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Bond Yield Spreads and Country Risk: A Lasting Relationship?

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I. Introduction

Country risk draws attention in waves resulting from storms that batter the international financial markets. For example, the Latin American debt crisis of the early 1980s boosted research in this direction. This also occurred after the Mexico crisis of 1994/95 and the Asia crisis of 1997. As a result, the economic literature has come up with various methods to analyze country risk (see Saini and Bates, 1984). One of them is the analysis of country risk by using interest rate spreads. In primary market analysis, the spread is determined in the market where loans are issued and where loan terms are determined (e.g. see Easton and Rockerbie, 1999). In analyzing country risk by using bond prices or bond yields, we get secondary market analysis, as the bond prices and yields are determined in the secondary market where bond issues are traded. Boehmer and Megginson (1990) go into the pricing of developing country loans during 1985–1988. From their research, it appears the most significant variables affecting LDC loan sale prices were a country's debt service ratios, its import ratio, its accumulated debt arrears, and the amount by which banks had already made loan provisions against these loans. Prices of bank loans, bond prices and bond yields have substantial informative value in determining country risk as they may render a sensitive reflection of expected debt payments (Dropsy and Solberg, 1992). Edwards (1986) indicates that country risk does play an important role in the bond market. He finds evidence that bond yield spreads are positively associated with country risk. Stone (1990) also applies secondary market analysis. He finds that debt returns are insensitive to changes in country risk indicators. Chalal et al. (1996) use secondary market prices to examine integration between emerging and US debt and equity mar-

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kets. Their evidence suggests that the degree of integration varies with security types and the country of origin. But these differences between security types become less apparent over time.

Prior research already has focussed on the relationship between bond spreads and country risk (see Cantor and Packer, 1996; Cline and Barnes, 1997; Larraín et al., 1997; Min, 1998; Kamin and Von Kleist, 1999; Ramcharran, 1999). Cantor and Packer show there to be a clear relationship between secondary market spreads on bonds issued by emerging market countries and country risk ratings assessed by Moody's and Standard and Poor's. Larraín et al. (1997) go into the causality between the two. They find that spreads seem to explain somewhat better the level of credit ratings than vice versa. Eichengreen and Mody (1998) model both the determinants of the decisions by countries to enter the bond markets and the factors that influence the pricing of these bonds when launched. They find that the level of the interest rate spreads is higher when the maturity of the bonds increases, when the country has a high ratio of external debt to GNP, when the country has experienced debt rescheduling, when there is a high ratio of debt-service payments to exports, and when the bond is a private placement. In contrast, they find spreads are significantly lower when the country's credit rating is higher and when the size of the bond issue is larger. Mody and Eichengreen also find that most of the changes in spreads are to be associated with changes in market sentiment rather than in economic fundamentals. As such, it seems interesting to find out how stable the relationship between spreads and country risk actually is. Cline and Barnes (1997) use charts to find a shift in the relationship between bonds and spreads. Kamin and Von Kleist (1999) apply formal statistical analysis and find a shift in the relationship between spreads and ratings. They also find evidence of a subsequent reversal of this shift. Ramcharran (1999) finds that secondary market prices of LDC's debt can be used in forecasting country risk.

In this paper, we elaborate on the stability of the relationship between Eurobond yield differentials and country risk. We determine the relationship between yield spreads and country risk by calculating rank correlations for more than a dozen countries and we construct regression equations for the relationship between yield spreads and country risk. Furthermore, we investigate whether the relationship holds in time, *i.e.* the stability of the relationship. An innovation in relation to prior research is how we analyze stability and the type of data used, *i.e.* Eurobonds instead of Brady bonds. US collateral ultimately backs Brady bonds, in contrast to Eurobonds. This will affect the yield as well as the yield spread. Therefore, we regard the yield spread of US T-bonds *vis-à-vis* Eurobonds as a superior reflection of country risk than the yield spread of these T-bonds against the Brady bond which is being used in most research up to date.

The structure of this paper is as follows. Section two gives the methodology we use in secondary market analysis. The data are introduced in section three. The correlation and regression results are presented in section four. Section five goes into the stability question. Section six is a brief conclusion. Some details concerning the stability analysis are given in two appendices.

II. Methodology

In using yield spreads for country risk analysis, a basic assumption is that one country is free of country risk. For most investors, the US suits this purpose best. From this assumption we derive that the US T-bond yield is the country risk free interest rate. This seems reasonable as non-US bond yields often are regarded as the risk-free US bond yield plus an extra risk premium (see Fabozzi 1996). We proceed as follows. First, we estimate the US T-bond yield curve. Then, we estimate the US T-bond yield curve as a function of remaining life, using a semi loglinear OLS regression (see Hull, 1997). For each remaining life, a corresponding US T-bond yield can be estimated by substituting the remaining life into the yield curve function. For each point in time, we estimate the following regression specification:

$$y = \beta_0 + \beta_1(\ln x) + \beta_2(\ln x)^2 + \beta_3(\ln x)^3 + \varepsilon \quad (1)$$

where y is the US Treasury bond yield, estimated for six months periods, x is the remaining life in six months periods, and ε is the disturbance term. The β's are the coefficients to be estimated. As we want to know the US T-bond yield estimated for a remaining life period in years, x has to be multiplied by two. The next step is to calculate the yield spreads for every international bond. Then, we calculate the yield spreads for every international bond. These spreads are found by subtracting the yield of the US T-bond from the yield of the bond of the country in question. The remaining life of these two bonds has to be identical. Thirdly, using loglinear regression, we come up with a spread curve. Then, we have to estimate yield spreads, given a remaining life, which is identical for all bonds from all countries in our sample. For example, if we estimate the yield spread for Brazil for a bond with a remaining life of four years, yield spreads for the other countries must be estimated for four years too.

For each country in our sample, yield spreads are estimated by using the loglinear estimation function of the spread curve:

$$y = \beta_1(\ln x) + \varepsilon \tag{2}$$

Then, yield spreads can be estimated by substituting the remaining life for x in the spread function. Yield spreads are estimated by substituting the remaining life in the spread function in the estimated spread curve. Estimated – not calculated – yield curves are used to compare with the country ratings, as the actual, calculated, yield spreads of the countries in the sample carry different remaining lives, which hampers comparison. Note that any relationship between the yield spread and country risk rating is a stochastic one. Therefore, changes in the country risk ratings implied by changes in the yield spread can only be estimated. We will use regression analysis in section four to describe the relationship between yield spreads and country risk.

We opt for the country rating as the independent variable and for the yield spread as the dependent variable. As we have chosen the US as the risk free country, it follows that the yield spread cannot be in the negative. This requires a functional form of the regression equation where the y value is always at least 0. It turns out that with the data at hand (see section three) a loglinear estimation results in the highest R^2. The general form of the equation is:

$$\ln YS_i = \alpha + \beta \text{ Rating} + \varepsilon \tag{3}$$

where ln YS_i is the natural logarithm of the yield spread given a constant remaining life of years i. Rating is the country risk rating that is assumed to be reflected in bond yield spreads. α and β are coefficients to be estimated. And ε is a disturbance term.

III. Data

We use the *Institutional Investor* country risk rating as the country risk indicator. This rating appears twice a year. The higher the country rating, the lower the probability of restructuring or default and, therefore, the lower the country risk. This score is determined separately for each country and it includes country risk specific fundamentals. Ul Haque et al. (1996) find that the Institutional Investor country risk rating is the best reflection of country risk if compared to other country risk indicators. In contrast with the existing country risk models, there is no interdependence among independent variables as there is only one independent variable. Another reason for choosing this indicator is that it is easily available and at relatively low cost.

We estimate the yield spreads twice a year across the entire dataset in the first week of May and November. It is in these periods that the Institutional Investor inquires with their panel what views they hold on country risk. Thus, the country risk scores published in March and September are the markets' view on country risk in the first week of November and May respectively. Therefore, we compare the March and September country scores with the yield spreads of the first week of November and May. The ordinary yields to maturity of the bond are yields reported by the *International Securities Market Association* (ISMA) in their *Weekly Eurobond Guide*. Reported yields are yields calculated by using bid prices and are transformed to mid yields. The bonds chosen and included in the dataset satisfy the following selection criteria: They are Eurobonds and they are US$ denominated. We use Eurobonds instead of the more commonly used Brady bonds because, ultimately, Brady bonds carry US T-bonds as collateral. This enters into the pricing and to the yield of the Brady bonds. In contrast, no such collateral exists for the Eurobonds. The result is that the Eurobond yield spread is a more straightforward reflection of the country risk premium than the yield spread between Brady bonds and US T-bonds. The Eurobonds are issued by a sovereign or by a sovereign related authority. The bonds carry a fixed interest rate, bullet maturity, and have no further provisions such as call or put provisions. The bonds are not extremely liquid and are of recent issue (*i.e.* post-1990). They do not carry extraordinary guarantees. Given the criteria, we have bonds for 25 countries.[1]

The dataset shows that most of the bonds in our sample have a remaining life between two and five years. But across our entire dataset, yield spreads cannot be estimated for exactly the same remaining life. Therefore, extrapolation of the spread curve might be needed in some cases, making the estimated spread curve less reliable. For almost all countries, yield spreads can be estimated for remaining lives of 3 and 4 years, without relying on extrapolation of the yield curve. At each moment in time

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1 Argentina, Austria, Brazil, Belgium, Canada, Chile, Colombia, Denmark, Finland, Greece, Hungary, Iceland, Ireland, Italy, South-Korea, Lebanon, Malaysia, Mexico, New Zealand, Philippines, Spain, Sweden, Trinidad & Tobago, Turkey, and Uruguay.

and for each country, yield spreads are estimated for three different remaining lives: 3 year, 4 year, and 2 or 5 year. The choice between 2 year and 5 year depends on the extent of extrapolation needed to estimate the yield curve for this maturity.

IV. Results

Here, we analyze the relation between Eurobond yield spreads and country risk on the basis of the methodology explained in section II. We present the rank correlations between the yield spreads and the country credit ratings. The higher the rank correlation, the better yield spreads reflect country risk and the more likely it is that yield spreads are useful to analyze country risk. We use regression analysis to estimate the relation between country risk and yield spreads.

Rank Correlation

When there really is a relation between yield spreads and country risk, we expect that higher yield spreads will correspond to higher country risk (see Fabozzi, 1996). Thus, we ask ourselves: do lower Institutional Investor scores (*i.e.* higher country risk) correspond with higher bond yield spreads?

The rank correlation is calculated as the ordinary correlation between two rankings on the basis of our data from the Institutional Investor for country risk and from the ISMA Eurobond Guide for yield spreads. The statistic used to calculate a rank correlation is the Spearman rank correlation (R_s). For four years, three Spearman rank correlations are calculated twice a year, as this is the broadest possible range of integer remaining lives for which the yield spreads can be estimated without having to rely on extrapolation of the yield spread curves. The rank correlations are calculated for May and November during 1993–1996. The results for all countries (see footnote 1) are in table 1. Note that this group is not always the same as some countries only had a country risk rating and/or a Eurobond issued during part of the period under review. Table 2 gives the results for a group of countries that is exactly the same in the whole period under review.[2]

[2] *I.e.* Argentina, Austria, Belgium, Brazil, Canada, Finland, Italy, Ireland, Mexico, Sweden, Venezuela; only in 1996 Ireland and Venezuela are excluded because the bonds of those countries did not meet the required specifications. The net effect of leaving those two countries out in 1996 will be ignored.

Table 1

Rank Correlations of Yield Spreads and Country Risk Scores, all Countries

	2 year	3 year	4 year	5 year
May 1993	$n = 14$ $R_s = 0.9253$	$n = 14$ $R_s = 0.9209$	$n = 14$ $R_s = 0.9165$	
November 1993		$n = 14$ $R_s = 0.9077$	$n = 15$ $R_s = 0.9071$	$n = 14$ $R_s = 0.9385$
May 1994	$n = 14$ $R_s = 0.8857$	$n = 14$ $R_s = 0.8725$	$n = 14$ $R_s = 0.8901$	
November 1994	$n = 16$ $R_s = 0.8706$	$n = 16$ $R_s = 0.9206$	$n = 16$ $R_s = 0.9382$	
May 1995	$n = 18$ $R_s = 0.8968$	$n = 18$ $R_s = 0.9092$	$n = 18$ $R_s = 0.8968$	
November 1995		$n = 19$ $R_s = 0.8877$	$n = 19$ $R_s = 0.8860$	$n = 19$ $R_s = 0.8860$
May 1996		$n = 18$ $R_s = 0.8803$	$n = 18$ $R_s = 0.8927$	$n = 18$ $R_s = 0.8968$
November 1996		$n = 23$ $R_s = 0.8765$	$n = 23$ $R_s = 0.8854$	$n = 23$ $R_s = 0.8923$

n = number of countries

Table 2

Rank Correlations of Yield Spreads and Country Risk Scores for Eleven Countries

	2 year	3 year	4 year	5 year
May 1993	$R_s = 0.9455$	$R_s = 0.8909$	$R_s = 0.9273$	
November 1993		$R_s = 0.9000$	$R_s = 0.8909$	$R_s = 0.8909$
May 1994	$R_s = 0.8545$	$R_s = 0.8455$	$R_s = 0.8273$	
November 1994	$R_s = 0.8182$	$R_s = 0.9091$	$R_s = 0.9091$	
May 1995	$R_s = 0.8636$	$R_s = 0.8909$	$R_s = 0.8091$	
November 1995		$R_s = 0.7455$	$R_s = 0.7273$	$R_s = 0.7455$
May 1996		$R_s = 0.9394$	$R_s = 0.9152$	$R_s = 0.8667$
November 1996		$R_s = 0.7152$	$R_s = 0.7030$	$R_s = 0.7636$

Table 1 and 2 provide us with the following results. First, from May 1993 up to May 1994, we have the smallest number of countries included in table 1, namely 14. The critical value for which the H_0 hypothesis of no association between yield spreads and country risk is rejected with $n = 14$ observations and a significance level of 5%, is 0.457. When R_s is larger than 0.457, the H_0 hypothesis of no relation between yield spreads and country risk can be rejected. The rank correlations range from 0.87 to 0.94 and thus significantly differ from zero at the 5%-significance level. Table 2 gives similar results, but here the rank correlations range from 0.70 to 0.95. Thus, our results clearly reject the H_0 hypothesis of no relation between yield spreads and country risk scores. They confirm strong and positive associations between yield spreads and country risk. Second, increasing the number of countries does not result in any decrease in the Spearman rank correlation coefficient. To the contrary, it appears that the correlations in table 1 for an increasing number of countries are somewhat higher than those in table 2. Third, table 1 and 2 show that country risk is independent of the remaining life of the bonds. When we calculate R_s's of different remaining lives and compare these for the same period, they appear not to differ very much.

Regression Analysis

We estimate the regressions for (1) for different moments in time; the same moments as analyzed previously in this section. Furthermore, we estimate the regressions for all countries and for the group of eleven countries. The results are in table 3 and 4.

Because all rank correlations are strongly significant (tables 1 and 2), it is not surprising that the regression results (tables 3 and 4) are strongly significant too. The estimates of α and β are significant at the 5%-level in both instances. The H_0 hypothesis, $\alpha = 0$ and $\beta = 0$, is strongly rejected by the F-statistic across the entire dataset. The R^2's for all eight moments in time and for all remaining lives are higher than 0.80 in table 3, except for one (namely for bonds with a remaining life of 2 years in November 1994, where the R^2 is 0.78). Comparing column 3 (N) and column 7 (R^2) in table 3 suggests that the estimated regressions, which include a larger number of countries, do not have lower reported R^2's. This suggests that the relation between yield spread and country risk remains robust, independent of the number of countries included in the regression estimation. Table 4 also shows that the hypothesis of α and β being simultaneously equal to zero is strongly rejected by the

Table 3

Regression Results for all Countries

Period	remaining life	N	α	t-stat. of α	β	t-stat. of β	R^2	F
May 1993	2	14	2.9848	10.46	–.0528	–11.95	.923	142.86
	3	14	2.9489	13.36	–.0492	–14.40	.945	207.54
	4	14	2.9433	14.72	–.0473	–15.29	.951	233.91
November 1993	3	14	2.5033	10.17	–.0437	–11.08	.911	122.79
	4	14	2.5668	13.06	–.0427	–13.23	.931	175.08
	5	14	2.4809	12.36	–.0403	–12.54	.929	157.33
May 1994	2	14	2.2993	6.84	–.0420	–7.80	.835	60.85
	3	14	2.4609	8.82	–.0423	–9.45	.881	89.24
	4	14	2.5526	10.10	–.0424	–10.45	.901	109.29
November 1994	2	16	2.0801	5.70	–.0415	–7.00	.779	49.03
	3	16	2.5247	9.07	–.0450	–9.96	.876	99.28
	4	16	2.7447	10.82	–.0465	–11.27	.900	126.98
May 1995	2	18	3.2449	10.21	–.0571	–10.78	.879	116.16
	3	18	3.4600	11.80	–.0579	–11.87	.898	140.91
	4	18	3.5801	12.35	–.0583	–12.08	.901	145.94
November 1995	3	19	3.2811	12.54	–.0564	–12.68	.904	160.84
	4	19	3.3309	13.01	–.0554	–12.75	.905	162.48
	5	19	3.3664	13.02	–.0548	–12.48	.902	155.81
May 1996	3	18	2.5825	8.52	–.0518	–10.03	.863	100.56
	4	18	2.6333	9.75	–.0502	–10.90	.881	118.91
	5	18	2.6559	10.00	–.0489	–10.81	.879	116.77
November 1996	3	23	2.8452	8.07	–.0564	–9.52	.812	90.59
	4	23	2.9589	10.41	–.0554	–11.55	.865	134.49
	5	23	3.0393	11.59	–.0549	–12.46	.881	155.18

Table 4

Regression Results for a Homogeneous Group of Countries

Period	remain-ing life	α	t-stat. of α	β	t-stat. of β	R^2	F
May 1993	2	2.8686	11.19	–.0501	–12.68	.947	160.90
	3	2.8429	15.98	–.0467	–17.05	.970	290.56
	4	2.8403	18.53	–.0449	–19.03	.976	362.17
November 1993	3	2.7644	15.58	–.0476	–17.47	.971	305.35
	4	2.6927	16.48	–.0448	–17.85	.973	318.76
	5	3.7542	11.13	–.0429	–17.46	.971	304.83
May 1994	2	2.7746	8.95	–.0489	–10.27	.921	105.40
	3	2.8208	10.58	–.0476	–11.62	.938	135.03
	4	2.8487	11.45	–.0469	–12.25	.943	150.13
November 1994	2	2.3973	7.52	–.0465	–9.52	.910	90.55
	3	2.7630	10.78	–.0490	–12.48	.945	155.87
	4	2.9568	12.57	–.0501	–13.91	.956	193.41
May 1995	2	3.7325	13.74	–.0633	–15.21	.963	231.44
	3	3.9403	14.57	–.0640	–15.44	.964	238.27
	4	4.0640	14.55	–.0643	–15.02	.962	225.63
November 1995	3	3.7542	11.61	–.0627	–12.64	.947	159.84
	4	3.7537	13.24	–.0628	–15.58	.972	242.69
	5	3.7542	11.13	–.0609	–13.61	.964	185.12
May 1996	3	3.6398	14.04	–.0657	–17.06	.977	291.37
	4	3.5944	13.24	–.0628	–15.58	.972	242.69
	5	3.5705	11.85	–.0609	–13.61	.964	185.12
November 1996	3	3.6031	8.96	–.0662	–11.22	.947	125.89
	4	3.6682	9.74	–.0652	–11.80	.952	139.17
	5	3.7177	9.81	–.0646	–11.61	.951	134.81

F-statistic. The reported R^2's are higher than 0.90 in all instances here. A White-test was applied to test for heteroskedasticity in table 3 and 4, however, no evidence with respect to heteroskedasticity was found (see appendix 1). This suggests that our least squares estimators are efficient, *i.e.* the standard errors are correct.

In comparing our regression results with the findings elsewhere in the literature (Angeloni and Short, 1980; Feder and Ross, 1982; Edwards, 1986; Eichengreen and Mody, 1998), it turns out that the relationship between yield spread and country risk is stronger than that between loan spread and country risk. This also applies to our regression results for the relation between yield spread and country risk. As such, it appears that yield spreads are a better reflection of country risk than loan spreads. Thus, the relation between yield spreads and country risk in this paper is much stronger than that found elsewhere in the literature. Our homogeneous dataset of Eurobonds, the coming of age of the international bond market for developing country debt in the period under review, and the use of Institutional Investor country risk scores instead of macroeconomic country risk variables, probably are the main reasons behind this observation.

V. Stability

Stability is of crucial importance when interpreting the relationship between country credit risk ratings and bond yield spreads. Of course, the estimated regressions and the calculated confidence intervals only are reliable for the period for which the regressions are estimated. Country risk analysts must be aware that analyzing country risk after November 1996 by using the regression results of November 1996 is the same as out-of-sample estimation, which can be unreliable. To avoid the unreliability that might stem from out-of-sample estimation, country risk analysts would have to wait until March 1997 to perform their country analysis along the lines shown in this paper. Furthermore, as the yield spreads of November 1996 are used, they can only analyze country risk of November 1996. Therefore, it seems worthwhile to analyze the stability of the estimates of the α's and the β's during a couple of years. If the α's and the β's would turn out to be stable during a certain period of time, a pooled regression estimation can be used. As such, more observations can be included, which increases the reliability of the regression results and leads to a narrower confidence interval.

We choose to test stability for the group of eleven countries for which we have observations throughout the entire period under review (see footnote 2) and for bonds with a remaining life of 3 and 4 years. As such, we have a homogeneous group of countries and we do not need to rely on extrapolation of the spread curve. Using the regression results of table 4, we investigate whether the estimate of both the α's and the β's differ between yield spread regressions with remaining life of 3 and 4 years. If so, the outcome of the stability tests also might differ for 3 and 4-year yield spreads. We create subsamples of observations for the 3 and 4-year yield spreads in each period of observation. It is verified whether the estimated coefficients of equation (3) significantly differ for the 3 and 4-year yield spread estimations. Since the dataset reveals that most spread curves are upward sloping as a function of remaining life, it is expected that the α's increase with remaining life. Furthermore, it is verified whether the β's differ for the two groups of remaining life. If both the α's and the β's do not significantly differ between the 3 and 4 year yield spread regressions, it doesn't matter which of the two is chosen to carry out the stability tests. A dummy is used for both the intercept and the slope coefficient. The following regression equation is estimated:

$$(4) \qquad \ln YS = \alpha_1 + (\alpha_2 - \alpha_1)D1c + \beta_1 \text{ Rating} + (\beta_2 - \beta_1)D2s + \varepsilon$$

$D1c = 0$ for all 3 year yield spreads
$D1c = 1$ for all 4 year yield spreads
$D2s = 0$ for all 3 year yield spreads
$D2s =$ Rating_2 for the respective values of country ratings for all 4 year yield spreads.

Subscript 1 refers to the first subset, *i.e.* the 3 year yield spread in each observation period; subscript 2 refers to the second subset: 4 year yield spreads. The '*c*' refers to the intercept dummy, the '*s*' to the slope dummy. Furthermore, in (2) we have $YS = YS_2$, if $D2s = \text{Rating}_2$ and $D1c = 1$. In this case, Rating_2 has to be substituted for Rating in the third term on the right hand side of the equation. If both dummies are zero, $YS = YS_1$ and $\text{Rating} = \text{Rating}_1$. Then, we get the estimated coefficients of the 3-year yield spreads; these already were reported in table 4. The group for which the regression is estimated, setting both dummies equal to zero, is called the reference group. The results of the estimates of regression (2) are in table 5.

Table 5 shows that the p-values of the dummies are much larger than 0.05. The null hypothesis of all dummies, each separately, being equal to zero cannot be rejected. All dummies are insignificant. Applying a White-test for heteroskedasticity shows that no heteroskedasticity was found. As α and β are not statistically different for 3 and 4-year yield

Table 5

Differences in α and β; 3 Year Versus 4 Year Yield Spreads

Period	estimate of $D1$ ($= \alpha_2 - \alpha_1$)	estimate of $D2$ ($= \beta_2 - \beta_1$)	t-stat. of $D1$	t-stat. of $D2$	p-value $D1c$	p-value $D2s$
May 1993	–.0026	.0018	–.01	.49	.9912	.6299
November 1993	–.0718	.0028	–.30	.76	.7694	.4549
May 1994	.0279	–.0008	.08	.13	.9400	.8942
November 1994	.1938	–.0011	.56	–.21	.5843	.8367
May 1995	.1237	–.0003	.32	–.06	.7540	.9548
November 1995	–.0006	.0018	–.001	.25	.9990	.8050
May 1996	–.0455	.0029	.53	–.12	.9053	.6070
November 1996	.0317	.0022	.06	.28	.9533	.7833

spread regressions, it doesn't make a difference whether the stability tests are carried out for 3 or for 4-year yield spreads. We test for 3-year yield spreads. Whether the estimated α and β are stable in time is tested by Chow breakpoint estimation and by the dummy variable technique. The Chow breakpoint estimation tests for stability of both intercept and slope parameters between two or more populations. A disadvantage is that it might reject the hypothesis of stability but not reveal which particular coefficients are unstable. Therefore, we also employ the dummy variable technique.

We analyzed three different subperiods with respect to the stability of the estimated coefficients α and β. Appendix 2 describes how we proceeded in this respect. We found that the estimates of α and β are constant during 1993 and 1994. However, the parameter estimates seem to differ in the 1993/94 subset from the estimates of the 1995/96 subset. But during 1995 and 1996, the parameter estimates appear to be constant once again. Thus, we have a breakpoint in our dataset. This indicates that bond investors' attitude towards country risk significantly changed in the period under review. Especially, it appears to be the Mexico-crisis of late 1994 that must be held responsible for this switch. This implies that the use of secondary bond market data, more specifically the use of bond yield spreads, is limited and warrants caution in the case of country risk analysis and country risk forecasting. This contrasts with earlier findings of, e.g., Ramcharran (1999).

VI. Conclusion

Secondary bond market analysis is useful to analyze country risk as bond yield spreads tend to reveal country risk. Secondary market analysis might be superior to primary market analysis from a theoretical perspective as it continuously reflects the changes in perceptions and expectations of bond traders and investors. Primary market analysis, in contrast, is bound to one single moment in time in this respect, namely the moment of the issue of the bond. This paper attempts to establish a link between country risk and bond yield spreads, and to test for the stability of this link. We use data from the Institutional Investor and the ISMA Eurobond Guide for the 1993–1996 period.

Compared with the secondary market analyses of Edwards (1986), Stone (1990), and Dropsy and Solberg (1992), our results suggest a much stronger relation between bond yield spreads and country risk. The high rank correlations between the bond yield spreads *vis-à-vis* US T-bonds and the country ratings indicate that bond yield spreads may be a better reflection of country risk than loan spreads in the secondary market. Our more homogeneous dataset of Eurobonds, the coming of age of the international bond market for developing country debt, and the use of country risk ratings instead of macroeconomic country risk variables probably are the main reasons behind the fact that the relation between yield spreads and country risk in this paper is stronger than that found elsewhere in the literature. As such, secondary market analysis appears to be a promising additional tool in country risk analysis. However, stability could not be found for the complete dataset (May 1993–November 1996), as there is a breakpoint with the Mexico crisis of December 1994. After this crisis, it seems that country risk was perceived quite different. Investors began to attach more weight to country risk than before when trading in bonds, as is evidenced by the significant rise in bond yield spreads. In this respect, our results confirm the findings of Eichengreen and Mody (1998), who find that changes in market sentiment, even more than those in fundamentals, can result in regime shifts that translate into substantial shifts of the spread. This outcome is not supporting the reliable use of yield spreads as indicators of country risk.

In all, the methodology developed in this paper appears to be valuable in analyzing the behavior of investors in international financial markets. Our findings also show that country risk analysts must be very careful in applying this methodology and in deriving policy consequences from it because of shifts in the risk perception that may inflict upon the stability of the relationship between bond yield spreads and country risk ratings.

Appendix 1

We use the White-test to check for heteroskedasticity for the two country groups. Table A-1 gives, as an example, for each remaining life the checks of two regressions, namely the regressions for the first and the last observation period. With a significance level of 5%, we find high p-values. This suggests that the null hypothesis of homoskedasticity cannot be rejected. As there is no evidence of heteroskedasticity, no correction for heteroskedasticity is needed. Although not reported here, all other regression results in table 3 and table 4 in the main text lack evidence of heteroskedasticity too.

Appendix 2

We investigate stability for the estimated coefficients α and β for: 1st 1993 versus 1994; 2nd 1993/94 versus 1995/96; 3rd 1995 versus 1996. The dummy variable technique is used to detect which α and β of the different subsets differ from their May 1993 levels. A dummy was included for both the intercept term (α) and the slope coefficient (β). The dummy for the intercept is denoted with *Dc* and the dummy for the slope coefficient with *Ds*. Based on the dummy test, it can be shown whether the α and β estimates of 1993 and 1994 are stable, whether the α and β estimates between 1993/94 and 1995/96 significantly differ, and whether the α and β estimates of 1995 and 1996 are stable. Furthermore, we apply a Wald test, as all estimated dummies must be simultaneously equal to zero for stability to be proven during a certain time period.

Table A-1

Results of the Heteroskedasticity Tests

Remaining life	Period	F-statistic	p-value
All countries			
2 years	May 1993	0.78	0.48
2 years	May 1995	1.08	0.37
3 years	May 1993	0.46	0.65
3 years	November 1996	1.02	0.38
4 years	May 1993	0.31	0.74
4 years	November 1996	0.99	0.39
5 years	May 1993	1.60	0.25
5 years	November 1996	0.77	0.48
11 countries			
2 years	May 1993	1.37	0.31
2 years	May 1995	0.11	0.90
3 years	May 1993	0.38	0.70
3 years	November 1996	1.01	0.42
4 years	May 1993	0.05	0.96
4 years	November 1996	0.79	0.50
5 years	May 1993	0.04	0.96
5 years	November 1996	0.48	0.64

1. 1993 Versus 1994

Using the dummy variable technique, we estimate the following regression specification:

$$\begin{aligned}
\ln YS = {} & \alpha_1 + (\alpha_2-\alpha_1)D1c + (\alpha_3-\alpha_1)D2c + (\alpha_4-\alpha_1)D3c + \\
& (\alpha_5-\alpha_1)D4c + (\alpha_6-\alpha_1)D5c + (\alpha_7-\alpha_1)D6c + (\alpha_8-\alpha_1)D7c + \\
& + \beta_1 \text{ Rating} + (\beta_2-\beta_1)D1s + (\beta_3-\beta_1)D2s + (\beta_4-\beta_1)D3s + \\
& (\beta_5-\beta_1)D4s + (\beta_6-\beta_1)D5s + (\beta_7-\beta_1)D6s + (\beta_8-\beta_1)D7s + \varepsilon
\end{aligned} \tag{A.1}$$

$D1c = 1$ for all November 1993 observations and = 0 otherwise, etc.

$D1s =$ Rating$_2$ for all November 1993 observations and = 0 otherwise, etc. subscript 1 refers to the 1st group, *i.e.* all 3 year yield spreads of May 1993, etc.

Furthermore, in (A.1) we have $YS = YS_2$ if $D1s = \text{Rating}_2$ and $D1c = 1$. In this case, Rating_2 has to be substituted for Rating in the eighth term on the right hand side of the '=' sign. Thus, $YS = YS_n$ if $Dc(n-1) = 1$ and $Ds(n-1) = \text{Rating}_n$. If all dummies are zero, $YS = YS_1$, and Rating_1 has to be substituted for Rating in the eighth term on the right hand side of the '=' sign. In this case, the regression is estimated using only the 3-year spreads of May 1993, which acts as the reference group. The results are in table A-2. Using the White test, no evidence of heteroskedasticity is found (p-value of 0.66). Please remember that the estimated coefficient of $D1c$ is equal to $\alpha_2 - \alpha_1$; the estimated coefficient of $D1s$ is equal to $\beta_2 - \beta_1$.

Table A-2

Estimation Results of the Stability Tests 1993/94 vs. 1995/96 (Dependent Variable: ln *YS*)

independent variables	coefficient estimate	t-statistic	p-value
α	2.8429	11.55	.0000
$D1c$	–.0785	–.22	.8255
$D2c$	–.0221	–.06	.9508
$D3c$	–.0799	–.22	.8351
$D4c$	1.0974	3.10	.0028
$D5c$	.9113	2.59	.0118
$D6c$	.7969	1.99	.0499
$D7c$	.7599	1.88	.0642
β	–.0467	–12.32	.0000
$D1s$	–.0009	–.17	.8676
$D2s$	–.0009	–.17	.8659
$D3s$	–.0023	–.42	.6789
$D4s$	–.0173	–3.17	.0023
$D5s$	–.0159	–2.95	.0044
$D6s$	–.0190	-3.16	.0023
$D7s$	–.0195	-3.23	.0019
R^2	.959		
F-statistic	106.15		
N	84		

Table A-2 suggests that the estimates of α and β are constant during 1993 and 1994, since both the intercept and slope dummies of 1993 and 1994 are insignificant ($D1c$, $D1s$, $D2c$, $D2s$, $D3c$, $D3s$), as can be concluded from the high p-values or the low t-statistics. To prove the stability of α and β in 1993 and 1994, we apply a Wald test. The corresponding null hypothesis is: $D1c = D1s = D2c = D2s = D3c = D3s = 0$. We found an F-statistic of 0.77 with probability 0.5997. This high p-value suggests that the null hypothesis cannot be rejected; the estimated dummies of 1993 and 1994 do not significantly differ from zero. This indicates that the estimated coefficients of α and β of November 1993, May 1994, and November 1994 do not significantly differ from the reference group May 1993. A constant α and β allow us to perform a pooled regression, including all subsets of 1993 and 1994 (t-statistics in parentheses):

$$\ln YS = \underset{(25.20)}{2.7989} \; \underset{(-27.98)}{-0.0478} \text{ Rating} \tag{A.2}$$

$R^2 = 0.949$; F-stat. $= 783.16$; $n = 44$.

No evidence of heteroskedasticity is found (p-value of 0.14). When we compare the pooled regression with the separate regressions (table 4 in the main text), it turns out that the t-statistics are much higher for the pooled regression. This implies smaller standard errors than for the separate regressions. Furthermore, the higher F-value in table A-2 indicates that the null hypothesis of both α and β being simultaneously equal to zero is even stronger rejected than for the separate regressions. The pooled regression results are more significant, more reliable, than those of the separate regressions, probably due to the fact that more observations are included.

2. 1993/94 Versus 1995/96

We divide the observations in two subsets (subset 1: 1993 + 1994, and subset 2: 1995 + 1996). The H_0 hypothesis corresponding to the Chow breakpoint estimation is:

$$\alpha_{\text{subset1}} = \alpha_{\text{subset2}}, \beta_{\text{subset1}} = \beta_{\text{subset2}}.$$

We find a F-value of 16.55 with probability 0.0000. The probability of incorrectly rejecting the H_0 hypothesis is very small (0.000001). The F-statistic is large enough to reject the H_0 hypothesis of constant coefficients across the dataset. Thus, the parameter estimate indeed seems to differ in the 1995/96 subset from the parameter estimates of the 1993/94 subset. No evidence of heteroskedasticity was found (p-value 0.17). We also use the dummy test to investigate stability. As such, we look at the estimated regression coefficients of the dummies of 1995 and 1996 in table A-2. Except for $D7$, all intercept and slope dummies are significant. These results too suggest that for the 1995/96 subset, the estimates of both α and β deviate from their 1993/94 level. We apply a Wald test to investigate the instability of α and β in 1995/96 compared to 1993/94. The corresponding null hypothesis is: $D4c = D4s = D5c = D5s = D6c = D6s = D7c = D7s = 0$. We find an F-statistic of 6.11 and a p-value of 0.0000. Therefore, the null hypothesis must be rejected: the dummies of 1995/96 simultaneously differ from zero. This indicates that the bond market has changed its attitude towards country risk! A very likely reason for this change lies in the Mexico crisis of December 1994 (see IMF, 1997).

3. 1995 Versus 1996

Here, we proceed in a similar manner to that in the case of *1993 versus 1994*. Of course, now the regression is estimated using only 3-year yield spreads of May 1995. This subsample is the reference group. Table A-3 gives the results.

The high p-values (and the low t-statistics) reveal that the null hypothesis of each dummy separately being equal to zero cannot be rejected. This suggests that the dummies of 1995 and 1996 do not statistically differ from the reference group (May 1995). No evidence of heteroskedasticity was found (p-value 0.76). A Wald test gave the F-statistic of 3.29 and a p-value of 0.0124. Therefore, the null hypothesis is rejected. This contrasts with the result found in the dummy test. Multicollinearity is a possible explanation here. We apply the Chow breakpoint estimation to detect differences in both α and β between 1995 and 1996. All observations in 1995 belong to subset 1; all observations in 1996 belong to subset 2. The H_0 hypothesis corresponding to the Chow breakpoint estimation is: $\alpha_{\text{subset1}} = \alpha_{\text{ubset2}}, \beta_{\text{subset1}} = \beta_{\text{subset2}}$.

We found the F-statistic 10.20 and a p-value of 0.0003. The probability of incorrectly rejecting the null hypothesis is very small (0.000310%). The F-statistic is large. The null hypothesis is rejected. This means that the α and β of the two 1995 subsets differ from the α and β of the two 1996 subsets. Next, we test whether the estimates of α and β are constant in 1995 and 1996. The results for 1995 are in table A-4.

Table A-3

Estimation Results of the Stability Tests 1995 Versus 1996 (Dependent Variable: ln *YS*)

Independent variables	coefficient estimate	t-statistic	p-value
α	3.9403	13.70	.0000
$D1c$	–.1860	–.46	.6489
$D2c$	–.3004	–.66	.5156
$D3c$	–.3375	–.73	.4705
β	–.0640	–14.51	.0000
$D1s$	.0013	.22	.8329
$D2s$	–.0017	–.25	.8016
$D3s$	–.0022	–.32	.7480
R^2	.960		
F-statistic	110.14		
N	40		

Table A-4

Estimation Results of the Stability Tests 1995 (Dependent Variable: ln *YS*)

Independent variables	coefficient estimate	t-statistic	p-value
α	3.9403	13.14	.0000
$D1c$	–.1860	–.44	.6645
β	–.0640	13.92	.0000
$D1s$	.0013	.20	.8406
R^2	.955		
F-statistic	128.04		
N	22		

We conclude that both dummies are insignificant; no evidence of heteroskedasticity is found (p-value 0.86). To prove that α and β are constant during 1995, a Wald test is applied (null hypothesis: $D1c = D1s = 0$). We found the F-statistic 0.35 and the p-value 0.7105. Therefore, the null hypothesis cannot be rejected, meaning that α and β are constant during 1995. This allows for a pooled regression, which results in:

$$\ln YS = \underset{(18.86)}{3.8467} - \underset{(-20.25)}{0.0633}\ \text{Rating} \tag{A.3}$$

$R^2 = 0.954$; F-stat. $= 410.14$; $n = 22$.

The p-values of the estimated parameters are close to zero, *i.e.* both coefficients are strongly significant. Again, no evidence of heteroskedasticity was found (p-value 0.92). Testing for stability in the α and β for 1996 by using the dummy variable technique leads to the results presented in table A-5.

We conclude that both dummies are insignificant; no evidence of heteroskedasticity is found (p-value 0.37). To prove that the α and β are constant during 1996, a Wald test is applied (null hypothesis: $D1c = D1s = 0$). We found the F-statistic 0.15 and the p-value 0.8636. Therefore, the null hypothesis cannot be rejected, meaning that α and β are constant during 1996. This allows for a pooled regression, which results in:

$$\ln YS = \underset{(16.05)}{3.6239} - \underset{(-19.81)}{0.0660}\ \text{Rating} \tag{A.4}$$

$R^2 = 0.961$; F-stat. $= 392.26$; $n = 18$.

Table A-5

Estimation Results of the Stability Tests 1996 (Dependent Variable: ln *YS*)

Independent variables	coefficient estimate	t-statistic	p-value
α	3.6398	10.87	.0000
$D1c$	–.0371	–.08	.9393
β	–.0657	–13.22	.0000
$D1s$	–.0005	–.07	.9450
R^2	.962		
F-statistic	116.93		
N	18		

The p-values of the estimated parameters are close to zero, *i.e.* both coefficients are strongly significant. Again, no evidence of heteroskedasticity was found (p-value 0.17). In all, it turns out that α and β are constant during both 1995 and 1996.

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Summary

Bond Yield Spreads and Country Risk: A Lasting Relationship?

This paper investigates whether bond yield spreads reflect country risk. As bond prices and bond yields are determined in the secondary market, bond yields and their spread *vis-à-vis* US Treasury bonds might provide a continuous and a more reliable information base than traditional measures of country risk. We show that there is a strong relation between changes in the bond yield spread and changes in country risk. Secondary market analysis appears to be a promising additional tool in country risk analysis. However, an important drawback of the bond yield spread as an indicator of country risk appears to be the limited stability of the relationship through time. (JEL F21, F34, G14, G15)

Zusammenfassung

Renditespannen bei Wertpapieren und Länderrisiko: Eine dauerhafte Beziehung?

In diesem Beitrag wird untersucht, ob sich bei Wertpapieren die Renditespannen für eine Darstellung des Länderrisikos eignen. Da Wertpapierkurse und -renditen auf dem Sekundärmarkt bestimmt werden, könnten sich im Vergleich zu US-Treasury-Bonds die Wertpapierrenditen und ihre Spannen für das Länderrisiko als eine dauerhafte und als eine zuverlässigere Informationsgrundlage erweisen als die traditionellen Maßstäbe. Wir zeigen, daß bei Wertpapieren eine starke Beziehung zwischen sich verändernden Renditespannen und Veränderungen des Länderrisikos besteht. Analysen der Sekundärmärkte scheinen ein vielversprechendes zusätzliches Werkzeug für die Länderrisikoanalyse zu sein. Jedoch scheint in der zeitlich begrenzten Stabilität der Beziehung ein wichtiger Nachteil der Wertpapierrenditespanne als Indikator des Länderrisikos zu liegen.

Résumé

Marges de rendement des obligations et risque-pays: une relation durable?

Cet article examine si les marges de rendement des obligations reflètent le risque-pays. Etant donné que les prix des obligations et le rendement des obligations sont déterminés sur le marché secondaire, le rendement des obligations et leur marge vis-à-vis des obligations de la trésorerie américaine pourraient livrer une base d'informations permanente et plus fiable que les mesures traditionnelles du risque-pays. L'auteur montre qu'il y a une forte relation entre les changements des marges de rendement des obligations et les changements du risque-pays. L'analyse du marché secondaire apparaît comme un instrument supplémentaire prometteur pour analyser le risque-pays. Cependant, un obstacle important à utiliser la marge de rendement des obligations comme indicateur du risque-pays semble être la stabilité limitée de la relation dans le temps.