# RISK PREMIA IN THE TERM STRUCTURE OF SWAPS IN PESETAS[1]

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**Abstract**: Some characteristics of the term structure in interest rate swap (*IRS*) markets are influenced by the own idiosyncrasy of this financial instrument, which could 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 *risk premia*. We then focus on characterizing some properties of realized, *ex-post term-premia*, and provide explanatory variables for them. 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, risk premium.

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## 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 made of the *IRS* term structure 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 favor of the non-stationarity of spot and forward rates, we explore the possibility 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 term-*premia*, the paper focuses on characterizing their time behavior 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 term premia*.

We briefly review in Section 2 the Expectations Theory on the formation of the term structure of interest rates as well as that of risk-premia, and 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 risk-premia, and their main characteristics are analyzed. In Section 6 we analyze 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 results from an equilibrium in which, given current expectations of future interest rates, the investor is indifferent between short- and long-term positions. In that case, *term-premia* are 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 on 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 $m<n$, ($f_{t,t+n,m}$):

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

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. Writing again the definition of the *forward* premium under the assumption that agents form their expectations rationally[1]:

$$r_{t+n,m}=-\frac{P^f_{t,n,m}}{m}+f_{t,t+n,m}+u_{t+n} \tag{2}$$

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

$$r_{t+n,m} = a + b f_{t,t+n,m} + u_{t+n} \tag{3}$$

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 risk premia. Then, characterizing the determinants of the sign and level of risk 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 Fama (1976, 1984a, 1984b). Fama finds positive premia, increasing with maturity, similarly to findings 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 risk premia in fixed income markets was Kessel (1965), who works under the assumption that the relationship between risk-premia and its determinants is linear. Empirical results on 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 term-premia, but with coefficients of 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 year, 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 seems to be in the literature a consensus on the fact that interest rate volatility is a main determinant of risk-premia. Fama (1976) shows evidence consistent with that 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 previous authors.

## 3. The data

We have used data from two markets. To test the *EH* and study term-premia in the market for *swaps* in pesetas, we have used the *TSIR* of *IRS* denominated in pesetas. To quantify the level of credit and liquidity risk involved in *IRS* , we have used the *TSIR* from the secondary market for Spanish public debt[3]. The *TSIR* for the *IRS* market was estimated through 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*[TM], 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 by nine zero coupon rates, observed daily from January 4, 1991 to December 31, 1998. There is 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]. As a consequence, we 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].

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[6].

## 4. Testing the expectations hypothesis in the market for *swaps*

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

Therefore, this evidence 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, contradicting the Expectations Hypothesis. As already indicated, this can be provoked by the presence of time-varying risk premia in this market. This is the question we analyze in the next section.

## 5. Computing *ex-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 risk premium [equation (1)]. The resulting premia are usually known as *ex-post* premia. We have computed them 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 risk premia, as pointed out by Evans and Lewis (1994) in fixed income markets, *ii*) term-premia are positive over the time period considered, except for the March1993 to March 1994 interval, and *iii*) term-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 most intense point in the process of monetary easing, the term structure adopted again a decreasing shape at the shorter maturities.

That *ex-post* premia are not stationary is ratified by unit root tests in Table 3. Furthermore, Table 3 also shows some descriptive statistics for term premia at each maturity[7]. Average term premia are positive, significantly different from zero, and increasing with maturity, in consistency with the intuition that uncertainty increases with the horizon of a given investment. On the contrary, daily changes in term-premia are not different from zero for any maturity. In both cases, dispersion increases with maturity.

Since unit root tests suggest that term-premia follow integrated processes of order one, we formulated dynamic models in first differences of *ex-post term 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 heteroskedastic and autocorrelated residuals. These results indicate that daily changes in *ex-post* premia follow autoregressive structures of up to order 9.

## 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 term-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 solvency 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 have therefore considered risk as a possible determinant of observed *ex-post* term-premia. Following Kessel (1965), who represent premia as linear functions of potential explanatory variables. That way, we have included in the models specified in previous sections two variables intended to capture the risk involved in an *IRS* contract, that we define next.

## 6.1. *Market risk*

Interest or market risk in *IRS* contracts is analogue to that involved in fixed income investments and, as indicated above, there is a broad consensus on the fact that the level of risk as perceived by investors explains the time evolution of term-premia in public debt markets. Following the existing literature, we approximate interest rate risk through the volatility of zero coupon *IRS* rates. Nevertheless, there is not a single way to compute unobserved volatility[8], and we consider several volatility proxies. Two of them belong to the class of *historical volatility* or Fama-type volatility measures. Specifically, we have used an unconditional standard deviation, measured as the sample standard deviation of spot rates for the last 15 days, and an exponential smoothing, with decay factor of $\lambda= 0{,}94$[9]. A third measure computes risk through autoregressive conditional heteroskedasticity (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 estimation, since it 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, showing 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 shown 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.

## 6.2. *Credit risk* and *liquidity risk*

On the contrary, credit and liquidity risks are specific to assets trading in *OTC* markets[10]. Since investments on public debt are exposed just 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.

Figure 2 shows the dynamic evolution of market spreads for each maturity, while Table 5 contains their main descriptive statistics. It can be seen that the dynamic evolution of these spreads is similar for the different maturities considered, suggesting that the term structure of spreads does not change significantly over time. It displays a *U*-shape pattern over the whole sample period, being more stable once premia became positive after March 1994. 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. Nevertheless, spreads are neither increasing nor decreasing on maturity, probably because liquidity in *swap* markets is unrelated to maturity. Average spread volatility seems to decrease with maturity.

## 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 6 show that to be the case for *credit/liquidity* risk, although not for *market* risk. Coefficients associated with the proxies 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. Furthermore, the effect is increasing with maturity.

On the contrary, the coefficient associated to *market* risk turns out not to be significant, suggesting that this type of risk may not influence *ex-post* premia. Even though we just present results for the model that includes the standard deviation of interest rates calculated on rolling windows as a proxy for *market* risk, results are robust to the use of alternative proxies. The consensus that market risk is relevant is strong enough that our results should be interpreted as a 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 previous estimates do not consider explicitly the fact that *ex-post* premia change sign from the first to the second part of our sample period. Because of that, we could be just 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 changes sign. Figure 2 shows that volatility was high in most of the first period, and the results in Table 7 suggest that *market risk* has then a significant positive effect on term-premia, except at the 2-year maturity, and the effect of *market risk* is increasing in maturity. On the contrary, in the more stable second subsample, *ex-post* premia becomes positive under a more credible monetary policy, and we do not detect a significant effect for *market risk*. It looks 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 term premia.

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 that 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 term-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, although 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*) , who might hold just 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 that end, we have assigned numerical measures to the different types of risk involved in *swap* positions, to estimate the extent to which observed premia are a consequence of risk perceptions among market participants.

As mentioned, our results suggest that the *EH* does not adequately explains the price formation mechanism in *swap* markets. The *EH* assumes that any information currently available which is of any use to predict 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, additional to that contained in forward rates, which is also useful to predict future spot rates. In particular, we have shown that *ex-post term-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.

Relative to *ex-post* premia in the *IRS* market in pesetas, we have shown that they present some characteristics which 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 it is first observed when the loosening 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.

## References


Benito, S. (2000), “Volatilidad de los rendimientos cupón cero de la deuda pública. Estudio de transmisión de volatilidades”, Mimeo, Universidad Complutense de Madrid.

Bollerslev, T., Engle, R.F. and Wooldridge, J.M. (1988), “A Capital Asset Pricing Model with Time Varying Covariances”, *Journal of Political Economy*, 96: 116-131.

Campbell, J.Y. and Shiller, R. (1987), “Cointegration and Tests of Present Value Models”, *Journal of Political Economy* , 95: 1062-1088.

Engle, R.F. and Granger, C.W.J. (1987), “Cointegration and Error Correction: Representation, Estimation and Testing”, *Econometrica*, 55: 251-276.

Engle, R. F., Lilien, D. M. and Robins, R. P. (1987), "Estimating Time Varying Risk Premia in the Term Structure: The Arch-M model", *Econometrica*, 55: 391-407.

Evans, M. and Lewis, K. (1994), “Do Stationary Risk Premia Explain it all? Evidence from the Term Structure”, *Journal of Monetary Economics*, 33: 285-318.

Fama, E.F. (1976), “Forward Rates as Predictors of Future Spot Rates”, *Journal of Financial Economics*, 3: 316-377.

Fama, E.F. (1984a), ”The Information in the Term Structure”, *Journal of Financial Economics*, 13: 509-528.

Fama, E.F. (1984b), ”Term Premiums in Bond Returns”, *Journal of Finantial Economics*, 13: 529-546.

Fama, E.F. and Bliss, R.R. (1987), “The Information in Long-maturity Forward Rates”, *American Economic Review*, 77: 680-692.

Gómez, I. and Novales, A. (1997), “Estrategias de inmunización ante posibles desplazamientos de la Estructura Temporal”, *Análisis Financieros Internacionales*, Enero, 1997: 15-39.

Johansen, S. (1988), “Statistical Analysis of Cointegration Vectors”, *Journal of Economic Dynamics and Control*, 12: 231-254.

Johansen, S. (1991), “Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive models”, *Econometrica*, 59: 1551-1580.

Kessel, R. A. (1965), *The Cyclical Behavior of the Term Structure of the Interest Rates*, New York NBER.

McCulloch, J.H. (1987), “The Monotonicity of the Term Premium: A closer look”, *Journal of Financial Economics*, 18: 158-192.

Modigliani, F. R. and Shiller, R. J. (1973), “Inflation, Rational Expectations and the Term Structure of Interest Rates”, *Economica*, 40: 12-43.

Nelson, C.R. (1976), “Inflation and Rates of Return on Common Stocks”, *The Journal of Finance*, 31: 471-483.

Nelson, C.R. and Siegel, A. (1987), “Parsimonious Modelling of Yield Curves for US Treasury Bills”, *Journal of Business*, 60: 473-489.

Newey, W. and West, K. (1987), "A Simple Positive Semi-definite, Heteroscedasticity and Autocorrelation Consistent Covariance Matrix", *Econometrica*, 55: 703-708.

Shiller, R. (1979), “The Volatility of Long-term Interest Rates and Expectations Models of the Term Structure”, *Journal of Political Economy*, 87: 1190-1219.

Shiller, R., Campbell, J.Y. and Schoenholtz, K.L. (1983), “Forward Rates and Future Policy: Interpreting the Term Structure of Interest Rates”, *Brookings Papers on Economic Activity*, 1: 173-217.

# Appendix

**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: $r_{t+2,m} = a + b f_{t,t+2,m} + u_t$

| | Engle-Granger tests | | | | | Reduced rank tests | | |
|---|---|---|---|---|---|---|---|---|
| m | a | b | $R^2$ | ADF $u_t$ | PP $u_t$ | Hypothesis | $\lambda_{MAX}$ | $\lambda_T$ |
| 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 y 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($\lambda_{MAX}$) and trace ($\lambda_T$) 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 were included in this model. An asterisk denotes rejection of the null hypothesis at the 90% confidence level.

**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 |
| Curtosis | 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 |
| Curtosis | 11.657 | 18.458 | 14.312 | 18.870 | 12.254 | 9.276 | 8.994 |
| Observations | 1563 | 1563 | 1563 | 1563 | 1563 | 1563 | 1563 |

Sample period: 1/4/1991toa 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 90% confidence.

**Table 4.** Dynamic models for *ex-post* premia

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

| 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) | | | [0.916] | [0.948] |
| 3 year | 0.012 | -- | -0.092* | 0.061* | 0.012 | -13.581* | 6.124 | 13.643 |
| | (1.088) | | (-2.819) | (1.881) | | | [0.805] | [0.553] |
| 4 year | 0.017 | -- | -0.077* | -- | 0.006 | -13.469* | 10.387 | 16.398 |
| | (1.188) | | (-2.341) | | | | [0.407] | [0.356] |
| 5 year | 0.021 | -- | -0.060* | -- | 0.004 | -13.628* | 11.168 | 15.100 |
| | (1.191) | | (-1.776) | | | | [0.345] | [0.444] |
| 6 year | 0.023 | 0.053* | -0.064* | -- | 0.007 | -14.088* | 7.575 | 11.244 |
| | (1.108) | (1.568) | (-1.907) | | | | [0.670] | [0.735] |
| 7 year | 0.026 | 0.062* | -0.066* | -- | 0.008 | -14.014* | 6.668 | 10.004 |
| | (1.045) | (1.859) | (-1.969) | | | | [0.756] | [0.819] |
| 8 year | 0.029 | 0.064* | -0.065* | -- | 0.008 | -14.004* | 6.177 | 9.706 |
| | (1.013) | (1.921) | (-1.924) | | | | [0.800] | [0.838] |

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. 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 90% confidence level. Q(10) y Q(15) are Ljung-Box statistics for residual autocorrelation. *p*-value in square brackets.

**Table 5.** Descriptive statistics 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 |
| Curtosis | 4.146 | 6.574 | 12.242 | 6.010 | 6.114 | 5.916 | 5.006 |
| Observations | 1197 | 1197 | 1197 | 1197 | 1197 | 1197 | 1197 |

Sample period: 6/1/1993 to 12/31/1997.

**Table 6.** Determinants of *ex-post* premia: the role of risk

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

| *m* | a | $b_5$ | $b_6$ | $b_9$ | $c_1$ | $c_2$ | $R^2$ | ADF $u_t$ | Q(10) | Q(15) |
|---|---|---|---|---|---|---|---|---|---|---|
| 2 year | -0.039* | -- | -0.074* | 0.041 | 0.242* | 0.085 | 0.016 | -14.110* | 4.352 | 6.853 |
| | (-2.297) | | (-2.615) | (1.498) | (2.769) | (0.796) | | | [0.930] | [0.962] |
| 3 year | -0.133* | -- | -0.120* | 0.046* | 1.552* | -0.122 | 0.098 | -13.228* | 8.857 | 17.149 |
| | (-4.934) | | (-3.646) | (1.543) | (7.127) | (-0.615) | | | [0.546] | [0.310] |
| 4 year | -0.116* | -- | -0.110* | -- | 2.454* | -0.023 | 0.119 | -13.641* | 11.396 | 20.492 |
| | (-3.699) | | (-3.239) | | (6.121) | (-0.089) | | | [0.327] | [0.154] |
| 5 year | -0.007 | -- | -0.074* | -- | 2.881* | 0.043 | 0.105 | -12.501* | 18.530 | 29.987 |
| | (-0.191) | | (-2.229) | | (8.013) | (0.121) | | | [0.047] | [0.012] |
| 6 year | 0.006 | 0.060* | -0.071* | -- | 4.041* | 0.033 | 0.129 | -12.890* | 15.503 | 29.428 |
| | (0.119) | (1.796) | (-2.221) | | (8.957) | (0.074) | | | [0.115] | [0.014] |
| 7 year | -0.139* | 0.067* | -0.073* | -- | 4.035* | 0.012 | 0.116 | -13.080* | 12.316 | 23.783 |
| | (-2.089) | (2.046) | (-2.279) | | (7.731) | (0.022) | | | [0.264] | [0.069] |
| 8 year | -0.107 | 0.065* | -0.075* | -- | 4.740* | -0.489 | 0.124 | -12.865* | 13.032 | 25.005 |
| | (-1.526) | (2.039) | (-2.303) | | (7.972) | (-0.906) | | | [0.222] | [0.050] |

Note: Sample period: 6/1/1993 to12/31/1996. Least squares estimates, with Newey-West standard deviations, robust to the presence of heteroscedasticiy and autocorrelation. *t*-ratios in parentheses. $P_t^m$ is the realized ex-post premia at maturity *m*. $S_t^m$ denotes the spread between the *IRS* and public debt term structures at maturity *m*. $V_t^m$ 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 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.

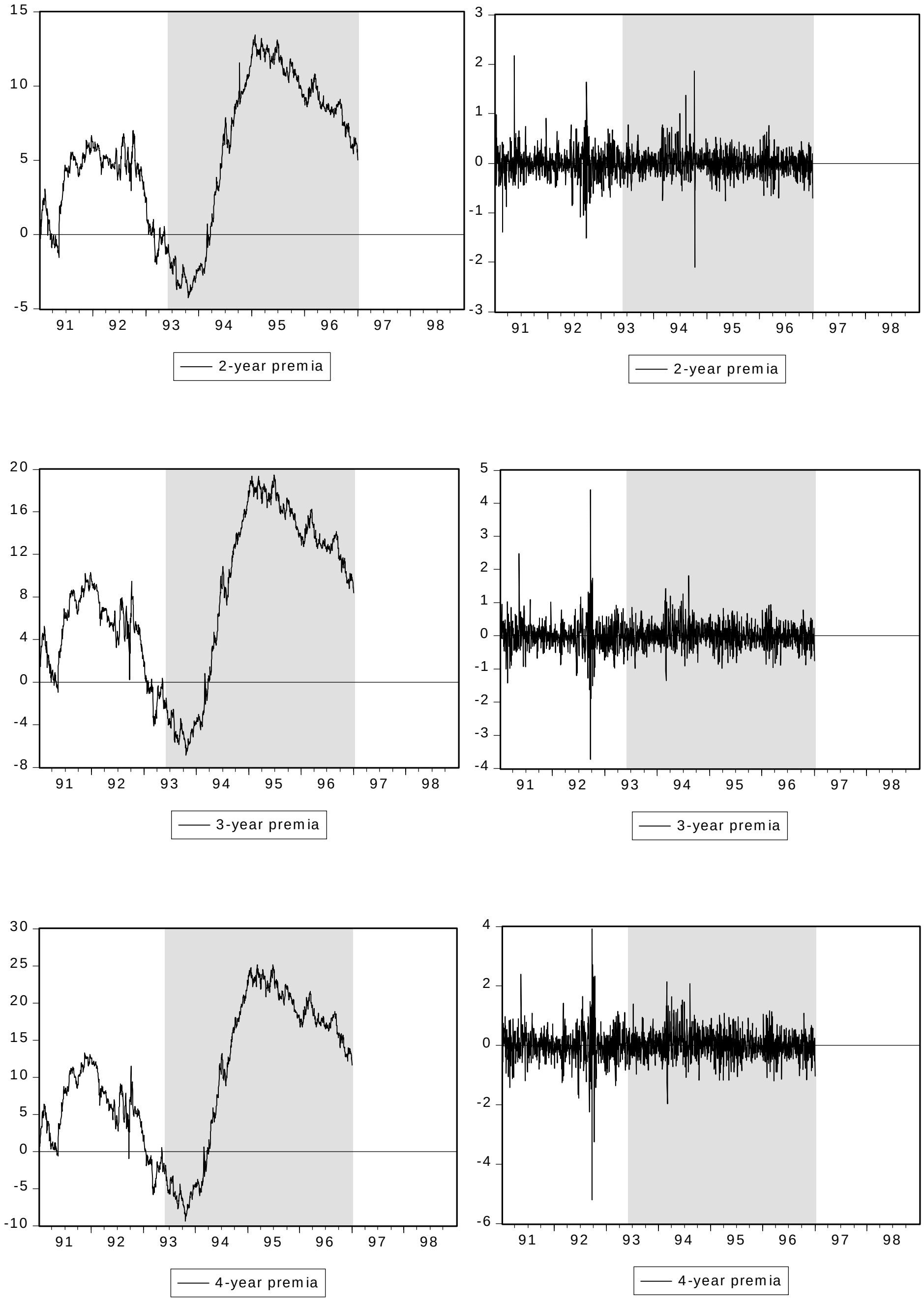
**Table 7.** Determinants of *ex-post premia*: Two subsamples

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

| *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 | -0.048* | -- | -0.071* | 0.044* | 0.306* | 0.006 | 0.309 | 0.018 | -13.935* | 4.103 | 6.649 |
| | (-2.621) | | (-2.459) | ( 1.573) | ( 2.858) | ( 0.045) | ( 1.197) | | | [0.943] | [0.967] |
| 3 year | -0.143* | -- | -0.113* | 0.050* | 1.701* | -0.325 | 0.738* | 0.107 | -13.030* | 8.621 | 16.543 |
| | (-5.549) | | (-3.444) | ( 1.647) | ( 7.560) | (-1.537) | ( 2.548) | | | [0.568] | [0.347] |
| 4 year | -0.122* | -- | -0.105* | -- | 2.605* | -0.204 | 0.857* | 0.126 | -13.596* | 10.710 | 20.094 |
| | (-4.115) | | (-3.104) | | ( 6.092) | (-0.745) | ( 2.242) | | | [0.381] | [0.168] |
| 5 year | -0.008 | -- | -0.067* | -- | 3.387* | -0.325 | 1.802* | 0.122 | -12.477* | 15.499 | 24.901 |
| | (-0.232) | | (-2.039) | | ( 9.275) | (-0.890) | ( 3.478) | | | [0.115] | [0.051] |
| 6 year | -0.007 | 0.069* | -0.064* | -- | 4.457* | -0.193 | 1.905* | 0.141 | -12.959* | 12.817 | 24.769 |
| | (-0.156) | ( 2.069) | (-2.010) | | ( 9.900) | (-0.423) | ( 3.077) | | | [0.234] | [0.053] |
| 7 year | -0.198* | 0.078* | -0.064* | -- | 4.657* | -0.112 | 2.696* | 0.131 | -13.144* | 10.974 | 20.640 |
| | (-3.013) | ( 2.382) | (-2.020) | | ( 8.312) | (-0.205) | ( 3.187) | | | [0.360] | [0.149] |
| 8 year | -0.179* | 0.077* | -0.066* | -- | 5.612* | -0.720 | 3.496* | 0.143 | -12.993* | 10.899 | 19.826 |
| | (-2.514) | ( 2.395) | (-2.039) | | ( 8.598) | (-1.293) | ( 3.538) | | | [0.365] | [0.179] |

Note: Sample period: 6/1/1993 to12/31/1996. Least squares estimates, with Newey-West standard deviations, robust to the presence of heteroscedasticiy and autocorrelation. *t*-ratios in parentheses. $P_t^m$ is the realized ex-post premia at maturity *m*. $S_t^m$ denotes the spread between the *IRS* and public debt term structures at maturity *m.* $V_t^m$ 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 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.

**Figure 1**. *Ex-post* premia and first differences
Sample period: 1/4/1991 to 12/31/1996.



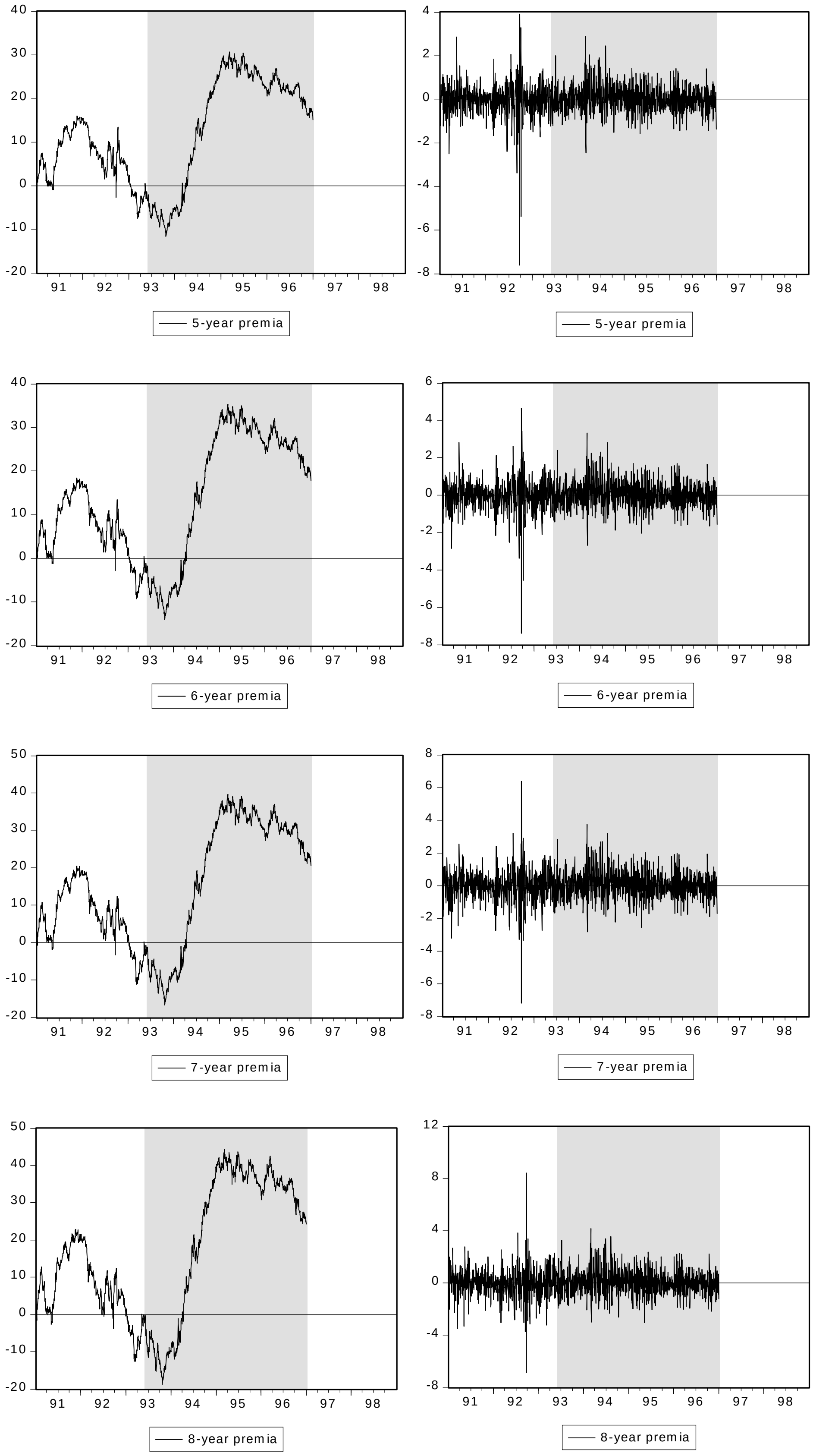
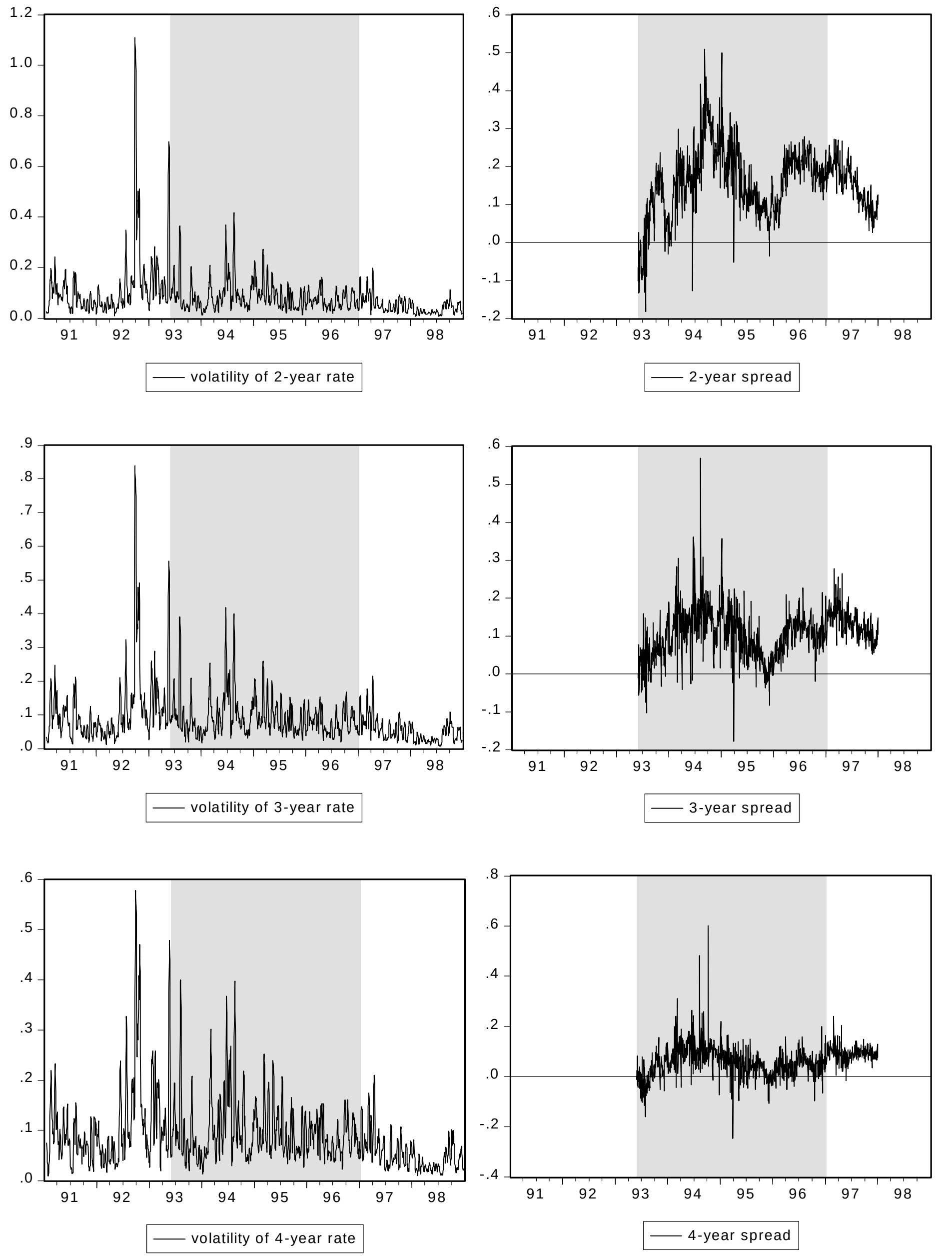
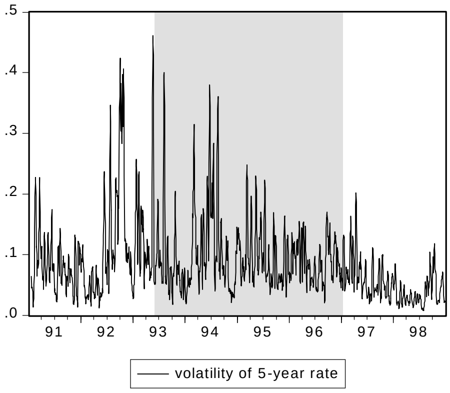
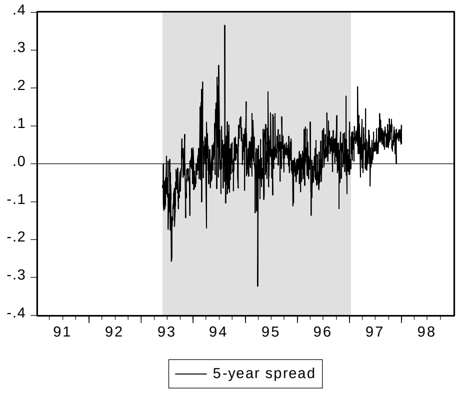
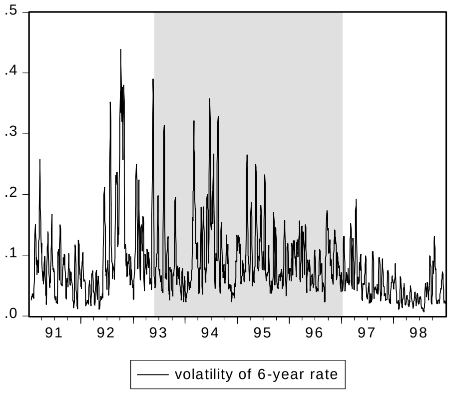
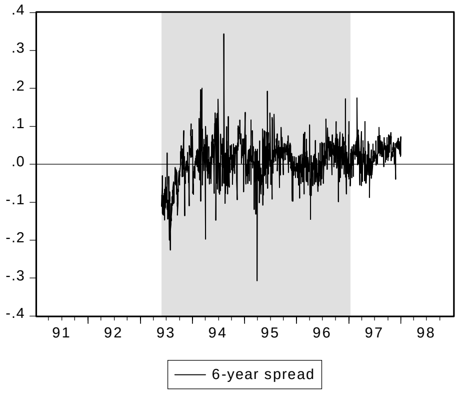
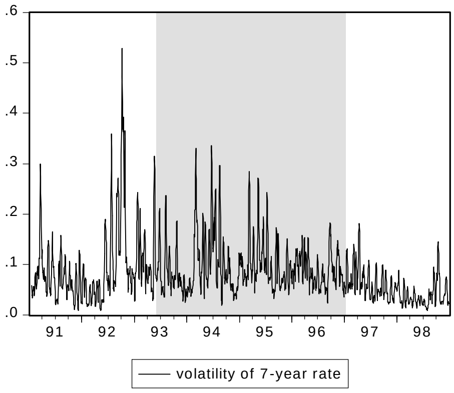
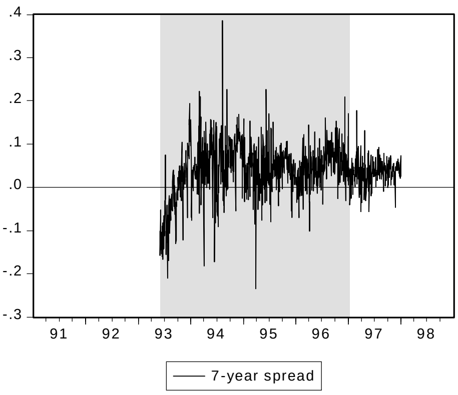
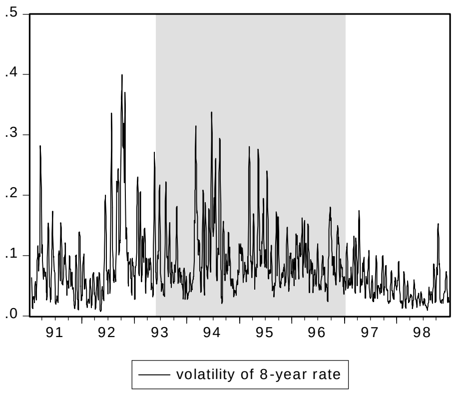
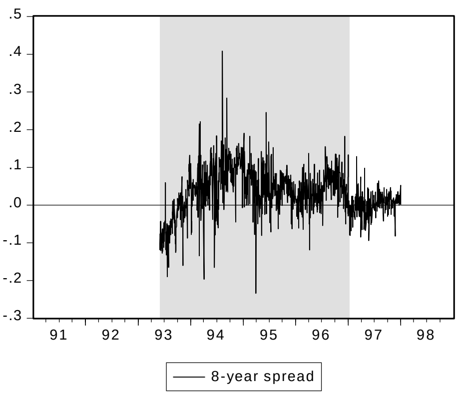

5-year premia
5-year premia
6-year premia
6-year premia
7-year premia
7-year premia
8-year premia
8-year premia

**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: 6/1/1993 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

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 information available at time $t$.

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. All of them 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, $f_{t,t+n,m}$, its computed from market rates observed at time $t$: $mf_{t,t+n,m} = (n+m)r_{t,n+m} - nr_{t,n}$.

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

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

8. There is 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 favor of a single volatility measure.

9. In the exponential smoothing method, the standard deviation is estimated by: $d_t(r_t) = \sqrt{(1-\lambda)(r_{t-1} - \bar{r})^2 + \lambda d_{t-1}^2}$ . The decay factor λ is chosen *a priori*. *JPMorgan* has developed *RiskMetrics*, where λ=0.94 is used to forecast volatility from daily data.

10. As it 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.