Sovereign Credit Default Swaps, Sovereign Debts and Volatility Transmissions across Emerging Markets

Francis In*

Department of Accounting and Finance, Monash University

Phone: 61-3-9905-1561

E-mail address: Francis.In@BusEco.monash.edu.au

Byoung Uk Kang

Graduate School of Management, Korea Advanced Institute of Science and Technology

Phone: 82-2-958-3689

E-mail address: wayfarer@kgsm.kaist.ac.kr

Tong Suk Kim

Graduate School of Management, Korea Advanced Institute of Science and Technology

Phone: 82-2-958-3018

E-mail address: tskim@kgsm.kaist.ac.kr

June, 2006

* Corresponding author, Francis In, Department of Accounting and Finance, Monash University, Clayton 3168, Victoria, Australia Tel : 61-3-9905-1561, Fax : 61-3-9905-5475, E-mail address : Francis.In@BusEco.monash.edu.au

Abstract

In this paper, we use sovereign credit default swap (CDS) prices and emerging market government bond credit spreads to study and compare their abilities in information processing and transmission across nations. In contrast to approaches used in previous studies, first we identify the long-run co-movements between the sovereign CDS prices and their corresponding credit spreads using the Canonical Cointegration Regression (CCR), and carry out an analysis on short-run dynamics including price discovery and volatility spillovers using the bivariate CCR-ECM with an EGARCH specification. We then investigate the credit spreads inter-linkages across the major emerging market debtors located in Latin America. Advantage of the multivariate EGARCH model enables us to detect the volatility transmission mechanisms, which gives valuable insights into market participants' perception towards emerging market credit worthiness.

1. Introduction

Recent improvement in fundamentals of emerging economies has strengthened investors' enthusiasm towards emerging markets. Cash inflows have been rallied and in 2005 have $160 billion been invested through debt securities, well above $60 billion that emerging equity market attracted in the same year. This underlines the importance of bonds as major financing vehicle for emerging economies. Also the emerging market bonds have been widely perceived by portfolio managers in the U.S. as a primary mean for diversifying into emerging markets.

As the volume of emerging market bonds has soared to its historical high both in terms of trading and issuance during the last decade, the world credit market has witnessed a new financial innovation, namely sovereign credit default swaps (CDS). After the standardisation of contract format and definitions in the late 1990s and the successful settlement following recent defaults, the sovereign CDSs are now an actively traded credit derivative to the extent that for some sovereign entities, the CDS offers more liquidity than their underlying sovereign bonds.

Despite the emerging bond market's considerable proportions and its significance to emerging economies as well as U.S. mutual, pension, and endowment funds, much of the latest theoretical as well as empirical works have analysed equity rather than bond markets. Moreover, research on emerging sovereign bonds' empirical relationship with sovereign CDSs is still miniscule compared to the literature on corporate level analysis. To make a contribution on this direction, we use sovereign CDS prices and emerging market government bond credit spreads to study and compare their abilities in information processing and transmission across nations.

Considering the recent remarkable growth of the global credit derivative market, there are also very few empirical studies on the relationship between credit derivatives and bond markets.[1] In analysing the dynamic behaviour of CDS prices and credit spreads, one of the important issues is to evaluate the information content of indicators of the price of credit risk. Hence, it is important to understand which market provides more efficiently and timely information. Since this price discovery is one of the central functions of derivative market, recently Blanco et al. (2005) and Zhu (2004) use corporate bonds to investigate their long and short-run relationship with corresponding CDS prices and concluded that short-term deviations from the theoretical parity are brought about by a lead for CDS prices over credit spreads. Similarly, Chan-Lau and Kim (2004) examine equilibrium price relationships and price discovery between credit default swaps (CDS), bond, and equity markets for emerging market sovereign issuers. Using sovereign CDS and JP Morgan Chase Emerging Market Bond Index Plus (EMBI+), they find long-run equilibrium price relationship between CDS and bond spreads but do not find any equilibrium price relationship between bond spreads and equity prices in most countries.

Beyond the price discovery, the emergence of CDS market opens up new pathway for empirical research on credit risk. As Hull et al. (2004) acknowledged, the CDS spreads are an interesting alternative to bond prices since the CDS spread data consists of firm bid and offer quotes from dealers. And also, unlike in the case of bond

[1] Duffie (1999) derived the theoretical equivalence of credit default swap prices and credit spreads under a number of simplifying conditions. With some of the assumptions relaxed, an alternative formula is proposed by Hull and White (2000a, b).

yields where an assumption should be made about the appropriate benchmark risk-free rate, the fact that the CDS spreads are already credit spreads is another attraction. In addition, we recognise two more advantage in using CDS spreads over bond yields for the purpose of credit risk research. When using bond yields to extract credit spreads over risk-free rate, researchers should be aware that some features embedded in bonds could deteriorate the correct reflection of bond issuer's credit worthiness. Usual practice in this line of research, therefore, excludes floating-rate bonds and bonds with step-up coupon, sinking funds, and embedded options such as callable, puttable or convertible bonds. As a result, very few bonds remain available for a given issuer, and this in many cases restricts the scope of analysis. These difficulties, of course, can be bypassed by the use of CDS prices and therefore it enables inclusion of a wider cross section of reference entities.[2] Moreover, a majority of CDS trading volumes is concentrated on relatively small number of product for each reference entity whereas sovereign or corporate bond issuers typically have several outstanding issuances so that analysis using CDSs can be more resilient to illiquidity problems.[3] With these advantages that CDS prices can offer, previous research on credit risk using bond yields can be extended, possibly with improved results. Hull et al. (2004), for example, used CDS prices to examine the relationship between the CDS market and credit rating announcements.

[2] Blanco et al. (2005) ended up with only 33 reference entities for their analysis, which is a small subset of 157 U.S. and European reference banks and companies in the database that they had started with. Of these, more than twice as many reference entities are dropped due to lack of bond data rather than insufficient CDS data.

[3] For corporate entities, trading has been concentrated largely in the five-year maturity contract, and for sovereign reference entities, CDSs with two or three different maturities are mainly traded.

This paper aims to contribute to the literature in the three *innovative* ways: First, the paper investigates volatility spillovers *between* sovereign CDS prices and government debt spreads in emerging markets. In addition to the price discovery process utilizing common factor models such as Hasbrouck (1995) and Gonzalo and Granger (1995), this will foster depth in our understanding of information transmission between the two indicators of credit risk. Second, the study examines volatility spillovers *across* the credit spreads of the four major Latin American debtors, namely Brazil, Colombia, Mexico and Venezuela. To the best of our knowledge, volatility spillover effects *across* the emerging credit derivative markets have not yet been recorded in the literature. Yet it is an intriguing area of study, because it provides insight into information transmission, volatility impacts, and the pricing of CDS. So, this paper, using CDS data, is the *first* to investigate the issue of volatility spillovers across emerging market credit spreads. The result of the analysis shed some lights on investors' perceived risks towards emerging markets; whether they are treating all countries in the same region as equal or are adept at distinguishing between nations with and without sound fundamentals. Third, to meet our specific research objective, we apply a multivariate exponential generalized autoregressive conditional heteroskedasticity (EGARCH) model for the volatility spillovers of CDS premiums. Our multivariate EGARCH methodology incorporates two error correction components in the mean equation: one from long-run relationship with corresponding bond credit spread, the other from long-run relationship with CDS prices of the neighbouring countries. Unlike other previous research, both error correction terms are estimated from the Canonical Cointegrating Regression (CCR) technique of Park (1992). In fact, this paper is the *first* to estimate the long-run relationships between CDS and bond spreads in emerging credit derivative markets using the CCR technique developed by Park (1992). The multivariate CCR-VECM-EGARCH modelling will capture the time-varying nature of CDS price volatility and provide evidence on the volatility transmission mechanism. The use of the CCR-VECM with an EGARCH specification is well-designed to accommodate the more subtle relationships within the volatility structure. One minor improvement of our research comes from the unique dataset on bond credit spreads. After filtering out inappropriate bonds for the analysis, we interpolate bond yield to match the constant maturity of the CDS contracts. By doing so, we eliminate the potential weakness in using bond spread indices such as EMBI+ as the EMBI+ country sub-indices have much higher durations than maturities of CDS contracts.

The rest of the paper is organised as follows. Section 2 discusses the data and provides descriptive statistics. Section 3 investigates price discovery and volatility spillovers between sovereign CDS market and sovereign debt market. Section 4 presents the empirical tests on volatility transmission across the emerging markets and discusses the findings. Finally, section 5 provides a summary and conclusion.

2. Data and descriptive statistics

For CDS prices, we use daily indicative mid-market price data on U.S. dollar denominated sovereign CDSs with maturity of five years. The data are purchased from

CreditTrade, one of the main inter-dealer brokers.[4] To match the constant five-year maturity of the CDS contract, we calculate five-year bond yields by interpolation. Bonds used in interpolation are restricted to the most standard category of bonds: those with a fixed interest rate, no collateral or third-party guarantor and no warrants or embedded options. This category practically excludes Brady bonds from the data set since the face value of Brady bonds is guaranteed by the U.S. government.[5] We, therefore, use only government issued Eurobonds denominated in U.S. dollars. Having passed the filtering process, daily mid-rate yields of selected bonds are downloaded from Datastream, and used to construct the time series of interpolated five-year yields. They are matched and interpolated by using the rule outlined in the footnote.[6]

From the constructed series of five-year yields, we then compute the credit spread of emerging market bonds by subtracting five-year U.S. treasury mid-market

---

[4] They provide reference prices for marking-to-market existing transactions, based on averages of prices supplied by dealers and/or on trade prices in the inter-dealer market. See Rule (2001)

[5] Kamin and von Kleist (2005) have shown that the presence of collateral causes additional and possibly distortive complexity and higher transaction costs.

[6] The rules are numbered in order of priority.

1) At each given date, two bonds are selected. One of them should have shorter and the other should have longer remaining term to maturity than the maturity of CDS contract. If the remaining terms to maturity of both fall between 3.5 years to 6.5 years, five-year yields are calculated by linear interpolation.
2) When no interpolation is possible, bond whose maturity is between 4.5 years and 5.5 years is used as proxy for five-year yield.
3) If at least one of the bonds has remaining term to maturity between 3.5 and 6.5 years, calculate the 5-year yield by linear interpolation.
4) If no interpolation is possible, use bond whose maturity is between 4 years and 6 years as proxy for five-year yield.

At each level of hierarchy, if a choice is available, we select bonds traded close to par. If a choice remains, bonds whose maturity is closer to five years are selected.

yields.[7] The combined time series of sovereign CDS prices and emerging market bond credit spreads constitute data used in our analysis.

The data run from May 1, 2003 through April 1, 2006, giving a sample size of 756 observations for each sovereign entity. Our coverage for sample reference sovereign entities and maturities is limited mainly due to lack of eligible bonds rather than that of CDS prices, leaving us 7 out of total 20 emerging market debtors available in our original CDS database. Those seven sovereign names are Brazil, Colombia, Mexico, Philippines, Russia, Turkey and Venezuela. Argentina is the only exception as the lack of CDS prices, not bond yields is to be blamed for its exclusion. There has been a period of no trading following its default in late 2001. Consequently, in Argentina, sovereign CDSs suffered a temporary setback in 2002 and 2003[8]. Other notable absentees include China and South Korea. For Chian, we observe time period when the price seises of two bonds does not change, which can be interpreted that the bonds do not traded at all during the period. After eliminating such observations, we are not able to compute the interpolated five-year yields during the period from 5th July, 2003 through 20th October, 2004. Also, we are reluctant to exclude South Korea considering its economic status in emerging Asian markets. To cover up the lack of Korean government bond issuance, we have tried bonds issued by Korea Development Bank in addition to government issues, but there seem to be non-negligible shifts in yield when

[7] Unlike in the case of corporate CDSs, we find that both average basis and average absolute basis is smaller when U.S. Treasury bond is used than in the case where five-year swap rate is used as benchmark risk free rate. This finding may be in line with Singh and Andritzky (2005)'s view of which sovereign CDSs are overpriced. See table 2.

[8] There was no quote on Argentine sovereign CDSs in 2002 and there were only 6 quotes in 2003. In the indicative BenchMark price data of CreditTrade, the Argentine name does not appear until June 2005.

we move from Korean government bonds to KDB bonds or *vice versa*.[9] Such jumps in yield with no change in the underlying riskiness of the nation could mislead implications.

The descriptive statistics for the both CDS prices and bond credit spreads for seven emerging countries are reported in Table 1. The preliminary analysis for establishing the time-series properties of individual CDS prices and bond credit spreads involves undertaking the Park-Choi (1998) unit root test. These statistics are presented in Table 1. The PC test statistics show no evidence against the null hypothesis that there is a unit root in the series, however, though it is not reported in the table, the data clearly reject the null hypothesis that there is a unit root in the first differences. A resealable conclusion from these results is that each indicator of sovereign credit risk is an *I(1)* process with all of tests supporting the unit root hypothesis at the any level of significance for most of the data series.

The measures for skewness and kurtosis are also reported to indicate whether the first differences in CDS prices and bond credit spreads are normally distributed or not. The sign of skewness varies between countries. The Jarque-Bera (denoted by JB) statistic rejects normality at any level of statistical significance in all cases. The Ljung-Box statistics for 24 lags applied on the series (denoted by Q(24) and squared term (Q^2(24)) indicate that significant linear and nonlinear dependencies exist. Nonlinear dependencies can be captured satisfactorily by autoregressive conditional heteroskedasticity models.

[9] Korea Development Bank is supported by government guarantee. So their bonds were expected to be a good proxy for sovereign bonds, which turns out not to be the case.

To test for cointegration between the two indicators of sovereign credit risk, we employ a procedure developed by Park-Ouliaris-Choi (POC). The $J_1(p,q)$ test statistics of POC is designed to reject the null hypothesis of cointegration. In other words, the test rejects the null hypothesis of cointegration in favour of the alternative of no cointegration when $J_1(p,q)$ is large, under a limiting chi-square distribution (χ^2) with degree of freedom equal to $q - p$.[10] The results of the cointegration test are reported in Table 3. In the table, $J_1(0,1)$ statistics show that there is no deterministic time trend in cointegration equations except in the case of Russia. Therefore, $J_1(1,5)$ statistics is used for Russian case while $J_1(0,3)$ are used for the rest of the sovereign names. Every J_1 statistics supports existence of the long run equilibrium relationship between sovereign CDS prices and bond credit spreads.

3. Price discovery and volatility spillovers

Price discovery between CDS prices and bond credit spreads has been investigated by recent literature. Blanco et al. (2005) and Zhu (2004), using investment-grade corporate reference entities, found that the CDS market leads the bond market in processing the information of the credit risk. However, as it is mentioned in Blanco et al. (2005), the results are representative of their relatively short span of time and only the investment-grade corporate reference entities have been analysed. Chan-Lau and Kim

[10] Formulating the test with the presence of cointegration as the null hypothesis causes no problem, contrary to other existing residual-based tests such as Phillips test. The J_1 test is in fact the first non-parametric test for the null of cointegration. No other tests of cointegration as general as Park (1992) have yet been developed in the literature.

(2004) is the only study so far on this issue with sovereign reference entities. However, their result shows, in many cases, wide gaps between the Hasbrouck lower and upper bounds, and also the average of two bounds do not provide similar results with Granger and Gonzalo's statistic. These problems seem to be brought about by highly correlated residuals from a vector error-correction model (VECM), which the both measures are based on.[11] Thus, for a given sovereign entity, the implication about which market provides more timely information is not clear.

To overcome the previous spurious results, this paper attempts to improve the analysis in three ways. First, the error correction term in the VECM is generated following the CCR estimation procedure of Park (1992). As Table 1 reports, all series have unit roots and cointegrated within each sovereign name. Therefore, standard Ordinary Least Squares (OLS) estimates would be superconsistent, but their limiting distributions would be biased and inefficient. Park and Ogaki (1991) and Park (1992) proved that OLS estimates with nonstationary regressors are inefficient and their distributions are asymptotically biased and contain nuisance parameters. In order to avoid the loss of power and nuisance parameter problems, we employ the CCR method and report the estimated coefficients in Table 3. Second, instead of relying on indices of credit risk such as EMBI+ country sub-indices, we interpolate yields of carefully selected bonds, as outlined in the previous section, to match the maturity of CDS. Ignoring this may render different results because the EMBI+ country sub-indices have much higher durations than maturities of CDS contracts, and credit spreads might vary with the maturity if investors have expectations about the probable timing of any default

[11] Ballie et al. (2002) and De Jong (2002) showed that Hasbrouk (1995) and Gonzalo and Granger (1995) models provide similar results if the residuals are uncorrelated with similar varianes.

as well as its likelihood.[12] Third, we examine the volatility spillover mechanism with a bivariate CCR-VECM-EGARCH model to deepen our understanding of information transmission between the two indicators of credit risk.

The bivariate CCR-VECM-EGARCH model is written as follows:

$$\Delta p_{CDS,t} = \lambda_1 \hat{u}_{t-1} + \sum_{j=1}^{p} b_{1j} \Delta p_{CDS,t-j} + \sum_{j=1}^{p} c_{1j} \Delta p_{CDS,t-j} + \varepsilon_{1,t}$$

$$\Delta p_{CS,t} = \lambda_2 \hat{u}_{t-1} + \sum_{j=1}^{p} b_{2j} \Delta p_{CDS,t-j} + \sum_{j=1}^{p} c_{2j} \Delta p_{CDS,t-j} + \varepsilon_{2,t}$$

$$h_t^{CDS} = \exp\left\{\alpha_{1,0} + \alpha_{1,1} f(Z_{t-1}^{CDS}) + \alpha_{1,2} f(Z_{t-1}^{CS}) + \gamma_1 \ell n(h_{t-1}^{CDS})\right\}$$
$$h_t^{CS} = \exp\left\{\alpha_{2,0} + \alpha_{2,1} f(Z_{t-1}^{CDS}) + \alpha_{2,2} f(Z_{t-1}^{CS}) + \gamma_2 \ell n(h_{t-1}^{CS})\right\}$$

$$f(Z_{t-1}^{j}) = \left(|\, Z_{t-1}^{j} \,| - E(|\, Z_{t-1}^{j} \,|) + \delta_i Z_{t-1}^{j}\right) \ \ for \ \ j = CDS, Credit\ Spread$$

$$\sigma_{j,k,t} = \rho_{j,k} \sigma_{j,t} \sigma_{k,t} \qquad for \ \ j,k = CDS, Credit\ Spread \ \ and \ \ j \neq k$$

The error correction term, $\hat{u}_t$ is measured by the CCR method and it is the very mechanism through which the two measures price credit risk equally in the long-run. The number of lags in the mean equation is determined using the AIC for each sovereign entity, which sometimes to be as large as 21 lags. We find that estimating the mean and variance equations simultaneously is impractical with this large number of coefficients. Thus, two-stage approach is employed, where in the first step the CCR-VECM is estimated and then in the second step, the bivariate EGARCH is estimated using the uncorrelated residuals from the CCR-VECM. As Tse (1999) argues, this two-step approach is asymptotically equivalent to a joint estimation of the CCR-VECM and

[12] See Cunningham et al. (2001)

EGARCH models because the least squares estimator used in the VECM is still unbiased and consistent in the presence of heteroskedasticity. The price discovery statistics are reported in Panel A of Table 4, and estimated result for the EGARCH is in Panel B of Table 4.

We find that in most emerging markets the upper and the lower bounds of Hasbrouck measure lead to the same conclusions. Moreover, for Russian and Turkish names where the two bounds are parted by 0.5 in between, the average of the two bounds provides the very similar result as the Gonzalo and Granger statistics. Therefore, unlike in the previous research with sovereign CDSs, the result of our analysis gives very clear implication about which market provides more timely information for a given sovereign entity. Our result shows that for three Latin American sovereign names such as Brazil, Mexico and Colombia, the CDS market plays a leading role in price discovery, while for Turkish and Venezuelan cases the CDS market moves afterwards to correct for price disequilibrium. Our finding coincides with the fact that Brazil, Mexico and Colombia are the top three reference entities that sovereign CDS trading activities are most concentrated on.[13] Another notable finding is that in contrast to the previous studies on corporate reference entities, where CDS market was found to be dominant in most of price discovery processes, four out of seven emerging markets see the information about their credit worthiness are, either first captured by bond markets or nearly equally reflected in both cash and derivatives markets. This might be explained in line with the notion that price discovery will occur in a more liquid market since on average sovereign issues are likely to be more liquid than corporate counterpart. In fact,

[13] See Packer (2003)

our sample countries constitute some of the major emerging market debtors with very liquid bond markets.

As Ross (1989) argues, volatility rather than the change in price is the one related to the rate of information flow to the market. Therefore, the coefficients α_{12} and α_{21} are of great importance as they describe the information transmission from one market to the other. The estimation results support the existence of volatility spillover between the two markets of credit indicator. In fact, for Mexican and Russian credit risks, the volatility in more informative market spills over to less informative one in the price discovery. However, for four emerging credit markets even where there is a strong lead of one market ahead of the other in the price discovery process, reciprocal spillovers between CDS and bond markets are identified. (for example, Brazil, Colombia, Turkey and Venezuela.) That is, innovations in one market can predict the future volatility in another market regardless of which market is the main forum for price discovery. Moreover, the asymmetric volatility coefficients δ_1 in many cases are significantly positive, implying that previous bad news (represented by positive innovations) in the CDS market will increase its own volatility (i.e., Philippines and Russia) or increase the volatility in both the CDS and the bond market (i.e., Brazil and Mexico) more than good news (represented by negative innovations). Also, the market-specific volatility clustering coefficients α_{11}, α_{22}, γ_1 and γ_2 are positively significant in both markets across all emerging nations.

4. Volatility Transmission across Emerging markets

Traditional window for viewing investors' perceived risks towards emerging markets has been expanded with a successful introduction of sovereign CDS contracts. Since the standardisation of contract format and definitions in the late 1990s and the successful settlement following recent defaults, the sovereign CDS market now offers more liquidity than their underlying sovereign bonds for some emerging market sovereign entities. Moreover, as it is found in the previous section, the CDS market leads the bond market in determining the price of credit risks for some major Latin American entities. Despite the CDS market's popularity and efficiency in processing the credit risk information, there has been very few research on emerging market credit risk using sovereign CDS data. To the best of our knowledge, co-movement of credit spreads and volatility spillovers across emerging markets have not yet been studied using CDS prices in the literature.

In this section, we examine the credit risk inter-linkage across four major Latin American debtors, namely Brazil, Colombia, Mexico and Venezuela. The region is the centre of interest for many emerging market investors since it accounts for around 60 percent of all emerging sovereign debts outstanding.[14] Investigation in this issue is of importance because it provides an insight in the nature of any shock to the Latin American capital markets. More specifically, we seek some evidence for or against the view that investors perceive a shock as a common emerging market event, as if they treat all countries in the region as equal, or more creditworthy countries with sound

[14] Source: BIS website

macroeconomic fundamentals are insulated from the shock.[15] We also attempt to compare the nature of credit spread inter-linkage across the Latin American sovereigns through two alternative credit indicators; the sovereign CDS prices and the government bond credit spreads. If one of the two markets is found to be less prone to the investors' sentiments, and as a result, more immune to the contagion-like phenomenon, it would give an important implication for investors who are seeking a better way to diversify their credit exposure into the emerging markets.

In pursuit of our goal, the following multivariate CCR-VECM-EGARCH model is estimated using the two-step procedure outlined in the previous section:

$$\Delta p_{i,t}^{BRA} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{BRA} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{COL} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{COL} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{MEX} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{MEX} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{VEN} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{VEN} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

[15] Calvo and Mendoza (1995, 2000) argue that investors tend to follow the market rather than investigate market fundamentals due to the cost of information acquisition in emerging nations.

$$h_{i,t}^{BRA} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t-1}^{COL}) + \alpha_{1,3} f(Z_{t-1}^{MEX}) + \alpha_{1,4} f(Z_{t-1}^{VEN}) + \gamma_i \ell n(h_{t-1}^{BRA})\right\}$$

$$h_{i,t}^{COL} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t-1}^{COL}) + \alpha_{1,3} f(Z_{t-1}^{MEX}) + \alpha_{1,4} f(Z_{t-1}^{VEN}) + \gamma_i \ell n(h_{t-1}^{COL})\right\}$$

$$h_{i,t}^{MEX} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t}^{COL}) + \alpha_{1,3} f(Z_{t}^{MEX}) + \alpha_{1,4} f(Z_{t}^{VEN}) + \gamma_i \ell n(h_{t-1}^{MEX})\right\}$$

$$h_{i,t}^{VEN} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t}^{COL}) + \alpha_{1,3} f(Z_{t}^{MEX}) + \alpha_{1,4} f(Z_{t}^{VEN}) + \gamma_i \ell n(h_{t-1}^{VEN})\right\}$$

for i= *CDS, Credit Spread(CS)*

$$f(Z_{t-1}^{j}) = \left(|Z_{t-1}^{j}| - E(|Z_{t-1}^{j}|) + \delta_i Z_{t-1}^{j}\right) \quad \text{for } j = BRA, COL, MEX, VEN$$

$$\sigma_{j,k,t} = \rho_{j,k}\sigma_{j,t}\sigma_{k,t} \quad \text{for } j, k = BRA, COL, MEX, VEN \text{ and } j \neq k$$

The mean equations is augmented with two types of error correction components; one denoted as $\hat{u}_t$ from long-run relationship between the CDS prices and bond credit spreads within each sovereign, and the other from long-run relationship of CDS prices (or bond credit spreads when i = *CS*) across the four countries, denoted as $\hat{\varepsilon}_{i,t}^{Latin1}$ and $\hat{\varepsilon}_{i,t}^{Latin2}$. [16] Both error correction terms are generated from the Canonical Cointegrating Regression (CCR) technique. We optimise the number of lag terms in the underlying vector autoregression using the AIC. The estimated coefficients in the EGARCH are reported in Table 5.

Regarding the volatility spillovers, first note that the conditional variance in each sovereign credit risk is significantly (positively) affected by its own past standardized

[16] We find two cointegration equations between the four CDS price as well as bond credit spread series. Johansen's (1988) multivariate procedure is used to find the number of cointegration equations.

innovations in both sovereign CDS prices and bond credit spreads. (for example, we find that for the CDS prices, $\alpha_{1,1}$ = 0.2063, $\alpha_{2,2}$ = 0.4839, $\alpha_{4,4}$ = 0.2885, for the bond credit spread, $\alpha_{1,1}$ = 0.1858, $\alpha_{2,2}$ = 0.1456, $\alpha_{3,3}$ = 0.1782, and $\alpha_{4,4}$ = 0.3231).

Second, another important empirical finding is that there are more number of significant volatility spillovers originated in Brazil and Venezuela than in Colombia and Mexico. In fact, if we line up the four countries in the order of the highest number of one percent-level significant volatility spillovers originated from the country, it would be the exactly the opposite as to the order of the highest credit rating.[17] For example, we find the robust evidence of volatility spillovers from Venezuela to other countries, (i.e., $\alpha_{1,4}$ = 0.2874, $\alpha_{2,4}$ = 0.2320, and $\alpha_{3,4}$ = 0.3606 for CDS prices and $\alpha_{1,4}$ = 0.1296 and $\alpha_{2,4}$ = 0.0951 for bond credit spreads), but in contrast there is not many cases where volatility from Mexico that spills over to the neighbours is found to be significant.(i.e., $\alpha_{2,4}$ = 0.3606 for CDS prices and none for bond credit spreads.) Overall, there is strong tendency of one-way volatility spillovers from nations with lower credit rating to nations with higher credit rating. The intuition behind this empirical finding is that the nations with lower credit rating are more prone to external shocks so that they are likely the first to react in response. Or they themselves may be the source of credit-risk-sensitive information.

As shown in Panel A.1 and B.1 of Table 6, the degree of volatility persistence (measured by γ_i) is negatively related with the sovereign credit ratings.(for sovereign CDS prices, γ_1 = 0.5851, γ_2 = 0.2213, γ_3 = 0.0734 and γ_4 = 0.4850 and for sovereign bond credit spreads, γ_1 = 0.9338, γ_2 = 0.8914, γ_3 = 0.6458 and γ_4 = 0.9172.) This implies that

[17] See Table 5.

Mexico or Colombia recovers more quickly than Brazil or Venezuela once hit by a shock. This finding, together with the previous one, suggests that creditworthy emerging economies are more robust to the market turbulence emanating from outsides, not in the sense that they are insulated from the shock but in the sense that they revive faster. Another important finding with regard to the coefficient γ_i is that volatility persistence in CDS prices is much smaller than that in bond credit spreads. This might be interpreted as an evidence for the Latin American CDS markets' efficiency in processing the credit related information.

Lastly, we find that the CDS and bond market have another notable difference in the way that credit risks interact with each other. Though the estimated coefficients in the mean equation are not reported for the sake of space, the sum of significant coefficients of neighbouring bond credit spreads (in the form of first difference) is found to be negative in all four sovereign cases. (For example, with the first difference of Brazilian bond credit spreads as dependent variable, $\sum b_{1,j}^* + \sum c_{1,j}^* + \sum d_{1,j}^*$ is found to be negative.) This finding is consistent with Valdes (1997) and Calvo (1998)'s argument of liquidity constraint. They point out that when an investor is faced with liquidity needs in one particular asset, he will withdraw liquidity form another country or asset. In other words, in order to purchase a country's sovereign bond, the investor with liquidity constraint will need to sell the bond issued by neighbouring countries to fund the position. As a result, the credit spread of the neighbours from which the liquidity is withdrawn will be pushed upwards while the credit spread of sovereign bond purchased moves the opposite direction. However this is not always the case with sovereign CDS prices. This intriguing difference may be due to the nature of derivatives; unlike in the case of bonds, the CDS does not require any funding to take a credit exposure so that the investors are not constrained by liquidity.

5. Concluding remarks

The main purposes of this empirical research are twofold. First, we investigate the issue of price discovery and volatility spillovers *between* the sovereign CDS and the sovereign debt market. Second, we examine volatility spillovers *across* the credit risks of the four major Latin American debtors. In this paper, we have developed a multivariate CCR-VECM-EGARCH model which incorporates error correction terms estimated by CCR method and have tested for price discovery (from the mean equation) as well as volatility spillovers (from the conditional variance equation).

Regarding the issue of price discovery, our empirical results support that for major Latin American sovereign debtors, namely Brazil, Mexico and Colombia, the CDS market plays a leading role in price discovery, while for Turkish and Venezuelan cases the bond market is found to be more efficient in information processing. Both the Hasbrouck (1995) and Gonzalo and Granger (1995) common trend model provide compatible evidences to support this finding. The bivariate EGARCH model shows that although there are one-directional volatility spillovers that are consistent with the result of price discovery, evidence of reciprocal volatility spillovers are also found. Moreover, bad news (positive innovations) in the CDS markets tends to increase the future volatility more than good news (negative innovation).

Empirical analysis on volatility spillovers *across* the credit risks of the four major Latin American debtors supports four conclusions. First, there is a strong tendency of one-way volatility spillovers from nations with lower credit rating to ones with higher credit rating, but *not* vice versa. Second, the degree of volatility persistence is negatively related with the sovereign credit ratings. These two empirical findings support the notion that market participants are growing adept at distinguishing between emerging market nations with and without sound fundamentals. Third, the volatility persistence in CDS markets is much smaller than that in bond credit spreads. Finally, there is a supporting evidence for the role of liquidity constraint in the fluctuation of emerging market credit spreads, the argument asserted by Valdes (1997) and Calvo (1998).

References

Baillie, R.T., Booth, G.G., Tse, Y. and Zabotina, T., 2002, Price discovery and common factor models, *Journal of Financial Markets,* 5, 309-321.

Calvo, G., 1998, Varieties of capital market crises, in: Calvo, G.,King, M.(Eds.), The Debt burden and its consequences for monetary policy, International Economic Association Proceedings, Macmillan, London.

Calvo, G.A. and Mendoza, E.G., 1995, Reflections on Mexico's balance-of-payments crisis: a chronicle of a death foretold. Mimeo, University of Maryland.

Calvo, G.A. and Mendoza, E.G., 2000, Rational contagion and the globalisation of securities markets, *Journal of International Economics*, 51, 79-113.

Chan-Lau, J. and Kim, Y., 2004, Equity prices, credit default swaps, and bond spreads in emerging markets, IMF Working Paper 04/27.

Cifarelli, G. and Paladino, G., 2002, An empirical analysis of the co-movement among spreads on emerging-market debt, University of Florence Economics Working Paper No. 127.

Cunningham, A., Dixon, L., and Hayes, S., 2001, Analysing yield spreads on emerging market sovereign bonds, *Financial Stability Review (Bank of England)*, December, 175-186.

De Jong, F., 2002, Measures of contribution to price discovery: a comparison, *Jouranl of Financial Markets*, 5, 323-327

Duffie, D., 1999, Credit swap valuation, *Financial Analysts Journal,* 55, 73-87.

Gonzalo, J. and Granger, C., 1995, Estimation of common long-memory components in cointegrated systems, *Journal of Business and Economics Statistics*, 13, 27-35

Hasbrouck, J., 1995, One security, many markets: Determining the contributions to price discovery, *Journal of Finance*, 50, 1175-1199.

Hull, J., Predescu, M., and White, A., 2004, The relationship between credit default swap spreads, bond yields, and credit rating announcements, *Journal of Banking and Finance,* 28, 2789-2811.

Hull, J., and White, A., 2000a, Valuing credit default swaps I: No counterparty default risk, *Journal of Derivatives,* 8, 29-40.

Hull, J., and White, A., 2000b, Valuing credit default swaps II: Modelling default correlations, *Journal of Derivatives,* 8, 12-22.

In, F., 2006, Volatility spillovers across international swap markets: the USA, Japan and the UK, *Journal of International Money and Finance*, forthcoming.

Johansen, S., 1988, Statistical analysis of cointegrating vectors, *Journal of Economic Dynamics and Control*, 12, 231-54.

Jostova, G., 2006, Predictability in emerging sovereign debt markets, *The Journal of Business,* 79, 527-565.

Kamin, S.B. and von Kleist, K., 2005, The evolution and determination of emerging market credit spreads in the 1990s, BIS Working Paper No. 68

Packer, F., 2003, Sovereign credit default swaps, *BIS Quarterly Review*, December, 79-88. Valdes, R., 1997

Park, J. Y., (1992), Canonical cointegration regressions, *Econometrica*, 60, 119-143.

Park, J. Y. and Choi, B., 1988, A new approach to testing for a unit root, CAE Working Paper, 88-23, Department of Economics, Cornell University.

Park, J. Y. and Ogaki, M., 1991, Seemingly unrelated canonical cointegrating regressions, Working Paper 280. Rochester, NY: Rochester Centrer for Economic Research.

Rule, D., 2001, The credit derivatives market: its development and possible implications for financial stability, *Financial Stability Review (Bank of England)*, June, 117-140.

Ross, S., 1989, Information and volatility: the no-arbitrage, martingale approach to timing and resolution irrelevancy, *Journal of Finance*, 44, 1-17.

Singh, M. and Andritzky, J., 2005, Overpricing in emerging market credit-default-swap contracts: some evidence from recent distress cases, IMF Working Paper 05/125.

Valdes, R., 1997, Emerging markets contagion: evidence and theory, Documento de Trabajo del Banco Central de Chile, March.

Tse,Y., 1999, Price discovery and volatility spillovers in the DJIA index and futures markets, *Journal of Futures Markets,* 19, 911-930.

Zhu, H., 2004, An empirical comparison of credit spreads between the bond market and the credit default swap market, BIS Working Paper No 160.

Table1. Descriptive Statistics: Panel A

Panel A: CDS	$J_1(0,3)$	$J_1(1,5)$	Skewness	Kurtosis	JB	LB(24)	$LB^2(24)$
Brazil	7.291	0.449	1.131	11.033	2194.1	61.377	196.18
					(0.00)	(0.000)	(0.000)
Colombia	7.628	0.463	0.725	29.518	22217.7	54.783	94.843
					(0.00)	(0.000)	(0.000)
Mexico	7.246	0.106^{**}	0.794	11.820	2530.1	61.768	97.879
					(0.00)	(0.000)	(0.000)
Philippines	8.164	4.016	-0.324	11.583	2333.7	39.156	105.69
					(0.00)	(0.026)	(0.000)
Russia	7.066	0.586	-0.521	14.907	4500.4	33.112	83.959
					(0.00)	(0.102)	(0.000)
Turkey	7.910	1.520	0.588	12.829	3087.0	20.882	21.677
					(0.00)	(0.646)	(0.656)
Venezuela	8.427	1.084	-1.507	47.190	61795.8	48.024	116.37
					(0.00)	(0.003)	(0.000)

Note: $J_1(0,3)$ and $J_1(1,5)$ is the Park-Choi (1988) test statistic for unit roots. The null hypothesis is that a unit root exists. The critical values of Park-Choi's $J_1(0,3)$ test are 0.3385 (0.1118) at 5% (1%) significance level. Park-Choi's $J_1(1,5)$ test shows the critical value 5% (1%) as 0.295 (0.12), respectively. Also provided is the Jarque-Bera statistic, LB and LB^2 Ljung-Box statistics for 24 lags of the covariances of the residuals and squared residuals. $J_1(0,3)$ and $J_1(1,5)$ are computed from the level of daily credit default swap prices. The other statistics are based on first differences. P-values are provided in parentheses for the statistics.

Table1. Descriptive Statistics: Panel B

Panel B: Credit Spread	$J_1(0,3)$	$J_1(1,5)$	Skewness	Kurtosis	JB	LB(24)	$LB^2(24)$
Brazil	7.431	0.386	0.731	8.000	853.4	23.835	461.00
					(0.00)	(0.471)	(0.000)
Colombia	7.137	0.364	1.751	18.204	7668.2	43.654	168.32
					(0.00)	(0.008)	(0.000)
Mexico	6.734	0.519	0.420	4.984	146.17	46.285	170.04
					(0.00)	(0.004)	(0.000)
Philippines	8.541	2.935	-0.475	13.271	3351.2	36.862	178.91
					(0.00)	(0.045)	(0.000)
Russia	3.991	6.379	0.188	4.853	112.5	39.827	183.15
					(0.00)	(0.022)	(0.000)
Turkey	7.697	1.346	0.993	10.926	2103.1	21.092	462.38
					(0.00)	(0.633)	(0.000)
Venezuela	8.437	1.071	-0.304	6.184	330.93	23.849	492.33
					(0.00)	(0.470)	(0.000)

Note: $J_1(0,3)$ and $J_1(1,5)$ is the Park-Choi (1988) test statistic for unit roots. The null hypothesis is that a unit root exists. The critical values of Park-Choi's $J_1(0,3)$ test are 0.3385 (0.1118) at 5% (1%) significance level. Park-Choi's $J_1(1,5)$ test shows the critical value 5% (1%) as 0.295 (0.12), respectively. Also provided is the Jarque-Bera statistic, LB and LB^2 Ljung-Box statistics for 24 lags of the covariances of the residuals and squared residuals. $J_1(0,3)$ and $J_1(1,5)$ are computed from the level of daily credit spreads of bonds. The other statistics are based on first differences. P-values are provided in parentheses for the statistics.

Table2.

Discrepancies in the Average Pricing of Credit Risk in sovereign CDS and Bond Markets

	Treasury rates		Swap rates	
	Average basis	Average absolute basis	Average basis	Average absolute basis
Brazil	66.198	66.474	106.417	106.417
Colombia	46.154	71.776	86.374	97.446
Mexico	-7.309	17.749	32.911	32.939
Philippines	62.998	63.635	103.217	103.378
Russia	-56.845	73.548	-16.626	60.511
Turkey	55.846	57.056	96.065	96.095
Venezuela	49.283	52.935	89.502	89.556
Mean	30.904	57.596	71.123	83.763
Median	49.283	63.635	89.502	96.095

Notes: This table provides descriptive statistics of the basis, defined to be the difference between the credit default swap price and the credit spread, for each reference sovereign and expressed in basis points. The credit spread in calculated as the difference between the interpolated five-year yield on the sovereign bonds and either the five-year treasury bond rate or the five-year swap rate.

Table 3. Estimated disequilibrium error terms obtained from the CCR model

	$J_1(0,1)$	*$J_1(0,3)$*	*$J_1(1,5)$*	*CCR estimated β*
Brazil	0.438	3.315		1.2434139
				(219.92)
Colombia	0.521	5.088		1.8829265
				(30.93)
Mexico	3.377	4.712		1.5489263
				(21.52)
Philippines	0.779	4.770		1.1729028
				(49.70)
Russia	7.033	8.115	7.644	0.99292675
				(5.70)
Turkey	0.245	6.364		1.2313875
				(176.84)
Venezuela	0.5154	2.100		1.1350116
				(150.85)

Note: The estimated values from the Canonical Cointegration Regression (CCR) were obtained using the parzen window 150 to estimate the long-run variances. Series of estimated disequilibrium errors are calculated from $\hat{u}_i = CDS_i - \hat{\alpha}_{CCR} - \hat{\beta}_{CCR} CreditSpread_i$ for each reference entity *i,* except for the case of Russia where a deterministic time trend is found to be present in the cointegration regression equation. The disequilibrium errors for Russia is estimated from $\hat{u}_{Russia} = CDS_{Russiai} - \hat{\alpha}_{CCR} - \hat{\gamma}_{CCR}\, t - \hat{\beta}_{CCR} CreditSpread_{Russia}$.

J_1 is a test statistic for cointegration. The null hypothesis is that a cointegration exists. The critical values of $J_1(0,3)$ test are 7.81 (11.3) at 5% (1%) significance level and those of $J_1(1,5)$ are 9.49 (13.3) at 5% (1%) significance level. $J_1(0,3)$ test is conducted to check if there exist a deterministic trend in cointegrating regression. The critical values are 3.84 (6.63) at 5% (1%) significance level.

Table 4. Price discovery and Volatility spillovers

Mean:

$$\Delta p_{CDS,t} = \lambda_1 \hat{u}_{t-1} + \sum_{j=1}^{p} b_{1j} \Delta p_{CDS,t-j} + \sum_{j=1}^{p} c_{1j} \Delta p_{CDS,t-j} + \varepsilon_{1,t}$$

$$\Delta p_{CS,t} = \lambda_2 \hat{u}_{t-1} + \sum_{j=1}^{p} b_{2j} \Delta p_{CDS,t-j} + \sum_{j=1}^{p} c_{2j} \Delta p_{CDS,t-j} + \varepsilon_{2,t}$$

Variance:

$$h_t^{CDS} = \exp\left\{\alpha_{1,0} + \alpha_{1,1} f(Z_{t-1}^{CDS}) + \alpha_{1,2} f(Z_{t-1}^{CS}) + \gamma_1 \ell n(h_{t-1}^{CDS})\right\}$$

$$h_t^{CS} = \exp\left\{\alpha_{2,0} + \alpha_{2,1} f(Z_{t-1}^{CDS}) + \alpha_{2,2} f(Z_{t-1}^{CS}) + \gamma_2 \ell n(h_{t-1}^{CS})\right\}$$

$$f(Z_{t-1}^{j}) = \left(|Z_{t-1}^{j}| - E(|Z_{t-1}^{j}|) + \delta_i Z_{t-1}^{j}\right) \qquad \text{for } j = CDS, CS$$

$$\sigma_{j,k,t} = \rho_{j,k} \sigma_{j,t} \sigma_{k,t} \qquad \text{for } j, k = CDS, CS \text{ and } j \neq k$$

Table 4.A. Contributions to Price Discovery

Panel A.			Hasbrouck			GG
	λ_1	λ_2	Lower	Upper	Mid	
Brazil	0.0062	0.1031**	0.641	0.998	0.820	1.063
	(0.219)	(4.147)				
Colombia	-0.0065	0.02934**	0.837	0.977	0.907	0.819
	(-0.519)	(3.103)				
Mexico	-0.0050	0.0680**	0.855	0.993	0.924	0.931
	(-0.439)	(4.818)				
Philippines	-0.0414*	0.0531**	0.554	0.778	0.666	0.562
	(-2.495)	(3.942)				
Russia	-0.0304*	0.0295*	0.321	0.661	0.491	0.493
	(-2.017)	(1.961)				
Turkey	-0.0564**	0.0311	0.112	0.671	0.392	0.356
	(-2.685)	(1.569)				
Venezuela	-0.0855**	0.0063	0.016	0.177	0.096	0.069
	(-2.794)	(0.384)				

Note: The two bounds of Hasbrouk's measures and the Gonzalo and Granger measures are reported. The upper and lower bound of Hasbrouck's measure is calculated as,

$$Upper = \frac{\lambda_2\left(\sigma_1^2 - \frac{\sigma_{12}^2}{\sigma_2^2}\right)}{\lambda_2^2\sigma_1^2 - 2\lambda_1\lambda_2\sigma_{12} + \lambda_1^2\sigma_2^2}, \; Lower = \frac{\left(\lambda_2\sigma_1 - \lambda_1\frac{\sigma_{12}}{\sigma_1}\right)^2}{\lambda_2^2\sigma_1^2 - 2\lambda_1\lambda_2\sigma_{12} + \lambda_1^2\sigma_2^2} \text{ and, } GG = \frac{\lambda_2}{\lambda_2 - \lambda_1}$$

where the covariance matrix of ε_{1t} and ε_{2t} is represented by the terms of $\sigma_1^2, \sigma_{12,}$ and σ_2^2.

Table 4.B. Volatility Spillovers between CDS markets and sovereign bond markets

Panel B.	Brazil	Colombia	Mexico	Philippines	Russia	Turkey	Venezuela
$\alpha_{1,0}$	0.0407**	0.2908**	0.1279**	2.6678**	0.2068**	0.6165**	0.2610**
	(5.420)	(6.697)	(8.851)	(6.796)	(9.941)	(6.395)	(10.643)
$\alpha_{2,0}$	0.0895**	0.3650**	0.1853**	0.5230**	0.1928**	0.3338**	0.3311**
	(6.460)	(6.568)	(3.015)	(5.228)	(9.069)	(6.191)	(9.1717)
$\alpha_{1,1}$	0.0359**	0.1150**	0.0934**	0.2350**	0.0853**	-0.0034**	0.1127**
	(6.210)	(10.480)	(8.162)	(5.470)	(4.446)	(-5.289)	(11.110)
$\alpha_{1,2}$	0.0931**	0.2113**	0.008	0.2155**	0.3416**	0.4505**	0.2160**
	(9.344)	(8.529)	(0.072)	(4.425)	(17.199)	(9.854)	(9.859)
$\alpha_{2,1}$	0.0418**	0.1482**	0.0349*	0.0122	0.0255	-0.0054**	0.1425**
	(4.876)	(4.935)	(2.428)	(0.614)	(1.906)	(-5.713)	(5.300)
$\alpha_{2,2}$	0.7120**	0.3093**	0.1236**	0.4250**	0.1887**	0.4244**	0.3083**
	(10.095)	(8.146)	(3.234)	(10.635)	(10.470)	(18.392)	(9.414)
δ_1	1.2724**	-0.0433	0.9074**	0.5915**	0.9271**	25.1144**	-0.0972
	(3.970)	(-0.490)	(8.336)	(4.023)	(3.938)	(6.365)	(-0.922)
δ_2	-0.1603*	0.0550	-0.2381	0.1200*	-0.1966**	0.0954**	0.0630
	(-2.215)	(1.145)	(-1.350)	(2.543)	(-5.433)	(2.745)	(0.256)
γ_1	0.9931**	0.9479**	0.9617**	0.4708**	0.9478**	0.8713**	0.9533**
	(722.3)	(115.6)	(208.7)	(6.039)	(157.6)	(43.888)	(154.770)
γ_2	0.9838**	0.9254**	0.9476**	0.8888**	0.9513**	0.9271**	0.9325**
	(348.9)	(78.404)	(52.76)	(39.751)	(150.9)	(79.443)	(112.420)
$\rho_{1,2}$	0.6032**	0.2180**	0.2755**	0.2701**	0.2833**	0.4661**	0.2223**
	(342.5)	(55.874)	(53.04)	(32.167)	(90.989)	(70.333)	(76.444)
LB(20) for $Z_{CDS,t}$	18.581	9.435	20.226	10.208	18.396	24.101	9.632
	(0.353)	(0.926)	(0.263)	(0.895)	(0.364)	(0.117)	(0.918)
LB(20) for $Z_{CS,t}$	16.370	17.190	13.003	21.154	19.807	13.279	17.394
	(0.498)	(0.442)	(0.736)	(0.219)	(0.284)	(0.717)	(0.428)
LB(20) for $Z^2_{CDS,t}$	8.094	12.422	26.837	13.124	32.622	23.511	11.381
	(0.965)	(0.774)	(0.060)	(0.728)	(0.013)	(0.133)	(0.836)
LB(20) for $Z^2_{CS,t}$	42.692**	29.692*	18.353	8.668	18.124	22.345	28.382*
	(0.000)	(0.029)	(0.367)	(0.950)	(0.381)	(0.172)	(0.041)

Notes: This table reports the estimated coefficients for the variance equations. * (**) indicates significance at 5% (1%) level. T-statistics are reported in parentheses. Bottom of this table illustrates the summary statistics for the standardized innovations. The standardized innovations are calculated by $Z_t^j = \varepsilon_t^j / \sqrt{h_t^j}$ for j = CDS, CS. LB(n) is the Ljung-Box statistic for up to n lags, distributed as χ^2 with n degrees of freedom. P-values for Ljung-Box statistic are given in the parentheses.

Table 5.A. Sovereign Credit Ratings (as of April 2003)

Moody's				Standard and Poor's			
Rating	Americas	Europe	Asia	Rating	Americas	Europe	Asia
Baa1				BBB+			
Baa2	Mexico			BBB			
Baa3				BBB-	Mexico		
Ba1			Philippines	BB+			
Ba2	Colombia	Russia		BB	Colombia	Russia	Philippines
Ba3				BB-	Brazil		
B1		Turkey		B+			
B2	Brazil			B			
B3				B-		Turkey	
Caa1	Venezuela			CCC+	Venezuela		
Caa2				CCC			

Table 5.B. Sovereign Credit Ratings (as of March 2005)

Moody's				Standard and Poor's			
Rating	Americas	Europe	Asia	Rating	Americas	Europe	Asia
Baa1	Mexico			BBB+			
Baa2		Russia		BBB	Mexico	Russia	
Baa3				BBB-			
Ba1		Turkey		BB+			
Ba2	Colombia			BB	Colombia Brazil		
Ba3	Brazil			BB-	Venezuela	Turkey	Philippines
B1			Philippines	B+			
B2	Venezuela			B			
B3				B-			
Caa1				CCC+			
Caa2				CCC			

Table 6. Price discovery and Volatility spillovers

Mean:

$$\Delta p_{i,t}^{BRA} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{BRA} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{COL} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{COL} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{MEX} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{MEX} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

$$\Delta p_{i,t}^{VEN} = \beta_{i,0} + \beta_{i,1}\hat{u}_{i,t-1}^{VEN} + \beta_{i,2}\hat{\varepsilon}_{i,t-1}^{Latin1} + \beta_{i,3}\hat{\varepsilon}_{i,t-1}^{Latin2} + \sum_{j=1}^{p} a_{1j}\Delta p_{i,t-j}^{BRA} + \sum_{j=1}^{p} b_{1j}\Delta p_{i,t-j}^{COL} + \sum_{j=1}^{p} c_{1j}\Delta p_{i,t-j}^{MEX} + \sum_{j=1}^{p} d_{1j}\Delta p_{i,t-j}^{VEN} + \varepsilon_{i,t}$$

Variance:

$$h_{i,t}^{BRA} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t-1}^{COL}) + \alpha_{1,3} f(Z_{t-1}^{MEX}) + \alpha_{1,4} f(Z_{t-1}^{VEN}) + \gamma_i \ell n(h_{t-1}^{BRA})\right\}$$

$$h_{i,t}^{COL} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t-1}^{COL}) + \alpha_{1,3} f(Z_{t-1}^{MEX}) + \alpha_{1,4} f(Z_{t-1}^{VEN}) + \gamma_i \ell n(h_{t-1}^{COL})\right\}\backslash$$

$$h_{i,t}^{MEX} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t}^{COL}) + \alpha_{1,3} f(Z_{t}^{MEX}) + \alpha_{1,4} f(Z_{t}^{VEN}) + \gamma_i \ell n(h_{t-1}^{MEX})\right\}$$

$$h_{i,t}^{VEN} = \exp\left\{\alpha_{i,0} + \alpha_{1,1} f(Z_{t-1}^{BRA}) + \alpha_{1,2} f(Z_{t}^{COL}) + \alpha_{1,3} f(Z_{t}^{MEX}) + \alpha_{1,4} f(Z_{t}^{VEN}) + \gamma_i \ell n(h_{t-1}^{VEN})\right\}$$

$$f(Z_{t-1}^{j}) = \left(| Z_{t-1}^{j} | - E(| Z_{t-1}^{j} |) + \delta_i Z_{t-1}^{j}\right)$$ for j = *BRA, COL, MEX, VEN*

$$\sigma_{j,k,t} = \rho_{j,k}\sigma_{j,t}\sigma_{k,t}$$ for j, k = *BRA,COL,MEX,VEN* and $j \neq k$

Panel A1.Estimated coefficients for CDS prices

Brazil		Colombia		Mexico		Venezuela	
$\alpha_{1,0}$	2.3686**	$\alpha_{2,0}$	4.1664**	$\alpha_{3,0}$	2.7546**	$\alpha_{4,0}$	3.2708**
	(10.918)		(29.548)		(18.825)		(20.956)
$\alpha_{1,1}$	0.2063**	$\alpha_{2,1}$	0.4807**	$\alpha_{3,1}$	0.1800**	$\alpha_{4,1}$	0.0129
	(7.394)		(10.112)		(4.535)		(0.659)
$\alpha_{1,2}$	-0.0178	$\alpha_{2,2}$	0.4839**	$\alpha_{3,2}$	0.0835	$\alpha_{4,2}$	-0.1585**
	(-0.517)		(13.992)		(1.771)		(-4.611)
$\alpha_{1,3}$	0.0085	$\alpha_{2,3}$	-0.0079	$\alpha_{3,3}$	0.0102	$\alpha_{4,3}$	0.0077
	(0.661)		(-0.713)		(0.688)		(0.671)
$\alpha_{1,4}$	0.2874**	$\alpha_{2,4}$	0.2320**	$\alpha_{3,4}$	0.3606**	$\alpha_{4,4}$	0.2885**
	(7.021)		(5.771)		(9.435)		(15.054)
δ_1	0.3689**	δ_2	-0.1377**	δ_3	10.585	δ_4	-0.0473
	(6.064)		(-2.626)		(0.687)		(-1.181)
γ_1	0.5851**	γ_2	0.2213**	γ_3	0.0734	γ_4	0.4850**
	(15.793)		(8.230)		(1.437)		(20.062)
$\rho_{1,2}$	0.4115**	$\rho_{1,3}$	0.3909**	$\rho_{1,4}$	0.4256**	$\rho_{2,3}$	0.2784**
	(10.762)		(11.494)		(3.340)		(7.121)
$\rho_{2,4}$	0.2987**	$\rho_{3,4}$	0.1466*				
	(3.547)		(2.387)				

Notes: This table reports the estimated coefficients for the CDS spreads of four countries. *(**) indicates significance at 5% (1%) level. The t statistics are reported in parentheses. For the sake of space, only the estimation results in volatility equation are reported.

Panel A.2 Diagnostic tests for standardized innovations

	Brazil	Colombia	Mexico	Venezuela
Mean	-0.0137	-0.0167	-0.0114	0.0127
Variance	1.0475	1.0560	1.0519	1.1936
Skewness	0.5926	-0.7945	0.4982	1.7820
Kurtosis	4.9103	13.9540	7.0750	36.968
LB(20) for $Z_{i,t}$	31.601	20.489	16.916	24.880
	(0.017)	(0.250)	(0.460)	(0.097)
LB(20) for $Z_{i,t}^2$	68.851	23.248	29.574	31.198
	(0.000)	(0.141)	(0.030)	(0.019)

Notes: This table illustrates the summary statistics for the standardized innovations. The standardized innovations are calculated by $Z_t^j = \varepsilon_t^j / \sqrt{h_t^j}$ for j = US, JP, UK. LB(n) is the Ljung-Box statistic for up to n lags, distributed as χ^2 with n degrees of freedom. Skewness and kurtosis are defined as $E[(R_t - \mu)]^3$ and $E[(R_t - \mu)]^4$, respectively, where μ is the sample mean.

Panel B 1.Estimated coefficients for bond credit spreads

Brazil		Colombia		Mexico		Venezuela	
$\alpha_{1,0}$	0.3590**	$\alpha_{2,0}$	0.5046**	$\alpha_{3,0}$	1.2132**	$\alpha_{4,0}$	0.4355**
	(5.191)		(8.611)		(43.084)		(5.751)
$\alpha_{1,1}$	0.1858**	$\alpha_{2,1}$	0.1508**	$\alpha_{3,1}$	0.1872**	$\alpha_{4,1}$	-0.0393
	(6.146)		(5.930)		(5.308)		(-1.061)
$\alpha_{1,2}$	0.0788*	$\alpha_{2,2}$	0.1456**	$\alpha_{3,2}$	0.0631	$\alpha_{4,2}$	0.0565*
	(2.522)		(5.333)		(1.850)		(2.309)
$\alpha_{1,3}$	0.0399	$\alpha_{2,3}$	0.1332**	$\alpha_{3,3}$	0.1782**	$\alpha_{4,3}$	0.0603*
	(1.175)		(5.473)		(6.005)		(2.275)
$\alpha_{1,4}$	0.1296**	$\alpha_{2,4}$	0.0951**	$\alpha_{3,4}$	0.0319	$\alpha_{4,4}$	0.3231**
	(3.444)		(2.980)		(0.802)		(7.592)
δ_1	-0.0835	δ_2	0.3644**	δ_3	0.0253	δ_4	-0.2350**
	(-1.385)		(3.111)		(0.266)		(-5.151)
γ_1	0.9338**	γ_2	0.8914**	γ_3	0.6458**	γ_4	0.9172**
	(72.717)		(68.926)		(118.844)		(65.767)
$\rho_{1,2}$	0.5328**	$\rho_{1,3}$	0.4559**	$\rho_{1,4}$	0.5586**	$\rho_{2,3}$	0.5336**
	(36.465)		(8.614)		(3.540)		(8.965)
$\rho_{2,4}$	0.5196**	$\rho_{3,4}$	0.4642**				
	(3.501)		(3.519)				

Notes: This table reports the estimated coefficients for the CDS spreads of four countries. *(**) indicates significance at 5% (1%) level. The t statistics are reported in parentheses. For the sake of space, only the estimation results in volatility equation are reported.

Panel B.2 Diagnostic tests for standardized innovations

	Brazil	Colombia	Mexico	Venezuela
Mean	-0.0016	-0.0046	-0.0038	0.0178
Variance	1.0408	1.0293	1.0312	1.0546
Skewness	0.3617	0.3500	0.2142	-0.1336
Kurtosis	1.7719	1.4245	0.9628	1.1366
LB(20) for $Z_{i,t}$	51.746	32.189	20.995	10.852
	(0.0000)	(0.0143)	(0.2265)	(0.8641)
LB(20) for $Z_{i,t}^2$	16.852	42.081	50.640	23.053
	(0.4645)	(0.0006)	(0.0000)	(0.1475)

Notes: This table illustrates the summary statistics for the standardized innovations. The standardized innovations are calculated by $Z_t^j = \varepsilon_t^j / \sqrt{h_t^j}$ for j = US, JP, UK. LB(n) is the Ljung-Box statistic for up to n lags, distributed as χ^2 with n degrees of freedom. Skewness and kurtosis are defined as $E[(R_t - \mu)]^3$ and $E[(R_t - \mu)]^4$, respectively, where μ is the sample mean.