War and Emerging Market Default Risk: The Case of India and the Iraqi Invasion of Kuwait

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ABSTRACT

We use the performance of Indian Eurobonds over the period 1990-1992 to examine the sensitivity of India's creditworthiness to the Iraqi invasion of Kuwait on August 2, 1990. We also explore the related question of whether the changes in creditworthiness, measured as the effect of changes in default probabilities on bond prices, were accurately assessed by the market in a timely manner. We find that the markets systematically mis-estimated these effects. They anticipated no effects on India's default probabilities in the invasion quarter. All the change in Indian bond prices in the quarter that the invasion took place was due to changes in the risk free term structure of interest rates. In the quarter following the invasion, effects of changes in default probabilities were significant and caused a fall of nearly 3 points in Indian Eurobond prices. In the quarter when the Gulf War took place changes in default probabilities caused a further fall of 1.34 points in Indian bond prices. We find evidence of market over-reaction to country specific invasion effects.

JEL: O530, O160, G150, P330, F340

Keywords: Term structure of interest rates; Duration; Spline

I.   INTRODUCTION

The purpose of this paper is to use the performance of Indian Eurobonds over the period 1990-1992 to examine the sensitivity of India's financial creditworthiness to the Iraqi invasion of Kuwait on August 2, 1990. To this end, we divide bond price changes into those induced by general market conditions and those induced by changes in default probabilities. We also explore the related question of whether the changes in creditworthiness, measured as the effect of changes in default probabilities on bond prices, were accurately assessed by the market in a timely manner.

Extreme political events such as terrorism and war have become a major concern of the international capital markets since September 11, 2001. Events like these have worldwide as well as country specific economic, financial, political and social consequences and their effect on markets can be dramatic. They also threaten to become more frequent. Besides the recent war in Afghanistan and the coalition invasion of Iraq, India and Pakistan are on the warpath, Turkey is threatening the Iraqi Kurds, the US is threatening the Axis of Evil, Al Queda is threatening the West and North Korea is threatening any country within range of its missals. Given that many emerging economies are saddled with structural imbalances, social and political fragility, and financial dependence, they may be particularly vulnerable to such events, whose effects can highlight and exacerbate certain weaknesses above and beyond what is warranted by the fundamentals.[1] This would be the case, for example, if information assymetries, such as those discussed by Calvo (1998) and Calvo and Mendoza (2000) in the context of contagion, hindered timely and accurate analysis. The first contribution of this paper, then, is that we develop a methodology that makes it possible to distinguish between general market effects and country specific changes in default probabilities. The second contribution is that we apply this methodology to determine the particular case of country specific effects that the Iraqi invasion of Kuwait had on the prices of Indian Eurobonds.

India and the invasion of Kuwait is an interesting case study. First of all, India was not directly involved in the conflict and was far enough from the war theater that it would not be threatened directly by the fighting. It did, however, have close links to the region through its large contingent of emigrants working in the Middle-East and their sizeable contribution to Indian foreign exchange earnings.[2] Thus, its situation was likely to be directly related to the invasion's effects but limited to the type of events outside the realm of physical and human destruction that lend themselves to evaluation. Secondly, India had a relatively large amount of Eurobonds outstanding in a wide range of currencies, coupons and maturities. This makes econometric testing feasible and guarantees that the case of India will reflect the market in general and not a specific feature of a particular bond, currency coupon or maturity. Third, at the beginning of the period under consideration, India's structural and political difficulties were longstanding and well known. Its credit rating was still a longstanding and respectable A2. Furthermore, over the period in question India was undertaking structural reforms urged by the IMF (International Monetary Fund) that were designed to reduce default probabilities. Thus, it was in a well-known, sensitive but solid and improving position, vulnerable to the effects of war but far from desperate. Thus, if invasion effects above

and beyond those generated by overall market conditions are present, they will be reflected in the data without being contaminated by bias due to conditions generated by extreme situations. In fact, although India experienced financial distress and its official credit rating was downgraded three times, it did not default or reschedule.

We proceed in three steps:

1. We present a simple default risk model that separates changes in bond prices into two categories: those caused by changes in the risk free term structure of interest rates and those caused by changes in country specific default risk.
2. We compute the riskless term structure of interest rates and use it to calculate the price of the theoretical bond presented in the model.
3. We use regression analysis to estimate the relationship between the risky and theoretical bonds to compute the price changes in the risky bond due to changes in the riskless term structure (market risk) along with dummy variables timed to measure the effect of changes in country specific default risk.

When this methodology is applied to Indian Eurobonds 1990-1992, we find that the markets anticipated no country specific effects of the invasion on India's default risk in the quarter that the invasion took place. All the changes in Indian bond prices in this quarter were due to changes in the risk free term structure of interest rates. In subsequent quarters this assessment was revised and default risk is found to account for a fall of over 4 points in Indian bond prices. Interestingly, we find that the markets systematically mis-estimated the effect that the events set off by the invasion would eventually have on India's default risk.

The rest of the paper is organized as follows. In section II we develop the relationship between the risky and theoretical riskless bonds. Section III presents the data and methodology. Section IV presents the results and section V concludes.

II.   BOND PRICES, THE TERM STRUCTURE OF INTEREST RATES AND DEFAULT RISK

The price of a risky bond can be represented as the difference between the expected loss from default on the risky bond and the price of a theoretical bond identical in every way to the risky bond except that the theoretical bond has no default risk and its price is determined by the riskless zero coupon term structure of interest rates.[3]

Consider the following notations:

P_0 = observed price of the risky bond at time 0

T_0 = the price of theoretical bond at time 0.

$t = 1,2…n$ = payment dates where n = maturity date of the bond.

$r_t = 1$ + the riskless zero coupon rate for period t.

C_t = the cash flow for time t.

R_t = the given (constant) recovery rate at time t in the case of default as a percent of the expected value of the theoretical bond at time t.

$F_{0,t}$ = the forward price of the theoretical bond for delivery at time t.

$K_t = \sum\limits_{i=0}^{t} C_i r_i^{-i}$ = present value of the coupons paid out up to the delivery date at time t.

λ_t = the risk neutral probability of default at time t.

To simplify the exposition, we assume that default can only occur immediately before each payment date. Thus, the difference between the theoretical riskless bond and the risky bond can be written as

$$T_0 - P_0 = \sum_{t=1}^{n} \lambda_t \left[F_{0,t} - R_t F_{0,t} \right] r_t^{-t} \tag{1}$$

where $F_{0,t}$ represents the expected value of the riskless asset. The value of $F_{0,t}$ is equal to $\left[T_0 - K_t \right] r_t^t$. Substituting this value into (1) and rearranging gives

$$P_0 = \sum_{t=1}^{n} \lambda_t K_t (1 - R_t) + T_0 \left[1 - \sum_{t=1}^{n} \lambda_t (1 - R_t) \right] \tag{2}$$

Equation (2) breaks the riskless cash flows into two parts. The first term on the right hand side (RHS) of equation (2) represents the present value of the coupons that will be paid out before default occurs at each default date. The second term represents the certainty equivalent of the uncertain cash flows. In the absence of default risk, the two bonds are equivalent. This can be seen if we set the λ_t's equal to zero. It is also clear that in the absence of changes in the default probabilities (λ_t), any change in the price of the risky bond is due to changes in the term structure of interest rates. To see this, we can write the value of T_0 as

$$T_0 = C_1 r_1^{-1} + C_2 r_2^{-2} + \ldots + C_n \ r_n^{-n} \tag{3}$$

Then take the total differential of T

$$dT_0 = -1 C_1 r_1^{-1} \frac{dr_1}{r_1} - 2 C_2 r_2^{-2} \frac{dr_2}{r_2} - \ldots - n \ C_n \ r_n^{-n} \ \frac{dr_n}{r_{n_i}} \tag{4}$$

The differential of P in equation (2) is equal to

$$dP_0 = \left[1 - \sum_{t=1}^{n} \lambda_t (1 - R_t) \right] dT + \Delta\Lambda \tag{5}$$

where

$$\Delta\Lambda = \sum_{i=1}^{n}\frac{\partial P}{\partial\lambda_i}d\lambda_i \qquad\qquad (6)$$

refers to price changes due to changes in default probabilities.

Substituting from (4) for *dT* in (5) shows that in the absence of changes in the bond's default risk probabilities, changes in the risky bond's price are due to changes in the term structure of interest rates, which we can call market risk.

Changes in the default probabilities change both the present value of the coupons that will be paid out before default occurs at each default date and the present value of the certainty equivalent. It is important to note that changes in the term structure are continuous changes driven by the market whereas changes in default probabilities are discrete and relatively rare as evidenced in credit ratings and migrations from one category to another. In the regressions that follow, we will use this fact to test for changes in default probabilities.

III.   DATA AND METHODOLOGY

A.   Data Set

The market prices of Indian bonds and data for modelling the term structure were obtained from the Handbooks published by the International Securities Market Association (ISMA), formerly known as the Association of International Bond Dealers (AIBD).

Our data is quarterly and the observation period runs from June 29, 1990 to September 30, 1992, a total of 10 observations for each bond. The quarterly window was chosen, based on the timing of the invasion, as the smallest window wide enough to encompass price variations due to changes in the term structure as well as changes in default probabilities. Our sample is the subset of eight Indian bonds with varying amounts and maturities issued by public sector and quasi public sector borrowers - 3 in USD, 4 in DEM and 1 in JPY - that remained outstanding over the entire observation period. We also considered the Fung and Rudd (1986) argument that the time period should be not be too close to the issue date of any bond, since these prices often mirror issue costs along with interest-rate driven price movements. There were no direct sovereign issues made but all the above issuers were under the control, management and ownership of the Government of India and were guaranteed by it. Apart from ONGC, they are all financial institutions. The details of these bonds are given in Table 1.

To estimate the riskless term structure in the unregulated, tax-free Eurobond market of 1990-1992, we constructed sample sets for each observation date of not less than 50 bonds[4] issued by officially backed supranationals, for each of the three currencies constituting India's external debt - US dollars, German marks, and Japanese yen.[5] We use the supranationals to estimate the international riskless term structure rather than the corresponding treasuries in order to avoid biases that can creep into national credit markets through taxes, regulations, government intervention and the like. The supranationals included in our sample are guaranteed at least *de facto* by their member governments and

borrow at terms equivalent to, and at times better than, the treasuries of the currencies in question. Thus, they are effectively riskless and give the best picture of the international riskless term structure of interest rates. The large number of bonds in each sample was necessary to ensure the desirable asymptotic qualities of consistency and sufficiency. To ensure a balanced sample over the whole term structure, bonds in equal numbers were chosen with term left to maturity of less than three years, between three and six years, and over six years.

Table 1
Bonds issued in the Euromarket by India (1980-92)

Name	Date of Issue	Currency	Amount	Maturity	Coupon
Industrial Development Bank of India	6/1989	dollar	100 million	6/6/1996	10%
Oil and Natural Gas Commission	12/1988	dollar	125 million	16/11/1993	9.75%
Oil and Natural Gas Commission	3/1990	dollar	125 million	16/03/1997	10%
State Bank of India	6/1988	yen	15 billion	21/06/1993	5.25%
Industrial Development Bank of India	3/1987	DM	200 million	21/12/1994	6.375%
Industrial Development Bank of India	9/1988	DM	250 million	1/9/1995	6.625%
Industrial Development Bank of India	2/1986	DM	100 million	1/2/1993	7%
Oil and Natural Gas Commission	2/1987	DM	150 million	25/02/1994	6.375%

B.   The Methodology

We proceed in three steps.

In step 1, we estimate the riskless term structure for each time period, developed from McCulloch's cubic spline methodology, on the cross section of supranational bonds in the USD market, the JPY market and the DEM market.[6] This gives us three time series for the riskless term structure, one in USD, one in DEM, and one in JPY.

In the estimation of the riskless yield curve, we used two spline[7] knot points of three and six years. The choice of these two points was based on the observation that the Eurobond market typically deals in shorter maturities than their respective domestic bond markets. Thus, we reasoned that the break points for investor perceptions of uncertainty,

liquidity and risk in the Eurobond market could reasonably be represented as relatively short term: up to three years, relatively medium term: between three and six years, and relatively long term: above six years.

Bond prices are quoted clean in the Eurobond market, i.e. they are quoted free from any accrued coupon in order to facilitate yield comparisons but the actual sale is on the basis of the dirty price, i.e. the clean price cum accrued interest. Thus, we computed the dirty prices based on the number of days the bond was not held by the buyer.[8] The ask prices were used to compute the dirty prices.[9]

We used this information in ordinary least square (OLS) regressions to estimate the parameters of the cubic spline model using observed values of prices, coupons and times to maturity. These parameters were then used to compute the discount curves and the risk-free spot rate curves. The discount curve was computed for twelve years to allow comparability between data sets.

In step 2, we apply the corresponding riskless term structure for each time period to each of the Indian Eurobonds in our sample to estimate their "theoretical riskless price". This gives us eight time series, one for each bond, of the theoretical riskless price of the Indian Eurobonds in the sample.[10]

Finally in step 3, we use the "theoretical riskless prices" in the relationship with the observed risky prices developed in section 2 along with dummy variables timed to the invasion and its aftermath to test the effects of the invasion of Kuwait on Indian bond prices.

IV.   RESULTS

A.   Estimates of the Term Structure

Five parameters were estimated for the cubic spline model. The results, not reported here, of the 30 regression coefficients, i.e., the three currency markets over ten time periods, are very good. The linear coefficient is always significant and always negative for all three currencies. The results are best for the dollar. The quadratic and cubic coefficients are usually significant at the 5% level. Otherwise, except in one case, they are significant at the 10% level. The curvature coefficients are also often significant at the 5% level and usually significant at the 10% level, more so for the first knot than for the second, thereby indicating more curvature effect at the short end of the structure than the long end. For the yen, the quadratic and cubic coefficients are usually significant at the 10% level. However, the curvature coefficients are clearly significant together in only three periods: June 1990, March 1991, and December 1991. In March 1992 short term curvature is significant and in June 1992 long term curvature is significant. For the mark no parameters except the linear coefficient are significant at the beginning of the observation period. However, starting in March 1991 the quadratic coefficient becomes significant at the 5% level and the cubic at the 10% level. Except for June 1991, they stay significant until the last period, September 1992. The curvature coefficients are only significant in December 1991 and June 1992.

We used the parameters estimated above to compute the yield curves and discount functions and compared the results with those estimated with the Cox, Ingersoll and Ross

(1985) model.[11] They are almost indistinguishable, which is strong evidence for their accuracy. The shape of the yield curve showed a consistently upward sloping curve for the dollar market. The yen long-term rate fell below the short-term rate from September 1991 to March 1992, down to almost 0, in the last two-mentioned time periods. In two cases it was almost the same as the short-term rate i.e. in September 1990 and March 1991. During this period i.e. from June 1990 to March 1991, it was only marginally above the short rate. During the remaining last period however, the difference widened. The DEM yield curves showed either the interest rate at the long end to be roughly of the same magnitude as the short, or inverted. During this period, Germany was also undergoing the event of re-unification and the yield curve was inverted. In summary, the shape of the two ends of the yield curves seems satisfactory.

B.   Regression Results on Indian Bonds

We then applied the foregoing riskless term structures to each of the Indian bonds in our sample for each observation period in order to estimate their theoretical prices. In all, 80 observations were collated (10 quarters × 8 bonds) from the thirty yield curves (3 currency markets × 10 quarters).

We test for stationarity in the panel data series for P and T using the Im, Pesaran, and Shin (1995) T-bar test as applied in Wu and Chen (1999). The Z scores are -0.5297 and -0.3548, respectively.[12] The corresponding 95% critical values are ±3.09177 generated by the Monte Carlo simulations. Thus we cannot reject non-stationarity. Differencing both series once and applying the T-bar test gives Z scores of -27.58 and -21.44 for the first differences of P and T, respectively. Thus, based on the 95% critical values, we reject non-stationarity for the differenced series.

Using first differenced data, we start testing with the benchmark case where there are no changes in default probabilities and price changes are due exclusively to changes in the risk free term structure, that is, equation 5 with $\Delta \Lambda = 0$

$$\Delta P_{i_t} = a_1 + a_2 \Delta T_t + a_3 e_{t-1} + u_t \qquad (7)$$

where Δ denotes the first difference e_{t-1} is the error correction term,[13] u_t is the error term, and a_1, a_2 and a_3 are estimated coefficients. We include the error correction term because from equation (5) we can see that in its absence, the coefficient a_2 is likely to be time varying, thereby causing the regression to suffer from omitted variable bias.

To overcome problems of heteroskedasticity and cross sectional correlation arising from the data pooled over 8 bonds and 10 time periods, we used the Kmenta (1990) full cross-sectionally correlated and time-wise autogressive model.[14] We expect a_1 to be equal to zero and a_2 to be positive.[15] The results are reported in table 2. They show that a_1 is very small and not statistically different from zero with a *p*-value of 0.564.[16] On the other hand, a_2 is highly significant with a *p*-value of 0.00 and, as expected, it is positive.[17] The coefficient a_3 of the error correction term is also highly significant.

Furthermore, the overall equation is very good with an adjusted R^2 of over 55% and no evidence of autocorrelation in the residuals.

Table 2
Regression results of Equation 6 on first differences with the error correction term

	Coefficient Value	t-statistic	p-value
a_1	-0.19622	-0.5644	0.564
a_2	0.70642	8.29	0.000
a_3	-0.2974	-3.337	0.001

Adjusted R square = 0.5513

To test for changes in default probabilities, we add three dummy variables: $D1$ for the third quarter of 1990 when the invasion took place, $D2$ for the fourth quarter of 1990, the period in between the invasion of Kuwait and the Gulf War, and $D3$, the first quarter of 1991 when the Gulf War was fought. Each dummy takes the value of 1 for the quarter in question and zeros everywhere else and is designed to capture the effect of country specific changes in default probabilities on bond prices. We then test equation 5 where the dummy variables capture the term $\Delta\Lambda$:

$$\Delta P = a_1 + a_2\Delta T + a_3 e_{t-1} + b_1 D1 + b_2 D2 + b_3 D3 + u_t \qquad (8)$$

With the inclusion of the dummy variables, we have no expectations about a_1, which will capture any constant effects associated with effects that were not anticipated by the market. We expect that a_2 will be similar to a_2 in Table 2. If the markets anticipate or perceive changes in default probabilities, the coefficients b_1, b_2, and b_3 will be statistically significant. If, on the other hand, no changes were anticipated or perceived, they will not be statistically significant.

Table 3 shows the results.[18] The overall results are much improved with respect to those of the benchmark in table 2. The adjusted R^2 has increased to 0.9710. As expected, a_2 is similar to a_2 in table 2, changing by only 0.0306 but the t-statistic is much higher. The coefficient b_1 is not significant, which is evidence that no change in default probabilities was perceived or anticipated in the quarter that the invasion took place. However, coefficients b_2 and b_3 are highly significant. This is evidence that the market perceived changes in default probabilities in these two quarters. It is interesting that changes are significant in both quarters. If the full extent of the change in default probabilities had been accurately assessed in the period following the invasion, b_3 would not be significant. The significance of b_3 suggests that the market either

underestimated the full extent of the invasion's effects on bond prices in the preceding period or over-reacted and overestimated them in the following period. We attribute the increase in default probabilities to invasion effects and not to a weakening of India's intrinsic position because India's structural problems and economic inefficiencies were well known for many years and basically unchanged over the 7 months following the invasion. Thus, the consequences of the invasion, the buildup to war and the war itself seem to have fragilized India's financial position beyond all expectations. The over-reaction that this implies is reflected in the enormous loss of four grades in its credit rating in the space of less than 6 months.

Table 3
Regression results of Equation 6 on first differences with the error correction term

	Coefficient Value	t-statistic	p-value
a_1	0.31532	2.274	0.026
a_2	0.73702	36.72	0.000
a_3	-1.3681	-16.96	0.000
b_1	0.46582	1.288	0.202
b_2	-2.7901	-6.986	0.000
b_3	-1.3363	-3.436	0.001

Adjusted R square = 0.9710

V.   CONCLUSIONS

In the current international climate of conflict and confrontation, political events such as invasions and war have become important factors in the performance of international capital markets. Given the structural imbalances, social and political fragility, and financial dependence of many emerging economies, these economies may be particularly vulnerable to such events. Furthermore, whilst the events themselves are often anticipated far in advance, it is unclear how accurately the effects of these events can be assessed by the markets. We use the case of Indian Eurobonds to examine these questions with respect to Iraq's invasion of Kuwait on August $2^{nd,}$ 1990. India was chosen because it was far enough away from the immediate destruction of the war itself but its close links with Indian emigrants in the Middle East and its large dependence on oil imports made it vulnerable to events in the Middle East. India also had a wide enough range of Eurobonds outstanding to make testing feasible. Furthermore, India's pre-invasion credit rating of A2 indicated a healthy financial position such that its financial troubles subsequent to the invasion could be attributed in large part to the invasion and its aftermath.

We test a simple default risk model that uses the prices of theoretical riskless bonds in USD, DEM and JPY calculated from cubic spline estimates of the international term structure of interest rates over the period 1990-1992. We find that in the quarter that the invasion took place, the markets anticipated no country specific effects of the invasion on India's default risk. All the changes in Indian bond prices in that quarter were due to changes in the risk free term structure of interest rates. However, in the quarter following the invasion, increased default risk, reflected in a two-notch downgrade of India's credit rating, caused a fall of nearly 3 points in Indian Eurobond prices. A further increase in perceived default probabilities, reflected in a further two notch downgrade of India's credit rating, caused a further fall of 1.34 points in Indian bond prices in the succeeding period. This suggests that effects were either underestimated in the preceding period or over-estimated in the succeeding period. Over the entire seven-month period following the invasion, India's credit rating fell by four grades, a huge amount. This is strong evidence of India's extreme vulnerability to the invasion's effects and suggests that over-estimation was present. The lagged reaction of the market to invasion effects on India's default probabilities is strong evidence that the markets were unable to effectively assess this vulnerability in a timely manner.

NOTES

1. Questions such as these resemble those in the literature on contagion that looks at things such as the transmission of a crisis from one country to another that is unwarranted by the fundamentals (Eichengreen et al., 1996) and excess co-movement of credit spreads (Doukas, 1989) or returns across countries (Valdes, 1997).
2. See,various annual issues of Economic Survey published by Govt. of India during this period.
3. See, for example, Jarrow and Turnbull (1995), Madan and Unal (1998), and Duffie and Singleton (1999).
4. Most studies like Brown and Dybvig (1986) use the same dataset. However, even when the currency market was the same, in this study, the data set was varied to enable the use of very short bonds. This was to prevent the underestimation of the very short end of the yield curve to the extent of the time between June 1990 and September 1992. The June 1990 observations would have had to otherwise include observations at least 27 months away from maturity.
5. These issuers include the World Bank, Eurofima, the European Investment Bank, the African Development Bank, the Asian Development Bank etc.,
6. Although several models such as Carleton and Cooper (1976), Schaefer (1981), Vasichek and Fong (1982), Chambers, Carleton and Waldman (1984), Mastronikola (1991) exist to estimate the term structure, Shea (1985) compares them and finds McCulloch's (1971, 1975a 1975b) cubic spline model empirically tractable, easily computable by OLS and parsimonious. Furthermore, Litzenberger and Rolfo (1984), Luther and Matatko (1992), Deacon and Derry (1994 a and b)

and Bradley (1991) have successfully applied this model in their empirical studies. In this study we use the McCulloch cubic spline

7. A spline is a model which incorporates switching coefficients of regression in two or more periods of time. To make this a smooth transition and to estimate it, it is essential that two regression lines meet at a switching point (knot) in a manner that in the example of a cubic spline satisfies the following:

$$Y_t = \alpha_1 + \beta_1 X_t + \gamma_1 X_t^2 + d_1 X_t^3 + e_t \quad (t = 1, 2 ... , t^*)$$
$$Y_t = \alpha_2 + \beta_2 X_t + \gamma_2 X_t^2 + d_2 X_t^3 + e_t \quad (t = t^* + 1, t^* + 2 ... , n)$$

The requirement is that at point $t = t^*$ the first and second derivative of these curves be the same.

8. Unlike more imprecise methods like Chambers et al (1984) who computed interest accrued to the nearest quarter, dirty prices used in this study were precise to the day.

9   There were two other choices in this matter :
    1)Bid prices could have been used on the argument that they are prices the market makers are ready to buy at, or
    2)The midpoint of the two, i.e. the mean of the bid and ask price could have been used. However it was found that the bid-ask spread was very narrow in the market. In view of this we found it reasonable to calculate the prices on the basis of ask quotations along with the accrued coupon.

10. We also estimate the yield curve using the Cox, Ingersoll and Ross (1985) model, which gives similar results. These results are not reported here but are available on request.

11. Results available on request.

12. Details of the simulations are available on request.

13. The error correction term is the error term e_t in the regression $P_t = c_1 + c_2 T_t + e_t$.

14. We found that the Kmenta model worked best with no correction for autocorrelation.

15. See Equation 5.

16. Tests for bond specific fixed effects were also negative at conventional levels of significance

17. Tests for bond specific slope effects confirm that the slopes for the individual bonds are all positive at conventional levels of significance.

18. We also controlled the dummy variables for currency specific and maturity specific effects and found that none were present.

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