

Default and Recovery Implicit in the Term Structure of Sovereign *CDS* Spreads

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Abstract

This paper explores in depth the nature of default arrival and recovery implicit in the *term structures* of sovereign *CDS* spreads. We argue that, in principle, a term structure of spreads reveals not only the parameters of the market-implied mean arrival rates of credit events ($\lambda^{\mathbb{Q}}$), but also the implicit loss rates ($L^{\mathbb{Q}}$) given credit events. Applying our framework to Mexico, Turkey, and Korea, three countries with different geopolitical characteristics and credit ratings, we show that a single-factor model in which $\lambda^{\mathbb{Q}}$ follows a lognormal process captures most of the variation in the term structures of spreads. Our models imply highly persistent $\lambda^{\mathbb{Q}}$ under the pricing measure, and economically significant risk premiums associated with unpredictable future variation in $\lambda^{\mathbb{Q}}$. We document significant correlations among these risk premiums and several economic measures of global event risk, financial market volatility and macroeconomic policy, both across maturities and countries. A potential role for (il)liquidity underlying the (small) mispricings of our model is explored along with the properties of the bid/ask spreads on the sovereign *CDS* contracts.

1 Introduction

The burgeoning market for sovereign credit default swaps (*CDS*) contracts offers a nearly unique window for viewing investors' risk-neutral probabilities of major credit events impinging on sovereign issuers, and their risk-neutral losses of principal in the event of a restructuring or repudiation of external debts. In contrast to many "emerging market" sovereign bonds, sovereign *CDS* contracts are designed without complex guarantees or embedded options. Trading activity in the *CDS* contracts of several sovereign issuers has developed to the point that they are more liquid than many of the underlying bonds. Moreover, in contrast to the corporate *CDS* market, where trading has been concentrated largely in the five-year maturity contract, *CDS* contracts at several maturity points between one and ten years have been actively traded for several years. As such, a full *term structure* of *CDS* spreads is available for inferring default and recovery information from market data.

This paper explores in depth the time-series properties of the risk-neutral mean arrival rates of credit events (λ^Q) implicit in the term structure of sovereign *CDS* spreads. Applying our framework to Mexico, Turkey, and Korea, three countries with different geopolitical characteristics and credit ratings, we find that single-factor models, in which country-specific λ^Q follow lognormal processes,¹ capture most of the variation in the term structures of spreads. The maximum likelihood estimates suggest that, for all three countries, there are systematic, priced risks associated with unpredictable future variation in λ^Q . Moreover, the time-series of the effects of risk premiums on *CDS* spreads covary strongly across countries. There are several large concurrent "run-ups" in risk premiums during our sample period (March, 2001 through August, 2006) that have natural interpretations in terms of political, macroeconomic, and financial market developments at the time.

A more formal regression analysis of the correlations between risk premiums and the CBOE U.S. VIX option volatility index (viewed as a measure of event risk), the spread between ten-year return on U.S. BBB-rated industrial corporate bonds and the six-month U.S. Treasury bill rate (viewed as a measure of both U.S. macroeconomic and global financial market developments), and the volatility in the own-currency options market corroborates our economic interpretations of the temporal changes in risk premiums in the sovereign *CDS* markets. The evidence is consistent with risk premiums in sovereign markets for credit risk being influenced by spillovers of real economic growth in the U.S. to economic growth in other regions of the world. Equally notable is that our findings suggest that, during some subperiods, a substantial portion of the co-movement among the term structures of sovereign spreads across countries was induced by changes in investors' appetites for credit exposure at a global level, rather than to reassessments of the fundamental strengths of these specific sovereign economies.

While most of our focus is on the economic underpinnings of the dynamic properties of the arrival rates of credit events, an equally central ingredient to modeling the credit risk

¹In the literature on corporate *CDS* spreads, λ^Q was modeled as a square-root process in Longstaff, Mithal, and Neis (2004), while Berndt, Douglas, Duffie, Ferguson, and Schranz (2004) argue that corporate *CDS* spreads are better described by a lognormal model. Zhang (2003) had λ^Q following a square-root process in his analysis of Argentinean *CDS* contracts.

of sovereign issuers is the recovery of bond holders in the face of a credit event. Standard practice in modeling corporate *CDS* spreads is to assume a fixed risk-neutral loss rate $L^{\mathbb{Q}}$, largely because the focus has been on the liquid five-year *CDS* contract.² We depart from this literature and exploit the term structure of *CDS* spreads to separately identify both $L^{\mathbb{Q}}$ and the parameters of the process $\lambda^{\mathbb{Q}}$. That we even attempt to separately identify these parameters of the default process may seem surprising in the light of the apparent demonstrations in Duffie and Singleton (1999), Houweling and Vorst (2003), and elsewhere of the infeasibility of achieving this objective. We show that in fact, in market environments where recovery is a fraction of face value, as is the case with *CDS* markets, these parameters can in principle be separately identified through the information contained in the term structure of *CDS* spreads.

The maximum likelihood (*ML*) estimates of the parameters governing $\lambda^{\mathbb{Q}}$ imply that its risk-neutral (\mathbb{Q}) distribution shows very little mean reversion and, in fact, in some cases $\lambda^{\mathbb{Q}}$ is \mathbb{Q} -explosive. In contrast, the historical data-generating process (\mathbb{P}) for $\lambda^{\mathbb{Q}}$ shows substantial mean reversion, consistent with the \mathbb{P} -stationarity of *CDS* spreads. This large difference between the properties of $\lambda^{\mathbb{Q}}$ under the \mathbb{Q} and \mathbb{P} measures implies, within the context of our models, that an economically important systematic risk is being priced in the *CDS* market.

Our *ML* estimates are obtained both with fixed $L^{\mathbb{Q}}$ at the market convention 0.75, and by searching over $L^{\mathbb{Q}}$ as a free parameter. In the latter case, the likelihood functions call for much smaller values of $L^{\mathbb{Q}}$ for Mexico and Turkey, more in the region of 0.25, whereas for Korea the constrained and unconstrained estimates of $L^{\mathbb{Q}}$ are comparable. An extensive Monte Carlo of the small-sample distributions of various moments reveals that many features of the implied distributions of *CDS* spreads for Mexico and Turkey are similar across the cases of $L^{\mathbb{Q}}$ equal to 0.75 or 0.25. It is only over long horizons the differences in \mathbb{P} -mean reversion in the two cases manifest themselves. Accordingly, given that a much longer time series of *CDS* spreads than is currently available would be needed to select between the constrained and unconstrained models, we impose the market convention of $L^{\mathbb{Q}} = 0.75$ for our analysis of risk premiums.

Throughout this analysis we maintain the assumption that a single risk factor underlies the temporal variation in $\lambda^{\mathbb{Q}}$, consistent with most previous studies of *CDS* spreads that have allowed for a stochastic arrival rate of credit events. In the case of our sovereign data, this initial focus is motivated by the high degree of comovement among spreads across the maturity spectrum within each country. For our sample period, this comovement is even greater than the high degree of correlations between yields in highly liquid treasury markets documented, for example, in Litterman and Scheinkman (1991). We investigate the potential role for a second risk factor, particularly for understanding *CDS* spreads at the short end of the maturity spectrum. As part of this analysis, the behaviors of bid/ask spreads are also examined, particularly in the context of examining a potential role for liquidity factors.

To our knowledge, the closest precursor to our analysis is the study by Zhang (2003) of *CDS* spreads for Argentina leading up to the default in late 2001. Our sample period begins

²See, for example, Berndt, Douglas, Duffie, Ferguson, and Schranz (2004), Hull and White (2004), and Houweling and Vorst (2003).

towards the end of his, is longer in length, and spans a period during which the sovereign *CDS* markets were more developed in breadth and liquidity. The complementary study of Mexican and Brazilian *CDS* spreads in Carr and Wu (2006) explores the correlation structure of spreads on contracts up to five years to maturity with implied volatilities on various currency options over the shorter period of January, 2002 through March, 2005. Relative to both of these studies, we examine a geographically more dispersed set of countries, and we explore in depth the economic underpinnings of the *comovements* of risk premiums for these countries. Toward this end, we allow for more flexible market prices of risk, and examine a broader array of economic factors underlying market risk premiums.

2 The Structure of the Sovereign *CDS* Market

The structure of the standard *CDS* contract for a sovereign issuer shares many of its features with the corporate counterpart. The default protection buyer pays a semi-annual premium, expressed in basis points per notional amount of the contract, in exchange for a contingent payment in the event one of a pre-specified credit events occurs. Settlement of a *CDS* contract is typically by physical delivery of an admissible bond in return for receipt of the original face value of the bonds,³ with admissibility determined by the characteristics of the reference obligation in the contract.

Typically, only bonds issued in *external* markets and denominated in one of the “standard specified currencies” are deliverable.⁴ In particular, bonds issued in domestic currency, issued domestically, or governed by domestic laws are not deliverable. For some sovereign issuers without extensive issuance of hard-currency denominated Eurobonds, loans may be included in the set of deliverable assets. Among the countries included in our analysis, Turkey and Mexico have sizeable amounts of outstanding loans, and their *CDS* contracts occasionally trade with “Bond or Loan” terms. The contracts we focus on are “Bond only.”

The key definition included in the term sheet of a sovereign *CDS* contract is the credit event. Typically, a sovereign *CDS* contract lists as events any of the following that affect the reference obligation: (i) obligation acceleration, (ii) failure to pay, (iii) restructuring; or (iv) repudiation/moratorium. Note that “default” is not included in this list, because there is no operable international bankruptcy court that applies to sovereign issuers.

Central to our analysis of the term structure of sovereign *CDS* spreads is the active trading of contracts across a wide range of maturities. In contrast to the U.S. corporate and

³Physical delivery is the predominant form of settlement in the sovereign *CDS* market, because both the buyers and sellers of protection typically want to avoid the dealer polling process involved in determining the value of the reference bond in what is often a very illiquid post-credit-event market place.

⁴The standard specified currencies are the Euro, U.S. dollar, Japanese yen, Canadian dollar, Swiss franc, and the British pound. The option to deliver bonds denominated in these currencies, and of various maturities, into a *CDS* contract introduces a cheapest-to-deliver option for the protection buyer. Our impression, from conversations with traders, is that usually there is a single bond (or small set of bonds) that are cheapest to deliver. So the price of the *CDS* contract tracks this cheapest to deliver bond and the option to deliver other bonds is not especially valuable. In any event, for the purpose of our subsequent analysis, we will ignore this complication in the market.

bank *CDS* markets, where a large majority of the trading volume is concentrated in five-year contracts, the three- and ten-year contracts have each accounted for roughly 20% of the volumes in sovereign markets, and the one-year contract has accounted for an additional 10% of the trading (see Figure 1).⁵ While the total volume of new contracts has been much larger in the corporate than in the sovereign market, the volumes for the most actively traded sovereign credits are large and growing. We focus our analysis on Mexico, Turkey, and Korea, three of the more actively traded names.⁶

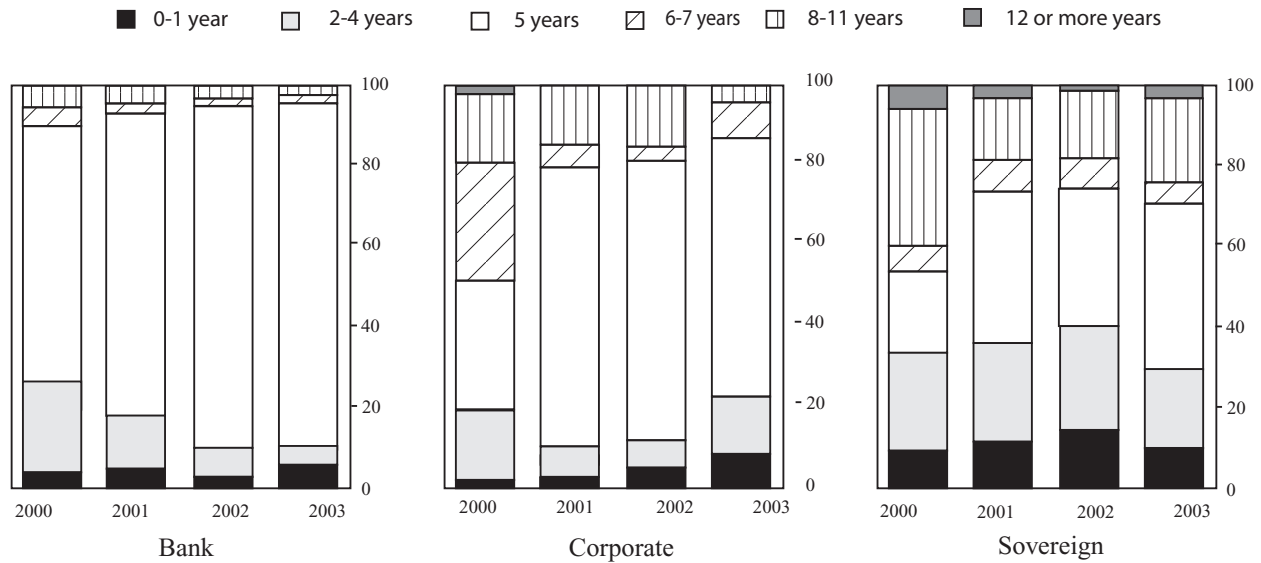


Figure 1: *CDS* volumes by maturity, as a percentage of total volume, based on *BIS* calculations from CreditTrade data. Source: *BIS Quarterly Review* [2003].

Our sample consists of daily trader quotes of bid and ask spreads for *CDS* contracts with maturities of one, three, five, and ten years. The sample covers the period March 19, 2001 through August 10, 2006. We focus on the data for three geographically dispersed countries— Mexico, Turkey, and Korea— displayed in Figure 2. (Descriptive statistics of these series are displayed on the left-hand side of Table 1.) At the beginning of our sample period (March, 2001), Mexico had achieved the investment grade rating of Baa3. In February, 2002, Mexico was upgraded one notch to Baa2, and it was subsequently upgraded again one notch to Baa1 in January, 2005. Turkey maintained the same speculative grade rating, B1, throughout most of our sample period. However, both in April, 2001 and July, 2002 it was put in the “negative outlook” category. Following the most recent negative outlook, Turkey returned to “stable outlook” in October, 2003. Moody’s changed its outlook for Turkey to positive in February, 2005, and then upgraded Turkish (external) government bonds to Ba3

⁵Figure 1 is a corrected version of the original appearing in Packer and Suthiphongchai (2003).

⁶Russia as well as several South American credits— Brazil, Columbia, and Venezuela— are also among the more traded sovereign credits. The behavior of the South American *CDS* spreads was largely dominated by the political turmoil in Brazil during the summer/fall of 2002. The co-movements among the *CDS* spreads of these countries is an interesting question for future research.

	CDS Price (bps)						CDS Bid Ask Spread (bps)					
	mean	med	std	min	max	a.c.	mean	med	std	min	max	a.c.
	Mexico						Mexico					
1yr	54.5	33	38.6	14	185	0.993	13.3	10	8.5	5	50	0.940
2yr	92.4	65	63.7	22	305	0.995	13.1	10	8.9	2	60	0.931
3yr	123.5	94	78.7	30	370	0.996	13.0	10	8.3	5	50	0.937
5yr	166.4	147	89.3	46	440	0.997	12.4	10	8.2	4	40	0.951
10yr	213.0	200	90.2	76	475	0.997	12.6	10	8.5	4	50	0.950
	Turkey						Turkey					
1yr	378.4	225	355.5	23	1700	0.993	61.1	50	62.3	8	850	0.875
2yr	458.1	315	357.0	45	1650	0.995	47.5	30	52.1	6	600	0.914
3yr	505.9	399	347.8	68	1600	0.995	44.3	30	49.6	6	575	0.889
5yr	563.1	504	327.7	116	1500	0.996	39.5	30	41.1	4	400	0.906
10yr	607.3	552	304.6	181	1450	0.996	39.4	30	39.0	4	300	0.935
	Korea						Korea					
1yr	33.7	31	25.0	4	165	0.991	9.2	10	1.0	8	10	0.998
2yr	41.7	38	27.8	9	176	0.994	9.2	10	1.0	8	10	0.998
3yr	48.6	45	29.8	13	184	0.995	9.2	10	1.0	6	10	0.995
5yr	62.0	58	33.2	22	197	0.996	9.2	10	1.0	5	10	0.993
10yr	81.3	78	38.5	32	212	0.996	9.2	10	1.0	5	10	0.993

Table 1: Summary Statistics. The sample period is March, 2001 until the beginning of August, 2006. med is the sample median; std is the sample standard deviation; a.c. is the first-order autocorrelation statistic.

in December, 2005. Korea was upgraded by Moody’s from Baa2 to A3 on March 28, 2002 and it maintained this rating throughout our sample period. However the outlook for Korea was negative towards the end of 2003 (due to concerns about North Korea), it was upgraded to stable in September 2004, and upgraded again to positive in April, 2006. Consistent with the relative credit qualities of these countries, the average five-year *CDS* spreads over our sample period are 89.9, 166.4, and 563.1 basis points, respectively, for Korea, Mexico, and Turkey (see Table 1).

In addition to the fact that they cover a broad range of credit quality, two important considerations factor into our choice of these three countries: their regional representativeness in the emerging markets and the relative liquidity and thus better data quality of their *CDS* markets compared to those of other countries in the same region. The first consideration is important for the economic interpretation of our results. These countries are geographically dispersed— being located in Latin American, Eastern Europe, and Asia— and each, in its own way, has been affected by significant local economic and political events. As such, we are interested in the degree and nature of the co-movements among *CDS* spreads for these countries. The second consideration plays a crucial role in our evaluation of our model’s implications for default and recovery implicit in *CDS* spreads, as we will assume that the levels of *CDS* spreads are largely reflective of credit assessments (as opposed to (il)liquidity,

for example).

As shown in Figure 2, the term-structures of *CDS* spreads exhibit interesting dynamics. One immediately noticeable feature present in all three countries is the high level of comovement among the 1y, 3y, 5y, and 10y *CDS* spreads. Indeed, a principal component (*PC*) analysis of the spreads in each country (see Section 5.2) shows that the first *PC* explains over 96% of the variation in *CDS* spreads for all three countries.⁷ It is these high levels of explained variation that motivate our focus on one-factor models.

Another prominent feature of the *CDS* data is the persistence of upward sloping term structures. This is especially true for the term structures of Mexican and Korean *CDS* spreads: throughout our sample period, the one-year *CDS* spreads were always lower than the respective longer maturity *CDS* spreads and, hence, the term structure was never inverted. For example, the difference between the five-year and one-year Mexican *CDS* spreads was 112 basis points on average, 31 basis points at minimum, and 275 basis points at maximum. Without resorting to institutional features that might separate the one-year from the longer maturity *CDS* contracts, this feature of *CDS* spreads implies an increasing term structure of risk-neutral one-year forward default probabilities.

The slope of the term structure of *CDS* spreads for Turkey was mostly positive. For example, the difference between the five- and one-year *CDS* spreads was on average 185 basis points with a standard deviation of 93 basis points. However, in contrast to the robust pattern of upward sloping spread curves in Mexico and Korea, the term structure of Turkish *CDS* spreads did occasionally invert, especially when credit spreads exploded to high levels due to financial or political crises that were (largely) specific to Turkey. For example, the differences between the five- and one-year *CDS* spreads were -250 basis points on March 29, 2001, -150 basis points on July 10, 2002, and -200 basis points on March 24, 2003. The related events were the devaluation of the Turkish lira, political elections in Turkey, and the collapse of talks between Turkey and Cyprus (which had implications for Turkey's bid to join the EU).

Sovereign credit default swaps trade, on average, in larger sizes than in the underlying cash markets: U.S. \$5 million, and occasionally much larger, against U.S. \$1 - 2 million. The liquidity of the underlying bond market is relevant, because traders hedge their *CDS* positions with cash market instruments and the less liquid is the cash market, the larger the bid/ask spread must be in the *CDS* market to cover the higher hedging costs. Comparing across sovereign *CDS* markets, a given bid/ask spread will sustain a larger trade in the market for Mexico (up to about \$40 million) relative to Turkey (up to about \$30 million) (Xu and Wilder (2003)).

For our sample of countries, the bid/ask spreads (in basis points for the five-year contract) ranged between 4 and 40 for Mexico, 4 and 400 for Turkey, and 2 and 20 for Korea (see Figure 3 and Table 1). Korea had the smallest and most stable bid/ask spreads. Notably, when Turkey's spreads widened out due to the "local" events chronicled above, so did the bid/ask spreads. For high-grade countries with large quantities of bonds outstanding like

⁷The only exception is the spread on the one-year contract for Mexico, and 90% of its variation is explained by the first *PC* of Mexican spreads.

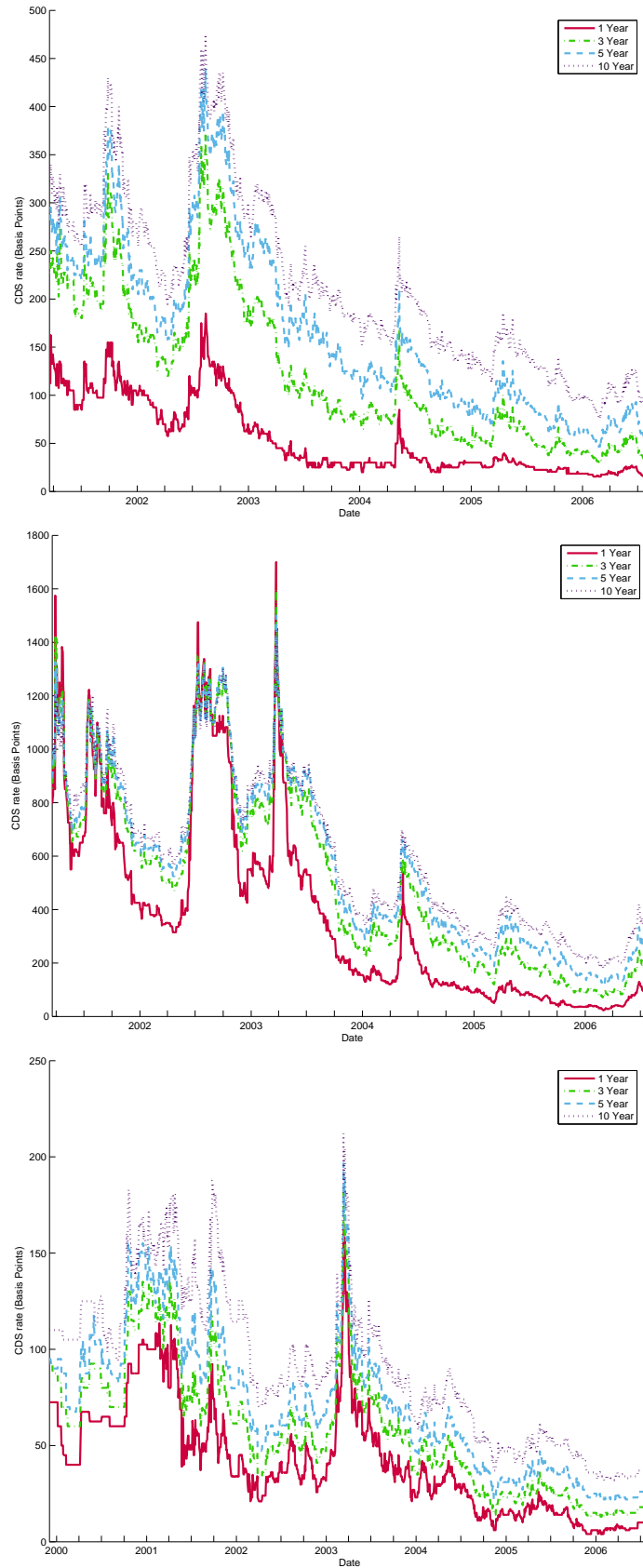


Figure 2: CDS Spreads: Mexico (upper), Turkey (middle), and Korea (lower), mid-market quotes.

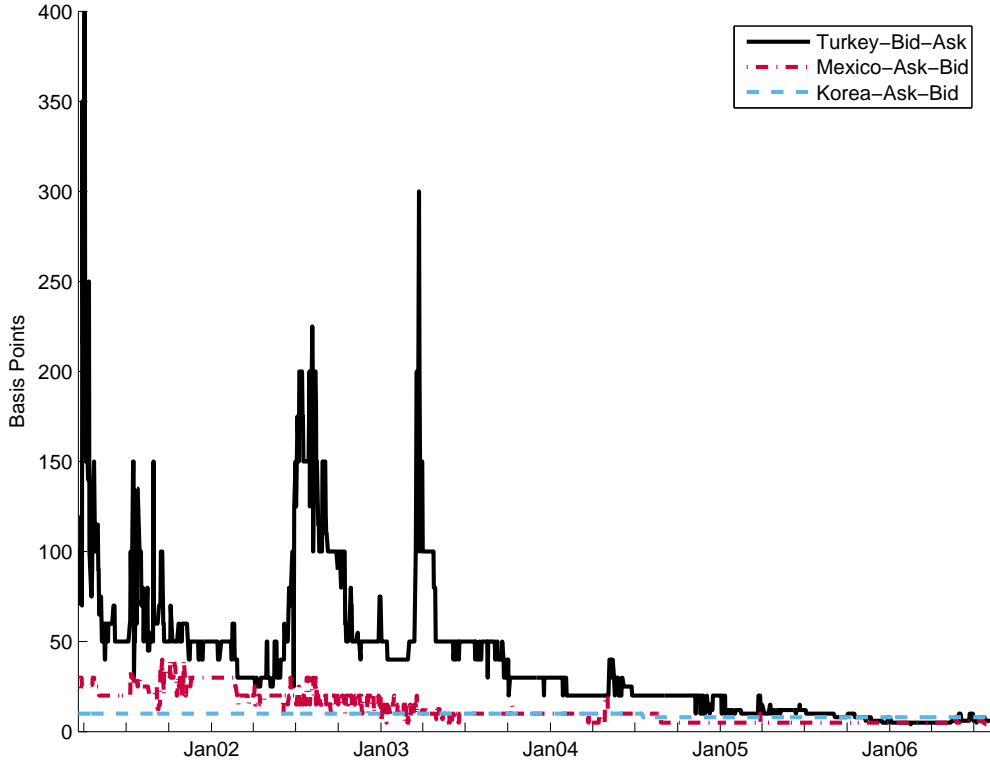


Figure 3: Ask-Bid Spreads (basis points) for five-year *CDS* contracts

Mexico and Korea, the magnitudes of the bid/ask spreads in the *CDS* markets are comparable to those for their bonds.

Particularly at the short end of the maturity spectrum, the *CDS* market is often more liquid than the underlying cash markets and, as such, market participants experience smaller bid/ask spreads. For our sample period, the bid/ask spreads for the one-year contracts (not shown) are comparable in magnitude to those of the five-year contracts. The primary exceptions are during turbulent periods, especially in Turkey, when the levels of *CDS* spreads are large. In such cases, the bid/asks spreads on the one-year contracts are larger than those on the five-year contracts. We examine the properties of the bid/asks spreads of our data in more depth in Section 5.2 in conjunction with our discussion of the challenges of fitting the one-year (and to a lesser extent the ten-year) spreads within our one-factor term structure model for *CDS* spreads.

3 Pricing Sovereign *CDS* Contracts

The basic pricing relation for sovereign *CDS* contracts is identical to that for corporate *CDS* contracts. Let M denote the maturity (in years) of the contract, $CDS_t(M)$ denote the (annualized) spread at issue, R^Q denote the (constant) risk-neutral fractional recovery of

face value on the underlying (cheapest-to-deliver) bond in the event of a credit event, and $\lambda^{\mathbb{Q}}$ denote the risk neutral arrival rate of a credit event. Then, at issue, a *CDS* contract with semi-annual premium payments is priced as (see, e.g., Duffie and Singleton (2003)):

$$\frac{1}{2} CDS_t(M) \sum_{j=1}^{2M} E_t^{\mathbb{Q}} \left[e^{-\int_t^{t+.5j} (r_s + \lambda_s^{\mathbb{Q}}) ds} \right] = (1 - R^{\mathbb{Q}}) \int_t^{t+M} E_t^{\mathbb{Q}} \left[\lambda_u^{\mathbb{Q}} e^{-\int_t^u (r_s + \lambda_s^{\mathbb{Q}}) ds} \right] du, \quad (1)$$

where r_t is the riskless rate relevant for pricing *CDS* contracts. The left-hand-side of (1) is the present value of the buyer's premiums, payable contingent upon a credit event not having occurred. Discounting by $r_t + \lambda_t^{\mathbb{Q}}$ captures the survival-dependent nature of these payments (Lando (1998)). The right-hand-side of this pricing relation is the present value of the contingent payment by the protection seller upon a credit event. We have normalized the face value of the underlying bond to \$1 and assumed a constant expected contingent payment (loss relative to face value) of $L^{\mathbb{Q}} = (1 - R^{\mathbb{Q}})$. In implementing (1), we use a slightly modified version that accounts for the buyer's obligation to pay an accrued premium if a credit event occurs between the premium payment dates.

How should $\lambda^{\mathbb{Q}}$ and $L^{\mathbb{Q}}$ be interpreted, given that default is not a relevant credit event, and *ISDA* terms sheets for plain vanilla sovereign *CDS* contracts reference four types of credit events? To accommodate this richness of the credit process for sovereign issuers, let each of the four relevant credit events have their own associated arrival intensities $\lambda_i^{\mathbb{Q}}$ and loss rates $L_i^{\mathbb{Q}}$. Then, following Duffie, Pedersen, and Singleton (2003) and adopting the usual "doubly stochastic" formulation of arrival of credit events (see, e.g., Lando (1998)), we can interpret the $\lambda_t^{\mathbb{Q}}$ and $L_t^{\mathbb{Q}}$ for pricing sovereign *CDS* contracts as:

$$\lambda_t^{\mathbb{Q}} = \lambda_{acc,t}^{\mathbb{Q}} + \lambda_{fail,t}^{\mathbb{Q}} + \lambda_{rest,t}^{\mathbb{Q}} + \lambda_{repud,t}^{\mathbb{Q}}, \quad (2)$$

$$L_t^{\mathbb{Q}} = \frac{\lambda_{acc,t}^{\mathbb{Q}}}{\lambda_t^{\mathbb{Q}}} L_{acc,t}^{\mathbb{Q}} + \frac{\lambda_{fail,t}^{\mathbb{Q}}}{\lambda_t^{\mathbb{Q}}} L_{fail,t}^{\mathbb{Q}} + \frac{\lambda_{rest,t}^{\mathbb{Q}}}{\lambda_t^{\mathbb{Q}}} L_{rest,t}^{\mathbb{Q}} + \frac{\lambda_{repud,t}^{\mathbb{Q}}}{\lambda_t^{\mathbb{Q}}} L_{repud,t}^{\mathbb{Q}}, \quad (3)$$

where the subscripts represent acceleration, failure to pay, restructuring, and repudiation. In a doubly stochastic setting, conditional on the pathes of the intensities, the probability that any two of the credit events happen at the same time is zero. Thus, $\lambda^{\mathbb{Q}}$ is naturally interpreted as the arrival rate of the first credit event of any type. Upon the occurrence of a credit event of type i , the relevant loss rate is $L_i^{\mathbb{Q}}$ and, given that a credit event has occurred, this loss rate is experienced with probability $\lambda_{it}^{\mathbb{Q}}/\lambda_t^{\mathbb{Q}}$. The corresponding $\lambda_i^{\mathbb{Q}}$ and $L_i^{\mathbb{Q}}$ may, of course, differ across countries.

For the purpose of valuing *CDS* contracts, the literature has typically assumed either that $\lambda^{\mathbb{Q}}$ follows a square-root diffusion (e.g., the Zhang (2003) analysis of Argentinean *CDS* spreads and the Longstaff, Mithal, and Neis (2004) analysis of U.S. corporate *CDS* spreads) or that that $\ln(\lambda^{\mathbb{Q}})$ follows an Ornstein-Uhlenbeck (Gaussian) process (e.g., the Berndt, Douglas, Duffie, Ferguson, and Schranck (2004) analysis of U.S. corporate *CDS* spreads). We explored in depth the relative goodness of fits of these two models, as well as a model in which $\lambda^{\mathbb{Q}}$ follows a "three-halves" diffusion, and concluded that, for the sovereign *CDS*

spreads and sample period examined, the lognormal model fits the best. Accordingly, we henceforth focus on the lognormal model for most of our analysis.

To set notation, we use the superscript \mathbb{Q} (\mathbb{P}) to denote the parameters of the process $\lambda^{\mathbb{Q}}$ under the risk-neutral (historical) distributions, respectively. We highlight a potential ambiguity in our notation here: we are discussing the properties of $\lambda^{\mathbb{Q}}$, as a stochastic process, under two different measures, \mathbb{Q} and \mathbb{P} . At this juncture, $\lambda^{\mathbb{P}}$, the arrival rate of default under the historical measure, is playing no role in our analysis. We comment briefly on the relation between $\lambda^{\mathbb{P}}$ and $\lambda^{\mathbb{Q}}$ in subsequent sections.

Under the historical measure \mathbb{P} , the risk-neutral mean arrival rate of a credit event is assumed to follow the process:

$$d \log \lambda_t^{\mathbb{Q}} = \kappa^{\mathbb{P}}(\theta^{\mathbb{P}} - \log \lambda_t^{\mathbb{Q}}) dt + \sigma_{\lambda^{\mathbb{Q}}} dB_t^{\mathbb{P}}. \quad (4)$$

The market price of risk η_t underlying the change of measure from \mathbb{P} to \mathbb{Q} for $\lambda^{\mathbb{Q}}$ is assumed to be an affine function of $\log \lambda_t^{\mathbb{Q}}$:

$$\eta_t = \delta_0 + \delta_1 \log \lambda_t^{\mathbb{Q}}. \quad (5)$$

This market price of risk allows κ and $\kappa\theta$ to differ across \mathbb{P} and \mathbb{Q} , while assuring that λ follows a lognormal process under both measures. Specifically, under the risk-neutral measure \mathbb{Q} , defined by the market price of risk η_t ,

$$d \log \lambda_t^{\mathbb{Q}} = \kappa^{\mathbb{Q}}(\theta^{\mathbb{Q}} - \log \lambda_t^{\mathbb{Q}}) dt + \sigma_{\lambda^{\mathbb{Q}}} dB_t^{\mathbb{Q}}, \quad (6)$$

where $\kappa^{\mathbb{Q}} = \kappa^{\mathbb{P}} + \delta_1 \sigma_{\lambda^{\mathbb{Q}}}$ and $\kappa^{\mathbb{Q}}\theta^{\mathbb{Q}} = \kappa^{\mathbb{P}}\theta^{\mathbb{P}} - \delta_0 \sigma_{\lambda^{\mathbb{Q}}}$.

Within this setting, closed-form solutions for zero-coupon bond prices and survival probabilities are not known. Accordingly, to price *CDS* contracts we assume that r_t and $\lambda^{\mathbb{Q}}$ are independent, and then construct a discrete approximation to

$$\int_t^{t_M} E_t^{\mathbb{Q}} \left[\lambda_u^{\mathbb{Q}} e^{-\int_t^u (r_s + \lambda_s^{\mathbb{Q}}) ds} \right] du = \int_t^{t_M} D(t, u) E_t^{\mathbb{Q}} \left[\lambda_u^{\mathbb{Q}} e^{-\int_t^u \lambda_s^{\mathbb{Q}} ds} \right] du$$

in terms of the price $D(t, u)$ of a default-free zero-coupon bond (issued at date t and maturing at date u) and the risk-neutral survival probabilities $E_t^{\mathbb{Q}} \left[e^{-\int_t^u \lambda_s^{\mathbb{Q}} ds} \right]$. The latter are then computed numerically using the Crank-Nicolson implicit finite-difference method to solve the associated Feynman-Kac partial differential equation.

Beyond the specification of the default arrival intensity, a critical input into the pricing of *CDS* contracts is the risk-neutral loss rate due to a credit event, $L^{\mathbb{Q}}$. Convention within both academic analyses and industry practice is to treat this loss rate as a constant parameter of the model. In the context of pricing corporate *CDS* contracts this practice has been questioned in the light of the evidence of a pronounced negative correlation between default rates and recovery over the business cycle (see, e.g., Altman, Brady, Resti, and Sironi (2003) and related publications by the U.S. rating agencies). A business-cycle induced correlation seems less compelling in the case of sovereign risk. Indeed, a theme we consistently heard in

conversations with sovereign *CDS* traders is that recovery depends on the size of the country (and the size and distribution of its external debt), but is not obviously cyclical in the same way that corporate recoveries are. In any event, we will follow industry practice and treat L^Q as a constant parameter of our pricing models, appropriately interpreted as the *expected* loss of face value on the underlying reference bond due to a credit event.

Traders are naturally inclined to call upon historical experience in setting loss rates in their pricing models. One source of this information is the agencies that rate sovereign debt issues. For example, Moody's (2003) estimates of the *recoveries* (weighted by issues sizes) on several recent sovereign defaults are: Argentina 28%; Ecuador 45%; Moldova 65%; Pakistan 48%; and Ukraine 69%. As stressed by Moody's, these numbers must be interpreted with some caution, because they are based on the market prices of sovereign bonds shortly after the relevant credit events. Moreover, just as in many discussions of corporate bond and *CDS* pricing, the setting of L^Q based on historical experience requires the assumption that there is no risk premium on recovery, $L^Q = L^P$.

That estimates of recovery may differ, depending on when market prices are sampled and perhaps also across measuring institutions, is confirmed by the recoveries estimated by *Credit Suisse First Boston (CSFB)*, as reported in the *Economist* (2004). The values at default of the bonds involved in Russia's default in May/June 1999 were 23.5% (15.9%) of face value, weighted (unweighted) by issue size. The corresponding numbers for Ecuador's default in October, 1999 were 23.4% (30.0%). Interestingly, at the time of restructuring, which in both of these cases was within a year of the default, the restructured values⁸ were substantially higher. For Russia they were 36.6% (38%), and for Ecuador they were 36.2% (49.3%). Singh (2003) provides additional examples of the market prices at the time of default being depressed relative to the subsequent amounts actually recovered, and that this phenomenon was more prevalent for sovereign than for corporate credit events. For valuing sovereign *CDS* contracts, it is the loss in value on the underlying bonds around the time of the credit event that matters for determining the payment from the insurer to the insured, regardless of whether or not these values accurately reflect the present values of the subsequently restructured debt.

At a practical level, to match a given day's term structure of new-issue *CDS* spreads, a range of combinations of L^Q and the set of parameters governing the Q -distribution of λ^Q will typically give a good fit. Several traders have told us that they set $L^Q = 0.75$ and then either bootstrap λ^Q or use a one-factor parametric model for the λ^Q process to match a day's cross-section of spreads. This particular standardized choice of L^Q (across maturities and countries) has, as we have just seen, some basis in historical experience. Whether it is in fact consistent with the historical behavior of spreads in the *CDS* contracts for a country is probably not material for the purpose of interpolating new-issue spreads across maturities.

On the other hand, the choice of L^Q is critical for marking to market seasoned *CDS* contracts (e.g., unwinding a seasoned position with a counterparty). In this situation, the price is not given by the market, but rather must be inferred from a model that requires

⁸This is the market value of the new bonds received as a percentage of of the original face value of the bonds.

as its inputs $L^{\mathbb{Q}}$ and the parameters of the stochastic \mathbb{Q} -process for $\lambda^{\mathbb{Q}}$. Accordingly, one is naturally led to inquire: Can $L^{\mathbb{Q}}$ and the conditional \mathbb{Q} distribution of $\lambda^{\mathbb{Q}}$ be separately identified from a *time-series* of market-provided spreads on newly issued *CDS* contracts? If the answer is yes, then the same pricing model can be used to mark to market the seasoned *CDS* contracts on the same issuer. We turn next to the challenges this separation presents for “reduced-form” *CDS* pricing models.

4 Can We Separately Identify $\lambda^{\mathbb{Q}}$ and $L^{\mathbb{Q}}$?

A common impression among academics and practitioners alike is that fixing $L^{\mathbb{Q}}$ at a specific value is necessary to achieve econometric identification. This is certainly true in an economic environment in which contracts are priced under the *fractional recovery of market value* convention (RMV) introduced by Duffie and Singleton (1999). In such a pricing framework, the product $\lambda^{\mathbb{Q}} \times L^{\mathbb{Q}}$ determines prices in the sense that the time- t spread on a defaultable bond takes the form

$$CDS_t^{RMV} = g(\lambda_t^{\mathbb{Q}} L^{\mathbb{Q}}), \quad (7)$$

for some function g . That $\lambda^{\mathbb{Q}}$ and $L^{\mathbb{Q}}$ enter symmetrically implies that they cannot be separately identified using defaultable bond data alone.

In the pricing framework of *fractional recovery of face value* (RFV) (see Duffie (1998) and Duffie and Singleton (1999)), which is the most natural pricing convention for *CDS* contracts, $\lambda^{\mathbb{Q}}$ and $L^{\mathbb{Q}}$ play distinct roles. Specifically, the *CDS* pricing relation in (1) takes the form

$$CDS_t = L^{\mathbb{Q}} f(\lambda_t^{\mathbb{Q}}). \quad (8)$$

Comparing equation (7) against (8), we can see that the joint identification problem in the RMV framework is no longer present for *CDS* prices. For example, the explicit linear dependence of CDS_t on $L^{\mathbb{Q}}$ implies that the ratio of two *CDS* spreads on contracts of different maturities does not depend on $L^{\mathbb{Q}}$, but does contain information about $\lambda^{\mathbb{Q}}$.

Now what is conceptually true need not be true in actual implementations of these pricing models, as is illustrated by the very similar prices for par coupon bonds under the pricing conventions RMV and RFV displayed in Duffie and Singleton (1999). To gauge the degree of numerical identification in practice, we perform the following analysis. Suppose that $\lambda^{\mathbb{Q}}$ follows a log-normal process⁹, $L^{\mathbb{Q}}$ is constant, and hence that $y_t = L^{\mathbb{Q}} \lambda_t^{\mathbb{Q}}$ also follows a log-normal process. More specifically, letting $X_t = \ln(\lambda^{\mathbb{Q}})$ and $Y_t = \ln(y_t)$, we have,

$$\begin{aligned} dX_t &= \kappa_x(\theta_x - X_t) dt + \sigma_x dB_t, \\ dY_t &= \kappa_y(\theta_y - Y_t) dt + \sigma_y dB_t, \end{aligned} \quad (9)$$

where $Y_t = X_t + \ln(L^{\mathbb{Q}})$, $\kappa_y = \kappa_x$, $\sigma_y = \sigma_x$, and $\theta_y = \theta_x + \ln(L^{\mathbb{Q}})$. Using this model we ask what happens to spreads as $L^{\mathbb{Q}}$ is varied holding y fixed. For this exercise, “fixed y ” means

⁹The particular dynamics of $\lambda^{\mathbb{Q}}$ is not crucial for the separate identification. For example, the same analysis goes through with the assumption that $\lambda^{\mathbb{Q}}$ follows a square-root process.

that the level of $y = L^Q \lambda^Q$ as well as its parameter values θ_y , κ_y , and σ_y are fixed. This, in turn, implies that any variation in L^Q is accompanied by an adjustment of $\lambda^Q = y/L^Q$ and its parameter values.

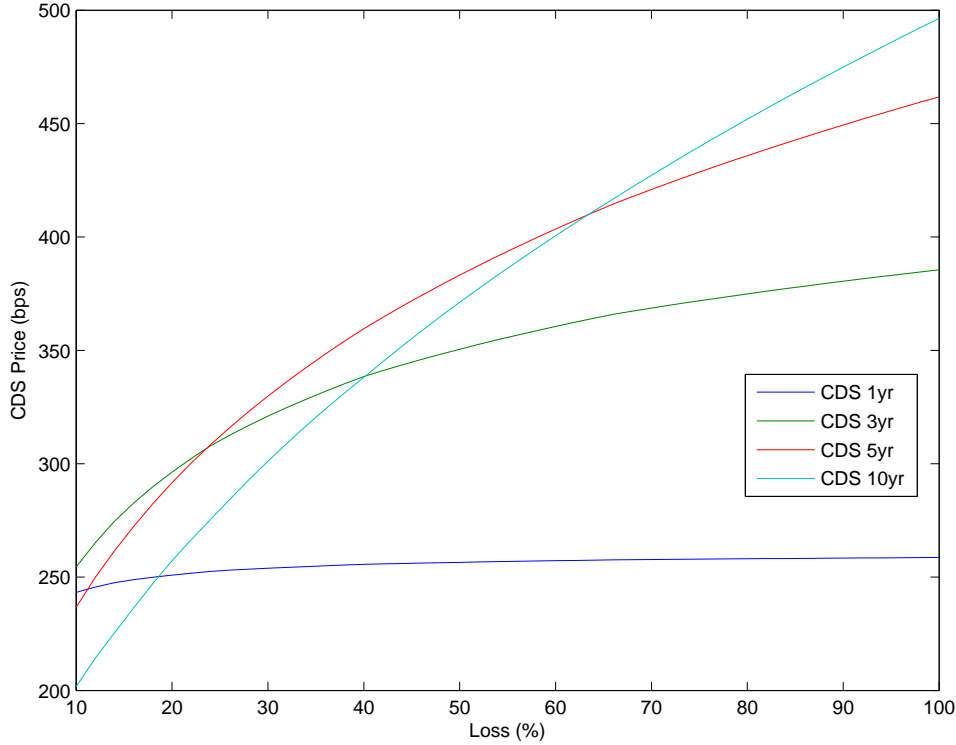


Figure 4: The sensitivity of *CDS* spreads to loss rate L^Q for fixed value of $y = L^Q \times \lambda^Q$. The level of y is fixed at $200bps$ and its parameter values are fixed at $\kappa_y = 0.01$, $\sigma_y = 1$ and $\theta_y = 200bps$.

Figure 4 illustrates the L^Q -sensitivity of *CDS* spreads, under the RFV convention, to variation in L^Q with $y = L^Q \times \lambda^Q$ fixed. The spreads clearly depend on L^Q and their sensitivity to changes in L^Q differs across maturities. This is to be contrasted against the RMV pricing framework in equation (7), under which the sensitivity of a defaultable bond to variation in L^Q is zero with fixed $y = L^Q \times \lambda^Q$. For these calculations we fixed the long-run mean of y to approximately reproduce the sample average of the five-year spread for Mexico, $e^{\theta_y^Q} = 200bps$; the volatility parameter was set at $\sigma_y = 1$, approximately the maximum likelihood estimate for this parameter; and the mean reversion parameter was set at $\kappa_y = 0.01$, between our maximum likelihood estimates for Mexico and Turkey (see Table 3).

Of course the degree of econometric identification may be sensitive to the choice of parameter values within the admissible regions of the parameter and state spaces. This is illustrated in Figure 5 by direct calculations of the partial derivatives $\partial CDS / \partial L^Q|_y$. Fixing

$L^Q = 75\%$, the top two panels of Figure 5 show that the $\partial CDS/\partial L^Q|_y$ are quite sensitive to changes in volatility (σ_y) and mean-reversion (κ_y). In particular, identification is strong when either volatility is relatively high or when the mean-reversion rate is low. Similarly, the bottom two panels of Figure 5 demonstrate that numerical identification is likely to be achieved over a wide range of values of $y = L^Q \times \lambda_t^Q$ and the loss rate L^Q . Moreover, the partial derivatives of the spreads are most sensitive to changes in the parameters for the longer maturity contracts. This is consistent with our prior that access to the term structure of *CDS* spreads enhances the numerical identification of L^Q separately from the parameters governing λ^Q .

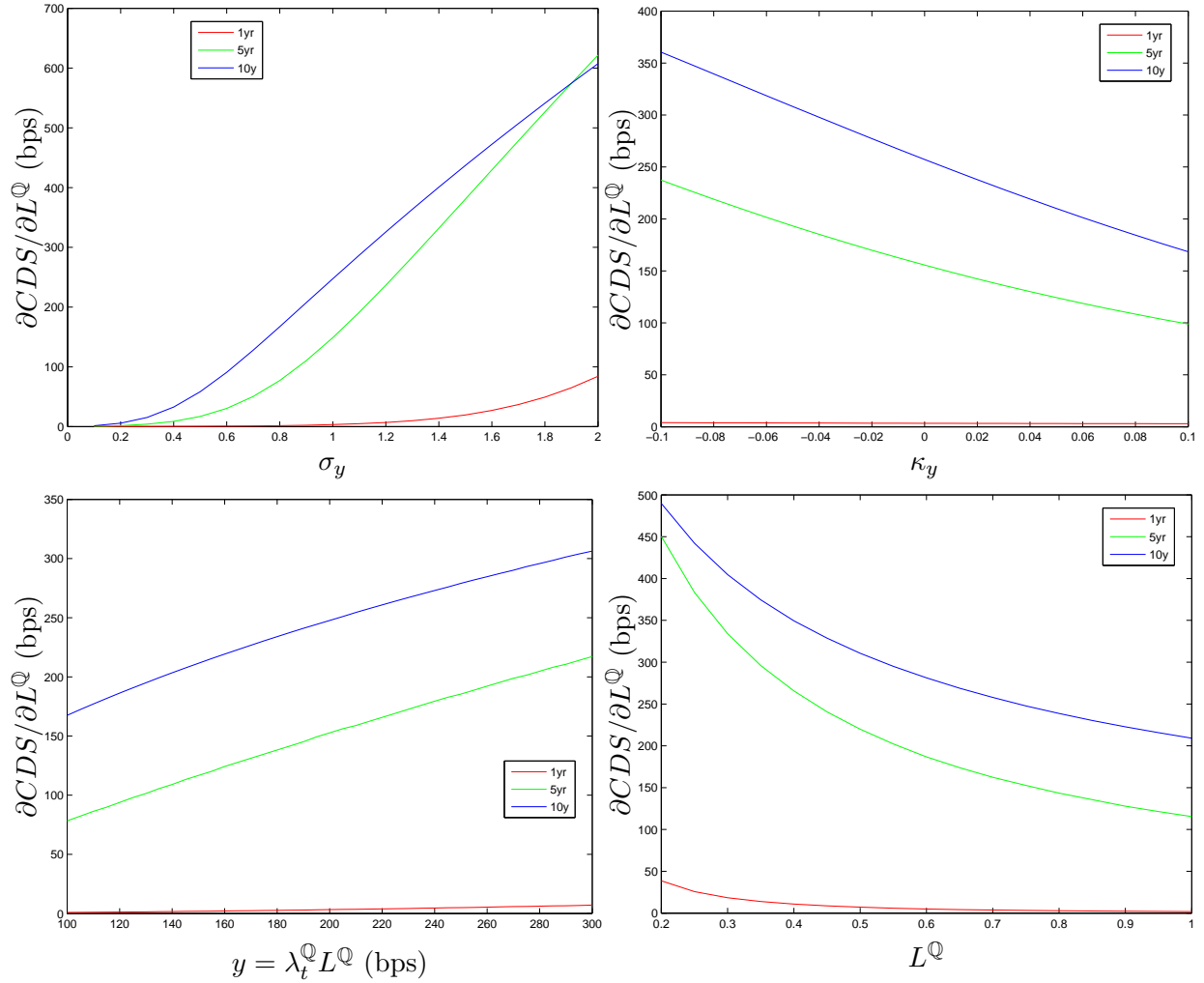


Figure 5: The partial derivative of CDS spread with respect to loss rate L^Q with fixed y : the level and parameter values of λ^Q are adjusted so that the process $y = L^Q \times \lambda^Q$ is kept fixed (both level and parameter values). In all figures, the base case parameters are: $\theta_y = 200$ bps, $\kappa_y = 0.01$, $\sigma_y = 1$, and $L^Q = 0.75$.

A natural question at this juncture is whether, with sample sizes that are available in the *CDS* markets, one can in fact reliably estimate L^Q in practice. To address this question we

conduct a small-scale Monte-Carlo exercise. Specifically, we simulate affine model-implied one-, three-, five-, and ten-year *CDS* spreads, and add normally distributed pricing errors to the one-, three- and ten-year *CDS* spreads.¹⁰ The resulting (noisy) simulated *CDS* data is then used to construct *ML* estimates of the underlying parameters. This was repeated one-hundred times, and the means and standard deviations of the *ML* estimates are displayed in Table 2. To gauge the effect of $\kappa^{\mathbb{Q}} < 0$, we consider two cases: one with explosive \mathbb{Q} -intensity ($\kappa^{\mathbb{Q}} < 0$), and the other with stationary \mathbb{Q} -intensity ($\kappa^{\mathbb{Q}} > 0$). To reduce the computational burden of estimation, we use a common coefficient $\sigma_{\epsilon}(M)$ for the volatilities of the one-, three-, and ten-year *CDS* pricing errors.

Table 2: Simulation results for the *affine* model

	$\theta^{\mathbb{P}}$	$\kappa^{\mathbb{P}}$	$\sigma_{\lambda^{\mathbb{Q}}}$	$\kappa^{\mathbb{Q}}$	σ_{ϵ}	$L^{\mathbb{Q}}$	$\theta^{\mathbb{Q}}\kappa^{\mathbb{Q}}$
explosive case							
true param	219bp	2.7880	0.1691	-0.3361	0.5000	0.7500	12bp
mean(estm)	224bp	3.1417	0.1704	-0.3458	0.5043	0.7265	12bp
std(estm)	41bp	0.8002	0.0007	0.0017	0.0069	0.0278	1bp
stationary case							
true param	219bp	2.7880	0.1691	0.1000	0.5000	0.7500	611bp
mean(estm)	232bp	3.2271	0.1711	0.0848	0.5046	0.7148	633bp
std(estm)	55bp	0.9935	0.0044	0.0073	0.0074	0.0135	7bp

Simulations are performed under the “true” parameter values with the same sample size as that of our *CDS* data. The mean and standard deviation of the estimates are calculated with 100 simulation runs.

The standard deviations of the simulated estimates are of the same orders of magnitude as the standard errors reported from the *ML* results for the affine model using the actual data, and the means of the simulated estimates are close in magnitude to the true parameter values. Moreover, for econometric identification, whether or not the default intensity is \mathbb{Q} -explosive appears to be inconsequential. The degree of persistence in $\kappa^{\mathbb{Q}}$ matters, of course, as was documented in Figure 5, but so long as $\lambda^{\mathbb{Q}}$ is reasonably persistent the likelihood function appears to exhibit sufficient curvature for reliable estimation of $L^{\mathbb{Q}}$.

5 Maximum Likelihood Estimates

The parameters were estimated by the method of maximum likelihood, with the conditional distribution of the spreads derived from the known conditional distribution of the state, which is lognormal. The five-year *CDS* contract was assumed to be priced perfectly, so that

¹⁰For reasons of tractability, we turn to an affine specification of $\lambda^{\mathbb{Q}}$. The components of the *CDS* prices can be computed analytically in this model and this substantially reduces the computational burden of our Monte Carlo analysis.

the pricing function could be inverted for $\lambda^{\mathbb{Q}}$.¹¹ The one-, two-, three-, and ten-year contracts were assumed to be priced with normally distributed errors with mean zero and standard deviations $\sigma_{\epsilon}(M)|Bid_t(M) - Ask_t(M)|$, where the $\sigma_{\epsilon}(M)$ are constants depending on the maturity of the contract, M . Time-varying variances that depend on the bid/ask spread allow for the possibility that the fits of our one-factor models deteriorate during periods of market turmoil when bid/ask spreads widen substantially. Conveniently, $\sigma_{\epsilon}(M)$ measures the degree of mis-pricing by the model relative to bid/ask spreads.

The risk-free interest rate (term structure) was assumed to be constant. We experimented with using a two-factor affine model (an $A_1(2)$ model in the nomenclature of Dai and Singleton (2000)) for $\lambda^{\mathbb{Q}}$, but we obtained virtually identical results to those for constant riskfree rate.¹² A simple arbitrage argument (see, e.g., Duffie and Singleton (2003)) shows that *CDS* spreads are approximately equal to the spreads on comparable maturity, par floating rate bonds from the same issuer as the reference bonds underlying the *CDS* contract. The prices of these bonds are not highly sensitive to the level of interest rates and this underlies the insensitivity of our findings to the introduction of a stochastic riskfree rate.

5.1 *ML* Estimates of One-Factor Models

The *ML* estimates of the parameters (expressed on an annual time scale) and their associated standard errors are presented in Table 3. Across all three countries, and regardless of whether $L^{\mathbb{Q}}$ is a fixed or free parameter, there is a striking contrast between the parameters governing the \mathbb{Q} - and \mathbb{P} -dynamics of $\lambda^{\mathbb{Q}}$. Indeed, in the cases of Mexico (constrained or unconstrained) and Turkey (unconstrained), the point estimates for $\kappa^{\mathbb{Q}}$ are negative, implying that the default intensity $\lambda^{\mathbb{Q}}$ is explosive under \mathbb{Q} ; while $\kappa^{\mathbb{P}} > 0$ so $\lambda^{\mathbb{Q}}$ is \mathbb{P} -stationary for all three countries. These large differences between the \mathbb{Q} and \mathbb{P} distributions are indicative of substantial market risk premiums related to uncertainty about future arrival rates of credit events.

From these parameter estimates, we can back out the coefficients for the market prices of risk, δ_0 and δ_1 , as defined in equation (5). The values for (Mexico, Turkey, Korea) are $d_0 = (-7.36, -2.29, -6.16)$ and $d_1 = (-1.35, -0.48, -0.98)$ in the constrained models with $L^{\mathbb{Q}} = 0.75$, and $d_0 = (-5.35, -2.03, -6.27)$ and $d_1 = (-1.16, -0.38, -0.98)$ in the unconstrained models. Recalling that $\kappa^{\mathbb{Q}} = \kappa^{\mathbb{P}} + \delta_1 \sigma_{\lambda^{\mathbb{Q}}}$ and $\kappa^{\mathbb{Q}} \theta^{\mathbb{Q}} = \kappa^{\mathbb{P}} \theta^{\mathbb{P}} - \delta_0 \sigma_{\lambda^{\mathbb{Q}}}$, the negative signs of δ_0 and δ_1 imply that the credit environment is much worse under \mathbb{Q} than under \mathbb{P} . More precisely, $\kappa^{\mathbb{Q}} \theta^{\mathbb{Q}} > \kappa^{\mathbb{P}} \theta^{\mathbb{P}}$ so, even at low arrival rates of credit events, $\lambda^{\mathbb{Q}}$ will tend to be larger under \mathbb{Q} than under \mathbb{P} . Moreover, for a given level of $\lambda^{\mathbb{Q}}$, there is more persistence under \mathbb{Q} than under \mathbb{P} (bad times last longer under \mathbb{Q}). It is this pessimism

¹¹The five-year contract was chosen because of its relatively liquidity. The liquidities of the five-year contracts are enhanced, for all three countries examined, by their inclusion in the Dow Jones CDX.EM traded index of emerging market *CDS* spreads.

¹²For checking the sensitivity our results to the presence of stochastic interest rates we once again shifted to an affine model for reasons of computational tractability. Within the affine setting we can allow for stochastic interest rates that are correlated with $\lambda^{\mathbb{Q}}$ and still obtain closed-form solutions for survival probabilities and zero-coupon bond prices.

	$L^{\mathbb{Q}}$ fixed at 0.75			$L^{\mathbb{Q}}$ unconstrained		
	Mexico	Turkey	Korea	Mexico	Turkey	Korea
$\kappa^{\mathbb{Q}}$	-0.0638 (0.0015)	0.0239 (0.0013)	0.0651 (0.0015)	-0.119 (0.003)	-0.0351 (0.0012)	0.0673 (0.0039)
$\theta^{\mathbb{Q}}\kappa^{\mathbb{Q}}$	0.268 (0.007)	-0.015 (0.004)	-0.384 (0.007)	0.661 (0.014)	0.480 (0.006)	-0.414 (0.043)
$\sigma_{\lambda^{\mathbb{Q}}}$	1.086 (0.004)	1.144 (0.004)	0.921 (0.007)	0.773 (0.015)	0.811 (0.006)	0.934 (0.018)
$\kappa^{\mathbb{P}}$	1.40 (1.15)	0.57 (0.56)	0.97 (0.66)	0.78 (0.67)	0.28 (0.31)	0.99 (0.68)
$\theta^{\mathbb{P}}$	-5.51 (0.59)	-4.61 (1.54)	-6.25 (0.69)	-4.45 (0.69)	-4.23 (2.44)	-6.35 (0.71)
$\sigma_{\epsilon}(1)$	1.436 (0.032)	1.056 (0.021)	0.619 (0.028)	1.472 (0.035)	1.069 (0.021)	0.618 (0.028)
$\sigma_{\epsilon}(2)$	1.084 (0.018)	0.858 (0.026)	0.442 (0.026)	1.057 (0.018)	0.839 (0.026)	0.442 (0.026)
$\sigma_{\epsilon}(3)$	0.933 (0.031)	0.595 (0.018)	0.296 (0.009)	0.935 (0.032)	0.586 (0.017)	0.296 (0.009)
$\sigma_{\epsilon}(10)$	0.838 (0.022)	1.350 (0.040)	0.869 (0.028)	0.855 (0.023)	0.885 (0.018)	0.867 (0.029)
$L^{\mathbb{Q}}$	=0.75 N/A	=0.75 N/A	=0.75 N/A	0.231 (0.010)	0.236 (0.004)	0.833 (0.129)
mean llk	32.030	27.213	36.626	32.126	27.700	36.626

Table 3: Maximum likelihood estimates based on daily data from March 19, 2001 through August 8, 2006. The sample size is 1357 for Mexico, 1377 for Turkey, and 1308 for Korea. llk is the sample average of log-likelihood.

about the credit environment that allows risk-neutral pricing to recover market prices in the presence of investors who are adverse to default risk.

Turning to the magnitudes of the pricing errors for the *CDS* contracts with maturities of one, two, three, and ten years, the estimates of $\sigma_{\epsilon}(M)$ in Table 3 measure the standard deviations of the pricing errors in units of the bid/ask spreads. Typically, $\sigma_{\epsilon}(M)$ is less than about one, the most notable exceptions being $\sigma_{\epsilon}(1)$ for Mexico (with or without $L^{\mathbb{Q}}$ constrained) and $\sigma_{\epsilon}(10)$ for Turkey with $L^{\mathbb{Q}} = 0.75$. Korea shows the best fit in that the $\sigma_{\epsilon}(M)$ are relatively small, as are the bid/ask spreads on these contracts (see Figure 3). For a given country, the $\sigma_{\epsilon}(M)$ tend to be smaller for the intermediate maturities, and the bid/ask spreads fall (on average, as seen from Table 1) with increasing maturity, so our models tend to fit somewhat better for $M = 3$ and 5 than for $M = 1$ or 10.

The time-series of *CDS* pricing errors, measured by the market minus the model-implied spreads and evaluated at the parameters with $L^{\mathbb{Q}} = 0.75$, are plotted in Figure 6. The high degree of comovement in the *CDS* spreads across maturities and countries is much less

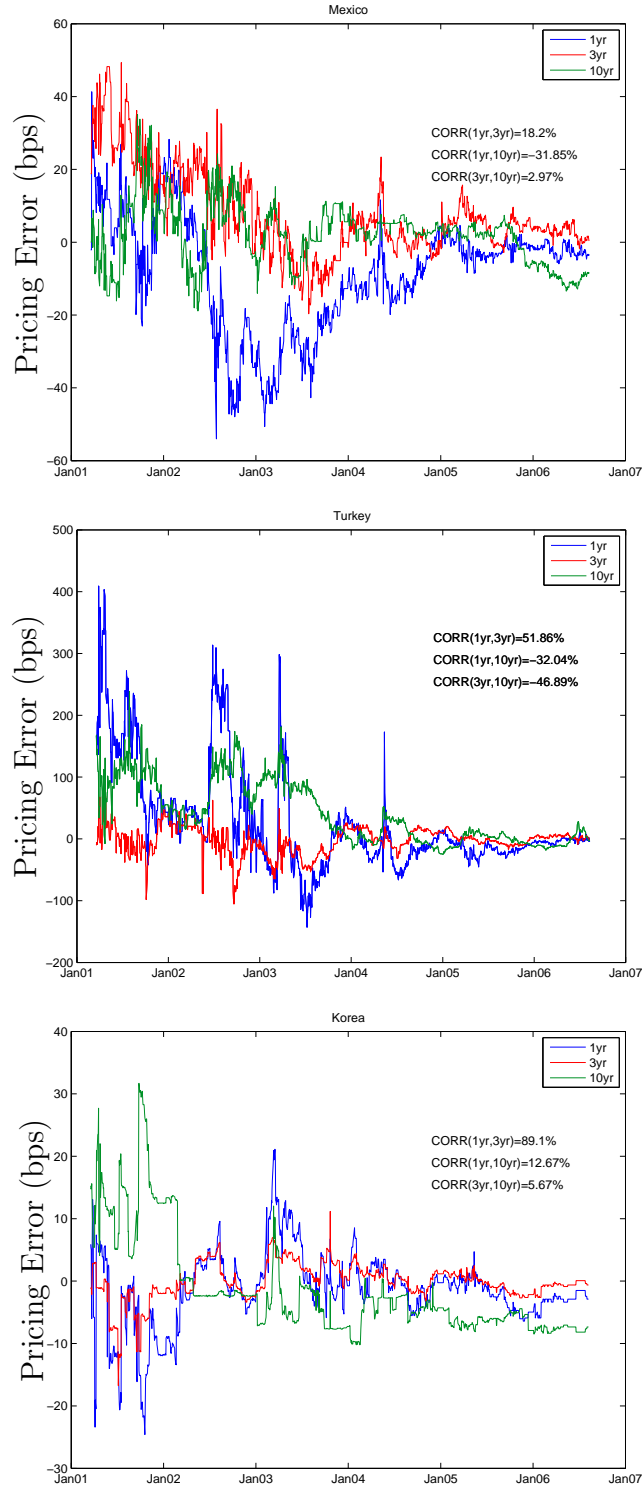


Figure 6: The *CDS* pricing errors, market *CDS* price minus the model implied, for maturities of 1yr, 3yr, and 5yr. These errors are evaluated the constrained maximum likelihood estimates with $L^Q = 0.75$.

evident in the corresponding pricing errors. In the cases of Mexico and Turkey, the pricing errors on the one- and ten-year contracts are negatively correlated suggesting that there is some tension in fitting both of these spreads simultaneously. For Korea, on the other hand, our one-factor model appears to price the short-dated contracts equally well in that $Corr(\epsilon(1), \epsilon(3)) = 0.89$. The pricing errors on long-dated Korean contracts move in a largely uncorrelated way with those at the short end. A more indepth analysis of these pricing errors and the potential role of a second factor is explored in Section 5.2. At this juncture we simply highlight the small magnitudes of the volatilities of these errors, typically less than one bid/ask spread.

There are several notable differences between the maximum likelihood estimates of the models with and without $L^{\mathbb{Q}}$ fixed. Perhaps most striking is the fact that the unconstrained estimates of $L^{\mathbb{Q}}$ for Mexico and Turkey are approximately 0.23, much smaller than market convention of 0.75. Standard likelihood ratio statistics reject the constraint $L^{\mathbb{Q}} = 0.75$ at conventional significance levels. On the other hand, for Korea the estimate is quite close to the market convention. Accompanying the relatively small values of $L^{\mathbb{Q}}$ for Mexico and Turkey are relatively larger values of $\kappa^{\mathbb{Q}}\theta^{\mathbb{Q}}$ and smaller values of both $\kappa^{\mathbb{Q}}$ and $\kappa^{\mathbb{P}}$ (compared to their counterparts in the models with $L^{\mathbb{Q}} = 0.75$). The larger values of $\kappa^{\mathbb{Q}}\theta^{\mathbb{Q}}$ are intuitive: to match spreads with a lower loss rate, the “intercept” of the $\lambda^{\mathbb{Q}}$ process under the \mathbb{Q} distribution must be larger.¹³

The log-likelihood function for our model has two components: the component associated with the dynamic properties of $\lambda^{\mathbb{Q}}$ under \mathbb{P} and the component associated with the pricing errors. It is the change in the first of these components that accounts for the increase in the likelihood function when the constraint $L^{\mathbb{Q}} = 0.75$ is relaxed. Accordingly, to gain further insight into both the different properties of the constrained and unconstrained models, and their relative goodness-of-fits, we examine the small-sample distributions of various moments of the *CDS* spreads and their first differences (time changes). Ten-thousand time series, each of length 1500 (the approximate length of our samples), are simulated and the means and standard deviations of the small-sample distributions of various moments are computed. Among the moments examined are the mean, standard deviation, skewness, and kurtosis, and the autocorrelations of the levels of *CDS* spreads and the slope of the *CDS* curve.

Table 4 displays the means and standard deviation of the small-sample distributions of mean, skewness, and kurtosis for Mexico and Turkey, along with their sample counterparts. For the first through fourth central moments, the differences between the means of the small-sample distributions across the corresponding models with and without $L^{\mathbb{Q}}$ constrained are small, certainly relative to the standard deviations of these distributions. Moreover, the means of the small sample distributions of the first, second (not shown), and third moments are quite close to their historical counterparts, particularly in the case of Mexico. There is a tendency for the sample kurtoses to be below their model-implied small-sample counterparts, but the former are within one standard deviation of the latter.

At first glance, we expected larger differences in the implied autocorrelations of *CDS*

¹³Conditional on $\lambda_t^{\mathbb{Q}}, \lambda_{t+1}^{\mathbb{Q}}$ will tend to be larger in the model with the lower estimate of $L^{\mathbb{Q}}$. Since $\kappa^{\mathbb{Q}} < 0$ in the unconstrained models for Mexico and Turkey, $\lambda^{\mathbb{Q}}$ does not have a finite \mathbb{Q} -mean.

Moment	Mexico			Turkey		
	Sample	MC^C	MC^U	Sample	MC^C	MC^U
$E[1yr]$	55	59 [18]	57 [21]	355	306 [183]	294 [168]
$E[5yr]$	166	155 [40]	151 [47]	563	504 [191]	495 [193]
$E[10yr]$	213	200 [39]	195 [47]	607	520 [152]	531 [175]
$Skew[1yr]$	0.95	1.28 [.56]	1.16 [.60]	1.09	1.50 [.69]	1.31 [.73]
$Skew[5yr]$	0.74	0.94 [.49]	0.84 [.54]	0.51	0.97 [.57]	0.88 [.61]
$Skew[10yr]$	0.62	0.71 [.45]	0.67 [.50]	0.48	0.89 [.54]	0.92 [.60]
$Kurt[1yr]$	2.64	4.86 [2.3]	4.34 [2.2]	3.24	5.53 [3.3]	4.75 [3.0]
$Kurt[5yr]$	2.65	3.75 [1.6]	3.44 [1.6]	2.10	3.75 [1.8]	3.49 [1.7]
$Kurt[10yr]$	2.56	3.26 [1.2]	3.11 [1.2]	2.02	3.58 [1.6]	3.60 [1.8]
$ACF1(5yr)$	0.996	0.989 [.005]	0.992 [.004]	0.995	0.992 [.004]	0.994 [.003]
$ACF2(5yr)$	0.991	0.978 [.009]	0.984 [.007]	0.991	0.985 [.007]	0.988 [.006]
$ACF1(slope)$	0.993	0.990 [.004]	0.993 [.003]	0.963	0.985 [.008]	0.991 [.006]
$ACF2(slope)$	0.988	0.981 [.008]	0.986 [.007]	0.940	0.970 [.016]	0.983 [.012]

Table 4: The means and standard deviations (in brackets) of the small sample distributions of the moments of the one-, five-, and ten-year CDS spreads. MC^C refers to Monte Carlo results for the model with $L^Q = 0.75$, and MC^U refers to the Monte Carlo results for the models with unconstrained L^Q . $ACF1$ and $ACF2$ refer to the first- and second-order autocorrelations, respectively, and $slope$ is the ten minus one-year spread.

spreads across the constrained (C) and unconstrained (U) models, because $\kappa^{\mathbb{P}^C} > \kappa^{\mathbb{P}^U}$ (see Table 3). However our models are parameterized on an annual time scale so, over moderate horizons, the differences in model-implied autocorrelations of CDS spreads are small. This can be seen from the last four rows of Table 4. The model-implied autocorrelations for the slope for Turkey are a bit larger than their sample counterparts, but otherwise the model and sample autocorrelations are very similar. Of course the higher degree of \mathbb{P} persistence with L^Q treated as a free parameter will manifest itself over sufficiently long horizons. However, the effects of $\kappa^{\mathbb{P}^C} > \kappa^{\mathbb{P}^U}$ on our analysis of risk premiums in Section 6 were negligible at the one-year horizon. At the five-year horizon, the differences were again negligible for Mexico, though they were material for Turkey.

In the light of all of these findings, how should we set L^Q ? Consistent with our theoretical and small-sample analyses in Section 4, the choice of L^Q does matter. Yet many features of the implied distributions of CDS spreads, and the model-implied pricing errors, are strikingly similar across a wide range of values of L^Q . Additionally, since the primary difference (at least as revealed by the moments we examined) was in the \mathbb{P} -persistence properties of λ^Q , it seems that a long time series of CDS spreads, surely much longer than our five-year sample, is needed to make a compelling case for setting L^Q other than at the market convention of 0.75.¹⁴ Additionally, there is the possibility of model-misspecification compromising our

¹⁴Simply because $L^Q = 0.75$ is market convention is not sufficient, in our minds for accepting this value as the best description of history. Market makers typically set L^Q in matching the cross-maturity prices of

models’ abilities to fit the highly persistent and volatile nature of spreads for Mexico and Turkey. Korean spreads are equally persistent, but they are smaller and less volatile, and it seems plausible that our lognormal model is a somewhat better approximation for these spreads. Given that our results for Korea are supportive of market convention and that most of our subsequent analysis is (qualitatively) robust to the choice of L^Q , we henceforth focus on the case of $L^Q = 0.75$.

5.2 Is One Factor Enough?

Up to this point we have chosen to focus on a single-factor model for λ^Q , largely because, for a given sovereign, the first *PC* of the *CDS* spreads explains a very large percentage of the variation for all maturities. However, the preceding discussion of pricing errors in one-factor models leads us naturally to inquire about the dimensions along which an additional factor might improve the fit of our model, if at all.

Mat.	Mexico				Turkey				Korea			
	PC1		PC2		PC1		PC2		PC1		PC2	
	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2
1yr	0.22	89.8%	0.59	8.4%	0.46	97.2%	0.78	2.8%	0.35	95.4%	0.56	4.3%
2yr	0.38	97.5%	0.49	2.1%	0.47	99.8%	0.13	0.1%	0.40	98.2%	0.39	1.7%
3yr	0.47	99.3%	0.25	0.4%	0.46	99.8%	-0.16	0.1%	0.43	99.5%	0.19	0.4%
5yr	0.54	99.4%	-0.31	0.4%	0.43	99.1%	-0.40	0.8%	0.48	99.8%	-0.10	0.1%
10yr	0.54	98.6%	-0.50	1.1%	0.40	98.6%	-0.44	1.2%	0.55	97.1%	-0.70	2.8%

Table 5: OLS Regressions of *CDS* Spreads on their First Two Principal Components. $\hat{\beta}$ is the estimated loading and R^2 is the coefficient of determination for the regression.

Table 5 displays the factor loadings and the percentage variation explained from projections of the *CDS* spreads onto the first two *PCs* of the data.¹⁵ As noted at the outset of our analysis, *PC1* explains a large percentage of the variation in spreads for all countries and all maturities. Indeed, for maturities of three years and longer, *PC1* accounts for at least 97% of the variation in all of the spreads. Moreover, parallel to the findings for the term-structures of the *US* treasury or swap markets (Litterman and Scheinkman (1991)), the first *PC* emerges as a “level” factor, as reflected in the roughly constant factor loadings across maturities (for a given sovereign). As expected, our one-factor model with default intensity λ^Q picks up this level factor: regressing the time series of model-implied λ^Q onto *PC1* yields an R^2 of 99.0% for Mexico, 98.6% for Turkey, and 98.7% for Korea.

CDS contracts on a given day. This does not require (or typically involve) calibrations to history or explicit analyses of the market prices of risk. Therefore, the question of what is the best setting of L^Q for matching the time-series properties of spreads, both in the *CDS* and associated bond markets, is a useful line of inquiry.

¹⁵This *PC* analysis was conducted using the covariance matrix of the levels of spreads.

Moment $X = 10\text{yr} - 1\text{yr}$	Korea		Mexico		Turkey	
	Sample	MC	Sample	MC	Sample	MC
$E[X]$	34	46 [.5]	158	141 [.21]	229	214 [.45]
$Corr(X, 1\text{yr})$	0.60	0.87 [.12]	0.77	0.96 [.02]	-0.60	-0.33 [.57]
$Corr(X, 2\text{yr})$	0.67	0.88 [.11]	0.87	0.96 [.02]	-0.48	-0.30 [.57]
$Corr(X, 3\text{yr})$	0.72	0.88 [.11]	0.90	0.97 [.02]	-0.43	-0.27 [.58]
$Corr(X, 5\text{yr})$	0.77	0.90 [.11]	0.95	0.98 [.01]	-0.37	-0.23 [.60]
$Corr(X, 10\text{yr})$	0.85	0.93 [.10]	0.96	0.99 [.01]	-0.35	-0.21 [.61]
Moment $X = \Delta 10\text{yr} - \Delta 1\text{yr}$	Korea		Mexico		Turkey	
	Sample	MC	Sample	MC	Sample	MC
$Corr(X, \Delta 1\text{yr})$	-0.36	0.58 [.26]	-0.04	0.88 [.08]	-0.77	-0.63 [.35]
$Corr(X, \Delta 2\text{yr})$	-0.09	0.60 [.26]	0.40	0.89 [.07]	-0.58	-0.58 [.36]
$Corr(X, \Delta 3\text{yr})$	-0.005	0.62 [.25]	0.52	0.90 [.06]	-0.50	-0.53 [.38]
$Corr(X, \Delta 5\text{yr})$	0.11	0.67 [.24]	0.67	0.94 [.05]	-0.39	-0.47 [.41]
$Corr(X, \Delta 10\text{yr})$	0.33	0.74 [.22]	0.80	0.97 [.04]	-0.16	-0.44 [.43]

Table 6: The means and standard deviations (in brackets) of the small sample distributions of the moments of the 10yr - 1yr slope. Ten thousand time series, each of length 1500, were simulated and the sample moments for each series were computed. The top (bottom) half sets X equal to the spread (changes in the spread) between the ten- and one-year CDS rates. Standard deviations of the small-sample distributions are given in brackets.

As an additional, more demanding check on the fit of our models, we display in Table 6 the correlations between the CDS spreads and the slopes of the CDS curves, using levels and first differences, for the historical sample and as implied by our models.¹⁶ Though the patterns in these correlations are quite different across countries (most notably the different signs for Turkey versus Korea and Mexico), our one-factor models match the correlations in levels quite closely. The models do less well at matching the correlations among the changes in levels and slopes, though this is to be expected as the changes are essentially daily innovations in these variables. Even for changes, the match is quite good for Turkey at all maturities and for Mexico and Korea at the longer maturities.

Our ML estimation results indicate that, among the various maturities, the one-factor model mis-prices the one-year contract most severely. As we have just seen, our models are also challenged by the low degree of correlation between innovations in the one-year CDS spreads and the slopes of the CDS curves. Taken together, these observations suggest that there are components of the short ends of the CDS curves that are not well captured by our one-factor models. Further support for this assessment comes from regressing the one-year pricing errors on $PC2$ which gives R^2 's of 67.6% for Mexico, 45.9% for Turkey, and 65.1% for Korea. The corresponding R^2 for the pricing errors on longer maturity contracts decline substantially with maturity in the cases of Mexico and Turkey, suggesting that what $PC2$ is

¹⁶The first row of Table 6 confirms that our models do a reasonable job of matching the average slopes of the CDS curves for our sample period.

picking up is primarily a short-maturity phenomenon in these markets.

Based on conversations with traders, it seems that the most likely explanation for this “anomalous” behavior of the one-year contract is due to a liquidity or supply/demand premium. We are told that large institutional money management firms often use the short-dated *CDS* contract as a primary trading vehicle for expressing views on sovereign bonds. The sizable trades involved in these transactions introduce an idiosyncratic “liquidity” factor into the behavior of the one-year contract.

Of interest then is whether or not there is a component of the bid/ask spreads that is orthogonal to the first *PC* of spreads; that is, whether there are large idiosyncratic components of the bid/ask spreads for specific maturities.¹⁷ This question is answered in Table 7 where we report the results from regressing the bid/ask spreads of the individual *CDS* contracts onto the first two principal components of the bid/ask spreads for Mexico and Turkey. There is a small role for a second factor in the bid/ask spreads, concentrated almost entirely at the one- and ten-year maturity points. These patterns suggest that there might indeed be something special about the one- and possibly ten-year contracts from a liquidity perspective. The roles of such illiquidity or trading pressures on *CDS* spreads are issues that we hope to explore in future research.

Mat.	Mexico				Turkey			
	PC1		PC2		PC1		PC2	
	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2	$\hat{\beta}$	R^2
1yr	0.44	89.7%	-0.79	8.2%	0.57	93.7%	0.65	4.9%
2yr	0.47	93.2%	-0.18	0.4%	0.48	95.2%	0.17	0.5%
3yr	0.44	93.8%	0.18	0.4%	0.45	94.6%	-0.25	1.1%
5yr	0.44	95.6%	0.37	1.9%	0.37	92.8%	-0.38	4.0%
10yr	0.45	93.9%	0.42	2.3%	0.33	81.6%	-0.59	10.4%

Table 7: OLS Regressions of *CDS* Bid/Ask Spreads on the First Two Principal Components of Bid/Ask Spreads for Mexico and Turkey.

6 On Priced Risks in Sovereign *CDS* Markets

The large differences between the parameters governing $\lambda^{\mathbb{Q}}$ under the risk-neutral and the actual measures suggest that there is a systematic risk related to changes in future arrival rates of sovereign credit events that is priced in the *CDS* market. To examine the economic underpinnings of the priced risks in the sovereign *CDS* markets, we take the *ML* estimates obtained in Section 5 and construct two measures of fitted *CDS* spreads. The first is the

¹⁷The bid/ask spreads are highly correlated with the corresponding levels of spreads. In particular, the correlations between *PC1* of the *CDS* spreads (contract prices) and *PC1* of the bid/ask spreads are 80.7% for Mexico and 86.3% for Turkey.

actual fitted spread $CDS_t(M)$ from (1). The second is

$$CDS_t^{\mathbb{P}}(M) = \frac{2(1 - R^{\mathbb{Q}}) \int_t^{t+M} E_t^{\mathbb{P}} \left[\lambda_u^{\mathbb{Q}} e^{-\int_t^u (r_s + \lambda_s^{\mathbb{Q}}) ds} \right] du}{\sum_{j=1}^{2M} E_t^{\mathbb{P}} \left[e^{-\int_t^{t+.5j} (r_s + \lambda_s^{\mathbb{Q}}) ds} \right]}, \quad (10)$$

obtained from (1) by replacing all of the expectations $E^{\mathbb{Q}}$ with expectations under the physical measure \mathbb{P} , $E^{\mathbb{P}}$. If market participants are neutral towards the risk of variation over time in $\lambda^{\mathbb{Q}}$, then $CDS_t^{\mathbb{P}}(M)$ should replicate the corresponding market price $CDS_t(M)$. Put differently, a mark-up in the CDS spread relative to the pseudo-spread implies that the buyer of the CDS contract is willing to pay a premium for holding the CDS contract, while the seller demands a premium. This is similar to what is found in equity options markets where the time-variation of volatility is a priced risk. To quantify the role of risk premiums regarding variation in $\lambda^{\mathbb{Q}}$, in percentage terms, we report¹⁸

$$CRP_t(M) \equiv (CDS_t(M) - CDS_t^{\mathbb{P}}(M)) / CDS_t^{\mathbb{P}}(M). \quad (11)$$

The percentage contribution of the risk premiums to spreads at the one-year maturity ($CRP_t(1)$) are displayed in Figure 7. The correlations between the CRP 's are 93.29% for (Mexico, Turkey), 82.48% for (Mexico, Korea), and 90.05% for (Turkey, Korea). This high degree of comovement in the CRP 's is striking given the very different credit qualities and geo-political features of the three countries examined. Risk premiums induced more volatility in the spreads during the early part of our sample, with the gap between CDS_t and $CDS_t^{\mathbb{P}}$ (on a percentage basis) being most volatile for Mexico. During the later period of our sample, when spreads in the credit markets were tight and when talks of “reaching for yield” were prevalent, the CRP 's (as seen through our lognormal model) turned negative. Figure 8 shows that $CRP_t(M)$ tends to increase with maturity.¹⁹ Evidently, not only does risk increase with horizon, but its effect on premiums increases on a *percentage* basis as the maturity of the contract increases. Additionally, unlike in the case of the one-year contract, the CRP 's do not become negative at the long end of the maturity spectrum.

To assist in interpreting the various “peaks” in the contributions of risk premiums to spreads during our sample period we have marked in Figure 8 the dates of several key economic events around the times of these peaks. The early part of our sample was dominated by economic and political events in South America. Argentina faced an economic crisis in the spring of 2001 and President de la Rúa removed his Minister of Economics and introduced a fiscal austerity program. This was followed in the summer of 2001 by a “zero-deficit” plan in an attempt to avoid major bank runs and reverse the depletion of foreign reserves (Zhang (2003)). A year later, in the summer of 2002, the prospect of the left-wing candidate Lula Da Silva winning the presidential elections in Brazil riled sovereign debt markets. He

¹⁸We stress that neither CDS_t nor $CDS_t^{\mathbb{P}}$ involve the physical intensity $\lambda^{\mathbb{P}}$. As emphasized by Jarrow, Lando, and Yu (2005) and Yu (2002), this information cannot be extracted from bond or CDS spread data alone.

¹⁹This measure of the effects of premiums on spreads is larger still when $M = 10$.

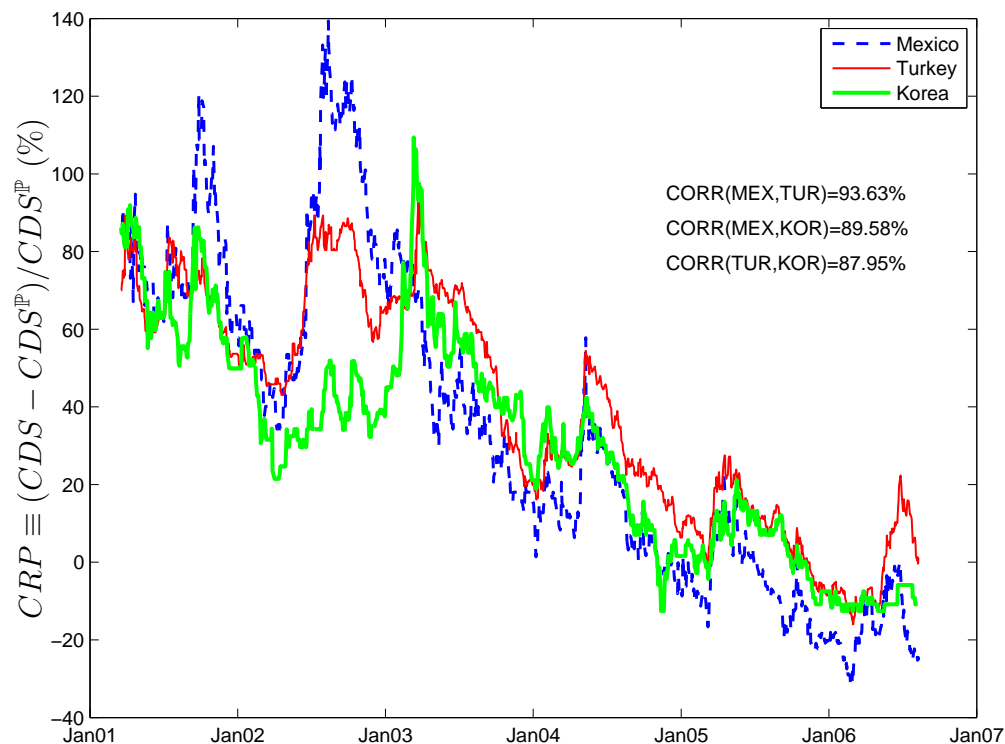


Figure 7: The percentage difference between the one-year *CDS* price and the one-year *pseudo-CDS* for Mexico, Turkey and Korea.

subsequently won the election in October of that year. Perhaps not surprisingly, all of these political developments in South America had much larger effects on the risk premiums for Mexico than for Turkey or Korea.

The simultaneous and large jumps both in *CDS* spreads and the *CRP*'s during May, 2004 have their roots in investors' portfolio reallocations due to macroeconomic developments in the U.S. During the second quarter of 2004 there was a substantial increase in non-farm payrolls in the U.S. This, combined with comments by representatives of the Federal Reserve, led market participants to expect a tightening of monetary policy. A reason that these concerns had large and widespread effects on spreads is that both financial institutions and hedge funds had substantial positions in "carry trades." They were borrowing short-term in dollars and investing in long-term bonds, often high-yield and emerging market bonds issued in various currencies. The unexpected strength in the U.S. economy led to an unwinding of some of these trades and, consequently, an across the board adjustment in spreads on corporate and sovereign credits.²⁰ This episode illustrates the importance of changes in investors' appetite for exposure to credit, as a global risk class, for co-movements

²⁰These concerns were widely noted in the media at the time. "In a single day, May 7, yields on Brazilian bonds jumped 1.52 percentage points as the unexpectedly strong jobs report in the U.S. increased the likelihood of higher short-term rates. (Henry (2004))." See also the discussion in Cogan (2005).

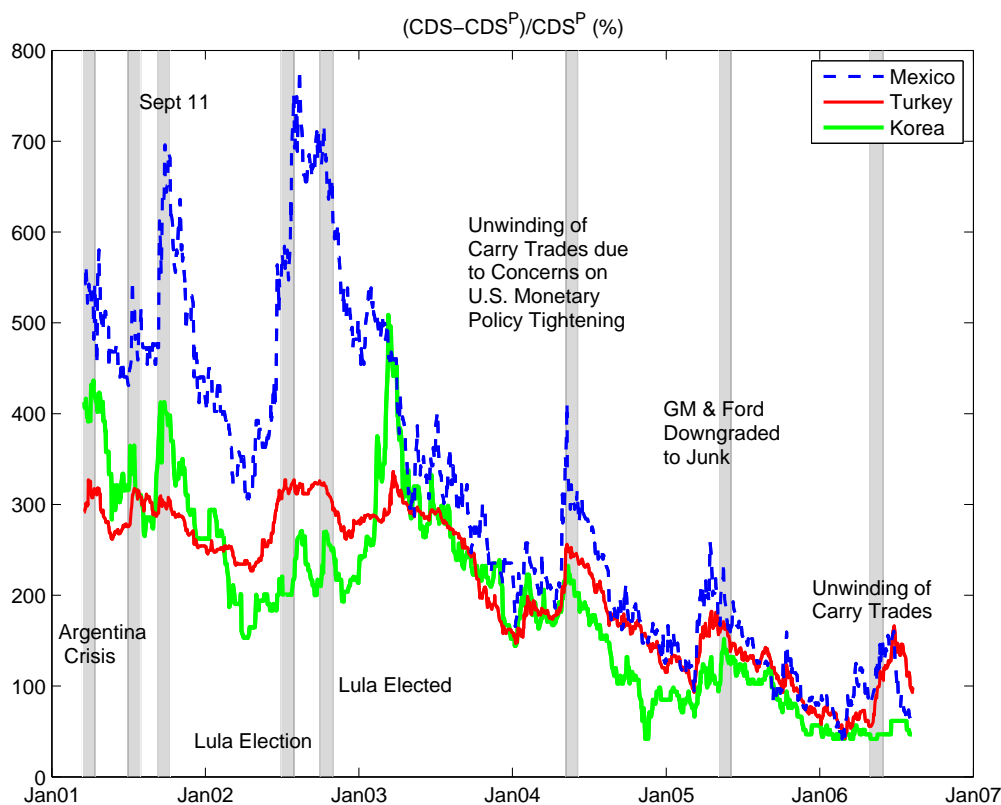


Figure 8: $CRP(5) \equiv (CDS - CDS^P) / CDS^P$ for Mexico, Turkey, and Korea, computed using the five-year CDS contract.

in yields. The induced changes in yields on the sovereign credits examined here (apparently) had nothing directly to do with the inherent credit qualities of the issuers.

In March of 2005 there were similarly sized run-ups in $CRP_t(5)$ associated with the deteriorating credit quality of General Motors and Ford in the U.S. In the middle of March Fitch downgraded GM, S&P changed its rating outlook to negative, and Moody's placed GM on review for a downgrade. These changes were followed with similarly negative outlooks on Ford in early April, 2005. Concurrently, there was a substantial widening of spreads not only on the individual-name CDS contracts for these issuers, but also on high-yield corporate indices (e.g., Packer and Wooldridge (2005)). Figure 8 shows that the retrenchment in high-yield positions extended to emerging markets as well.

Finally, $CRP(5)$ shows a sizable increase during the late spring of 2006. Once again the evidence supports an increased aversion to exposure to emerging market credit risk rather than reassessments of the fundamental economic strengths of individual countries. There was a broad sell-off in emerging market equities and a concurrent correction in foreign currency markets as hedge funds and other leveraged investors unwound carry trades in the emerging market currencies (e.g., IMF (2006)). During this episode Turkey in particular experienced

large balance of payments pressures on its currency, as well as domestic political uncertainties related to its EU accession.

An interesting feature of the time-series of $CRP(5)$'s in Figure 8 is that adjustments to Mexico's risk premiums had the largest percentage effects on spreads throughout most of our sample period. During the first half of our sample this is no doubt attributable to the political and economic upheavals in Latin America. The gaps between the countries' $CRP(5)$ are smaller during the second half of our sample, and the events in early 2006 had the largest effect on Turkey. As noted, this was most likely a manifestation of domestic policy and political issues in Turkey at the time.

Another striking country-specific episode in Figure 8 is the brief, but large, run-up in $CRP(5)$ for Korea in the early part of 2003. This was a period of rising delinquencies on credit card debts following a very rapid expansion in consumer borrowing. Concurrently, the financial stability of several credit card companies and investment trusts were called into question (Kang (2004)). In addition, the conglomerate SK Global reported material accounting irregularities in March, 2003 and this contributed to existing concerns about the stability of the Korean financial system (Cooper and Madigan (2003)).

Comparison of Figures 3 and 8 suggests that episodes when the risk premiums associated with variation in λ^Q were large (as measured by CRP) were also episodes when the bid/ask spreads on the CDS contracts were large.²¹ This is true of Mexico to some degree and, on an absolute basis, it is particularly true of Turkey over the early part of our sample. However, other than for a brief period in early 2002 for Mexico, the changes in bid/ask spreads for Mexico and Korea were much smaller and their ratios $(ask - bid)/bid$ remained below 10%. Thus, although the gradual increase in the liquidity of the sovereign CDS markets during our sample period no doubt contributed to the downward trend in spreads, changes in liquidity do not appear to have been a major source of variation in the $CRP_t(M)$.

With these observations about the economic events associated with peaks in the CRP 's in mind, we turn next to a more in depth exploration of the relationships between $CRP_t(M)$ and various measures of global risk and macroeconomic and financial developments. One widely watched measure of event risk in credit markets is the CBOE VIX volatility index. We view VIX as a central ingredient in investors' appetite for exposure to the high-yield bond credit class and, as such, expect VIX to correlate over time with changes in the effects of risk premiums on CDS spreads. As our second risk factor we include the spread between the US Industrial 10-year BBB Yield and the 6-month Treasury bill yield (US-Spread),²² and we view correlation between CRP and US-Spread as having two potential sources. First, we have seen from our discussion of Figure 8 that the unwinding of carry trades had large effects on sovereign CDS spreads, especially during 2004 and 2006. The "long long-dated corporate,

²¹Concurrent movements in liquidity and credit quality is often observed in credit markets. As shown by Duffie and Singleton (1999), the pricing formulas we use can be adapted to accommodate liquidity risk by adjusting the discount rate from $r_t + \lambda_t^Q$ to $r_t + \lambda_t^Q + \ell_t$, where ℓ_t is a measure of liquidity costs. Longstaff, Mithal, and Neis (2004) use this extended framework in their analysis of corporate bond and CDS contracts. They assume that $\ell_t = 0$ in their pricing of corporate CDS contracts or, equivalently, that CDS spreads are driven nearly entirely by variation in λ^Q .

²²The yield data were downloaded from Bloomberg Financial Services.

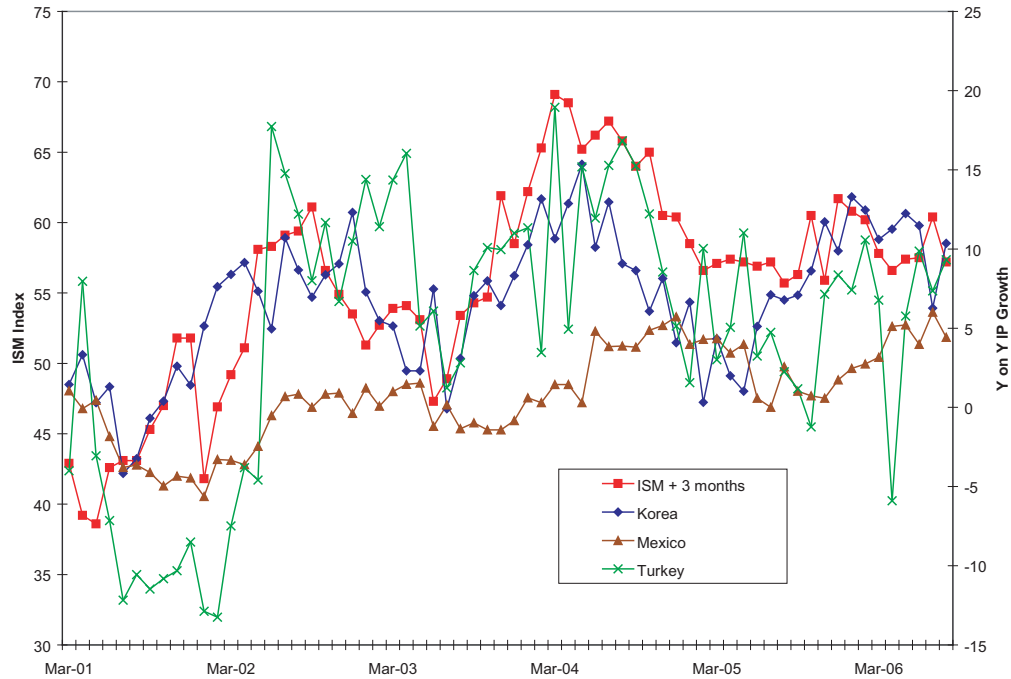


Figure 9: The year-on-year growth rates of industrial production in Korea, Turkey, and Mexico (right scale), and the ISM Index of U.S. Manufacturing led three months (left scale).

short short-dated Treasury” exposure captured by *US-Spread* is a domestic measure of the attractiveness of carry trades. *US-Spread* reflects, in addition, the slope of the U.S. Treasury yield curve which is widely watched as an indicator of the stance of U.S. monetary policy and, thus, of the state of condition of the U.S. economy.

The strengths of the economies in all three of the countries examined depend, to varying degrees and through various economic channels, on the strength of the U.S. economy. This is apparent from Figure 9 which displays the year-on-year growth rates of industrial production (right scale) and the Institute for Supply Management (ISM) index of U.S. manufacturing shifted one quarter ahead (left scale).²³ The ISM index in month t and the annualized emerging market economic growth rates measured at date $t + 3months$; the sample correlations are 0.66, 0.65, and 0.58 for Korea, Turkey, and Mexico, respectively. For Korea, the most persistent gap between these measures of economic growth occurred during 2004 when Korea experienced a marked slowdown in private consumption expenditures in part as a consequence of the consumer debt overhang from 2003 noted above. Turkey shows much more country-specific variation in growth, though one can visually see the secular co-movement with the U.S. economy.

Continuing our analysis of risk premiums, the risk factors *VIX* and *US-Spread* are plotted

²³The data on industrial production was obtained from the International Monetary Fund. The ISM index of manufacturing, based on a monthly survey of purchasing and supply executives throughout the U.S., is constructed by weighting seasonally adjusted new orders, production, employment, supplier deliveries and inventories.

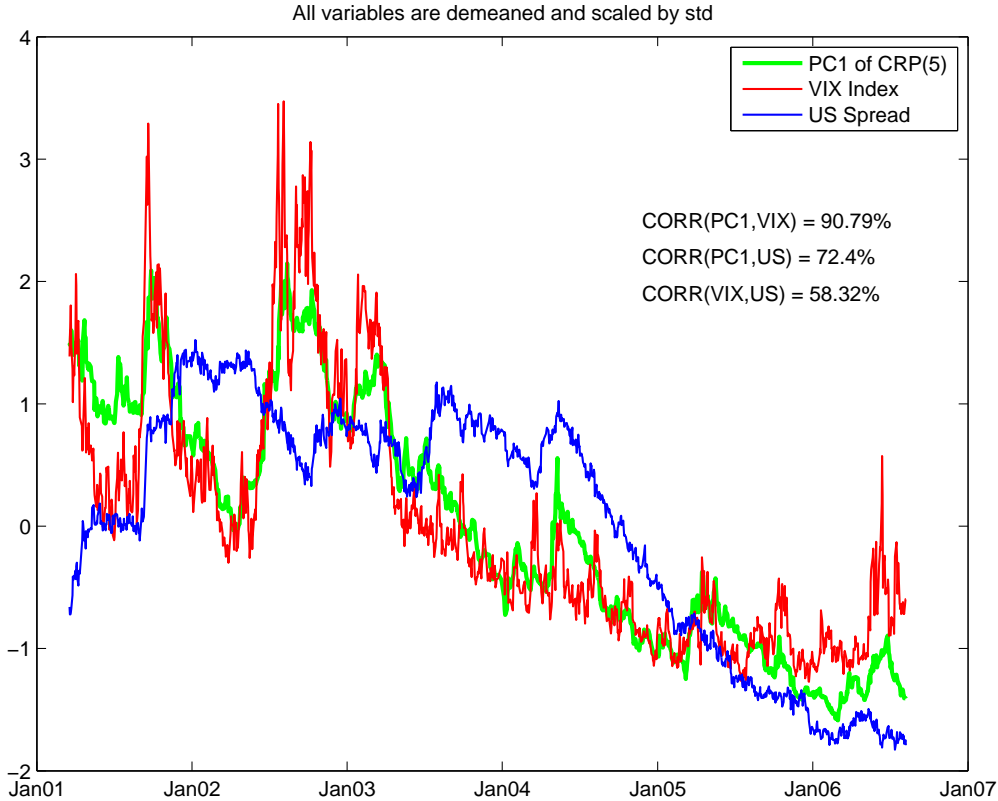


Figure 10: The CBOE VIX index, the spread between the U.S. BBB corporate yield the six-month Treasury bill rate (US-Spread), and the first principal component of the $CRP(5) = (CDS(5) - CDS^p(5))/CDS^p(5)$, for Mexico, Turkey, and Korea. All variables are demeaned and scaled by their respective sample standard deviations.

against the first principal component (CRP-PC1) of the $CRP_t(5)$ for Mexico, Turkey, and Korea in Figure 10. The comovement of CRP-PC1 with VIX is notable; their sample correlation is over 90%. A strong correlation between VIX volatility and U.S. corporate credit spreads has been extensively documented (see, e.g., Collin-Dufresne, Goldstein, and Martin (2001) and Schaefer and Strebulaev (2004)). That VIX, a domestic equity volatility index, is also highly correlated with spreads on sovereign entities as widely dispersed as Mexico and Turkey supports the view that VIX is a key factor in investors' appetite for global "event risk" in credit markets. Turning to US-Spread, the association with CRP-PC1 is relatively weak over the first half of our sample. However, particularly during the run-up in CRP(5)'s in the springs of 2004 and 2006, US-Spread and CRP-PC1 track each other closely. This appears to be a graphical depiction of the effects on risk premiums of the widespread unwinding of carry trades.

More formally, we next examine the relative contributions to the variation in the individual country $CRP(5)$'s of the risk factors VIX, US-Spread, and the own-country implied

currency option volatility (CVOL). The risk factor CVOL is included to assist in capturing the effects of capital flows induced both by external macroeconomic developments and their effects on the flows of goods and capital, and the effects of local political and economic events on the credit qualities of sovereign issuers. Table 8 displays the regression estimates with Newey-West t-Statistics reported in squared brackets, for our entire sample period and the second half of our sample.²⁴ Focusing first on the univariate regressions over the entire sample period, VIX has the most explanatory power for $CRP(5)$ for Mexico and Turkey, while CVOL has slightly more explanatory power than VIX for Korea. In the multivariate regressions both VIX and US-Spread have significant explanatory power for all three countries and, in the case of Korea, CVOL adds explanatory power.

While the reduced-form nature of our regressions introduces some ambiguity into the interpretations these regressions, the evidence is consistent with the view that much of the effect of risk premiums on CDS spreads for Mexico was associated with investors' appetite for exposure to event risk. At least over the early part of our sample, there is a notable comovement between VIX, the CRP 's, and major political/economic events throughout Latin America. The incremental explanatory power from US-Spread for Mexico is, at least partially, owing to investors' reallocations of capital through the unwinding of carry trades. That these "price pressure" effects of capital flows mattered for Mexico is further supported by the results for the second half of our sample during which US-Spread had by far the most explanatory power for $CRP(5)$ (see the lower half of Table 8). The correlation between US-Spread and $CRP(5)$ may also reflect real economic risk associated with the close trading relationship between Mexico and U.S. Notably, after accounting for VIX and US-Spread, the coefficient on CVOL is statistically insignificant, both for the entire and the second half of our sample period. This suggests that the currency option volatilities in Carr and Wu (2006)'s analysis of Mexican CDS spreads may have served as stand-ins for the more fundamental macroeconomic and event risks embodied in VIX and US-Spread.

Turning to the results for Turkey, over the entire sample period for which we only have data on VIX and US-Spread, these two risk factors had comparable explanatory power suggesting that many of the considerations discussed for Mexico were relevant for Turkey as well. Over the second half of our sample, for which we have data on all three risk factors, once again US-Spread is a key explanatory variable for variation in $CRP(5)$. However, unlike for Mexico, CVOL has significant explanatory power and the coefficient on VIX is statistically insignificant. The economic underpinnings of the substantial incremental explanatory power of CVOL for Turkey were the large current account deficits and substantial portfolio inflows during the latter half of our sample. Turkey experienced a consumption-led expansion that was partly financed by these large portfolio inflows. Many foreign investors hedged the currency risk of their local bond positions so it seems natural that the placement and subsequent unwinding of carry trades induced a significant correlation between CVOL and $CRP(5)$.

²⁴Reliable data for Turkey on the implied volatilities of currency options was available only for the second half of our sample and this explained the partial results for Turkey in the upper half of Table 8. Fifty lags were used in computing the Newey-West standard errors.

Table 8: OLS Regressions of the 5-year Risk Premium, measured by $(CDS - CDS^{\text{IP}})/CDS^{\text{IP}}$, on the CBOE VIX Index, the US Industrial 10-year BBB Yield minus the 6-month Treasury, and the 6-month currency implied volatility, VOL, for the respective countries.

Mexico (2001/4/24 – 2006/8/10)				Turkey (2001/3/19 – 2006/8/9)				Korea (2001/3/19 – 2006/8/8)			
VIX	US Spread	CVOL	R-sqr (%)	VIX	US Spread	CVOL	R-sqr (%)	VIX	US Spread	CVOL	R-sqr (%)
22.84			86.04	9.20			68.66	10.49			52.65
[19.06]				[9.54]				[4.73]			
	86.80		55.41		47.54		60.05		53.73		44.44
	[7.92]				[9.40]				[7.96]		
		80.73	61.27							55.28	59.78
		[7.44]								[9.87]	
18.61	32.47		90.85	6.36	27.40		82.05	7.38	29.75		61.64
[12.38]	[6.01]			[8.96]	[6.31]			[2.99]	[3.58]		
19.43	34.22	-5.43	90.93					4.27	22.46	34.85	77.68
[14.47]	[6.79]	[-1.03]						[2.21]	[4.18]	[5.47]	
Mexico (2003/12/12 – 2006/8/10)				Turkey (2003/12/12 – 2006/8/9)				Korea (2003/12/12 – 2006/8/8)			
VIX	US Spread	CVOL	R-sqr (%)	VIX	US Spread	VOL	R-sqr (%)	VIX	US Spread	CVOL	R-sqr (%)
17.95			37.50	14.44			39.73	12.27			29.66
[3.89]				[3.64]				[2.73]			
	47.87		74.17		37.01		72.41		37.40		74.93
	[8.96]				[7.12]				[9.41]		
		30.89	24.90			11.46	68.06			31.86	43.68
		[2.92]				[5.73]				[5.47]	
5.92	41.74		77.04	6.30	31.06		78.09	2.37	34.93		75.72
[3.05]	[8.21]			[4.13]	[7.39]			[1.18]	[8.25]		
7.58	44.31	-8.13	77.94	1.08	23.85	6.42	87.78	2.39	33.75	1.73	75.77
[3.68]	[7.89]	[-1.13]		[0.71]	[6.95]	[7.93]		[1.16]	[5.18]	[0.25]	

The results for Korea show that all three risk factors had significant explanatory power for $CRP(5)$ for the entire sample period. Interestingly the contribution of VIX is muted, certainly relative to the case of Mexico, and, like Turkey, the coefficient on VIX is statistically insignificant for the second half of our sample. We have been told by some investment bankers that investors in Korean bonds have a more “local” focus and that, consequently, spreads are not as highly correlated with VIX as for some other countries. Comparing our results for all three countries, it may simply be that, at least recently, VIX has served more as a measure of risk in the U.S. (or perhaps regionally in the Americas). The explanatory power of US-Spread and CVOL for Korea are surely in part a reflection of the dependence of Asian economies, through exports and imports, on the strength of the U.S. economy (see Figure 9).

7 Concluding Remarks

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