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Nonlinearities in CDS-Bond Basis

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Abstract: Theoretically, the risk premium captured by Credit Default Swap (CDS) and bond yield spreads should be equal. However, data reveals a significant difference between the two spreads. We explore the presence of a mean-reverting behavior in this difference (CDS-bond basis), for selected emerging markets, employing alternative threshold models (TAR, TAR-GARCH and ESTAR). Our results indicate a positive relationship between the speed of adjustment and the trading frequency of the sovereign CDS's and bonds. The TAR-GARCH model suggests that the adjustment of the CDS-bond basis is immediate for economies with more liquid CDS's and bonds, such as Argentina, Brazil and Mexico. The ESTAR model indicates that the adjustment displays a gradual pattern for the basis of the economies with less frequently traded bonds and CDS's.

JEL Classification: C32, G12

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I. Introduction

Two popular credit risk measures for emerging markets are the Credit Default Swap (CDS) spread and the Emerging Markets Bond Index Plus (EMBI+) spread. The former one denotes the annual premium, which is paid by the buyer of the CDS for protection, in case of default of the reference entity, as a percentage of the notional principal. The latter one is the difference between yields of the sovereign bond and a comparable US treasury bond, expressed in basis points. Theoretically, these two spreads should be the same to avoid arbitrage opportunity. However, more often than not, this theoretical equivalence fails to hold. Figure 1 displays the discrepancy between these two spreads which is defined as the *CDS-bond basis*, for ten emerging markets.

Despite this deviation from the parity between CDS and bond spreads, the literature suggests a long-run co-integrating relationship between these two. These studies, which are explored in the next section, assume a linear adjustment and mostly adopt vector error-correction models. We aim to extend this literature through investigating nonlinearities in the adjustment pattern of the CDS-bond basis.

The central conjecture put under scrutiny in our study is whether the behavior of the CDS premium and the bond yield depend on the *magnitude of the change* in the CDS-bond basis itself. If the basis were within a certain threshold, then the market participants would not respond. This band of inaction could be motivated through transaction costs or a combination of the structural or market-related factors, as will be discussed in a while. However, above this threshold, arbitrage opportunity would dominate the effects of all the other determinants, giving rise to an adjustment towards a long-run equilibrium value.

We show that the speed of this reversion depends on the liquidity structure of these instruments. Higher the trading frequency of the particular sovereign CDS or bonds, higher the speed of adjustment in basis, that is captured by the threshold autoregressive models which assume a sharp adjustment form (TAR and TAR-GARCH). Alternatively, the adjustment displays a gradual pattern for the basis of the sovereign assets with lower trade volumes, depicted by the exponential smooth transition (ESTAR) model.

In the next section, we briefly explain determinants of the change in CDS-bond basis. However, our focus is on the *adjustment process* of the basis, rather than exploring these determinants. Our emphasis could be motivated both from the point of view of investors and policymakers. First, market participants occasionally engage in *basis-trade*, exploiting the arbitrage opportunity based on the deviation in CDS-bond basis from its long-run value[1]. Therefore, the *speed* of this adjustment emerges as an essential issue for profit seeking investors**.** Second, a long-run co-integration between CDS and bond markets would be valuable information for policymakers, who are interested in the risk premium implied by the price of sovereign assets. Moreover, as will be explored in the second section, the level of the basis could be suggested as an indicator of market liquidity, especially during periods of financial market stress.

We examine the adjustment mechanism of the CDS-bond basis and the liquidity structure of ten selected emerging markets. First, by means of employing TAR and TAR-GARCH (1,1) models, we find an abrupt adjustment for sovereigns with liquid CDS and bonds such as Argentina, Brazil and Mexico. Second, our ESTAR model points out a smooth adjustment pattern for sovereign CDS and bonds with relatively lower frequency, such as Colombia, Panama, Peru and Philippines. Turkey is also in this latter group despite the recent high trading volume of its CDS contracts.

The rest of the paper is organized as follows. Next section examines the determinants of the CDS-bond basis, provides a literature review and further delves into our motivation.

[1] An interesting blog article on basis-trade is at http://zerohedge.blogspot.com/2009/01/was-merrill-casualty-3-of-basis-trade.html

Third section demonstrates our data on asset prices and alternative liquidity measures. Fourth section discusses the methodology and presents the results. Fifth and last section concludes.

II. Literature review

The CDS-bond basis deviates from zero as a result of structural as well as market-related factors. Structural factors are mainly related with the specification of contracts. One example is the cheapest-to-delivery (CTD) option that gives the right to the buyer to deliver the cheapest par-value bond to the seller among a basket of pre-defined bonds, in realization of a credit event[2]. The seller would ask for a higher premium as a compensation for this clause, which would create an upward bias on the CDS spread. Moreover, during times of financial stress, intense demand for these CTD bonds may result in an increase in the implicit recovery value for the buyer, which emerges as a significant determinant that drives the wedge between CDS and bond (Andritzky and Singh, 2006).

Another example is the counterparty risk embedded in the CDS contracts. This additional risk carried by the buyer should be compensated through a lower premium, as opposite to the previous example. Remarkably, increasing counterparty risk in the CDS market after the failure of Lehman Brothers was one of the determinants of the plunge in CDS-bond basis in October 2008, as can be seen in Figure 1. A recent survey conducted by European Central Bank (2009a) points out that counterparty risk causes distress for banks -as CDS buyers- due to high market concentration in the seller side of the CDS market, especially after Lehman Brother's default[3].

Among market factors, the most significant one is the relative market liquidity. ECB (2009b, pp. 77-78) suggests that a sudden movement into a negative basis might be a result of

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[2] In a *naked CDS* contract, the buyer does not have to hold the underlying bond at the time of the initial contract setting.

[3] See Kane and McAdie (2006) and DeWit (2006) for additional examples of determinants of the CDS-bond basis.

a liquidity shock[4]. As the argument goes, the negative relationship between the market depth and the premium on both CDS and bonds is more noticeable during distress times. Recently, October 2008 marks a period where worsening liquidity conditions resulted in higher bond yields while the effect on the CDS premiums was relatively limited. Funding costs has significantly increased during this period and market participants favored cash over bonds, which lowered bond prices and increased their yields. Moreover, for developing markets, the situation was worse because of the flight to safety behavior of the investors, which causes a shift away from their bonds.

The empirical relationship between CDS and bond markets has been explored in the literature. Blanco et al. (2005) documents evidence of co-integration between two markets. They show that CDS and bond markets price credit risk equally in the long run. They further suggest a leading role for CDS market over bond market in price discovery process. Zhu (2006) corroborates with these results and highlights the relatively higher responsiveness of the CDS market to volatile credit conditions in explaining the short-term deviation of the basis from zero. Norden and Weber (2009) also documents a long-run co-integration between CDS and bond markets for around sixty percent of the corporate issuers, where they argue that the leading role of the CDS market is more significant for US firms over their European counterparts. Adler and Song (2010) confirm the validity of the parity relationship for emerging market sovereigns by means of constructed par equivalent bond spreads making use of CDS premiums. Trapp (2009) argues that for financial firms, the long-run equilibrium relationship is not as strong as suggested by the literature due to credit risk and liquidity conditions.

Our study contributes to this literature with three main aspects. First, as discussed in the introductory section, our interest lies in exploring a non-linear adjustment mechanism in

[4] See Fontana (2010) for an exploration of the persistent negative basis during the recent crises.

the CDS-bond basis. We show that the adjustment could take place in two different forms depending on the trading frequency of the particular asset or the heterogeneity amongst market players. On the one hand, if the sovereign CDS or bond were intensely traded in the market, then we would expect a sharp correction, once the threshold is crossed. On the other hand, if the trading frequency of the bonds and corresponding CDS for a particular sovereign is relatively low, then the reversion towards the mean can be gradual. Another explanation for this kind of a smooth adjustment could be the heterogeneity among market players. If the market participants have different beliefs about the credit risk structure of the particular sovereign, they would respond at different times. This implies a continuum of thresholds where the adjustment would be gradual rather than a sharp one.

Second, our results also corroborate with the emerging literature that finds a positive relationship between the market liquidity and speed of the disappearance of the arbitrage opportunity, in other asset classes. Roll et.al (2007) explores the relationship between future contracts and a constructed measure of aggregate liquidity in US stock market, employing a vector auto-regression framework. Their results suggest that higher market liquidity leads to faster mean reversion in futures-cash basis. Using survival analysis, Deville and Riva (2007) confirms the same relationship for French index option market. Deville et.al (2009) argues that introduction of exchange trade funds as a liquidity enhancing financial innovation has improved this relationship. Andani et.al (2009) suggests that the performance of alternative hedging strategies in exploiting futures-cash basis is lower in markets with higher trading volume. Our study contributes to this literature by finding a similar positive relationship between trading volumes and the speed of mean reversion in CDS-bond basis.

Third, our framework is also novel from a technical perspective. A significant obscurity that affects the limiting theory of linearity tests is the presence of conditional heteroscedasticity, which is the case for most financial data. Conventional unit root tests may result in detecting spurious nonlinearity in case of neglected heteroscedasticity[5]. We adopted bootstrap procedure build by Gospodinov (2008) that studies the limiting distribution of the LM test for linearity in unit root TAR models with GARCH (1,1) errors. He shows that the size and power of his proposed test is high in presence of non-normal errors. We compared the results with the seminal Caner and Hansen (2001) TAR nonlinearity tests. Similarly, for ESTAR model, we employed heteroscedastic consistent standard errors as suggested by White (1980). In both models, the tests that take conditional heteroscedasticity into account suggest lower rejection rate for the null hypothesis of linearity. These results confirm with the recent literature which shows that omitting the time varying variance problem may lead to oversizing of the linearity tests, as in Gospodinov (2008) for immediate transition model and Pavlidis et.al (2010) for smooth transition models.

III. Data

We collected daily CDS and EMBI+ spreads, from Bloomberg database, for ten emerging markets: Argentina, Brazil, Bulgaria, Colombia, Mexico, Panama, Peru, Philippines, Russia and Turkey. Our data spans a nine years period from October 24, 2000 to January 8, 2010. However, for some countries we have shorter data coverage, as displayed in data appendix.

EMBI+ is an index developed by JP Morgan to provide a measure of emerging market default risk. It includes US dollar denominated debt instruments issued by emerging sovereigns such as Brady bonds, Eurobonds, traded loans and other local market debt instruments. The EMBI+ country spread is the difference of the yield on the sovereign portfolio over the risk-free US bonds with identical maturity.

We chose US dollar denominated CDS quotes, each with five years to maturity in order to obtain match with bond data. Both CDS and bond spreads are used in the natural

[5] See Ling et.al (2003) for a review of the theory of unit root processes with heteroscedastic disturbances.

logarithm form. The CDS-bond basis is measured as the difference between CDS spread and EMBI+ spread for each sovereign country in our sample.

We rely on two different liquidity measures to determine the trading frequency of a particular CDS. First one is the quoted bid-ask spreads presented in Table 1. A lower bid-ask spread indicates a higher trading volume for the asset. First column presents the sample average of the CDS bid-ask spreads. However, a ranking based solely on these averages would be misleading since the CDS spreads have different magnitudes for different sovereigns. Therefore, we divided these figures to the sample average of CDS quotes that are given in the second column. The resulting third column gives a ranking of our liquidity measure where the lowest figure on top indicates the highest trading volume. Parts (b) and (c) do the same exercise for pre-crises and post-crises samples. The results indicate that Argentinian CDS's lead in the pre-crises period while Mexican and Brazilian CDS's are the most frequently traded CDS contracts in the post-crises period.

Our second liquidity measure is a recent market survey conducted by Emerging Markets Trade Association (2011) for the trading volume of emerging market CDS contracts. The results of this survey, which is depicted on Figure 2, indicate that Brazilian, Turkish, Mexican and Russian sovereign CDS's, are the most frequently traded sovereign CDS contracts in 2010. These findings are in line with the results of our first proxy, as presented in the Part (c) of Table 1.

IV. Methodology and Results

We explore the possibility of a nonlinear adjustment pattern in the CDS-bond basis. We try two different threshold processes where the speed of adjustment depends on the liquidity structure of the assets and heterogeneity amongst investors. In the first subsection, we work through the case where the correction is sharp once the threshold is crossed, using TAR and TAR-GARCH (1,1) models. These types of models could better capture the behavior of the CDS and bond markets with high trading frequency where the arbitrage profits are immediately exhausted. Alternatively, when the asset market is relatively shallow, then the adjustment could display a more gradual pattern. ESTAR models capture this smooth transition process as will be presented in the second subsection.

a) Sharp adjustment (TAR model)

We first postulate an adjustment mechanism where market participants disregard small deviations from zero in basis, while large discrepancies are eliminated almost immediately. That nonlinear correction pattern could be captured by a simple version of TAR model given as:

$$\Delta y_t = I_t \rho_1 y_{t-1} + (1 - I_t)\rho_2 y_{t-1} + \zeta_t \qquad (1)$$

where ζ_t is an *i.i.d.* error and I_t is the indicator function with:

$$I_t = \begin{cases} 1 & if \quad z_{t-1} < \lambda \\ 0 & if \quad z_{t-1} \geq \lambda \end{cases} \qquad (2)$$

In a standard TAR model, the adjustment depends on the level of the variable where $z_t = y_t$. Instead, Enders and Granger (1998) select $z_{t-1} = y_t - y_{t-m}$, where the adjustment depends on the past changes in y_t against a threshold. Caner and Hansen (2001, hereafter CH) develop unit root tests for this self-exciting momentum threshold autoregressive (M-TAR) model that considers nonlinearity and nonstationarity in a joint manner. The null hypothesis is standard, H_o: $\rho_1 = \rho_2 = 0$, and corresponds to the nonstationary case where they consider two possible alternatives: H_1: $\rho_1 < 0$, $\rho_2 < 0$ and

$$H_2 : \begin{cases} \rho_1 < 0 \;\; and \;\; \rho_2 = 0 \\ or \\ \rho_1 = 0 \;\; and \;\; \rho_2 < 0 \end{cases}$$

The former alternative, H_1, implies overall stationarity where the latter one, H_2, suggests a partial unit root case with stationarity in one of the regimes and nonstationarity in the other. The optimal threshold level λ and the lag length, *m*, are determined endogenously as the values that minimizes the residual sum of squares of the least squares estimation of (1) [6,7]. They show that under a nonlinear setting, the power of their proposed Wald test is substantially higher than the conventional unit root tests.

Gospodinov (2008) extends the framework of CH by considering conditional heteroscedasticity with GARCH (1,1) errors, inherent in the data generating process. He suggests the model:

$$y_t = a_1 y_{t-1} + I_t \phi y_{t-1} + \zeta_t$$

$$\zeta_t = \sqrt{h_t}\varepsilon_t \qquad (3)$$

$$h_t = \omega + \alpha \xi^2{}_{t-1} + \beta h_{t-1}$$

where the indicator function I_t is the same as in (2), ε_t~iid (0,1), $\alpha \geq 0$ and $\beta \geq 0$.

LM test for linearity for the second model in equation (3), suggests the null hypothesis as H_o: $\phi = 0$. Gospodinov (2008) provides the asymptotic critical values using bootstrap methods and shows that, in presence of non-normal errors, the size and power of the test has been improved, once we allow for conditional heteroscedasticity.

We consider two different threshold variables, following Hansen (1997). Long difference implies a threshold variable where $z_{t-1} = y_{t-1} - y_{t-m-1}$ whereas the alternative, lagged difference, defines $z_{t-1} = y_{t-m} - y_{t-m-1}$. CH imposes strict stationarity and geometric

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[6] Given that the Wald test is a monotonic function of the residual variance, this is identical with the level of m that maximizes the Wald test.
[7] Selection of the threshold level includes elimination of outliers by trimming the series for the highest and lowest values.

ergodicity on these predetermined threshold variables. Third to sixth columns of Table 2 suggest that our threshold variables comply with this assumption, in accordance with two alternative unit root tests. Moreover, first two rows of Table 2 shows y_t is I(1) for almost all of the cases.

In order to provide a comparison of CH and Gospodinov (2008) studies, we first pick the endogenously selected optimal delay parameter, *m*, from the former exercise, and impose this threshold variable and delay parameter combination to the latter, Gospodinov (2008) estimation.

Table 3 presents preliminary diagnostic tests based on level values of CDS-Bond basis and residuals extracted from a simple AR(1) model. First part of the table suggests high dependence and persistence for CDS-Bond basis of each country. BDS Independence test (Broock et.al, 1996) results reveal that the residuals from AR(1) model have a time dependent structure, most probably due to a time varying variance. We test for the existence of any volatility structure for the variance conducting ARCH-LM tests and concluded that there are significant ARCH effects. Last two rows of the table show that the null hypothesis of parameter stability is not rejected almost for all cases, with respect to sup-LM and sup-Wald tests proposed by Andrews (1993). The results discard a significant state dependence; hence rule out a possible spurious nonlinearity problem.

The results of the TAR and TAR-GARCH (1,1) tests are presented in Table 4. Second column describes the threshold variable used in the model[8]. Third column reports the optimal delay parameter. W_u stands for the p-values of Wald tests for the unconstrained model in CH[9].

[8] Lagged thresholds are not valid for Argentina and Mexico due to non-stationarity reported in Table 3. Optimal delay parameter is unity for Turkey, Peru and Panama, which indicates a single threshold variable.
[9] We conducted both unconstrained and constrained models as explained in CH study. Since results are largely the same, only ones of the unconstrained model are displayed.

t_{1u} and t_{2u} stand for p-values of the *t1* and *t2* tests for nonstationarity for the partial unit root cases. LM_G is the p-value of the Sup-LM test for TAR-GARCH estimation.

The third column indicates a non-linear mean reversion of the CDS-bond basis, for nine out of ten countries. Partial unit root test results, displayed in the second and third columns of the TAR estimation section of the table are by and large in line with the linearity tests. We detect the presence of a partial unit root case for Colombia, Panama, Russia and Turkey. Alternatively, both regimes display a random walk behavior in Argentina, Bulgaria, Mexico, Peru and Philippines. The non-linearity in these latter cases comes forward as a result of the different types of unit root behaviors in these regimes. The results for Brazil are diverse for alternative threshold variables.

The results of the TAR-GARCH (1,1) estimation as suggested by Gospodinov (2008) are depicted on the last column of Table 4. This exercise reports substantially lower rate of rejecting the null of linearity compared to the CH study. Nonlinear mean reversion is suggested only for three countries Argentina, Brazil and Mexico, much less than the previous exercise[10]. These results are in line with the Gospodinov (2008) results, suggesting that the TAR test which neglects the time varying conditional variance problem display considerable size distortions compared to the robust version that includes heteroscedastic structure.

Table 5 reports diagnostic tests using the residuals of TAR Estimation. TAR model provides a significant improvement in skewness and kurtosis statistics in addition with BDS Independence test statistics, however there still exists a significant ARCH effect which constitutes a ground for the use of TAR-GARCH estimation. Table 6 reveals that GARCH

[10] Moreover, TAR-GARCH (1,1) estimation suggests weak evidence of immediate adjustment mechanism for Panama and Bulgaria.

effect is remedied by employing TAR-GARCH model. In addition to that, BDS Independence statistics indicate that the residuals from the TAR-GARCH model are i.i.d.

The results of the TAR-GARCH (1,1) estimation suggest that the adjustment in CDS-bond basis is sharp for Argentina, Brazil and Mexico. Note that, CDS contracts of Argentina have the highest trading frequency in the pre-crises period while those of Brazil and Mexico's are the most liquid ones after the crises, as presented both in Table 1 and Figure 2. This confirms our conjecture that the CDS-bond basis of emerging economies with relatively liquid CDS and bond markets display a sharp correction. Any arbitrage opportunity exploiting the CDS-bond basis (sometimes called as *basis trade* in the market) is wiped out immediately for Argentina, Brazil and Mexico.

These results also suggest that CDS and bonds markets follow a similar pattern to other asset markets at which a deeper liquidity structure implies that the arbitrage opportunity would quickly evaporate, as discussed in the second section.

b) Gradual adjustment (ESTAR model)

Second, we test the conjecture that the market participants would be relatively less responsive to deviations from zero, for sovereign CDS's and bonds with comparably shallower liquidity. We postulate an adjustment mechanism where the deviations from a long-run value are corrected in a gradual manner. To this end, an exponential smooth transition autoregressive (ESTAR) setting similar to Kapetanios et al. (2003) is proposed:

$$\Delta y_t = a_1 y_{t-1} + a_2 y_{t-1} G\left[1 - \exp(-\theta y^2{}_{t-d})\right] + \varepsilon_t \qquad (4)$$

where G(.) is the transition function including the speed of adjustment coefficient θ.

The transition function $G\left[1-\exp(-\theta y^{2}{}_{t-d})\right]$ determines the degree of nonlinearity as a function of the speed of adjustment coefficient θ. Kapetanios et al. (2003) impose a unit root behavior in the middle regime with a_1=0 and select d=1. Then equation (4) turns into:

$$\Delta y_t = a_2 y_{t-1} G\left[1-\exp(-\theta y^{2}{}_{t-1})\right] + \varepsilon_t \qquad (5)$$

The null hypothesis is H_0: θ=0 against the alternative H_1: θ>0. Since a_2 is not identified under the null, an auxiliary regression by employing a first order Taylor series approximation to the ESTAR model under the null can be obtained as:

$$\Delta y_t = \gamma y_{t-1}{}^{3} + error \qquad (6)$$

Then they observe the t-statistics for the null γ=0 against γ<0 by using the OLS estimation of $\gamma(\hat{\gamma})$ and provide asymptotic critical values of this *t* test. Moreover, we examine the general case with serially correlated errors in line with Kapetenios et al (2003), where the auxiliary regression (6) becomes[11]:

$$\Delta y_t = \sum_{j=1}^{p} p_j \Delta y_{t-j} + \gamma y_{t-1}{}^{3} + error \qquad (7)$$

Table 7 reports the results of the ESTAR test. Similar to the TAR exercise, when we take conditional heteroscedasticity into account and use White heteroscedastic consistent (WHC) covariance matrix, t_{NL} values goes down at most of the cases.

The results suggest that the mean reversion of the basis displays a smooth pattern for emerging market economies with relatively less liquid CDS' and bonds such as Colombia, Panama, Peru and Philippines. Turkey is also in that group even though Turkish sovereign

[11] Our selection of the number of autoregressive lags, $0 \leq p \leq 7$, is based on Akaike Information Criterion.

CDS is the most frequently traded sovereign CDS contract in the first and second quarter of 2010, as can be seen in Figure 2.

We could not find any evidence of a nonlinear adjustment pattern for Russia. The recent default episodes of Russia in 1998 might provide an explanation for this decoupling. The default experience might have resulted in a higher market risk sentiment for the country, which pushes CDS premiums to even higher levels in crises times. Therefore, market liquidity might act as a weaker determinant for the asset price fluctuations, compared with default considerations of default countries. Increasing prudence might dominate arbitrage concerns, which might widen the basis at crisis times, and make the mean-reversion much later than expected. Alternative explanations of this behavior deserve a more detailed analysis, which might be the subject of a further study.

V. Conclusions and Future Work

This paper examines the adjustment pattern of the CDS-bond basis, using nonlinear modeling techniques. Our results are in line with the literature that proposes a long-run equilibrium relationship between CDS and bond yield spreads. We further suggest that adjustment towards the mean for the CDS-bond basis follows a nonlinear pattern. Our results suggest that the deviation in CDS-bond basis has a tendency to be eliminated immediately for emerging market economies whose financial assets have a high trading volume. However, this adjustment is gradual for the economies with CDS's and bonds that have a lower trading frequency.

Our results that suggest a mean-reverting behavior for the CDS-bond basis have important implications both for market players and policymakers. From the point of view of investors, exploiting the arbitrage opportunity arising due to the deviation in CDS-bond basis from its fundamental level requires swift action, where the speed of adjustment plays an essential role. From the point of view of policymakers, CDS-bond basis contains information value for both projecting the liquidity conditions in the market and the risk premium implied by the asset prices, as discussed in the second section.

Our results also corroborates with the finding of another strand of literature that suggests a positive relationship between the liquidity structure of assets and speed of arbitrage behavior in various asset markets. It would be interesting to explore this issue with alternative techniques in more detail, once we have better liquidity indicators.

Moreover, our study confirms that the presence of conditional heteroscedasticity has a substantial impact on the linearity tests. Our results suggest that the TAR-GARCH representation and the use of WHC errors in ESTAR testing procedure reduce the size distortions and improve the power of these tests. However, one caveat that needs to be addressed is about the degree of robustness of these tests. A recent study by Pavlidis et.al (2010) suggests that while WHC errors generates better results than conventional tests, they still exhibit severe size distortions and reveal low power. They propose using bootstrap techniques. Future work aiming to develop these procedures would be pertinent and valuable for more robust linearity testing.

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Data Appendix

Tickers of the CDS and EMBI+ data are taken from Bloomberg database. Initial data value dates are provided in the table below. All series end at 01/08/2010.

	CDS Ticker	EMBI+ Ticker	Start	# of Obs.
Argentina	ct350188	jpssemar	06/30/05	1131
Brazil	cbrz1u5	jpssembr	10/16/01	1984
Bulgaria	cbulg1u5	jpssembu	10/25/00	2275
Colombia	ccol1u5	jpssemco	01/27/03	1719
Mexico	cmex1u5	jpssemme	10/16/01	1984
Panama	cpan1u5	jpssempa	11/04/03	1534
Peru	cperu1u5	jpssempe	10/21/03	1541
Philippines	cphil1u5	jpssemph	04/12/02	1770
Russia	cruss1u5	jpssemru	10/24/00	2283
Turkey	cturk1u5	jpssemtu	10/24/00	2282

Figure 1: CDS-Bond Basis

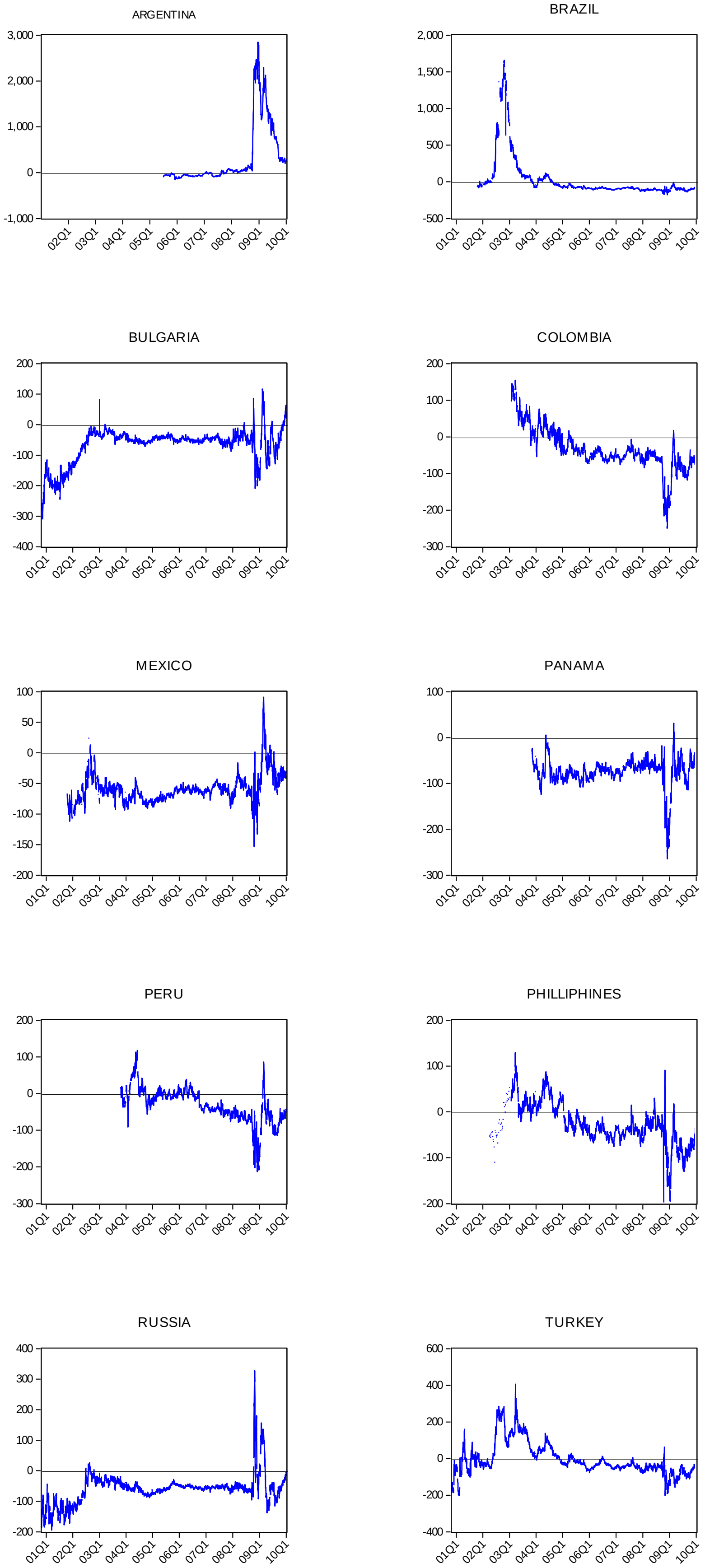
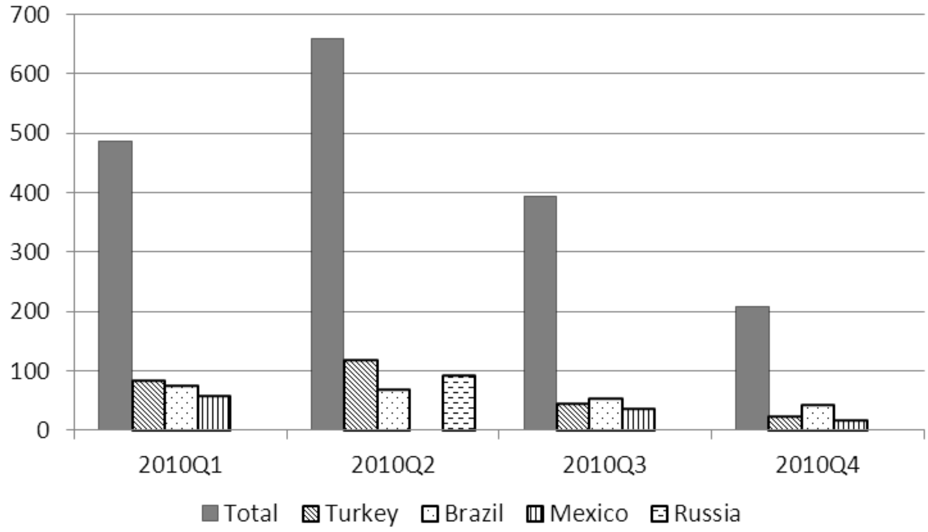


Figure 2: Trading Volume of the Emerging Market Sovereign CDS's (US$ billion).



Source: Emerging Market Trade Association (2011)

Table 1: CDS Bid-Ask Spreads

a) Whole Sample

	Liq. average	CDS average	Liq. / CDS %
Colombia	**12.0**	279.7	4.3
Philippines	**14.6**	314.6	4.6
Turkey	**22.0**	453.4	4.8
Bulgaria	**10.2**	205.1	5.0
Brazil	**32.5**	524.0	6.2
Russia	**18.9**	303.3	6.2
Peru	**14.1**	204.8	6.9
Argentina	**75.5**	961.3	7.8
Mexico	**13.0**	143.7	9.0
Panama	**20.1**	188.1	10.7

b) Pre-Crisis

	Liq. average	CDS average	Liq. / CDS %
Argentina	9.3	355.5	2.6
Bulgaria	7.6	177.3	4.3
Philippines	13.8	322.0	4.3
Colombia	12.9	285.9	4.5
Turkey	24.1	479.1	5.0
Brazil	38.2	589.5	6.5
Russia	19.7	285.2	6.9
Peru	14.6	194.9	7.5
Mexico	14.3	122.3	11.7
Panama	22.3	167.6	13.3

c) Post-Crisis

	Liq. average	CDS average	Liq. / CDS %
Mexico	7.3	241.9	3.0
Brazil	7.0	227.3	3.1
Turkey	10.6	315.1	3.4
Colombia	8.6	256.1	3.4
Russia	14.5	400.9	3.6
Panama	13.1	255.2	5.1
Peru	12.6	237.6	5.3
Philippines	17.9	284.8	6.3
Bulgaria	24.6	355.3	6.9
Argentina	218.7	2272.6	9.6

Table 2: Unit Root Tests

Series	Level		Level		Long Diff.		Long Diff.		Lagged Diff.		Lagged Diff.	
Test	ERS		PP		ERS		PP		ERS		PP	
Argentina	0.33		-0.75		-3.37	***	-16.97	***	-1.54		-31.93	
Brazil	-1.59		-2.15		-7.38	***	-24.69	***	-8.23	***	-38.36	***
Bulgaria	-1.33		-1.60		-8.14	***	-29.24	***	-21.99	***	-47.98	***
Colombia	-0.35		-3.14	**	-8.28	***	-44.83	***	-16.99	***	-80.83	***
Mexio	-0.98		-2.05		-2.14	**	-33.67	***	-1.58		-56.72	
Panama	-0.85		-1.33		-6.89	***	-30.43	***	-26.73	***	-55.54	***
Peru	-1.63	*	-2.43		-2.64	***	-36.89	***	-3.62	***	-64.22	***
Philippines	-2.19	**	-4.22	***	-5.07	***	-30.21	***	-40.30	***	-49.84	***
Russia	-1.14		-1.83		-5.40	***	-40.93	***	-2.20	**	-67.19	**
Turkey	-1.09		-2.51		-2.63	***	-39.05	***	-2.37	**	-66.86	**

*Note: ERS denotes Elliot, Rothenberg, and Stock ADF-GLS and PP denotes Phillips-Perron unit root test statistics. ***, ** and * stand for 1%, 5% and 10% significance levels respectively.*

Table 3: Preliminary Diagnostics

	Argentina	Brazil	Bulgaria	Colombia	Mexio	Panama	Peru	Philippines	Russia	Turkey
CDS-Bond Basis										
ρ_1 [Q-test p value]	0.99 [0.000]	0.99 [0.000]	0.99 [0.000]	0.96 [0.000]	0.99 [0.000]	0.99 [0.000]	0.98 [0.000]	0.94 [0.000]	0.99 [0.000]	0.99 [0.000]
ρ_{12} [Q-test p value]	0.97 [0.000]	0.92 [0.000]	0.98 [0.000]	0.91 [0.000]	0.96 [0.000]	0.98 [0.000]	0.93 [0.000]	0.83 [0.000]	0.97 [0.000]	0.95 [0.000]
ρ_{24} [Q-test p value]	0.95 [0.000]	0.84 [0.000]	0.95 [0.000]	0.83 [0.000]	0.92 [0.000]	0.96 [0.000]	0.89 [0.000]	0.72 [0.000]	0.95 [0.000]	0.88 [0.000]
ρ_{48} [Q-test p value]	0.90 [0.000]	0.67 [0.000]	0.91 [0.000]	0.80 [0.000]	0.85 [0.000]	0.91 [0.000]	0.83 [0.000]	0.50 [0.000]	0.88 [0.000]	0.76 [0.000]
Residuals of AR(1)										
Skewness	0.012	8.530	-6.332	-1.060	-6.593	-2.001	0.238	-1.689	-2.821	5.568
Kurtosis	4.467	163.845	166.742	190.838	400.358	169.690	175.120	233.553	220.480	144.756
BDS Indep. Test										
Dimension = 1	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Dimension [1,150]	[0.133]	[0.577]	[0.960]	[0.939]	[0.697]	[0.988]	[0.991]	[0.000]	[0.000]	[0.982]
	(19)	(123)	(98)	(143)	(62)	(94)	(95)	(150)	(150)	(78)
ARCH-LM (*p* value)	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
At Lag	1	1	1	1	1	1	1	1	1	1
ARCH parameter (SE)	0.990***	0.659***	0.485***	0.789***	0.731***	0.764***	0.577***	1.365***	0.242***	0.769***
	(0.263)	(0.103)	(0.053)	(0.063)	(0.092)	(0.043)	(0.056)	(0.063)	(0.032)	(0.091)
GARCH parameter (SE)	0.005	0.359***	0.538***	0.271***	0.275***	0.294***	0.440***	0.126***	0.783***	0.248***
	(0.137)	(0.091)	(0.033)	(0.041)	(0.053)	(0.037)	(0.031)	(0.026)	(0.017)	(0.058)
Andrews Sup-LM	6.062	1.439	22.052***	8.861	2.784	6.777	3.72	17.077***	11.300*	15.868***
Andrews Wald	1.468	2.417	8.808	17.899***	10.558*	9.498	1.797	10.208*	10.400*	12.777**

Table 4: TAR and TAR-GARCH Estimation Results (p-values)

	Difference	m	Tar Estimation W_u		t_{1u}	t_{2u}	Tar-Garch Estimation LM_G	
Argentina	**long**	**2**	0.000	***	0.845	0.164	0.014	**
Brazil	**long**	**2**	0.005	***	0.016	0.275	0.233	
	lagged	**1**	0.045	**	0.133	0.399	0.049	**
Bulgaria	**long**	**2**	0.000	***	0.219	0.943	0.065	*
	lagged	**1**	0.000	***	0.219	0.943	0.115	
Colombia	**long**	**2**	0.000	***	0.144	0.014	0.326	
	lagged	**1**	0.008	***	0.913	0.001	0.101	
Mexico	**long**	**2**	0.324		0.949	0.316	0.021	**
Panama	**long/lagged**	**1**	0.000	***	0.785	0.010	0.058	*
Peru	**long/lagged**	**1**	0.029	**	0.436	0.213	0.310	
Philippines	**long**	**2**	0.000	***	0.361	0.278	0.926	
	lagged	**1**	0.000	***	0.361	0.278	0.936	
Russia	**long**	**2**	0.000	***	0.016	0.001	0.503	
	lagged	**2**	0.001	***	0.000	0.956	0.438	
Turkey	**long/lagged**	**1**	0.002	***	0.829	0.000	0.227	

*Note: m is the optimal delay parameter that is endogenously selected. W_u and W_c stand for p-values of Wald tests for unconstrained and constrained models, respectively. Similarly t_{1u}, t_{2u}, t_{1c}, t_{2c} stands for p-values for the unconstrained and constrained models, respectively. LM_G is the p-value of the Sup-LM test for TAR-GARCH estimation. ***, ** and * stand for 1%, 5% and 10% significance levels, respectively.*

Table 5: TAR Model Diagnostics

	Argentina	Brazil	Bulgaria	Colombia	Mexio	Panama	Peru	Philippines	Russia	Turkey
Residuals of TAR Model (Threshold Variable is Lagged Difference)										
Skewness	5.797	0.300	0.398	-0.144	0.692	0.219	-0.129	0.098	0.356	0.196
Kurtosis	122.579	13.058	13.231	6.890	10.693	7.900	13.685	10.027	6.972	6.696
BDS Independence Test										
Dimension = 1	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Dimension [1,150]	[0.466]	[0.566]	[0.991]	[0.435]	[0.439]	[0.177]	[0.349]	[0.384]	[0.437]	[0.307]
	(27)	(31)	(72)	(25)	(26)	(20)	(25)	(27)	(26)	(25)
ARCH-LM (p value)	[0.015]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
At Lag	4	1	1	1	1	1	1	1	1	1
Andrews Sup-LM	3.032	4.751	1.440	3.547	1.926	1.734	5.089	2.918	3.072	4.828
Andrews Wald	3.588	5.225	4.552	4.793	2.488	3.589	7.049	8.587	4.274	7.078
Residuals of TAR Model (Threshold Variable is Long Difference)										
Skewness	3.475	0.483	0.244	0.029	0.462	0.172	-0.091	0.132	0.363	0.220
Kurtosis	62.582	13.511	11.603	6.014	7.818	7.335	9.799	12.264	6.638	6.121
BDS Independence Test										
Dimension = 1	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Dimension [1,150]	[0.569]	[0.480]	[0.992]	[0.492]	[0.653]	[0.351]	[0.224]	[0.382]	[0.431]	[0.335]
	(28)	(29)	(70)	(28)	(32)	(25)	(22)	(28)	(27)	(26)
ARCH-LM (p value)	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
At Lag	1	1	1	1	1	1	1	1	1	1
Andrews Sup-LM	2.451	3.791	5.121	5.999	2.104	10.015	4.792	4.702	1.015	2.547
Andrews Wald	2.538	2.197	3.082	2.512	3.206	3.353	4.681	6.855	1.011	2.686

Table 6: TAR-GARCH (1,1) Model Diagnostics

	Argentina	Brazil	Bulgaria	Colombia	Mexio	Panama	Peru	Philippines	Russia	Turkey
Residuals of TAR-GARCH(1,1) Model (Threshold Variable is Lagged Difference)										
Skewness	4.353	0.140	2.648	-0.032	0.328	0.242	0.113	-0.120	0.042	0.271
Kurtosis	101.236	6.543	52.455	4.473	5.738	5.477	7.628	5.420	4.587	4.825
BDS Independence Test										
Dimension = 1	[0.000]	[0.462]	[0.000]	[0.091]	[0.000]	[0.128]	[0.784]	[0.053]	[0.018]	[0.110]
Dimension [1,150]	[0.478]	-	[0.293]	[0.168]	[0.354]	-	-	[0.138]	[0.126]	-
	(12)	-	(25)	(2)	(15)	-	-	(2)	(13)	-
ARCH-LM (p value)	[0.910]	[0.691]	[0.577]	[0.512]	[0.178]	[0.276]	[0.804]	[0.136]	[0.120]	[0.983]
At Lag	1	1	1	1	1	1	1	30	1	1
Andrews Sup-LM	6.504	2.186	2.180	3.951	5.484	3.237	4.646	7.766	2.517	5.690
Andrews Wald	6.918	2.246	1.656	4.361	4.437	2.942	4.940	9.148	2.899	1.007
Residuals of TAR-GARCH(1,1) Model (Threshold Variable is Long Difference)										
Skewness	3.245	0.068	2.161	-0.006	0.233	0.172	-0.025	-0.055	0.116	0.147
Kurtosis	67.047	5.500	39.019	3.878	4.833	4.513	5.456	4.928	4.168	4.185
BDS Independence Test										
Dimension = 1	[0.631]	[0.766]	[0.000]	[0.320]	[0.660]	[0.649]	[0.961]	[0.110]	[0.382]	[0.587]
Dimension [1,150]	-	-	[0.223]	-	-	-	-	-	-	-
	-	-	(24)	-	-	-	-	-	-	-
ARCH-LM (p value)	[0.918]	[0.743]	[0.894]	[0.602]	[0.810]	[0.613]	[0.601]	[0.050]	[0.529]	[0.983]
At Lag	1	1	1	1	1	1	1	1	1	1
Andrews Sup-LM	1.745	1.338	1.711	1.873	5.068	2.070	2.838	3.048	1.495	5.687
Andrews Wald	1.638	1.222	1.590	1.980	3.981	1.990	3.015	3.429	1.707	1.007

Table 7: ESTAR Estimation Results

	ESTAR Estimation t_{NL}		ESTAR Estimation with WHC covariance matrix t_{NL}	
Argentina	-2.51		-2.23	
Brazil	-2.12		-1.60	
Bulgaria	-3.66	**	-2.31	
Colombia	-3.42	**	-4.10	***
Mexico	-2.68		-2.78	
Panama	-3.46	**	-3.65	**
Peru	-4.13	***	-4.55	***
Philippines	-4.21	***	-4.00	***
Russia	-2.62		-2.08	
Turkey	-3.91	**	-3.84	**

Note: Critical values for 1%, 5% and 10% are -3.93, -3.4 and -3.13 respectively.
****, ** and * stand for 1%, 5% and 10% significance levels respectively.*

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