates to fluctuations in shorter-term rates.

On the other hand, there does seem to be consensus in the literature as regards interest rate volatility being the main determinant of premia. Fama (1976) presents evidence consistent with such a view. Modigliani and Shiller (1973), as well as Shiller, Campbell and Schoenholtz (1983) obtain similar results using interest rate standard deviations computed on rolling-windows. More recently, Engle, Lilien and Robins (1987), as well as Bollerslev, Engle and Wooldridge (1988), using ARCH in the mean models and multivariate GARCH in the mean models to represent interest rate volatility, reach the same conclusion as the previously cited authors.

2. u_{t+n} is the error from predicting r_{t+n},m at time t and, therefore, it will have an MA(n-1) stochastic structure.

3. THE DATA

We used data from two markets. To test the *EH* and study ex-post premia in the market for swaps in pesetas, we used the *TSIR* of *IRS* denominated in pesetas. To quantify the level of credit and liquidity risk involved in *IRS*, we used the *TSIR* from the secondary market for Spanish public debt[3]. The *TSIR* for the *IRS* market was estimated via the recursive method from quoted rates for the fixed interest branch of a generic *IRS* of 2-, 3-, 4-,..., 9-, and 10-year maturity. Quoted rates were obtained from *Datastream™*, which collects them at 18:00 hours GTM. They are the average of *bid* and *ask* rates, as provided by *Dark Limited*, from *Intercapital Brokers Limited*. The *TSIR* is made up of nine zero coupon rates, observed daily from January 4, 1991 to December 31, 1998. There are a large number of implicit *forward* rates in the *IRS* term structure, but since our objective is to evaluate and explain observed premia, we only consider those maturities corresponding to estimated zero coupon rates[4]. We consequently considered *forward* rates as of time t for an investment starting at t+2 and lasting m periods, $f_{t,t+2,m}$, with m: 2, 3, 4, 5, 6, 7 and 8 years[5, 6].

The *TSIR* for the secondary market for Spanish public debt was obtained from a zero coupon interest rate curve as proposed by Nelson and Siegel (1987). Daily estimates of the curve were obtained from closing *bid* and *ask* prices for the more liquid references in the market. These estimates cover the June 1, 1993 to December 31, 1996 period.

4. TESTING THE EXPECTATIONS HYPOTHESIS IN THE MARKET FOR *SWAPS*

Tests of the Expectations Hypothesis must take into account the fact that spot and *forward* zero coupon rates in the term structure of swaps are all nonstationary (see Table 1), and hence (3) must be considered as a cointegration relationship between a spot rate and the associated, appropriately lagged forward rate. Hence, under the *EH*, (3) is a long-run equilibrium relationship, with the cointegration vector (1,-1). Estimation and hypothesis testing on this vector can be implemented through either the two-step least squares procedure proposed by Engle and Granger (1987) or the maximum likelihood method developed by Johansen (1988, 1991).

Table 2 contains the results from testing the *EH* by both methods. The first column presents the estimation of (3) by least squares with standard deviations robust to the presence of heteroscedasticity and autocorrelation, as suggested by Newey and West (1987). Augmented Dickey-Fuller (*ADF*) statistics on the residuals show that residuals in the estimated models are not stationary. Hence, according to this procedure, we do not detect an equilibrium long-run relationship between *forward* rates and future spot rates, contrary to the *EH*. In maximum-likelihood estimation (right panel in Table 2), the maximum eigenvalue and trace statistics reject, at the 90% confidence level and for all maturities, the hypothesis that *forward* and future spot rates are cointegrated.

3. All of these are continuously compounded interest rates.

4. This is done to avoid possible distortions that could arise when computing *forward* rates from interpolated spot rates.

5. The *forward* rate at time t for an investment at $t+n$ lasting m periods, $f_{t,t+n,m}$, is computed from market rates observed at time t: $m f_{t,t+n,m} = (n+m) r_{t,n+m} - n r_{t,n}$.

6. Two years are lost at the end of the sample when computing forward rates.

Table 1
Unit root tests on spot and *forward* interest rates

	Spot rates				*Forward* rates			
	Level		First difference		Level		First difference	
	ADF	PP	ADF	PP	ADF	PP	ADF	PP
2 year	-0.588	-0.652	-21.353*	-45.472*	-0.906	-0.690	-17.382*	-42.589*
3 year	-0.522	-0.546	-20.616*	-42.661*	-0.924	-0.881	-17.429*	-43.924*
4 year	-0.563	-0.449	-20.109*	-42.146*	-0.790	-0.852	-17.437*	-41.240*
5 year	-0.502	-0.446	-19.702*	-42.317*	-0.751	-0.962	-18.008*	-41.252*
6 year	-0.391	-0.361	-19.538*	-42.229*	-0.711	-0.901	-17.207*	-40.825*
7 year	-0.261	-0.302	-19.757*	-43.770*	-0.710	-0.907	-16.736*	-40.972*
8 year	-0.161	-0.218	-19.543*	-43.571*	-0.739	-0.976	-16.607*	-41.564*

Note: Sample period: 1/4/1991 to 12/31/1998. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) statistics in levels and first differences of spot and *forward* rates obtained from the term structure for *IRS* include a constant term, but no trend, and 4 lags of the dependent 4. Critical values at 90% confidence: ADF = -2.568, PP = -2.568. An asterisk denotes rejection of the corresponding null hypothesis at 90% confidence level.

Table 2
Long-run tests of Expectations Hypothesis: $r_{t+2,m}=a+bf_{t,t+2,m}+u_t$

	Engle-Granger tests					Reduced rank tests		
m	a	b	R^2	ADF u_t	PP u_t	Hypothesis	λ_{MAX}	λ_T
2 year	1.364	0.610	0.153	-1.378	-1.545	r ≤0	9.510	11.000
	(1.220)	(5.714)				r≤1	1.490	1.490
3 year	1.596	0.599	0.131	-1.184	-1.239	r≤0	8.020	9.710
	(1.310)	(5.153)				r≤1	1.690	1.690
4 year	1.423	0.628	0.127	-1.116	-1.177	r≤0	7.710	9.160
	(1.091)	(5.052)				r≤1	1.460	1.460
5 year	1.533	0.626	0.114	-1.067	-1.077	r≤0	7.420	8.910
	(1.099)	(4.717)				r≤1	1.490	1.490
6 year	1.415	0.648	0.117	-1.058	-1.029	r≤0	7.450	8.940
	(0.991)	(4.778)				r≤1	1.500	1.500
7 year	1.433	0.658	0.115	-1.048	-0.986	r≤0	7.580	9.080
	(0.973)	(4.710)				r≤1	1.500	1.500
8 year	1.545	0.652	0.107	-1.035	-0.951	r≤0	7.880	9.400
	(1.012)	(4.512)				r≤1	1.520	1.520

Note: Sample period: 1/4/1991 to 12/31/1996. Two-step least squares estimates of the cointegrating relationship [Engle and Granger (1987)], with robust standard deviations [Newey-West (1987)]. *t*-statistics in parentheses. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests on the residuals include a constant term, but no trend. The number of lags included was 4 in all cases. Critical values for both statistics at 10% significance are -2.568 and -2.568, respectively. Maximum eigenvalue (λ_{MAX}) and trace (λ_T) statistics are defined in Johansen (1988). Critical values at 10% significance for r=0 are 10.29 and 17.79, while for r=1 they are 7.50 and 7.50, respectively. The number of lags used in the VAR model in first differences was10. No constant or trend was included in this model. An asterisk denotes rejection of the null hypothesis at the 90% confidence level.

This evidence therefore overwhelmingly suggests that there is no equilibrium relationship between forward and future spot rates in the swap market in pesetas, between January 1993 and December 1996, thus contradicting the Expectations Hypothesis. As already indicated, this may be provoked by the presence of time-varying premia in this market. This is the question we analyze in the next section.

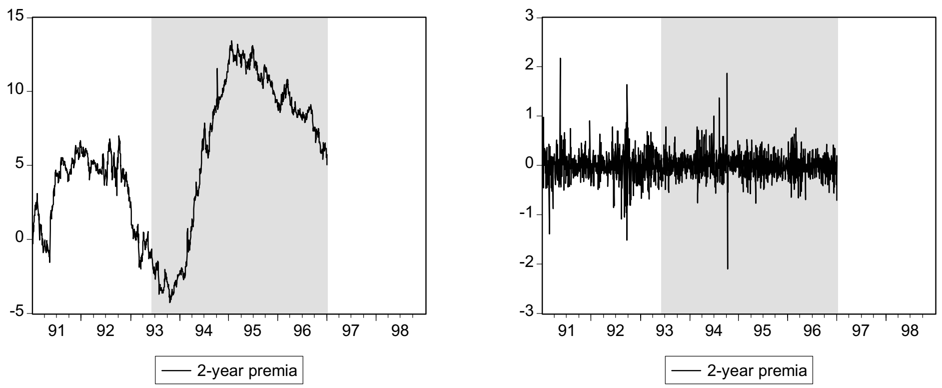
5. COMPUTING E*X-POST* PREMIA IN THE MARKET FOR *SWAPS* IN PESETAS

To examine the possible existence of premia in each of the maturities, we substitute $r_{t+n,m}$ for $E_t(r_{t+n,m})$ in the definition of premium [Equation (1)]. The resulting premia are usually known as ex-post premia. We computed these for the period between January 1991 and December 1996.

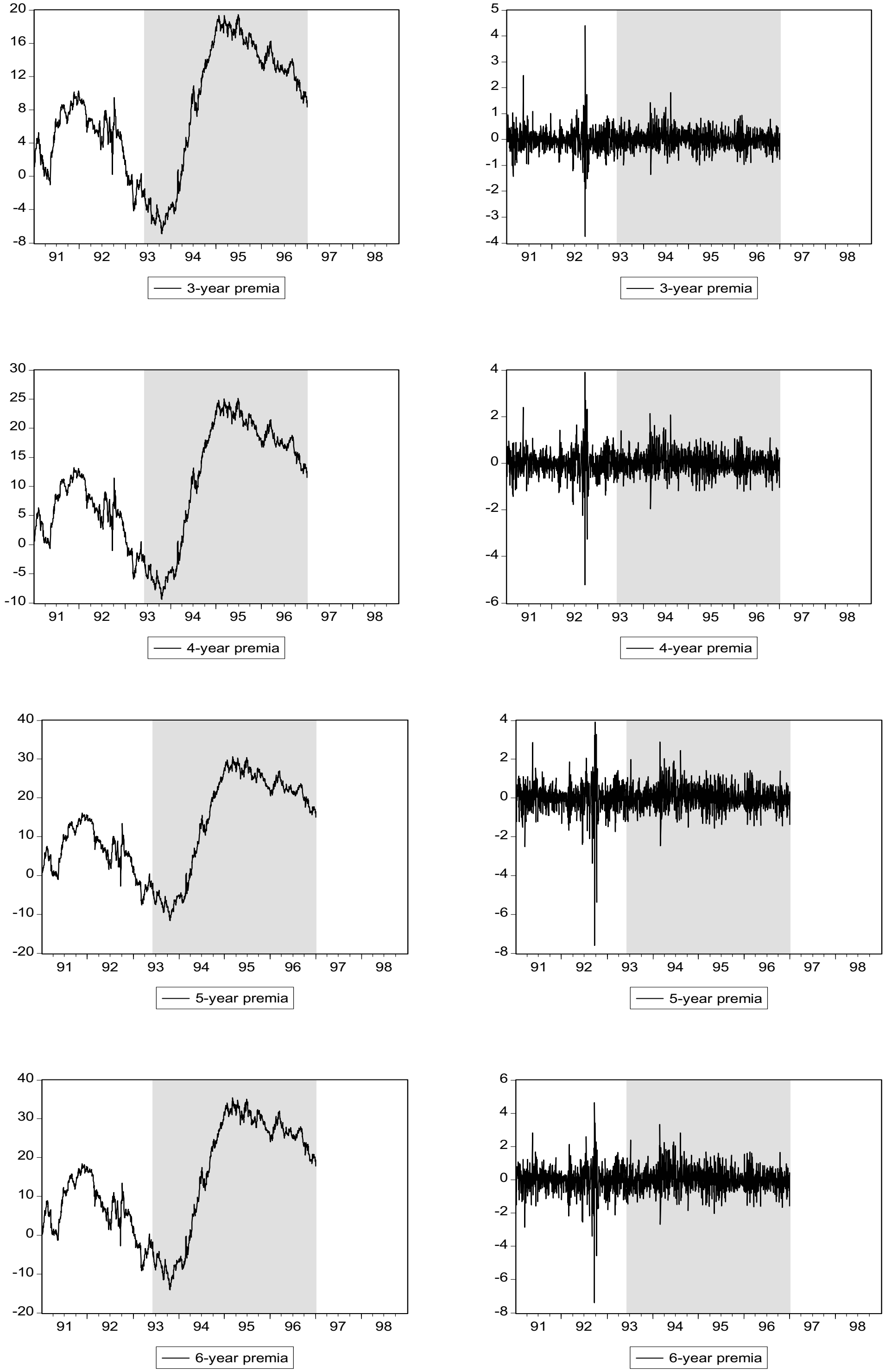
5.1. Descriptive analysis of ex-post premia

The dynamic behavior of *ex-post* premia is shown in Figure 1. Stylized facts are: *i*) a clearly non-stationary dynamic behavior in premia[7], *ii*) premia are positive over the time period considered, except for the March1993 to March 1994 interval, and *iii*) premia are increasing up to January 1995, decreasing from then onwards, and stabilizing towards the end of the observation period. This is a consequence of the implementation of monetary policy in Spain, as pointed out by Gómez and Novales (1997). These authors show that in June 1994 there was a drastic change in the shape of the term structure in the Spanish market for public debt, which went from being increasing to showing a decreasing shape in all maturities. At the end of 1995, at the pinnacle of the process of monetary easing, the term structure adopted again a decreasing shape at the shorter maturities.

Figure 1
Ex-post premia and first differences.
Sample period: 1/4/1991 to 12/31/199



7. Evans and Lewis (1994) also observe non-stationary premium in fixed income markets.


3-year premia
3-year premia
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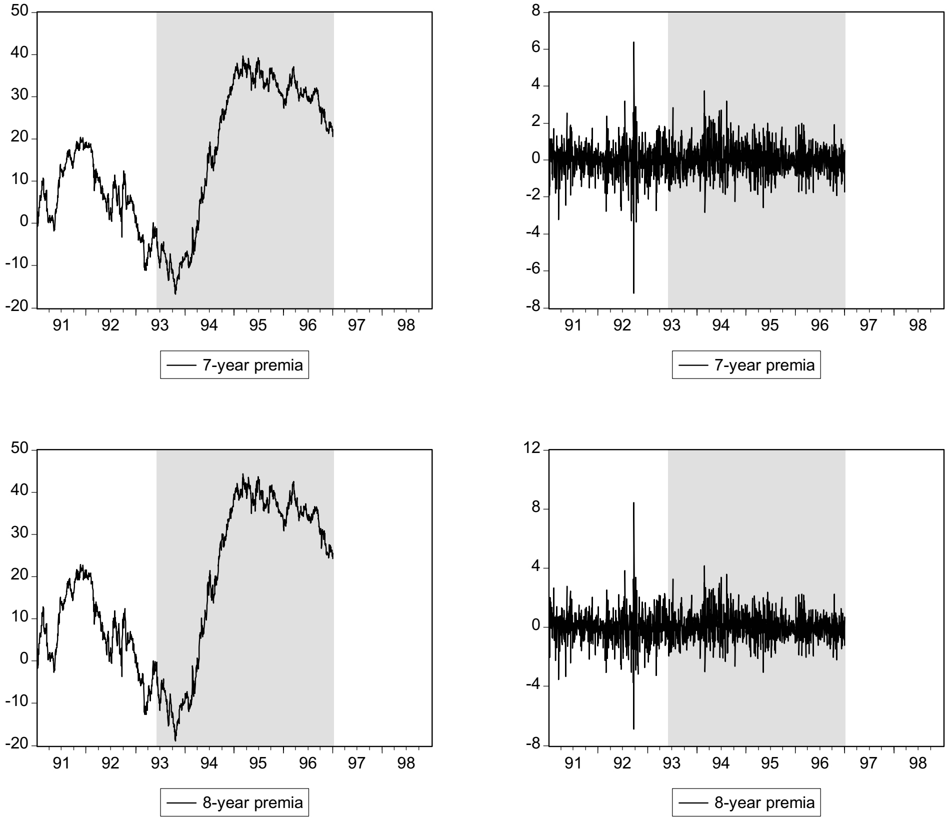
6-year premia
6-year premia



The fact that *ex-post* premia are not stationary is ratified by unit root tests in Table 3. Furthermore, Table 3 also shows a number of descriptive statistics for premia at each maturity[8]. Average premia are positive, significantly different from zero, and increasing with maturity, in consonance with the intuition that uncertainty increases with the horizon of a given investment. Daily changes in premia, in contrast, are not different from zero for any maturity. In both cases, dispersion increases with maturity.

Since unit root tests suggest that premia follow integrated processes of order one, we formulated dynamic models in first differences of *ex-post premia*. To detect autoregressive and moving average structures, we used the Box-Jenkins methodology. Least-squares estimation results are shown in Table 4, where we have used standard deviations robust to possibly heteroscedastic and autocorrelated residuals. These results indicate that daily changes in *ex-post* premia follow autoregressive structures of up to order 9.

8. Full interpretation of these statistics would only been justified under the assumptions of stationarity and lack of serial correlation.

Table 3

Unit root tests and descriptive statistics for *ex-post* premia

Level	**2 year**	**3 year**	**4 year**	**5 year**	**6 year**	**7 year**	**8 year**
ADF	-1.312	-1.208	-1.176	-1.170	-1.179	-1.198	-1.224
PP	-1.423	-1.254	-1.227	-1.198	-1.164	-1.149	-1.152
Average	5.195	7455	9.507	11.489	13.155	14.583	16.265
Maximum	13.443	19.461	25.116	30.662	35.373	39.680	44.337
Minimum	-4273	-6.863	-9.415	-11.626	-14.120	-16.745	-18.884
Standard Deviation	4.729	7.148	9.405	11.613	13.655	15.676	17.731
Skewness	-0.231	-0.202	-0.167	-0.139	-0.128	-0.112	-0.089
Kurtosis	2.011	1.946	1.859	1.778	1.748	1.713	1.675
Observations	1564	1564	1564	1564	1564	1564	1564
First difference	**2 year**	**3 year**	**4 year**	**5 year**	**6 year**	**7 year**	**8 year**
ADF	-17.506*	-17.983*	-17.834*	-18.232*	-17.564*	-17.161*	-16.990*
PP	-39.842*	-41.764*	-39.363*	-39.792*	-39.831*	-40.447*	-40.666*
Average	0.003	0.005	0.007	0.008	0.010	0.011	0.014
Maximum	2.178	4.401	3.918	3.905	4.648	6.378	8.433
Minimum	-2.106	-3.735	-5.207	-7.603	-7.396	-7.193	-6.895
Standard Deviation	0.274	0.406	0.517	0.664	0.742	0.851	0.979
Skewness	0.254	0.547	-0.293	-1.057	-0.544	-0.088	0.249
Kurtosis	11.657	18.458	14.312	18.870	12.254	9.276	8.994
Observations	1563	1563	1563	1563	1563	1563	1563

Sample period: 1/4/1991 to 12/31/1996. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests include a constant term, but no trend. The number of included lags is 4 in all cases. Critical values at 90% confidence: ADF = -2.568, PP = -2.568. An asterisk denotes rejection of the null hypothesis at the 90% confidence level.

Table 4

Unit root tests and descriptive statistics for *ex-post* premia

$$\Lambda P_t^m = a + b_5\,\Lambda P_{t-5}^m + b_6\,\Lambda P_{t-6}^m + b_9\,\Lambda P_{t-9}^m + u_t$$

m	a	b_5	b_6	b_9	R^2	ADF u_t	Q(10)	Q(15)
2 year	0.007	--	-0.068*	0.045*	0.007	-13.944*	4.608	7.316
	(0.899)		(-2.407)	(1.593)			[0.92]	[0.95]
3 year	0.012	--	-0.092*	0.061*	0.012	-13.581*	6.124	13.643
	(1.088)		(-2.819)	(1.881)			[0.81]	[0.55]
4 year	0.017	--	-0.077*	--	0.006	-13.469*	10.387	16.398
	(1.188)		(-2.341)				[0.41]	[0.36]
5 year	0.021	--	-0.060*	--	0.004	-13.628*	11.168	15.100
	(1.191)		(-1.776)				[0.35]	[0.44]
6 year	0.023	0.053*	-0.064*	--	0.007	-14.088*	7.575	11.244
	(1.108)	(1.568)	(-1.907)				[0.67]	[0.74]
7 year	0.026	0.062*	-0.066*	--	0.008	-14.014*	6.668	10.004
	(1.045)	(1.859)	(-1.969)				[0.76]	[0.82]
8 year	0.029	0.064*	-0.065*	--	0.008	-14.004*	6.177	9.706
	(1.013)	(1.921)	(-1.924)				[0.80]	[0.84]

Sample period: 6/1/1993 to 12/31/1996. Least squares estimation, with robust standard deviations, as in Newey-West (1987). *t-statistic* in parentheses. ΛP_t^m is the realized *ex-post* premia at maturity *m* in first differences. Augmented Dickey-Fuller (ADF) tests on the residuals include a constant term, but no trend. Four lags of the differenced residuals were included in all cases. Critical value at 10%significance level is -2.568. An asterisk denotes rejection of the null hypothesis at the 90% confidence level. Q(10) and Q(15) are Ljung-Box statistics for residual autocorrelation. *p*-value in square brackets.

6. IDENTIFYING FACTORS AFFECTING *EX-POST* PREMIA

Ex-post premia are positive for most of the time period considered, and increasing with maturity, which is consistent with investors having a preference for the short-term. Consequently, long-term interest rates are the sum of expectations of future short-term rates plus a premium that compensates for risk, since long-term rates involve greater uncertainty. This is because *IRS* are subject to diverse types of risk: *a) market or interest risk*, because of the uncertainty on future fluctuations in interest rates, *b) credit or default risk*, due to the possibility that one of the counterparts in the swap agreement will not fulfill his obligation, and *c) liquidity risk*, due to the difficulty in closing down the position in an *IRS* agreement.

We hence considered risk as a possible determinant of observed *ex-post* term-premia, following Kessel (1965), who represents premia as linear functions of potential explanatory variables. We thus included two variables in the models specified in previous sections in order to capture the risk involved in an *IRS* contract, which we define next.

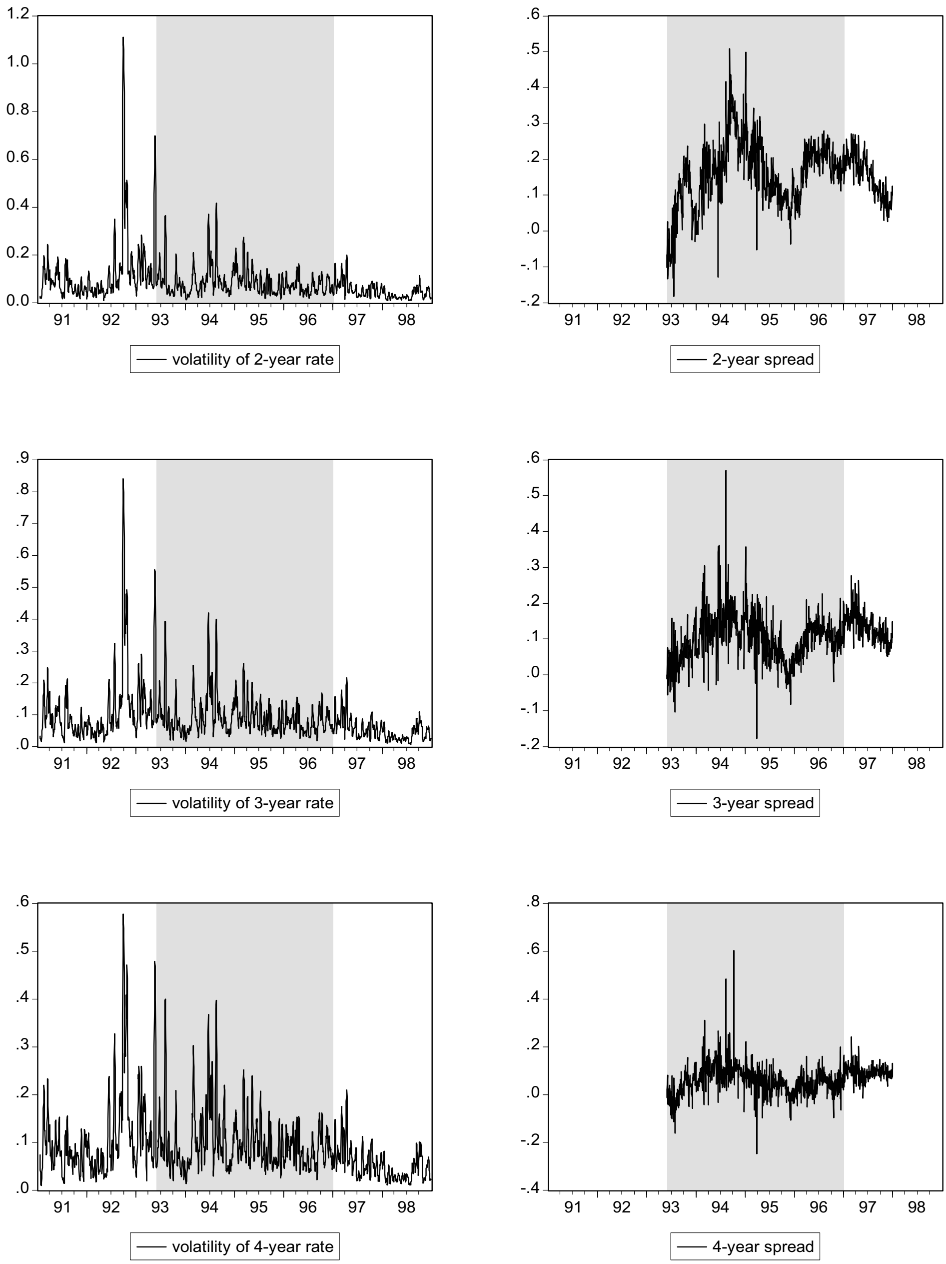
6.1. *Market risk*

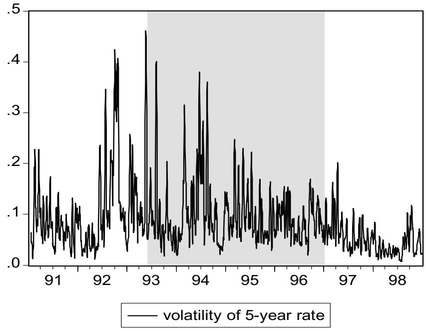
Interest or market risk in *IRS* contracts is analogous to that involved in fixed income investments and, as indicated above, there is a broad consensus that the level of risk as perceived by investors explains the time evolution of term-premia in public debt markets. In keeping with the existing literature, we approximate interest rate risk through the volatility of zero coupon *IRS* rates. However, there is no single way of computing unobserved volatility[9], and so we consider several volatility proxies. The first belongs to the class of *historical volatility* or Fama-type volatility measures. Specifically, we used an unconditional standard deviation, measured as the sample standard deviation of spot rates for the last 15 days. A second measure computes risk through autoregressive conditional heteroscedasticity (GARCH) models, which assume a specific data generating process for the level of interest rates as well as for their variance.

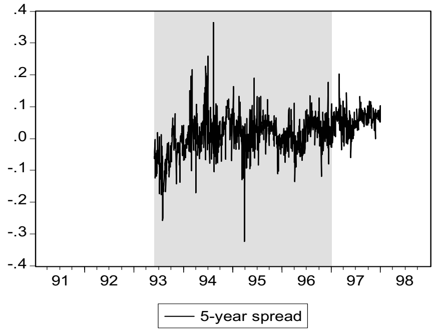
To obtain a first approximation to *market risk*, the left column in Figure 2 presents graphs of interest rate volatility for each maturity, computed as the standard deviation in a rolling window of 15 days of amplitude. The right column shows the interest rate spreads between the *IRS* and public debt markets. In both cases, the shaded area refers to the sample period used in the estimation, since this is the only period for which we have information on both premia and risk indicators. The variability in *swap* rates is similar for the different maturities, presenting almost the same pattern. Furthermore, interest rates exhibit greater volatility levels for the period before the end of 1995, becoming smoother after that point. Similar results are reported by Benito (2000), who stresses the significant reduction in the volatility of the term structure for the Spanish public debt market since the beginning of 1996. This is justified by the sharp increase in the probability assigned by market operators to the entrance of Spain in the European Monetary Union.

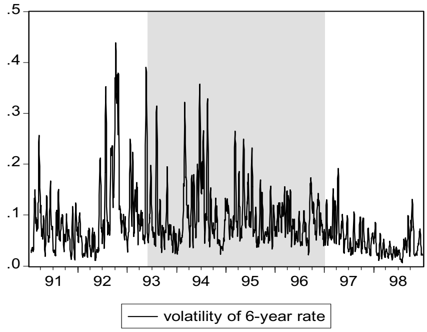
9. There are also a large number of papers comparing the ability of the different measures to predict future volatility. However, these results do not find significant evidence in favour of a single volatility measure.

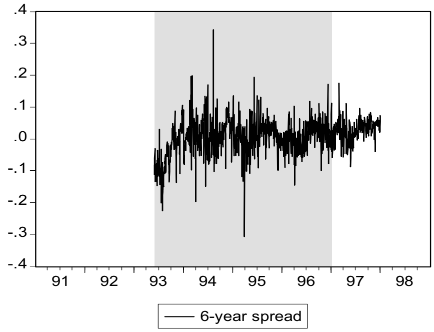
Figure 2.
Interest rate volatility indicator: half-month rolling-window standard deviation. Sample period: 1/4/1991 to 12/31/1998.
Spreads between term structure of *IRS* and public debt markets.
Sample period: 61/1993 to 12/31/1997.

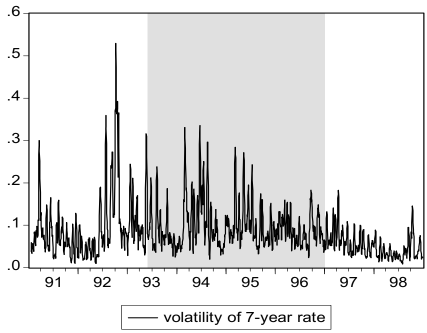


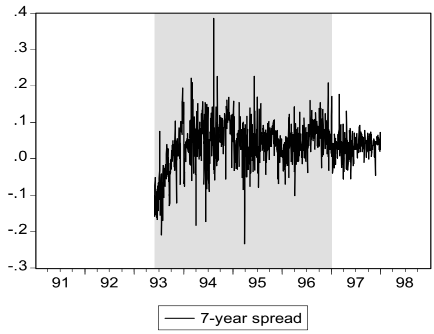

volatility of 5-year rate

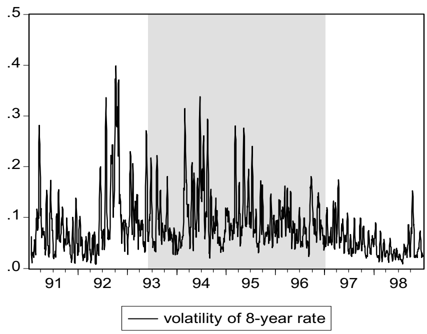

5-year spread

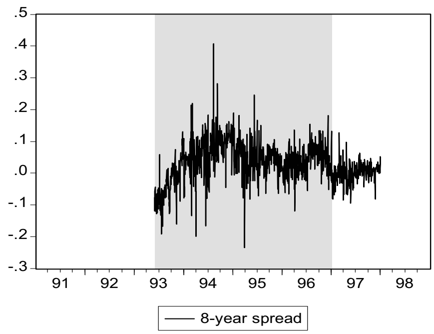

volatility of 6-year rate


6-year spread


volatility of 7-year rate


7-year spread


volatility of 8-year rate


8-year spread

6.2. *Credit risk* and *liquidity risk*

Credit and liquidity risks, in contrast, are specific to assets trading in *OTC* markets[10]. Since investment in public debt is exposed solely to interest rate risk, any difference between returns in both markets can be explained by the existence of credit and liquidity risk in the *IRS* market. Consequently, we propose a joint measure of these two sources of risk, as the spread between the estimated term structures for the *IRS* and the public debt markets.

Table 5

Descriptive statistics and unit root test for spreads between Spanish public debt and *IRS* markets

Spreads	2 year	3 year	4 year	5 year	6 year	7 year	8 year
Average	0.159	0.108	0.063	0.019	0.008	0.039	0.031
Maximum	0.508	0.570	0.602	0.365	0.343	0.387	0.408
Minimum	-0.183	-0.178	-0.247	-0.324	-0.307	-0.235	-0.235
Standard deviation	0.088	0.062	0.058	0.061	0.054	0.058	0.060
Skewness	-0.159	0.430	0.777	-0.419	-0.439	-0.514	0.028
Kurtosis	4.146	6.574	12.242	6.010	6.114	5.916	5.006
Observations	1197	1197	1197	1197	1197	1197	1197
ADF-Level	-4.421	-6.242	-6.862	-6.244	-7.324	-7.606	-6.419
PP- Level	-8.633	-18.099	-22.528	-19.853	-23.074	-21.858	-20.520

Note: Sample period: 6/1/1993 to 12/31/1997. Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root statistics. Critical values for both statistics: -3.439 (1%), -2.864 (5%), -2.568 (10%).

Figure 2 shows the evolution of market spreads for each maturity, while Table 5 contains some descriptive statistics. Average spreads are positive and statistically significant in all cases, reflecting that swap rates are usually above the zero coupon rates that emerge from the secondary public debt market. The volatility of spreads is generally decreasing with maturity. Spreads seem to be stationary for all maturities. Occasionally negative spreads can be explained by the lower liquidity of the public debt market relative to the swap market over the period considered.[11] Nevertheless, spreads are neither increasing nor decreasing on maturity, probably because liquidity in swap markets is unrelated to maturity.[12]

As can be seen in Table 6, correlations between spreads across maturities are substantial, generally decreasing as the time gap between maturities increases. This suggests the existence of a significant common factor producing co-movements across the term structure of spreads, which we estimate through the principal components technique. Principal components is a particular form of

10. As is well known, *over the counter (OTC)* trades take place outside organized markets, being made by financial intermediaries who trade directly among them through electronic systems. Their main differences with an organized market are: a) absence of a compensation chamber that could assume the counterpart risk and b) flexible contracts, which can be made to accommodate the needs of any specific trade.

11. Swap markets are viewed by many market participants as more liquid than the corresponding government bond markets, making swaps more efficient in reflecting changes in underlying interest rates. Consequently, there is a growing trend in financial markets to use the swap curve rather than the government curve for the pricing of bonds and other derivative securities.

12. Díaz and Navarro (2002) analyse the yield spreads between Treasury and non-Treasury Spanish fixed income assets and their relationship with the term to maturity. They observe a similar relation between the average spread and the term as that presented in this paper.

Table 6

Correlation coefficients between spreads

Spreads	2 year	3 year	4 year	5 year	6 year	7 year	8 year
2 year	1	0.714	0.534	0.390	0.398	0.539	0.587
3 year	0.714	1	0.795	0.626	0.583	0.566	0.544
4 year	0.534	0.795	1	0.779	0.748	0.650	0.605
5 year	0.391	0.626	0.779	1	0.935	0.741	0.656
6 year	0.398	0.583	0.748	0.935	1	0.908	0.811
7 year	0.539	0.566	0.650	0.741	0.908	1	0.928
8 year	0.587	0.544	0.605	0.656	0.811	0.928	1

Table 7

Principal components in the term structure of spreads

	2 year	3 year	4 year	5 year	6 year	7 year	8 year
Eigenvalues	0.0199	0.0049	0.0024	0.0008	0.0005	0.0002	2.18 E-5
First eigenvector	0.492	0.37	0.348	0.356	0.331	0.359	0.366
Second eigenvector	-0.76	-0.19	0.117	0.386	0.379	0.237	0.153
Third eigenvector	0.183	-0.505	-0.474	-0.246	0.056	0.394	0.518
Variance prop.	0.691	0.169	0.084	0.028	0.019	0.008	0.001
Cumulative prop.	0.691	0.860	0.944	0.972	0.991	0.999	1

Note: Eigenvalues and eigenvector for the variance-covariance matrix of spreads. The Variance prop. displays the variance proportion explained by each principal component (this value is simply the ratio of each eigenvalue to the sum of all eigenvalues). The Cumulative prop. displays the cumulative sum of the variance prop. from left to right and is the variance proportion explained by principal components up to that order.

factor analysis aimed at producing a few linear combinations of a set of variables explaining as much fluctuation as possible in the whole vector of variables. Principal components analysis starts from the variance-covariance matrix of the variables, in differences to their sample means. This is a semi-positive definite matrix with non-negative eigenvalues. This methodology constructs a specific linear transformation to the original variables, each transformed variable being a linear combination of the original, observed variables. The linear combinations are, by construction, uncorrelated with each other. The characteristic property of principal components is that they are obtained in a specific order: the first principal component is the linear combination that best explains fluctuations in the set of original variables, the second one is the linear combination of the original variables which best explains fluctuations among observed variables orthogonal to those experienced by the first principal component, and so on. The ratio between the largest eigenvalue and the sum of them all is the percentage of variance being explained by the first principal component.

Relative values of the eigenvalues of the covariance matrix of the vector of standardized spreads show that the first principal component explains 69% of the fluctuations in the whole vector (Table 7). This component, defined by the eigenvector associated to the largest eigenvalue, turns out to be an approximate average, assigning similar weights to all the spreads. It can therefore be interpreted as capturing the general level of the spreads between markets. The second eigenvector gives weights of an opposite sign to short- and long-term maturities, smoothly increasing as we move from one end of the term structure to the other. It can hence be interpreted as a general measure of steepness along the term structure of spreads. This second component captures changes in the slope of the term structure of spreads. The third eigenvector assigns positive weights to maturities at both the short- and the long-end of the term structure, and negative weights to intermediate maturities. Changes in this component will therefore generally be associated with changes of the opposite sign at short- and long-term maturities compared with intermediate maturities, thereby capturing changes in the curvature of the term structure of spreads. This interpretation of the principal components of the vector of spreads is very similar to that obtained in previous work on the zero coupon curve.

Two significant considerations emerge from this analysis. Firstly, slope and curvature components jointly explain 25% of the fluctuations in the vector of spreads. Hence, the slope and curvature of the term structure in the *IRS* and bond markets are generally different, which may be corroborated by simple visual inspection of both curves. Determinants of slope and curvature of the zero coupon curve in the bond and swap markets are significantly different. A similar conclusion is reached in Abad (2004), where it is shown that one of the common trends in the term structure of swap market can be identified with the trend of the term structure in the bond market, while the remaining trends are specific to the swap market. Moreover, this result can explain the disparate evidence regarding the validity of the Expectations Hypothesis in the bond and swap markets: evidence on the term structure for the Spanish bond market is broadly consistent with the expectations hypothesis (Massot and Nave (2003) and Martínez and Navarro (2002)), while the hypothesis cannot explain the behavior of interest rates across the term structure in the swap market in pesetas.

Secondly, this result allows us to decompose spreads, and hence default and credit risk, in two components: one, which is specific to the swap market, that does not depend on maturity and is captured by the first principal component; and a second one that depends on maturity and is identified with components capturing changes in slopes and curvature. Since our goal is to determine whether the risk specific to the swap market is a determinant factor of ex-post premia, we use the first principal component, which captures the level of the vector of spreads, as a proxy for *credit/liquidity* risk.

An interesting issue is the attempt to separate the impact of two different sources of risk on these spreads: liquidity and default risk. Recent papers analyze this issue in corporate spreads, for instance, Díaz and Navarro (2002), who examine the yield spreads between Treasury and non-Treasury bonds on the Spanish fixed income markets. They use a variables vector to explain corporate spreads: proxies of liquidity and proxies of default. Unfortunately, it is not possible to analyse both sources of risk separately on the *IRS* market since there is no information on the commonly used explanatory variables.

6.3. Is there any a risk premium incorporated in swap rates?

Once we have proxies for the different types of risk involved in an *IRS* portfolio, we can search for their possible effects on observed premia. Regression estimates in Table 8 show an effect to exist for credit/liquidity risk, though not for market risk. Coefficients associated with the proxy for credit/liquidity risk are positive and statistically significant, suggesting that an increase in either one of these two types of risk increases ex-post premia at all maturities.[13] Furthermore, the effect is increasing with maturity.

13. We considered an alternative model specification in which the second principal component (which captures changes in the slope of the term structure of spreads) was included as an additional proxy for credit/liquidity risk. This variable was significant for some maturities. However, the high correlation (higher than 0.96) between the residuals obtained from this specification and those corresponding to the specification in Table 8 indicates that the second principal component is not a relevant explanatory factor.

Table 8

Determinants of *ex-post* premia: the role of risk

$$\Lambda P_t^m = a + b_5\,\Lambda P_{t-5}^m + b_6\,\Lambda P_{t-6}^m + b_9\,\Lambda P_{t-9}^m + c_1\,S_t + c_2\,V_t^m + u_t$$

m	a	b_5	b_6	b_9	c_1	c_2	R^2	ADF u_t	Q(10)	Q(15)
2 year	0.003	--	-0.091*	0.036	0.489*	0.085	0.100	-13.307*	6.02	13.63
	(0.221)		(-3.146)	(1.435)	(6.871)	(0.656)			[0.81]	[0.55]
3 year	-0.001	--	-0.120*	0.046	0.714*	0.195	0.113	-12.942*	13.11	26.67
	(-0.045)		(-3.619)	(1.543)	(7.270)	(0.969)			[0.22]	[0.03]
4 year	0.018	--	-0.102*	--	0.940*	0.059	0.108	-12.784*	17.74	30.49
	(0.620)		(-3.154)		(8.115)	(0.236)			[0.06]	[0.01]
5 year	0.032	--	-0.081*	--	1.163*	-0.030	0.100	-12.889*	18.32	25.66
	(0.906)		(-2.487)		(8.008)	(-0.096)			[0.12]	[0.04]
6 year	0.059	0.045	-0.083*	--	1.390*	-0.278	0.105	-13.149*	11.65	22.06
	(1.305)	(1.357)	(-2.567)		(7.996)	(-0.676)			[0.31]	[0.11]
7 year	0.087	0.057*	-0.082*	--	1.627*	-0.521	0.107	-13.080*	10.31	19.76
	(1.562)	(1.709)	(-2.545)		(7.910)	(-1.038)			[0.41]	[0.18]
8 year	0.113*	0.059*	-0.080*	--	1.874*	-0.746	0.108	-13.086*	9.02	18.62
	(1.806)	(1.826)	(-2.446)		(7.813)	(-1.331)			[0.53]	[0.23]

Note: Sample period: 6/1/1993 to12/31/1996. Least squares estimates, with Newey-West standard deviations, robust to the presence of heteroscedasticity and autocorrelation. *t*-ratios in brackets. ΛP_t^m is the realized *ex-post* premia at maturity *m* in first differences. S_t denotes the first principal component that captures the general level of spread between the *IRS* and public debt term structures. V_t^m is the rolling-window standard deviation of interest rates at maturity *m*. Augmented Dickey-Fuller (ADF) unit root tests on the residuals include a constant term, but no trend, and 4 lagged residuals. Critical value at 10% significance is -2.568. In all cases, an asterisk denotes a rejection of the null hypothesis at the 90% confidence level. Q(10), Q(15) stand for Ljung-Box statistics on the residuals. *p*-values for the null hypotheses of lack of autocorrelation are shown in square brackets.

In contrast, the coefficient associated with market risk turns out not to be significant, suggesting that this type of risk may not influence ex-post premia. Even though we only present results for the model that includes the standard deviation of interest rates calculated on rolling windows as a proxy for *market* risk, these results are robust to the use of alternative proxies. The consensus that market risk is relevant is strong enough for our results to be interpreted as failure to detect a significant effect in the available data, rather than suggesting that this type of risk is not important.

A possible explanation for this result is that the above estimates do not explicitly consider the fact that *ex-post* premia change sign from the first to the second part of our sample period. Consequently, we may just be averaging an effect, which was of a different size and/or sign in the two subperiods. We estimated the same model including a dummy variable to distinguish between the two time periods before and after March 1994, when *ex-post* premia change sign. Figure 2 shows that volatility was high in most of the first period, and the results in Table 9 suggest that *market risk* then has a significant positive effect on premia, and the effect of *market risk* is increasing in maturity[14]. In the more stable second subsample, however, *ex-post* premia becomes positive under a more credible monetary policy, and we do not detect a significant effect for *market risk*. It appears as if in volatile periods market participants extrapolate the currently high level of volatility when forecasting future spot rates. This higher forecast gets embedded in the term structure in the form of higher risk premia.

14. Before March 1994, the mean interest rate volatility is similar for different maturities (about 0.083).

Table 9

Determinants of *ex-post premia*: Two subsamples

$$\Lambda P_t^m = a + b_5\,\Lambda P_{t-5}^m + b_6\,\Lambda P_{t-6}^m + b_9\,\Lambda P_{t-9}^m + c_1\,S_t + c_2\,V_t^m + c_3\,V_t^m \bullet F_t + u_t$$

m	a	b_5	b_6	b_9	c_1	c_2	c_3	R^2	ADF u_t	Q(10)	Q(15)
2 year	0.006	--	-0.080*	0.045*	0.605*	-0.141	0.970*	0.116	-13.259*	3.85	7.84
	(0.516)		(-2.860)	(1.800)	(8.096)	(-1.046)	(4.123)			[0.95]	[0.93]
3 year	0.005	--	-0.109*	0.051*	0.881*	-0.121	1.342*	0.137	-12.879*	9.68	19.92
	(0.260)		(-3.329)	(1.728)	(8.639)	(-0.606)	(3.953)			[0.47]	[0.17]
4 year	0.024	--	-0.095*	--	1.136*	-0.284	1.619*	0.129	-12.802*	13.61	23.89
	(0.863)		(-2.934)		(9.358)	(-1.075)	(3.660)			[0.19]	[0.07]
5 year	0.038	--	-0.076*	--	1.381*	-0.421	1.851*	0.116	-12.911*	12.20	20.57
	(1.079)		(-2.337)		(9.043)	(-1.232)	(3.284)			[0.27]	[0.15]
6 year	0.052	0.054*	-0.076*	--	1.647*	-0.617	2.420*	0.123	-13.240*	8.62	17.05
	(1.203)	(1.631)	(-2.375)		(8.975)	(-1.449)	(3.568)			[0.57]	[0.32]
7 year	0.061	0.067*	-0.075*	--	1.934*	-0.753	2.972*	0.125	-13.115*	8.50	16.58
	(1.133)	(2.007)	(-2.318)		(8.754)	(-1.457)	(3.531)			[0.58]	[0.35]
8 year	0.086	0.069*	-0.072*	--	2.199*	-1.002	3.194*	0.123	-13.119*	7.57	15.58
	(1.401)	(2.109)	(-2.226)		(8.447)	(-1.715)	(3.266)			[0.67]	[0.41]

Note: Sample period: 6/1/1993 to12/31/1996. Least squares estimates, with Newey-West standard deviations, robust to the presence of heteroscedasticity and autocorrelation. *t*-ratios in parentheses. ΛP_t^m is the realized *ex-post* premia at maturity *m* in first differences. S_t denotes the first principal component that captures the general level of spread between the *IRS* and public debt term structures. V_t^m is the rolling-window standard deviation of interest rates at maturity *m*, F_t is a dummy variable, equal to 1 from 6/1/1993 to 3/1/1994, 0 otherwise. Augmented Dickey-Fuller (ADF) unit root tests on the residuals include a constant term, but no trend, and 4 lagged residuals. Critical value at 10% significance is -2.568. In all cases, an asterisk denotes a rejection of the null hypothesis at the 90% confidence level. Q(10), Q(15) stand for Ljung-Box statistics on the residuals. *p*-values for the null hypotheses of lack of autocorrelation are shown in square brackets.

From these results, we conclude that the level of risk involved in *IRS* positions is a relevant variable to explain *ex-post* premia, at least in periods of higher market volatility. Models explaining premia through the use of a *market risk* show a much better fit than without the proxy. As expected, in this case *market* and *credit/liquidity risk* have a positive effect on changes in premia, indicating that an increase in either type of risk implies an increase in risk premia. Consequently, observed premia in *swap* markets seem to partially compensate investors for the level of risk in their market positions.

7. CONCLUSIONS

Price formation at long maturities in swap markets (*IRS*) or public debt markets might be expected to be relatively comparable, though possibly different from interbank markets or markets for eurodeposits, where only maturities up to one year are negotiated. This difference is potentially relevant for tests of the Expectations Hypothesis (*EH*), which might hold only on some interval of the term structure. In fact, tests of the hypothesis on short maturities find generally favorable evidence, while those using longer maturities fare much worse. In this paper, we test the *EH* using estimated relationships between *forward* and future spot interest rates. After conclusively rejecting the hypothesis, we proceed to analyze *ex-post* premia and their determinants. To do so, we assigned numerical measures to the different types of risk involved in *swap* positions so as to estimate the extent to which observed premia are a consequence of risk perceptions among market participants.

As mentioned above, our results suggest that the *EH* does not adequately explain the price formation mechanism in *swap* markets. The *EH* assumes that any information currently available that is of any use in predicting future spot rates is contained in the *forward* rates implicit in the current term structure. Contrary to this view, we have shown that there is information available to the investor, which is additional to that contained in forward rates, that is also useful in predicting future spot rates. In particular, we have shown that *ex-post risk premia*, the difference between future spot rates and current forward rates, are partially predictable, since they present a non-trivial dynamic pattern, and their value depends on the levels of the different kinds of risk involved in this financial product. This should be taken into account when predicting future spot rates. However, a more explicit evaluation of the additional predictive ability is needed.

As regards *ex-post* premia in the *IRS* market in pesetas, we have shown that they present certain characteristics that are specific to this market: a) they change over time, b) they are relatively stable in sign, and c) their value depends on the level of risk in *IRS* positions. We have also shown that over most of our sample period investors in the *swap* markets display a preference for the short-term. This preference is stable over time and is first observed when the easing of monetary policy was most intense in Spain. These results have a clear potential for portfolio management in practice, for which risk premia determination is crucial.

Finally, it is worth noting that the slope and curvature components of the term structure of spreads between the *IRS* and the public debt markets (proxy for *credit/liquidity* risk) jointly explain 25% of the fluctuations in the vector of spreads. The slope and curvature of the term structure in the *IRS* and bond markets are hence generally different. A similar conclusion is reached in Abad (2004), and this result may explain the disparate evidence regarding the validity of the Expectations Hypothesis in the bond and swap markets in pesetas. However, we have not analyzed the determinant for these spreads. This question and the search for the best sets of variables as factors for separately explaining liquidity risk and credit risk remain as interesting issues for further research.

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