INTERNATIONAL JOURNAL OF FINANCE AND ECONOMICS
Int. J. Fin. Econ. 6: 115–126 (2001)
DOI: 10.1002/ijfe.152

MARKET BASED DEBT REDUCTION AGREEMENTS: A CASE STUDY ON MEXICAN AND POLISH BRADY BONDS

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ABSTRACT

This paper analyzes some aspects of the workings of the Brady bond (restructured Less Developed Countries debt) market. It concentrates on the effects of the December 1994 Mexican crisis on the risk assessment (as measured by the *stripped spread*) of Poland, another Brady country. The main findings are: (i) over the sample period, the unit root hypothesis on the risk premium (measured by the stripped spread) of Mexico and Poland cannot be rejected; this is consistent with the idea that the risk premium reflects new information accruing to the market; (ii) comovements in stripped spreads between Mexico and Poland were stronger during the period of the Peso crisis: we do not reject the null of cointegration for the year that includes the crisis (July 1994–July 1995), but we do reject the null for the year starting 6 months after the crisis (July 1995–July 1996); (iii) the crisis has had a strong permanent effect on the risk assessment of Mexico with respect to the one of Poland (550 basis points circa). Copyright © 2001 John Wiley & Sons, Ltd.

KEY WORDS: Brady bonds; comovements; Third World debt; risk premium

JEL CODE: F34; G14

1. BACKGROUND

In the autumn of 1982, the Mexican government announced that Mexico was no longer able to fulfill its external commitments. A run out of developing country debt ensued. It was the first major default of a sovereign country after the crisis of the 1930s, and the event was unexpected, despite the high real interest rate and falling commodity prices obtaining at the time. Most of the debt was held by only several hundred commercial banks.

One of the main inefficiencies associated with the negotiation process that followed was that the 'rules of the game' were neither clear nor predictable. Different strategies were proposed in the 1980s, ranging from the 'involuntary lending' of the early 1980s to the Brady agreements of the 1990s. One important milepost in the debt crisis was the introduction of a secondary market in 1985. As most claims were held by a relatively small group of banks, the market was initially thin. The process of securitization of the developing countries debt, though, accelerated in the last few years.

The Brady agreements are based on the exchange, on a voluntary basis, of commercial debt for a menu of options including bonds or new money—the bond option being the most popular one. Brady agreements have involved debt reductions, and have usually been implemented when the debtor country was already undergoing economic reforms. Debt reductions and economic reforms have relaxed the credit constraints that were binding for most of the debtors in the 1980s, thereby providing the new market instruments with a relevant degree of liquidity. Therefore, the *Brady bonds'* market developed very quickly in the past few years.[1]

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The impressive development of the secondary market largely resulted from the new approach introduced by the Brady plan. Brady agreements have aimed at, and actually led to, a securitization of the external debt, to a relaxation of credit constraints, and to a restored access to the international financial market. This approach has been consistent with the view that international financial market can be a fair and relatively certain environment to satisfy the financial needs of middle-income countries; probably a better environment than the lasting 'bargaining process' that prevailed in the 1980s.

On 20 December 1994, the Mexican peso was devalued by 15%, and 2 days later it was allowed to float. At the end of December, the Mexican government was forced to cancel a Tesobono refinancing auction because investors were no longer willing to bear the sovereign risk at rates acceptable to the Mexican authorities.[2] The perception of sovereign risk—as embedded, for example, in the Brady bonds prices—jumped dramatically during the crisis. Although the Mexican crisis reflected mainly internal problems, it immediately affected the perception of the sovereign risk in other debtor countries. In particular, the yield on most emerging market bonds jumped alongside that of Mexico.

In 1981, the choice not to differentiate among different debtors might have been a rational response to the crisis, considering the political and international dimension that the crisis soon reached. This statement could certainly be questioned in 1994. As long as the international community believes in the process of debt securitization, one would expect that, over time, only country specific elements or international shocks should affect the risk premium of a sovereign country. This paper originates from the initial observation that, for the 1994 Mexican crisis, there is evidence of *spillover effects* in the assessment of the sovereign risk premium for different countries.

The paper focuses on the effects on the Brady bond market of the 1994 Mexican crisis. By now, we have observed other crisis after the 1994 Peso collapse, such as in Asia in 1997–1998, in Russia in 1998, and more recently, in Brazil. With respect to the Mexican crisis, these more recent ones have developed in a situation of global uncertainty. In particular, the Asian crisis has hit, from the beginning, more than one country (Thailand, Indonesia, Malaysia) and it has quickly spread into more mature markets as well (such as South Korea). The Russia and Brazil crises, more limited in scope, have developed in a situation of uncertainty and tension in the international financial markets. For these, more recent crises, therefore, it might be more difficult to disentangle country specific from global shocks when assessing the degree of contagion.

On the other hand, the Peso crisis has the advantage of being a country specific shock that has happened during a period of normal functioning of international financial markets. Therefore, it is a good candidate for an event study approach: the December 1994 Mexican shock was substantial, but truly originating from the Mexican economy and, in this sense, exogenous to other economies. In particular, in this paper, we will focus on the effects of the Peso crisis on the risk assessment of Poland. The focus is on Poland for many reasons: several papers have shown that contagion tend to be regional; however, Poland is not in the same region, and has not relevant commercial linkages with Mexico; at the same time, already in 1994, the Polish Brady bonds were very liquid, and with an investment grade rating (see Table 1). Because, as we argue in Barbone and Forni (1997), the reaction to the Peso crisis was not homogeneous across emerging economies, it would be very difficult to find significant effects pooling several countries together in the analysis. Therefore, Poland looked good as a candidate for the exercise.

A number of studies have addressed the issue of contagion in relation to the recent emerging market crisis, among which Calvo and Reinhart (1996), Sachs *et al.* (1996), Cerra and Saxena (1998), Kaminsky and Reinhart (1998), just to quote few of them. Most of these studies usually compare the correlation between two stock markets during a relatively stable period with that during a period of turmoil. Contagion is then defined as a significant increase in the cross-market correlation during the period of turmoil. However, recently, some authors, among which are Forbes and Rigobon (1998), have highlighted the limits of such approach: in particular, Forbes and Rigobon show that, during periods of turmoil when market volatility increases, standard estimates of cross-market correlation are biased upward.

We adopt a test for contagion different from correlation comparisons, making use of cointegration analysis. We have data on weekly stripped spreads on Mexican and Polish Brady bonds for 2 years (July 1994–July 1996). From unit root tests on the series, we find that the null of unit root can not be rejected.

Table 1. Sample description

Country	Bond	Terms announced	Issue date	Maturity date	Ratings[a]	Liquidity[b]
Mexico (19%)	Pars (A, B),	January 1990	28 March 1990	31 December 2019	Ba2, BB	L1
	Discounts (A, B, C, D)		28 March 1990	31 December 2019		L2+
Poland (6%)	Pars,	March 1994	27 October 1994	27 October 2024	Baa3, BBB−	L2
	Discounts		27 October 1994	27 October 2024		L2+

[a] Moody's, Standard and Poor's, respectively. The ratings are as 21 June 1996. The following table summarizes the grade rank:

Investment grade		Speculative grade	
Moody	S&P	Moody	S&P
Aaa	AAA	Ba1	BB+
Aa1	AA+	Ba2	BB
Aa2	AA	Ba3	BB−
Aa3	AA−	B1	B+
A1	A+	B2	B
A2	A	B3	B−
A3	A−		
Baa1	BBB+		
Baa2	BBB		
Baa3	BBB−		

Note: Baa1/BBB + or below is the rating requirement designed by JPMorgan to define an 'emerging market' in the context of external debt market. Source: JPMorgan, *Emerging Market Analytics*, 21 June 1996.

[b] The liquidity rating for Brady bonds provided by JPMorgan is:

- L1 Benchmark — Average bid/offer < 3/8 and bond *quoted* by the *main brokers*.
- L2 Active — Average bid/offer < 3/4 and bond *quoted* by at least half of the *main brokers*.
- L3 Traded — Average bid/offer < 2 and bond *quoted* by at least one *main broker*.
- L4 Mostly Illiquid — Average bid/offer < 3 and bond *quoted* by at least one *main broker*.
- L5 Illiquid — Bond rarely or never *quoted* by *main brokers*.

Note: A bond is considered *quoted* in categories L1, L2 and L3 if it is priced 75% of the time; it is considered quoted in category L4 if it is priced 25% of the time. The *main brokers* designed by JPMorgan are: Eurobrokers, Tullets, Tradition, Cantor, Chapdelaine and RMJ. Based on JPMorgan, *Introducing the Emerging Markets Bond Index Plus*, 12 July 1995.

The finding is consistent with the efficient market hypothesis, where yield spreads move to reflect new information accruing to the market. We then test for cointegration, both during the period of the Peso collapse, and afterwards. As our working assumption is that the Mexican fundamentals were dramatically changed during the crisis, while Poland's were not, we do not control for fundamentals[3]. We say *there is* comovement (or contagion) if we cannot reject the null of stripped spreads cointegration. While usually cointegration is interpreted as a long run relationship, we interpret it as a 'test' of yield comovement restricted to the sample period of the analysis.

A useful way to understand cointegration relationships is contained in Stock and Watson (1988). They observed that cointegrated variables share common stochastic trends (i.e. random walk components). A unit root process can be decomposed into a random walk, plus a stationary (but not necessarily white noise) component. Hence, a linear combination of two non-stationary variables is stationary if, and only if, the random walk components are the same (up to a multiplication scalar) for both variables. Therefore, cointegration is not merely a test for comovements, it is a test for the hypothesis that different variables share the same innovation (random walk) process. In this sense, our test for comovements is much more than a test on the increase in the correlation coefficient. It is a test based on the fact that two variables share the same innovation process.

2. MEXICAN AND POLISH BRADY BONDS

We will focus on the sovereign credit risk as measured by the *sovereign (stripped) spread*.[4] The sovereign spread is a measure that analysts use to assess the risk of a bond and it is given by the yield differential with a comparable US Treasury bill return, which is considered a riskless asset[5]. We took par and discount Brady bonds, as shown in Table 1 and Figures 1 and 2.[6]

We concentrate on the *sovereign* (*stripped*) *spreads* for several reasons. First, we want to exclude the effects of changes in the (same maturity) US interest rates on yields comovements. Note that, if the solvency of a country is in doubt, an increase in the US interest rate can have an effect on the sovereign premium as well. However, we would expect this effect to be both of second order and not homogeneous among different countries. Moreover, the US interest rate has been much smoother than the stripped spreads over the period of our study.[7]

Second, even if the stripped spread is a rough measure of country risk, we are, in fact, interested in cross-country comparisons. For us, it is more important to have a very homogeneous measure, as we believe it is the stripped spread on the most liquid Brady bonds.

Third, it would be ideal to have time series of the risk spread for given maturity instead of following the same bonds over time. However, the Brady market has not reached the dimension to make this possible. The point is that yields' spreads may not provide a valid way of comparing the relative risks of bonds with different maturity. However, the maturity difference for the Mexican and Polish Brady bonds is less than 5 years over a 30 year maturity. The effect of a different maturity is important for short maturity, and it is not a relevant factor over such a long time horizon.

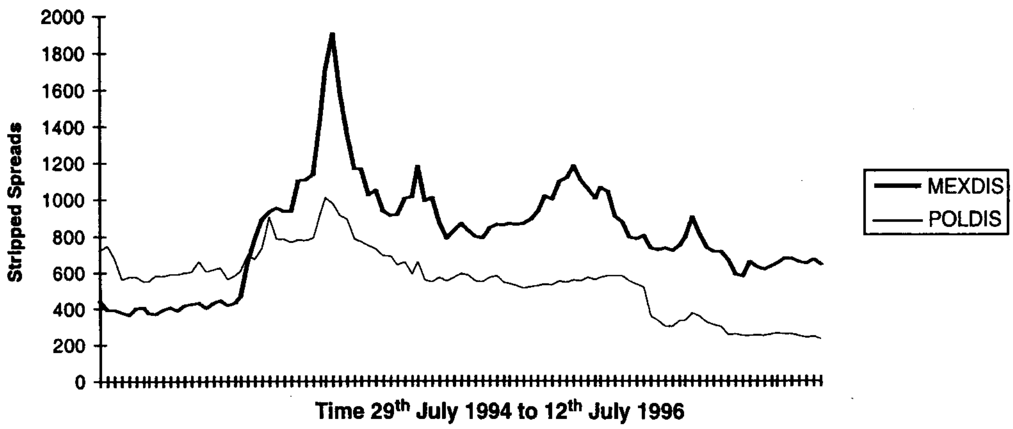

Figure 1. Mexico and Poland, discount bonds

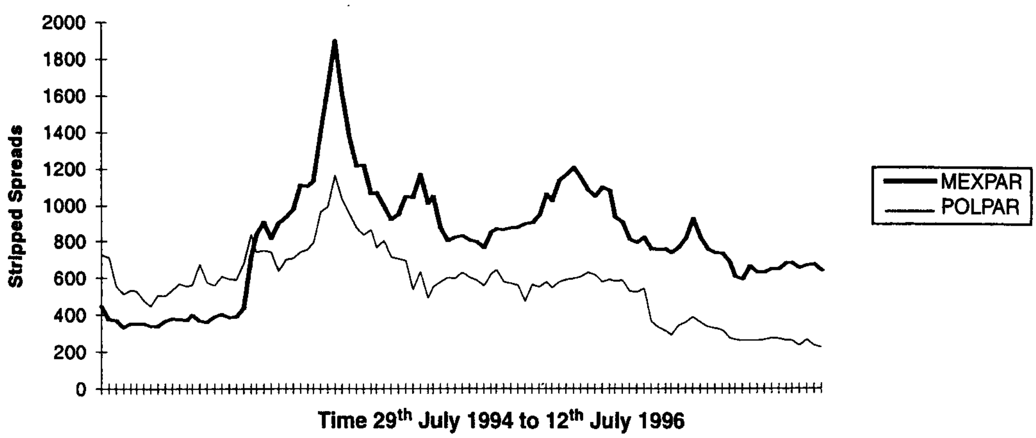

Figure 2. Mexico and Poland, par bonds

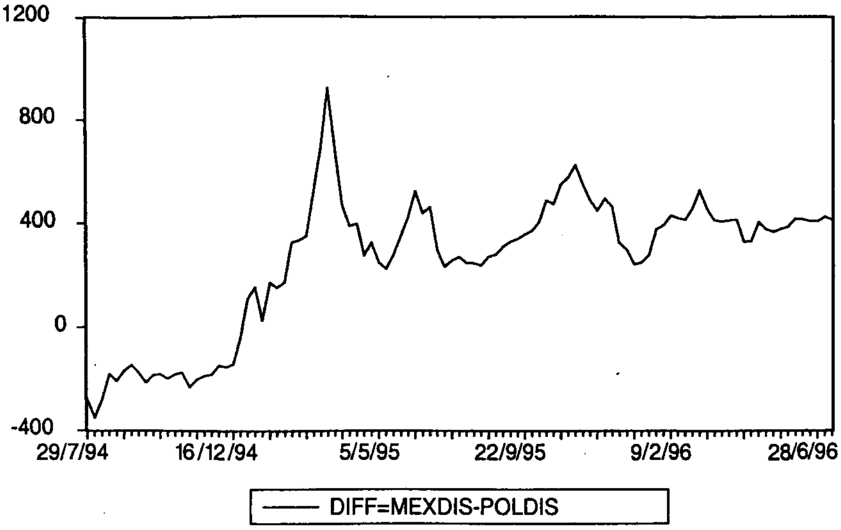

Figure 3. Difference between HEXDIS and POLDIS

We will concentrate on discount bonds. Discount bonds are 30 years' maturity floating market interest rate, issued at a discount to the original face value of the previously rescheduled loans. As they are floating market interest rate, discount bonds do not involve market risk (the risk of a change in the interest rate), so that the measure of country risk given by the stripped spread on discount bonds is more accurate than the one given by the stripped spread on par bonds.

In the following, we will test for comovements in the Mexican and Polish discount bond stripped spreads. As we could not reject the null of unit root on Mexican and Polish stripped spreads, we decided to measure comovements in terms of cointegration: that is, we say *there is* yields' comovement if we cannot reject the null of stripped spreads cointegration.[8] While usually cointegration is interpreted as a long run relationship, we interpret it as a 'measure' of yield comovement restricted to the sample period of the analysis.

With non-stationary time series, the standard asymptotic theory does not hold, and ordinary least squares (OLS) estimates can be misleading. Engle and Granger (1987) introduced the concept of *cointegration*: a group of non-stationary time series is cointegrated if there is a linear combination of them that is stationary; the parameters of the linear combination are called the *cointegrating vector*. If two or more variables are cointegrated, the cointegrating vector can be 'super-consistently' estimated by OLS. Usually, the cointegrating equation is interpreted as a long run equilibrium relationship. We interpret cointegration not only as a 'measure' for comovements, but also as a test for the hypothesis that different variables share the same innovation (random walk) process.

To begin our analysis on comovements of Mexican discount bonds (MEXDIS) and Polish discount bonds (POLDIS), let us consider Figure 3, which represents the difference (DIFF) between the two variables MEXDIS and POLDIS. The variable DIFF looks stationary with a shift in the mean from -200 bs. pts. to $+350$ bs. pts. *after* the Mexican crisis of December 1994. A preliminary estimate of the cointegration relationship is obtained regressing MEXDIS on a constant, POLDIS and a dummy $d2$ (dummy with ones after 20 January 1995, i.e. after the US support package was announced on 12 January 1995).[9] The results of the regression are shown in Table 2.

All variables are very significant, and the estimated coefficient of $d2$, 554, shows the dramatic change in the mean of the relation after 20 January 1995. The change in the mean of the relation can be interpreted as the change in the relation between Mexico and Poland's fundamentals induced by the Mexican crisis. Thus, we generated the variables $\text{MEXDIS1} = \text{MEXDIS} - 550 * d2$ and $\text{DIFF1} = \text{MEXDIS1} - \text{POLDIS}$ to perform tests of cointegration.

Table 2. Preliminary regression

Variable	Coefficient	S.E.	t-Statistic
C	−207.4990	49.24584	−4.213533
$d2$	554.0909	28.96765	19.12792
POLDIS	1.080204	0.067715	15.95221
R^2	0.834172	Mean dependent variable	809.9767
Adjusted R^2	0.830855	S.D. dependent variable	297.5443
S.E. of regression	122.3717	Akaike information criterion	9.642820
Sum squared residual	1 497 483	Schwartz criterion	9.719560
Log likelihood	−639.7559	F-statistic	251.5167
Durbin–Watson statistic	0.451944	Prob(F-statistic)	0.000000

Least squares dependent variable is MEXDIS.
Sample: 29 July 1994–12 July 1996.
Included observations: 103.

To test properly for cointegration, we consider the following ECM:[10]

$$\Delta\text{MEXDIS1}_t = \alpha_M(c + \text{MEXDIS1}_{t-1} + \beta\text{POLDIS}_{t-1}) + \delta_M\Delta\text{MEXDIS1}_{t-1} + \gamma_M\Delta\text{POLDIS}_{t-1} + \varepsilon_{Mt} \tag{1}$$

$$\Delta\text{POLDIS}_t = \alpha_P(c + \text{MEXDIS1}_{t-1} + \beta\text{POLDIS}_{t-1}) + \delta_P\Delta\text{MEXDIS1}_{t-1} + \gamma_P\Delta\text{POLDIS}_{t-1} + \varepsilon_{Pt}$$

The results of the Johansen cointegration test are reported in Table 3.

The cointegration hypothesis is not rejected, even at 1% critical level.[11] Moreover, we cannot reject the null that the (normalized) cointegration vector is (1, − 1), which implies that the series have exactly the same random walk (stochastic trend) component.

Next, we split the sample (2 years from 29 July 1994 to 12 July 1996) in two: the first year goes from 29 July 1994 to 28 July 1995 (51 observations), and includes the period of the Mexican crisis; the second one goes from 28 July 1995 to 12 July 1996 (51 observations). The split is done in order to see if the null of cointegration cannot be rejected also in the period starting 6 months after the crisis (28 July 1995). That is, we divide the sample into two sub-samples, such that the first one includes data in the neighborhood of the crisis, while the second refer to a period when the crisis was perceived as overcome. We report in Tables 4 and 5 the results of the cointegration tests on the two sub-samples.

Table 3. Johansen cointegration test

Eigenvalue	Likelihood ratio	5% critical value	1% critical value	Hypothesized no. of cointegrating equation(s)
0.259056	32.22498	19.96	24.60	None
0.019045	1.942089	9.24	12.97	At most 1
Unnormalized cointegrating coefficients				
MEXDIS1	POLDIS	C		
0.000880	−0.000924	0.152483		
−5.53E−05	−0.000259	0.084977		
Normalized cointegrating coefficients: 1 cointegrating equation(s)				
MEXDIS1	POLDIS	C		
1.000000	−1.050131	173.3165		
	(0.10657)	(62.6566)		

Sample: 29 July 1994–12 July 1996.
Included observations: 101.
Test assumption: no deterministic trend in the data.
Series: MEXDIS1, POLDIS.
Lags interval: 1–1.
LR test indicates one cointegrating equation(s) at 5% significance level.

Table 4. First sub-period (29 July 1994–28 July 1995)

Eigenvalue	Likelihood ratio	5% critical value	1% critical value	Hypothesized no. of cointegrating equation(s)
0.343390	24.04490	19.96	24.60	None
0.049535	2.590984	9.24	12.97	At most 1
Unnormalized cointegrating coefficients				
MEXDIS1	POLDIS	*C*		
0.001156	−0.001715	0.561364		
−1.78E−05	0.001175	−0.757838		
Normalized cointegrating coefficients: 1 cointegrating equation(s)				
MEXDIS1	POLDIS	*C*		
1.000000	−1.483337	485.6730		
	(0.19940)	(138.440)		

Sample: 29 July 1994–28 July 1995.
Included observations: 51.
Test assumption: no deterministic trend in the data.
Series: MEXDIS1, POLDIS.
Lags interval: 1–1.
LR test indicates 1 cointegrating equation(s) at 5% significance level.

The null hypothesis of cointegration is not rejected in the first sub-sample, but it is rejected on the second one. Note also that, for the first sub-sample, the estimated coefficient on POLDIS is significantly greater than one.[12] From these results, we can reach two main conclusions: first, *comovements were stronger during the high volatility period of the Mexican crisis.* In fact, we reject the null hypothesis of cointegration for the period 28 July 1995–12 July 1996, when the crisis was perceived as overcome; second, the Mexican crisis has caused a change in the mean of the cointegration relation of approximately 550 bs. pts. That is, *the crisis has had a strong permanent effect on the risk assessment of Mexico with respect to the one of Poland.* Note that this result does not depend on the sample period used: on the basis of more updated data, the gap in stripped spreads between Mexico and Poland that turned positive (from negative) after the Peso collapse, is still positive and significant.

We study next the properties of the ECM on the complete sample period in order to address the following questions: Can we quantify the response of POLDIS to a shock to MEXDIS? To which extent

Table 5. Second sub-period (28 July 1995–12 July 1996)

Eigenvalue	Likelihood ratio	5% critical value	1% critical value	Hypothesized no. of cointegrating error(s)
0.185747	11.99515	19.96	24.60	None
0.029278	1.515477	9.24	12.97	At most 1
Unnormalized cointegrating coefficients:				
MEXDIS1	POLDIS	*C*		
0.001493	−0.001499	0.176611		
−0.000671	7.73E−05	0.044013		
Normalized cointegrating coefficients: 1 cointegrating equation(s)				
MEXDIS1	POLDIS	*C*		
1.000000	−1.003506	118.2626		
	(0.20818)	(93.4870)		

Sample: 28 July 1995 to 12 July 1996.
Included observations: 51.
Test assumption: no deterministic trend in the data.
Series: MEXDIS1, POLDIS.
Lags interval: 1–1.
LR rejects any cointegration at 5% significance level.

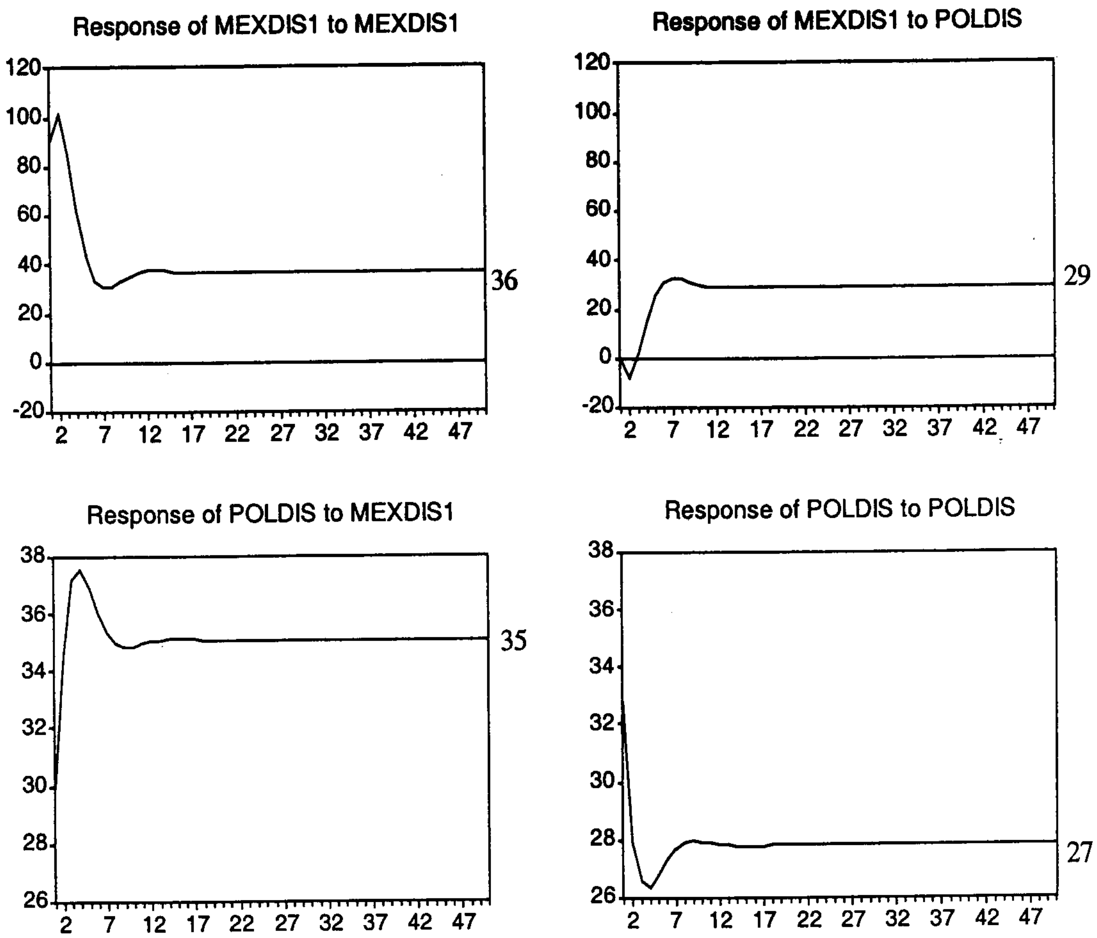

Figure 4. Response to one S.D. innovations

the volatility of POLDIS is owing to the volatility of MEXDIS? These questions can be answered through what is generally referred to as *innovation accounting*, i.e. the *impulse response function* and the *forecast error variance decomposition.*[13]

The impulse response function, plotted in Figure 4, shows that: (1) one S.D. innovation to MEXDIS1 (the S.D. of the residuals on MEXDIS1 being of 91 bs. pts.) has a permanent effect of *36 bs. pts* on MEXDIS1 and of *35 bs. pts.* on POLDIS; (2) one S.D. innovation to POLDIS (the S.D. of the residuals on POLDIS being of 44 bs. pts.) has a permanent effect of *29 bs. pts.* on MEXDIS1 and of *27 bs. pts.* on POLDIS. We can conclude that a shock to MEXDIS1 has a stronger permanent effect on both variables than a shock to POLDIS (although the residuals on MEXDIS1 are more volatile).

From the analysis of the *impulse response function*, we can reach another important conclusion: on the basis of the estimated coefficients of the vector error correction model, one S.D. innovation to MEXDIS1 has a permanent effect on POLDIS of 35 bs. pts. (out of 91 bs. pts.), or of 44 bs. pts. (out of 118 bs. pts.), depending on the sample period—in both cases, 38% of a shock to MEXDIS1 has a permanent effect on POLDIS.

The *forecast error variance decomposition*, reported in Figure 5, shows that: (1) the long run forecast error on MEXDIS1 is responsible for *70%* of innovations to MEXDIS1, and for *30%* of innovations to POLDIS; (2) the long run forecast error on POLDIS is responsible for *61%* of innovations to MEXDIS1, and the remaining *39%* for innovations to POLDIS. We note a significant asymmetry in the behavior of the two variables, with *61%* of the variability of POLDIS coming from the variability of MEXDIS1.[14]

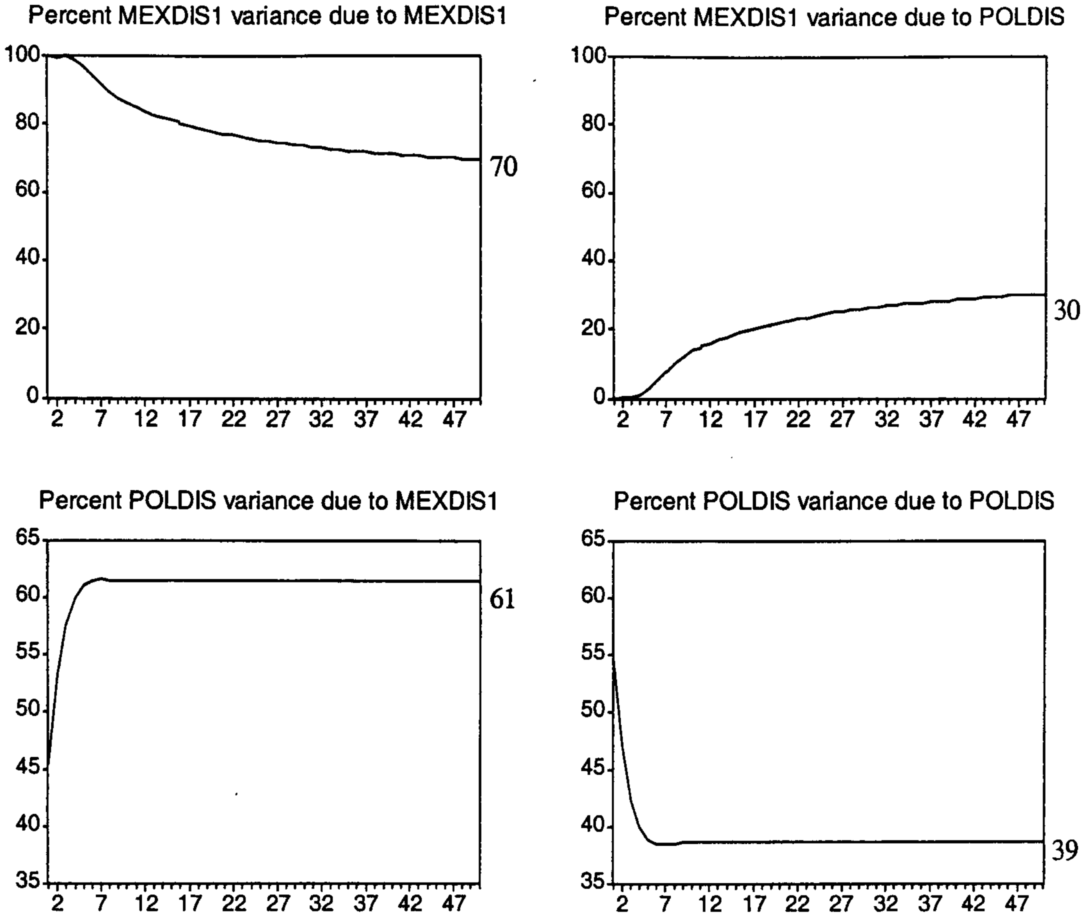

Figure 5. Variance decomposition

3. CONCLUSIONS AND IMPLICATIONS

This paper has studied the effects of the Mexican crisis on the sovereign risk assessment (as measured by the stripped spreads) of Mexican and Polish Brady bonds. It has adopted a test for contagion different from correlation comparisons, making use of cointegration analysis on stripped spreads. We have followed an event study type of approach: our working assumption is that the Mexican fundamentals were dramatically changed during the Peso crisis, while Poland's were not. We say *there is* comovement if we cannot reject the null of stripped spreads cointegration. While usually cointegration is interpreted as a long run relationship, we interpret it as a 'test' of yield comovement restricted to the sample period of the analysis. Our findings can be summarized as follows:

(i) over the sample period, the unit root hypothesis on the risk premium (measured by the stripped spread) of Mexico and Poland cannot be rejected; this is consistent with the idea that the risk premium reflects new information accruing to the market;
(ii) comovements in stripped spreads between Mexico and Poland were stronger during the period of the Peso crisis: we do not reject the null of cointegration for the year that includes the crisis (July 1994–July 1995), but we do reject the null for the year starting 6 months after the crisis (July 1995–July 1996);
(iii) the crisis has had a strong permanent effect on the risk assessment of Mexico with respect to the one of Poland (550 bs. pts. circa).

While it is very difficult to rationalize evidence of contagion, these results suggest that fund managers had poor experience and limited information on the fundamentals of Mexico and Poland. In fact, it is reasonable to assume that investment decisions are based mainly on private information on a country's fundamentals. Private information on fundamentals provides managers with a belief on the probability of default of a debtor country. Moreover, as long as fund managers believe in the process of developing countries debt securitization through the Brady agreements, one would expect that, over time, only country specific elements or international shocks affect the risk premium of a sovereign country. However, *spillover effects* can occur if news on a particular emerging country conveys some information on the fundamentals of other emerging countries. This could be owing, for example, to the fact that different countries are applying reform programs—like Brady plans—which are perceived to be similar. Thus, the Mexican crisis can be a signal that the Brady plan on Mexico is not feasible, or that the Mexican government finds the commitment to the plan not rewarding enough. Therefore, fund managers might think that, also, the Brady plan for Poland is not feasible, or that the Polish government may not have enough incentive to abide to the plan. To be more specific, assume that the probability of default, , is a function of 'observable' fundamentals Z_{obs} (as budget deficit, exchange rate, etc.) and 'not-observable' fundamentals $Z_{\mathrm{not\text{-}obs}}$ (as the government commitment to the reform program or the feasibility of the program itself). That is: $\pi = f(Z_{\mathrm{obs}}, Z_{\mathrm{not\text{-}obs}})$. The Mexican crisis has substantially changed Z_{obs} for Mexico, but not for other countries like Poland. Therefore, for Poland, the jump in the probability of default, and thus, in the risk premium after the Mexican crisis cannot be owing to a change in Z_{obs}. It has to come from a change in $Z_{\mathrm{not\text{-}obs}}$. The interesting point in analyzing the period of the 1994 Mexican crisis is that we can assume that, during the crisis, Z_{obs} did change only for Mexico. Therefore, it cannot be argued that comovements were owing to a similar change in observable fundamentals. The discussed change in the mean of the cointegrating relation between Mexico and Poland is consistent with the idea that the Mexican crisis changed the relation between Mexico and Poland's observable fundamentals. Also, the fact that the null of cointegration for the second year of our sample (July 1995–July 1996) is rejected goes against the idea that comovements during the Mexican crisis were a result of similar movements in observable fundamentals.

ACKNOWLEDGEMENTS

The authors would like to thank, for their comments and suggestions, Stijn Claessens, Daniel Oks and Enrico Perotti, as well as the participants in a seminar held at the World Bank. The views and opinions expressed herein are those of the authors' only, and not those of the World Bank, or the Bank of Italy.

APPENDIX A. BRADY BONDS: A PRIMER

Par and discount bonds have been, by far, the most popular options in several Brady agreements. Par bonds are 30 years' maturity, issued at 'par', i.e. for the original face value of the loans for which they were exchanged. Par bonds carry a fixed, below market interest rate (relative to the interest rate environment at the time of the issuance). Discount bonds are 30 years maturity floating market interest rate, issued at a discount to the original face value of the previously rescheduled loans.

In general, both security types are collateralized as to principal at maturity by US Treasury zero coupon bonds. Par and discount bonds may possess additional collateral in the form of cash collateral account usually maintained at the Federal Reserve Bank of New York. The funds in these accounts, which are allowed to be invested in securities rated AA—or better, are maintained in order to pay interest for a specific number of months (usually from 12 to 18 months, depending on the country) in the event that the sovereign debtor suspends the interest payments. In effect, the collateral provides a 'rolling' interest payments guarantee, meaning that, if each successive interest payment is made and the collateral is not utilized, it is 'rolled' or applied to guarantee the nearest dated unsecured semi-annual interest payment. If the collateral is drawn upon, it is not required to be replaced.

Certain par and discount bonds also carry 'value recovery' warrants, which give bondholders the opportunity to 'recapture' some of the debt and debt service reduction provided in the exchange if the future economic performance and/or the debt servicing capacity of the sovereign debtor improves. For example, Mexico, Venezuela and Nigeria Brady bonds carry warrants linked to indices of oil export prices, or of national oil export receipts.

The value assigned to the principal on par and discount bonds is simply the present value of the security used as collateral. For dollar denominated bonds, the principal is discounted by the risk-free rate of a comparable maturity US Treasury zero-coupon bond.

The interest collateral of a Brady bond is a more complicated component to value, as it is uncertain when, if ever, it will be used. This provides an occasion for differences in valuation approaches that arise from assumptions about the probable life of the 'roll' of the collateral. One method assumes the limiting case, which ascribes the minimal value to the collateral's current market value, effectively treating it as if there was no life to the 'roll' (i.e. default occurs today and collateral is utilized and exhausted to pay the nearest two or three semiannual interest coupons). Other methods rely on the use of probability models or option pricing formulas to assign probabilities to interest payment default, which are then incorporated in the valuation model.

To obtain the *stripped yield*, the value of the principal and interest collateral is calculated as described above, and this notional 'market value' of the collateral is subtracted from the current market price. After 'stripping' out the values of the principal and interest collateral from the price, an internal rate of return is computed on the non-collateralized contractual cash flows. These non-collateralized interest flows should be discounted by a rate, the *stripped yield*, that produces a present value which, when added to the collateralized portion of the bond, equals the current market price. The internal rate of return of a Brady bond cash flow with its collateral left intact is simply the bond's implied yield-to-maturity or *blended yield*.

The value of the 'value recovery' warrants, which are linked to the value of the country's oil exports, is usually ignored in the valuation of the bonds.

Swapping floating-rate coupons for fixed rate equivalents (using either the current LIBOR-based Eurodollar swap curve or forward rate curve) places all Bradies on equal footing with other sovereign instruments. In particular, it makes possible to compute the *stripped spreads* for all bonds.

NOTES

1. The Emerging Market Trading Association was founded in 1990 (a successor to the earlier 'LDC Debt Traders Association'), and by now, most bonds are traded continuously all over the world. Trading volumes in 1994 reached US$2.8 trillion (from US$100 billion in 1990), the total amount outstanding of the US dollar-denominated Brady market in 1994 had an approximate face value of US$138 billion, and future and option market are now traded for all major Bradies. The reader interested in knowing more (daily prices, futures and options, analysis and ratings, news and forum) can now browse the pages of the BradyNet (http://www.2020tech.com/bradynet).
2. For an assessment of the Brady deal on Mexico the reader can see Claessens *et al.* (1993). For an account of the 1994 Mexican debacle, see Calvo (1996).
3. Considering the fact that we work with weekly data, it would be rather difficult to control for fundamentals in a detailed way.
4. The reader not familiar with this term can refer to Appendix A.
5. To be rigorous, the credit risk should be split into two parts: one is the probability of default, and the other is the amount of loss in case of default. For example, it would be of interest to compute the internal rate of return (yield) on a bond based on what are the *expected payments* instead of the *contractual ones*. The differential to an analogous, riskless asset would be a proper measure to infer the probability of default. The authors who have attempted to differentiate between the two elements (i.e. the probability of default and the amount of loss in case of default), however, are forced to make strong assumptions on the probability of default, and on the expected repayment in case of default. For example, Cumby and Evans (1993) assume that in case of default the repayments are zero forever. Claessens and Pennacchi (1992) assume that the probability of default follows a Brownian motion. An interesting finding of Cumby and Evans (1993), though, is that the hypothesis of an expected decreasing default probability is consistent with their data set (Brady bonds prices starting 1990 for Mexico, Venezuela and Costa Rica).
6. The data are provided by Salomon Brothers, and are based on weekly (usually Wednesday) close of the day bid prices quoted in New York. We have data beginning on the week Salomon Brothers have started trading on the loans included in the Brady agreements, even though the bonds have been, in all cases, formally issued a few months later. The data set we use contains the longest time series as possible on Brady bonds as of July 1996, as it includes information on yield in some cases even before the Brady bonds were formally issued.
7. For a discussion of this point, see Barbone and Forni (1997).
8. We do not report in this version of the paper the tedious discussion of the non-stationarity of the series. On this issue, the reader can refer to Barbone and Forni (1997) (see http://www.worldbank.org/html/extpb/index.htm). Note that the finding of unit root on stripped spreads is itself an interesting result. For example, the IMF (1995) rejects the unit root hypothesis on Mexican Brady bond prices.

9. To test properly for cointegration, we consider below an error correction model (ECM).
10. Note that, if two or more variables are cointegrated, it is guaranteed that an ECM representation exists. This is the so called *Granger representation theorem*: for any set of $I(1)$ variables, cointegration and error-correction are equivalent representations. Moreover, the most common procedure to test for the lag length is to estimate a vector auto regression (VAR) using data in *levels*. We consider VARs with two or more lags: the VAR with one lag presents evidence of autocorrelation in the residuals. The Akaike information criterion suggests a specification with two lags and the Chi test suggested by Sims (1980) does not reject the null of a lag length of two versus three or more lags.
11. Note that we performed the tests for the lag length on a VAR in levels. The ECM is specified in first differences, so the lag length must be reduced by one.
12. To confirm the results, we perform the cointegration test on the original variable MEXDIS (instead of MEXDIS1). The results of the cointegration tests for the two sub-periods—from 20 January 1995 to 12 July 1996 (78 observations), and from 28 July 1995 to 12 July 1996 (51 observations)—confirm that there is strong evidence of cointegration on the whole sample, while the variables are not cointegrated for the period 28 July 1995 to 12 July 1996 (the last year).
13. The impulse response function uses the estimated parameters of the model to trace out the effects of one standard deviation (S.D.) change in the errors on the time path of MEXDIS and POLDIS sequences; the forecast error variance decomposition uses the estimated parameters, and shows which proportion of the forecast errors on a variable is owing to its own shocks (residuals) versus shocks (residuals) to the other variable. The only problem with this methodology is that an estimated VAR is underidentified: an additional restriction must be imposed on the two variable VAR systems in order to identify the impulse responses and variance decomposition. The usual restriction is given by the assumption that one error has a contemporaneous effect on both variables, while the other one has a contemporaneous effect only on its own variable. This assumption is said to imply an *ordering* of the variables. We assume that has a contemporaneous effect on both variables, while does not. This is not an innocuous assumption: the correlation coefficient on the estimated residuals is above 0.7, so the ordering matters. Our assumption is based on the following considerations. First, at least for the first year of our sample (when the evidence of cointegration is stronger), all the relevant volatility was coming from the news on Mexico, and not on Poland. Second, consistent with our belief that Mexico is a leader in the Brady market, the IMF (1995) finds evidence of Granger causality from Mexico to other countries using daily bond prices.
14. We have estimated the ECM for MEXDIS1 and POLDIS also for the period 29 July 1994 to 28 July 1995, i.e. the period comprising the Mexican crisis, and during which the comovements are stronger. The results are confirmed: in particular, the variability of POLDIS is now even more dependent on the variability of MEXDIS1 (68% versus 61%) and the amount of the permanent effect on MEXDIS1 owing to innovations on both variables has increased relative to the increase in the S.D.s of the residuals (the permanent relative response to MEXDIS1 innovations has shifted from 0.4 to 0.55, whereas the permanent relative response to POLDIS innovations has shifted from 0.66 to 0.82). Finally, the ECM estimates for MEXDIS (instead of MEXDIS1) and POLDIS for the sample 20 January 1995–12 July 1996 are substantially the same.

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