Do Dollar-Denominated Emerging Market Corporate Bonds Insure Foreign Exchange Risk?*

Stefanos Delikouras[†]
Robert F. Dittmar[‡]
Haitao Li[§]

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[†]Department of Finance, University of Miami

[‡]Department of Finance, Stephen Ross School of Business, University of Michigan

[§]Cheung Kong Graduate School of Business

Abstract

We examine the sensitivity of dollar-denominated emerging market corporate bond prices to currency risk. Investors in international markets overwhelmingly demand that emerging market corporate issuers float debt in major currencies; over 85% of emerging market debt is denominated in developed market currencies. Investors cite insurance against foreign exchange risk as the rationale for demanding developed market currency debt. However, in doing so, these investors may overlook the influence of foreign exchange risk on the probability that emerging market corporations will default on their debt due to a currency mismatch between revenues and liability payments. We find in our sample that on average 35% of hazard rate variability can be attributed to changes in exchange rate volatility. We propose a model incorporating currency risk in spreads and find significant impacts on spread sensitivity to foreign exchange risk and material impacts on prices of default risk. Our results suggest that investors in dollar-denominated emerging market bonds are substituting currency risk for default risk.

1 Introduction

The vast majority of emerging market debt is issued in a handful of developed market currencies. As shown in Figure 1, while the prevalence of international debt denominated in emerging market currencies has steadily increased over the past two decades, over 85% of the emerging market debt outstanding (in U.S. dollar terms) is denominated in developed market currencies.[1] Popular wisdom suggests that the prevalence of major currency-denominated emerging market debt is due to investors' desire to hedge currency risk. Indeed, as suggested in this article from Reuters Money, investors may view dollar-denominated emerging market bonds as free of currency risk:

> Those interested in emerging market bonds can choose from a growing roster of mutual funds that mine this space in different ways. Some skirt currency risk by investing exclusively in U.S. dollar-denominated bonds, while others seek to profit from a weakening dollar through bonds denominated in local currencies.[2]

A similar sentiment is echoed in this research memorandum from Morgan Stanley Smith Barney:

> For U.S. based investors, the key difference is foreign currency risk where local currency debt (if unhedged) exposes investors to currency fluctuations.[3]

These quotes suggest that, from the perspective a U.S.-domiciled investor, an emerging market bond denominated in U.S. dollars should be viewed as free of foreign exchange risk.

While the sentiment expressed in these quotations may reflect prevailing investment wisdom, economic arguments suggest that the freedom of dollar-denominated bonds of exchange risk is less clear-cut. Krugman (1999) suggests that the effects of currency crises on firms' balance sheets may result in intensification of the currency crisis through mechanisms modeled in Kiyotaki and Moore (1997) and Bernanke, Gertler, and Gilchrist (1999). He notes that devaluation results in an expansion of the value of the firms' foreign currency liabilities, resulting in deterioration in firm capital. This deterioration is further undermined by higher interest rates and lower revenues resulting from the devaluation. The resulting financial distress constrains firms' ability to invest, exacerbating the crisis. In the context of traditional finance default risk modeling in the vein of Merton (1974), the value of the firms' assets has decreased, the face value of debt has increased

[1] Data are taken from the Bank for International Settlements. Percentages are calculated by summing the dollar amount outstanding of international bonds and notes denominated in emerging market currencies (as designated by the BIS) from the BIS Quarterly Review Table 13B, and dividing by the total dollar amount outstanding issued by emerging markets issuers in BIS Quarterly Review Table 15B.

[2] "Investors warm up to emerging market bonds," *Reuters Money Online*, July 14, 2011

[3] 'Emerging Markets Debt: An Evolving Opportunity Set," by Steve Lee, CFA, Morgan Stanley Smith Barney Consulting Group Investment Advisor Research.

and, as a result, the firm has a lower distance to default, or equivalently, a higher probability of default. Thus, in the absence of perfect hedging, the presence of dollar-denominated bonds on a firm's balance sheet has the potential to generate default risk sensitivity to foreign exchange fluctuations.[4]

This sensitivity of default risk to foreign exchange fluctuations calls into question the prevailing wisdom prescribed in the opening paragraph. From an asset pricing perspective, the principal source of variation in risky bond spreads relative to Treasury yields is default risk, as exemplified in reduced form risky bond pricing models such as Duffie and Singleton (1997, 1999). Thus, if the default risk of companies issuing dollar-denominated bonds is affected by foreign exchange variation, the prices of these dollar-denominated bonds should also be indirectly affected by foreign exchange risk. As a result, dollar-denominated bonds will not be fully insured against fluctuations in exchange rates. This paper quantifies the exposure of dollar-denominated bond prices to foreign exchange rate risks, examines sources of cross-sectional variation in these exposures, and estimates a reduced-form model of defaultable bond prices that incorporates currency risks in dollar-denominated bonds. We show that incorporating this risk results in a reduction of pricing errors in many cases relative to a model that does not incorporate currency exposures.

There is a substantial literature investigating the effect of currency risk on corporate balance sheets in the presence of dollar-denominated debt. This literature focuses largely on the real effects of this risk on the economy and optimal monetary policy response to a currency crisis. In addition to Krugman (1999), discussed above, Aghion, Bacchetta, and Banerjee (2004) derive a model in which the currency composition of debt is endogenous and it can be optimal for firms to borrow in foreign currency, despite the fact that it exposes the country's economy to currency crises. Cèspedes, Chang, and Velcasco (2004) also show that dollarization can lead to currency crises but that while such an equilibrium is possible, it is empirically unlikely. The empirical evidence on the effect of dollarized liabilities on firms' investment behavior and accounting performance is not clear cut. Bleakley and Cowan (2008) conduct a study of firms in five Latin American countries and find that firms with more dollar debt do not invest less than local currency-indebted counterparts in response to devaluations, and that these firms operationally hedge their liabilities. Similarly, Allayanis, Brown, and Klapper (2003) find that foreign currency debt has little effect on East Asian firms' investment after the Asian financial crisis. In contrast, Aguiar (2005) finds that the Mexican peso crisis of 1994 resulted in decreased investment and revenue uncertainty.

In addition to this literature, a body of research asks why corporations issue so much dollar debt. As mentioned above, in Aghion, Bacchetta, and Banerjee (2004), currency composition

[4] A related idea is the increased default risk caused by deflation for nominally-denominated corporate bonds. Fisher (1933) suggests that deflation led to defaults and thus prolonged the Great Depression. In more recent work, Kang and Pflueger (2011) explore the extent to which fears about deflation are reflected in corporate bond prices.

of debt is endogenous. Caballero and Krishnamurthy (2003) show that when domestic agents undervalue insurance against exchange rate depreciation and lenders have limited participation in emerging markets, firms will borrow in dollars, providing a rationale for excessive dollar debt. The authors focus on corporate agents' decisions rather than investors, who are risk neutral in the model. Korinek (2011) shows that when agents rationally make borrowing decisions about foreign currency debt, but ignore the impact of their borrowing on the amplification of currency crises, they will borrow more debt than is optimal. Empirically, Kedia and Mozumadar (2003) investigate why firms issue foreign currency debt. They find that segmented capital markets and hedging motives are principal drivers of debt choice.

We differ from these strands of the literature in a number of ways. Most importantly, while the existing literature focuses on monetary policy responses to a currency crisis or optimal corporate policy in currency composition of capital structure, we focus on the issue of how *investors* price risks of currency fluctuation in dollar-denominated securities. The literature mentioned above largely takes investors' demands as given and assumes a preference for dollar-denominated securities, which makes borrowing in foreign currency cheaper than borrowing domestically. This preference imposes a potential welfare cost on the borrowing firms and their economies. If the benefits to investors of demanding dollar-denominated debt are somewhat illusory, it is possible that welfare might be enhanced by firms issuing more local currency-denominated debt, and allowing investors to find alternative mechanisms through which to hedge foreign currency exposures.

Our focus on investors also allows us to better understand the pricing of dollar-denominated debt. Here we complement the literature pioneered in Duffie and Singleton (1999, 1997), in which default spreads are a function of the level of default-free term structure variables and a firm-specific default intensity. We explicitly incorporate the impact of exchange rate risks in the context of their reduced-form models and find substantial cross-sectional improvements in pricing performance after accounting for these risks. Our model also complements structural models of defaultable dollar-denominated debt such as Chan-Lau and Santos (2006) and Galai and Wiener (2009). In addition to differing from these papers in pursuing a reduced form rather than structural approach, we contribute empirical evidence on the pricing impact of dollarization of firms' liabilities.

The remainder of this paper is organized as follows. In Section 2, we discuss the data and the empirical relation between corporate spreads and foreign exchange risk. In Section 3, we model risky debt with and without foreign exchange risk and explain the methodology for estimating the models. Estimation results are presented in Section 4, and Section 5 presents concluding remarks and directions for further research.

2 Exchange Rate Sensitivity of Dollar-Denominated Bonds

The central question of this paper is whether dollar-denominated bond prices exhibit sensitivity to foreign exchange rate risks. As discussed in the introduction, theoretical arguments suggest that the answer to this question is a qualified yes, since unhedged exchange rate variation may result in increased default risk. To our knowledge this question has not been investigated empirically. This section, therefore, conducts an investigation into the exchange rate sensitivity of the prices of dollar-denominated bonds.

2.1 Data

The starting point for our investigation is all companies with U.S. Dollar bonds in countries defined as emerging markets by MSCI at the beginning of calendar year 2000. We focus on corporate bonds rather than sovereign bonds to avoid "original sin" issues that may dominate the issuance of foreign currency-denominated sovereign bonds (see Eichengreen and Hausmann (1999)). In particular, our assumption is that corporate issuers do not have direct control over monetary policy that may affect the value of the domestic currency.

We eliminate all bonds that are not standard semiannual fixed coupon debentures. Additionally, we remove all obligations of quasi-government agencies, including subsidiaries of sovereign wealth funds, airport and port authorities, and toll roads. We delete obligations of companies in countries with exchange rates pegged to the U.S. dollar and companies that are wholly-owned subsidiaries of developed market firms. Many of the bonds in our sample trade infrequently; as a result, we screen bonds for liquidity. Since we have only price information, we use the liquidity measure proposed in Lesmond, Ogden, and Trzcinka (1999), the fraction of non-zero price change days. In order to balance between liquidity and the number of bonds in the sample, we somewhat arbitrarily choose bonds with at least 75% of days with non-zero price changes. Observations with prices that imply negative yields are also eliminated. Finally, we eliminate bonds with fewer than 250 trading days of data available. This data screening process results in a sample of 68 obligations from 24 companies in six countries; Brazil, Chile, Mexico, the Russian Federation, Singapore, and South Korea.

Descriptive information for these issues is presented in Table 1. There are 11 bonds issued by six companies in Brazil, 11 bonds issued by four companies in Chile, 11 bonds issued by three companies in Mexico, 10 bonds issued by four companies in Russia, 7 bonds issued by two companies in Singapore, and 18 bonds issued by five companies in South Korea. Thus, in terms of number of companies and number of bonds, each of the six countries is relatively well represented, with a slight skew in number of issues toward South Korea and away from Singapore. Median coupon rates are relatively high in Brazil and Russia, and lower in Mexico and South Korea. The maximum coupon in our sample is a 10.50% coupon for a Brazilian issue, and the lowest is 4.25% for a South Korean issue. In all countries except Chile, the minimum initial maturity is five years; in Chile the minimum initial maturity is 9.5 years. Median and maximum initial maturities are also similar across countries except for Russia, where the median and maximum life at issue are substantially shorter, at 6 and 10 years, respectively. The first bond issued in our sample was issued in December, 2000, and our sample extends through September, 2010.

In Figure 2, we depict the time series of yield spreads averaged within each country across bonds in our sample. Spreads are calculated relative to the constant maturity yield on a Treasury security with maturity closest to the maturity of the bond in question, obtained from the FRED database at the Federal Reserve. To facilitate comparison, we plot the averaged spreads on a common time and spread scale. The exception to our spread scaling is the Russian Federation, where average bond yields approach 30% during the global financial crisis, which is approximately twice as large as the next maximum average yield spread observed over our sample period. As shown in the plots, spreads exhibit a pronounced and sustained increase associated with the global financial crisis of 2007-2009. This increase is less pronounced in Chile and Mexico, with spreads increasing to approximately 6% during the height of the crisis in these countries, similar to the spread on Moody's Baa bonds in excess of 10-year Treasury constant maturity bonds, obtained from the FRED database at the Federal Reserve. The spreads of bonds in the remaining countries suggest a greater sensitivity of these bond prices to the global financial crisis than those of speculative grade issues. Brazilian and Singaporean corporate spreads exhibit approximately twice as large of an increase, and Russian spreads four times as large of an increase, as U.S. speculative grade issues.

2.2 Emerging Market Corporate Bond Spreads and Exchange Rate Risk

We speculate that foreign exchange dynamics may affect the magnitude of dollar-denominated corporate bond spreads in two ways. First, as alluded to in the introduction, unhedged level variation in exchange rates may affect default risk and, hence, dollar-denominated corporate bond spreads. Specifically, a depreciation in local currency results in an increase in dollar-denominated debt service from the perspective of a firm with local currency revenues. Moreover, since depreciations tend to occur in states of the world in which local currency revenues are depressed, a depreciation may have an accelerated impact on default risk. The second mechanism is volatility of foreign exchange rates. An increase in exchange rate volatility implies increased volatility in cash flows from a U.S. Dollar perspective. Since the value of a firm's assets depends on the value of its cash flows, increased volatility in Dollar cash flows results in increased volatility of Dollar asset value. In the context of Merton (1974), this increased asset volatility increases the probability of default and, as a consequence, the corporate bond spread.

In order to investigate the impact of these two sources of risk on corporate yield spreads, we conduct a simple regression analysis. Specifically, we estimate the parameters of the following regression,

$$s_{i,k,t} = a_i + b_{fx,i,k}\Delta fx_{k,t} + b_{v,i,k}v_{k,t} + \epsilon_{i,t}, \tag{1}$$

where $s_{i,k,t}$ is spread on bond i in country k at time t, the difference in the yield on bond i and a comparable Treasury, $\Delta fx_{k,t}$ is the change in the log level of exchange rate between the home currency of the issuer of bond and the U.S. Dollar, and Δv_{ikt} is the volatility of the first difference in the log exchange rate between the home currency of the issuer of bond and the U.S. Dollar. The comparable Treasury security yield used in computing the spread on bond i is the constant maturity Treasury yield on a Treasury security with time to maturity closest to that of bond i. Treasury yields for 1-, 2-, 3-, 5-, 7-, 10-, 20-, and 30-year maturities are obtained from the FRED database at the Federal reserve. Our data on exchange rates are taken from Datastream. We sample exchange rates in terms of foreign currency per U.S. Dollar at the daily frequency over the period January 3, 1994 through September 28, 2010. Since currency is expressed in terms of local currency per U.S. dollar, we hypothesize that $b_{fx,i,k} > 0$; that is, when the home currency depreciates relative to the dollar, bond spreads will rise. Similarly, we hypothesize that $b_{v,i,k} > 0$; when foreign exchange innovations are more volatile, default risk, and thus spreads, will increase.

Incorporating volatility of exchange rates into our regression specification, equation (1), requires us to measure volatility in exchange rates. This measurement, in turn, forces us to take a stand on modeling volatility, the subject of a vast literature. Arguably the state of the art for volatility modeling is the use of realized volatility, measured using intraday data. Unfortunately, we do not have access to intraday data, and must rely on the available daily foreign exchange data instead. Andersen and Bollerslev (1998) and Baillie and Bollerslev (1989) model exchange rates using an MA(1)-GARCH(1,1) model. The authors argue that this simple model delivers satisfactory performance in modeling exchange rate volatility. Since the focus of this paper is not the modeling of foreign exchange volatility, we follow these authors' advice and use an MA(1)-EGARCH(1,1) model to capture exchange rate volatility dynamics. We utilize the EGARCH volatility model since it leads to more stable parameter estimates in our data than standard GARCH or the asymmetric volatility model of Glosten, Jagannathan, and Runkle (1993).[5]

Results of the estimation of equation (1) are presented in Table 2. We present medians of parameter estimates, associated t-statistics, and adjusted R^2, where medians are calculated over all bonds and bonds within each of the six countries in our sample. The t-statistics are computed using standard errors calculated via the Newey-West correction for autocorrelation. In addition to the median statistics, we present the fraction of point estimates for which the null hypotheses

[5] For brevity, the results of the exchange rate volatility model estimation are not reported, but are available from the authors upon request.

$b_{fx,i,k} > 0$ and $b_{v,i,k} > 0$ hold at the 2.5% critical level. The main message of the table is captured in the set of results for all countries. The coefficient on level innovation risk, $b_{fx,i,k}$, is negative at the median (median point estimate of -2.04), but statistically insignificant (median t-statistic of -0.46). In contrast, the median coefficient on volatility risk, $b_{v,i,k}$ is positive (median point estimate of 192.60) and statistically significant (median t-statistic of 27.14). None of the point estimates for level innovation risk are statistically significantly greater than zero, whereas 94% of the point estimates for volatility risk are. The median regression explanatory power is 37%, indicating that for the median bond over one third of variation in yield spreads can be accounted for by variation in exchange rate risks. Given the preponderance of significant point estimates for volatility sensitivity, it follows that most of the explanatory power is due to variation in foreign exchange volatility rather than foreign exchange level innovations.

Each country exhibits similar results with varying degrees of sensitivity, statistical significance, and explained variation. In Mexico and Russia, the median point estimate for exposure of spreads to innovations in exchange rates is positive (median point estimates of 0.10 and 19.66, respectively), but in neither country is the median point estimate statistically different than zero. The median point estimate for exposure of spreads to volatility in foreign exchange rates is positive and statistically significant in all six countries; the highest median sensitivity is in Russia (median point estimate of 737.28) and the lowest is Brazil (median point estimate of 115.08). All bonds' spreads in Chile, Russia, and South Korea are positively and statistically significantly exposed to foreign exchange volatility; the lowest proportion of bonds with positive and statistically significant spread exposures to foreign exchange volatility is 82% in Mexico. Finally, there is considerable cross-sectional variation in the degree to which time series variation in spreads can be traced to time series variation in foreign exchange volatility. In Russia, 27% of the variation in the spread on the median bond can be accounted for by foreign exchange volatility, while in Singapore 62% of the median bond's spread variation is traced to foreign exchange volatility.

The results in Table 2 suggest that variation in dollar-denominated bonds' yield spreads are strongly and statistically significantly related to variation in exchange rate volatility. These results support the conjecture that dollar-denominated bonds are exposed to foreign exchange risks through the effect of foreign exchange volatility on asset volatility. However, caution is warranted before concluding that dollar bond yield spreads are exposed to foreign exchange rate volatility due its impact on default risk. In particular, exchange rate volatility is highly correlated with other measures of aggregate economic uncertainty, such as the VIX.[6] As such, while our evidence suggests that emerging market dollar-denominated bond spreads are affected by exchange rate volatility, the effect may be due to greater economic uncertainty rather than increased risk of default per se. We

[6]This fact was brought to our attention separately by Bo Becker and Pab Jotikasithra, whom we thank for their observations.

investigate this question further in the next section.

2.3 Determinants of Dollar-Denominated Yield Spreads

In reduced-form models of defaultable bond prices, the yield spread is driven by a default intensity variable as well as systematic variables determining the risk-free interest rate. The determinants of the default intensity are not modeled, but some empirical insight into their determinants can be gleaned from the results of Collin-Dufresne, Goldstein, and Martin (2001). The authors examine the impact of a broad menu of variables on the time series variation of individual U.S. corporate bond spreads. The authors find that a set of aggregate variables can explain approximately 25% of the variation in innovations in corporate bond spreads.[7]

In this section we conduct a similar analysis for the bonds in our sample. The variables that we examine, like those in Collin-Dufresne, Goldstein, and Martin (2001), are meant to capture common variation in sources of risk that determine credit spreads. We ask whether foreign exchange level innovations and volatility contribute marginally to determining credit spreads, controlling for these other variables. Our goal in doing so is to assess whether foreign exchange risks impact dollar-denominated bond prices beyond the effect of economic uncertainty captured in these other variables. That is, we ask whether there is something "special" about foreign exchange risks in explaining variation in yield spreads. If so, there is a stronger possibility that foreign exchange risks affect yield spreads not only due to their relation to aggregate economic uncertainty, but also due to their effect on default risk.

Collin-Dufresne, Goldstein, and Martin (2001) utilize a set of covariates in their analysis that are motivated by structural models of default risk in the spirit of Merton (1974). We use a similar set of variables in our analysis:

1. *The level of interest rates.* Longstaff and Schwartz (1995) note that a higher spot rate increases the risk neutral drift of the stochastic process for firm value. Because a higher spot rate results in a higher drift, it implies a higher risk-neutral conditional mean of the expected change in assets, reducing the probability of default. Following Collin-Dufresne, Goldstein, and Martin (2001), we measure this quantity using the 10-year constant maturity Treasury yield, $y_{10,t}$ from the FRED database at the Federal Reserve.

2. *The slope of the yield curve.* Litterman and Scheinkmann (1991) document the presence of three dominant latent factors in the term structure, of which the most important correspond

[7]The authors' focus is on not only the 25% of variation explained by these variables, but also the 75% of variation *not* explained by these variables. They find a common factor in the regression residuals that cannot be explained by macroeconomic or common liquidity variables. The authors conclude that most of monthly credit spread changes are due to local supply or demand shocks.

to a level and a slope factor. From the perspective of the expectations hypothesis of the term structure of interest rates, a widening slope implies an increase in expected future interest rates which, following the logic for the level above, implies a lower credit spread. Again, following Collin-Dufresne, Goldstein, and Martin (2001), we measure the slope, ts_t, as the difference in the yield on 10-year and 2-year constant maturity Treasuries from FRED.

3. *Volatility.* The central variable in an options-based approach to modeling credit risk is the variability of the firm's assets. In principle, this volatility would be measured using volatility implied by options on the firm's equity; unfortunately, these data are not available for the firms in our sample. Following Collin-Dufresne, Goldstein, and Martin (2001), we measure volatility using the VIX, vx_t, obtained from Datastream. An additional advantage to the use of the VIX is that it may capture aggregate economic uncertainty, controlling for this component of exchange rate volatility.

4. *Returns on the market.* Another measure of overall market and economic conditions is the return on the aggregate stock market. We measure this return using the CRSP value-weighted market index, $r_{m,t}$.

These variables are a subset of the variables used in Collin-Dufresne, Goldstein, and Martin (2001). The authors use three additional variables in their analysis; market leverage, the square of the (change in) the level of interest rates, and the magnitude and probability of jumps in firm value. We do not incorporate leverage because for our firms we have a relatively short time sample and our observations are sampled at the daily level. Therefore, the vast majority of variation in market leverage would be due not necessarily to leverage, but rather to movements in the firm's equity return. The squared change in level of interest rates is rarely and inconsistently statistically significant in Collin-Dufresne, et. al's analysis.

In addition to the variables used in Collin-Dufresne, Goldstein, and Martin (2001), we incorporate the return on the local stock market, $r_{l,t}$, to reflect local market conditions. These data are also obtained from Datastream and are the Datastream Brazil stock market index, the IGPA from Chile, the INMEX from Mexico, the MICEX from Russia, the MSCI Singapore index from Singapore, and the KOSPI from South Korea. We also considered two additional variables in our analysis; the level of local interest rates and a measure of local market volatility. We omit the local interest rate because under uncovered interest rate parity, log innovations in exchange rates and log interest rate differentials are collinear. Inclusion of local market implied volatility is desirable to capture additional sources of local market economic uncertainty. Unfortunately, local market implied volatility measures are not available for the countries in our sample. We experimented with including parametric measures of stock market volatility, but found that these measures contributed little to explaining spreads and therefore excluded them from the analysis.

Given this set of variables, we re-examine an augmented version of regression (1),

$$s_{i,t} = a_i + b_{v,i,k} v_{k,t} + b_{y,i} y_{10,t} + b_{ts,i} ts_t + b_{vx,i} vx_t + b_{r_m,i} r_{m,t} + b_{r_l,i} r_{l,t} + \xi_{i,t}, \tag{2}$$

where we include the covariates discussed above. As in the previous section, we report the median point estimate of each parameter, the median t-statistic, and the median R^2. We also report the proportion of the parameters $b_{v,i,k}$ that are positive and statistically different than zero. Results are displayed in Table 3, and are compiled for the set of all firms in all countries, as well as within each country.

Table 3 shows that, across all bonds, foreign exchange volatility retains power for explaining time series variation in default spreads. At the median, the point estimate is positive (median point estimate of 13.39) and statistically significant (median t-statistic of 1.77). However, some of the variation explained by volatility is absorbed by the other covariates. The table indicates that 53% of the point estimates for volatility are statistically significantly greater than zero at the 5.0% critical level, compared to 94% when the covariates are not included. These results suggest that at least some of the explanatory power of foreign exchange volatility is due to variation in economic conditions and uncertainty present in variables not directly linked to exchange rate variation. Also noteworthy is the fact that, like the earlier results, exchange rate level innovations are negatively, but not statistically significantly related to spreads at the median.

Across countries, median results and results for the proportion of bonds affected by foreign exchange level innovations and volatility vary dramatically. For instance, in Brazil, Chile, and Mexico, the median bond has negative exposure to foreign exchange volatility. In Chile and Mexico the median t-stat for foreign exchange innovations suggests significant negative foreign exchange exposure, while in Brazil and Mexico the median t-stat for volatility indicates significant negative exposure. In Brazil and Chile, however, a significant fraction (46% and 36%, respectively) of the bonds have positive and statistically significant exposure to foreign exchange rate volatility, controlling for covariates. In the three remaining countries, exposures to foreign exchange volatility are positive at the median, and the median t-statistic in Russia and Singapore indicates statistical significance. Further, 100% of the bonds in Russia and Singapore, and 44% of the bonds in South Korea exhibit positive and significant exposures of spreads to foreign exchange volatility.

The results in this section suggest that while there is some evidence of sensitivity of spreads to foreign exchange volatility at the median, the median number hides substantial variation in the cross section in volatility exposures. In some countries, such as Singapore and Russia, all bonds are exposed to foreign exchange volatility, whereas in Mexico, very few bonds are. It is likely that cross-sectional variation in exposure is driven by operational considerations. Some firms may have natural hedges against foreign exchange risk due to foreign revenues, or may hedge their exposure using currency swaps or forwards. In the next section, we analyze the relation between volatility exposure of spreads and these operational features of the businesses of the companies in our sample.

2.4 Sensitivity of Spreads to Exchange Rate Risks in the Cross-Section

In a recent paper, Bartram, Brown, and Minton (2010) investigate the sensitivity of firms' equity returns to exchange rates. The authors note that although many theoretical and empirical papers posit and document an exposure of firms' fundamentals to exchange rates, empirical studies have found little exposure of equity returns to exchange rates.[8] The authors derive a model of foreign exchange exposure and estimate exchange rate exposures based on measures that are suggested by the model. In this section, we analyze whether the empirical determinants of exposure used in their study help us explain cross-sectional variation in the exposure of dollar-denominated emerging market bond spreads to foreign exchange level innovations and volatility.

Bartram, Brown, and Minton (2010) utilize a number of empirical proxies for the inputs to their theoretical model. The variables that they utilize are meant to capture information about the degree of revenue derived from foreign sources, marginal costs in foreign currency, competitors' marginal costs in terms of foreign currency, and firms' market shares in foreign and domestic markets. These variables are used to estimate currency exposures. Since we are instead interested in the determinants of these exposures which have been pre-estimated in the previous section, we use a reduced set of variables guided by their analysis:

1. *Percent of sales in U.S. Dollars.* This variable, $sales_j$, is the fraction of total revenues denominated in U.S. dollars. When U.S. Dollar sales were not explicitly stated, we assumed that North American sales were U.S. dollar sales. For companies producing commodities that are sold in U.S. dollars, we assumed that 100% of sales were in U.S. dollars. Finally, if neither U.S. nor North American sales numbers were available, we utilized all non-domestic sales. We expect that, all other considerations constant, firms with more U.S. dollar sales will be less vulnerable to foreign exchange risks, as these firms' dollar revenues will offset risks induced by lower cash flows due to exchange rate fluctuations.

2. *Percent of U.S. dollar bond debt.* The percent of U.S. dollar debt, $debt_j$ reflects the importance of these bonds in the overall debt structure of the firm. The variable is calculated as the reported local currency value of U.S. dollar corporate debentures to total local currency debt.

[8]The authors cite Hung (1992) and Williamson (2001) as empirical papers documenting the sensitivity of firm fundamentals to exchange rates and Bodnar, Dumas, and Marston (2002) as theoretical papers that posit that fundamentals will be exposed to exchange rate movements. Empirical studies documenting weak relations between equity returns and exchange rates include Jorion (1990), Amihud (1993), Bodnar and Gentry (1993), Bartov and Bodnar (1994), Bartov, Bodnar, and Kaul (1996), Choi and Prasad (1995), He and Ng (1998), Chow, Lee, and Solt (1997), Griffin and Stulz (2001), and Dominguez and Tesar (2006).

If the firm's balance sheet is reported in U.S. dollars, the U.S. dollar corporate debentures are the face value of the U.S. dollar debt. We expect that for firms for which the U.S. dollar bonds are a more important fraction of their overall capital structure, that sensitivities to foreign exchange risk will be higher.

3. *Foreign currency derivative usage.* The variable $deriv_j$ takes the value 1 if the company reports the use of forwards or swaps to manage foreign currency risk exposure. Since this management will reduce the exposure of cash flows to currency risks, we expect a negative coefficient.

We hand-collect these data for the firms in our sample from three different sources. If available, we utilize 20-F filings on the EDGAR database at the SEC, and search the notes to the financial statements for the information needed to construct our variables. If we are unable to find the filings on EDGAR, we collect financial statements from the company's investor relations website. Finally, for two of the stocks in our sample, AK Transneft and Evraz Group, we were unable to locate financial statements on the company websites. For these firms we utilized financial statements found on the website `http://www.rustocks.com`. We collect data for the fiscal years 2006, 2007, 2008, and 2009, which represent the bulk of the time series and cross-section observations in our sample. In three cases, Telemar Norte, JBS, and Enersis, bonds were not issued until later in our sample. In these cases, we sample only data for fiscal years in which the bonds were outstanding.

Given the financial statement information, we estimate the following pooled regressions:

$$\hat{b}_{v,i,k} = d_{01} + d_{11} sales_{j,t} + d_{21} debt_{j,t} + d_{31} deriv_{j,t} + d_{41} coup_i + d_{51} mat_i + u_{1i} \quad (3)$$

where i indexes bonds, j indexes firms, and $t = 2006, 2007, 2008, 2009$ indexes fiscal years. In untabulated results, we estimate the regressions year-by-year and obtain similar results. The variables $coup_i$ and mat_i are the coupon rate and initial maturity of the bonds, respectively. These variables are included to control for bond-specific sources of sensitivity, as the variables are related to bond duration.

Results of the regressions are presented in Table 4. As shown in the table, sensitivities respond in the cross-section to variables that we expect to matter for exchange rate sensitivity in a manner consistent with our priors. Sensitivities are negatively related to the hedging dummy, indicating that firms that hedge interest rate risk are less sensitive to foreign exchange risks, but the estimate is not statistically different than zero. This result suggests that firms' hedging activities may reduce delta-sensitive, but not necessarily gamma-sensitive foreign exchange risks. However, this result should be interpreted with some caution, as nearly all of the firms in our sample use financial instruments to hedge foreign exchange risks. Sensitivities are negatively and significantly related to the proportion U.S. sales, suggesting that U.S. sales provide a natural hedge against the default risks assumed by issuing U.S. dollar debt. Finally, the proportion of debt contained in U.S. dollar issues is positively and significantly associated with both level and volatility exposures, suggesting that greater importance of the debt in the capital structure leads to increased foreign exchange exposure.

The evidence provided in the preceding three sections suggests the following conclusions. First, spreads on dollar-denominated emerging market debt are not immune from foreign exchange risks. The spreads respond to foreign exchange volatility, and thus investors are exposed to risks that are contained in currency dynamics. Second, a significant component of this volatility exposure remains after controlling for aggregate risks that might drive default spreads. In particular, there remains considerable significant cross-sectional variation in these sensitivities after allowing for common determinants of the spread. Third, the cross-sectional variation in sensitivities is significantly affected by cross-sectional variation in operational and financial hedges against exchange rates. Thus, investors in dollar-denominated bonds issued by firms with fewer financial and operating hedges are more exposed to foreign exchange risks than those with more financial and operating hedges. The overarching conclusion that can be drawn from these three results is that dollar-denominated emerging market bonds have significant exposures to foreign exchange risk, which vary substantially in the cross-section.

3 Pricing Dollar-Denominated Corporate Bonds

Given the evidence in the preceding section supporting the hypothesis that emerging market corporate bonds are sensitive to foreign exchange risks, we proceed to examine the impact of foreign exchange risk exposure on the pricing of these bonds. The lens through which we approach this question is a reduced-form risky term structure model in the spirit of Duffie and Singleton (1997, 1999). We first examine the pricing of bonds when we ignore the contribution of exchange rate risk, and then augment the model with exchange rate sensitivity and analyze the impact of this sensitivity on pricing.

3.1 Dollar-Denominated Bonds Without Foreign Exchange Risk

We first analyze the pricing of U.S. Dollar-denominated bonds assuming that they represent claims on corporate cash flows with default risk. That is, we model the price of a dollar-denominated bond as if it were a U.S. dollar corporate debenture and assume that pricing is not sensitive to the home currency of the issuer. Consequently, we can utilize well-developed tools for the pricing of the security. In particular, we rely on the reduced-form modeling approach of Duffie and Singleton

(1999), in which we assume that the price of a zero-coupon bond with default risk is given by

$$P_i(t,T) = E_t^Q\left[e^{-\int_t^T R_{i,s}ds}\right], \tag{4}$$

with R_s representing the instantaneous default-adjusted discount rate,

$$R_{i,t} = r_t + (1-\delta_i)\,\lambda_{i,t} \tag{5}$$

where r_t is the instantaneous risk free rate, i indexes bonds, δ is the rate of recovery on the debt, and $\lambda_{i,t}\,(1-\delta_i)$ is the spread in excess of the risk-free rate.

We specify the risk free term structure following Duffee (1999) as a two-factor term structure model in the affine class of models derived by Duffie and Kan (1996). We assume that the risk free rate can be expressed as an affine function of two state variables,

$$r_t = a_f + x_{1,t} + x_{2,t}, \tag{6}$$

where the state variables $x_{1,t}$ and $x_{2,t}$ follow square root dynamic processes under the risk-neutral probability measure Q as in Cox, Ingersoll, and Ross (1985),

$$dx_{1,t} = (\kappa_1\theta_1 - (\kappa_1+\eta_1)\,x_{1,t})\,dt + \sigma_1\sqrt{x_{1,t}}dw_{1,t}^Q \tag{7}$$

$$dx_{2,t} = (\kappa_2\theta_2 - (\kappa_2+\eta_2)\,x_{2,t})\,dt + \sigma_2\sqrt{x_{2,t}}dw_{2,t}^Q. \tag{8}$$

The parameters η_1 and η_2 represent prices of risk and $dw_{1,t}^Q$ and $dw_{2,t}^Q$ are independent Brownian motions under the risk neutral probability measure Q.

The credit spread, $(1-\delta_i)\lambda_{i,t}$, is modeled using the special case of Duffie and Singleton (1999) employed in Duffee (1999). The spread is assumed to be a function of the risk-free term structure state variables and a default risk variable,

$$(1-\delta_i)\,\lambda_{i,t} = a_i + h_{i,t} + \boldsymbol{\beta}_i'\,(\mathbf{x}_t - \bar{\mathbf{x}})\,. \tag{9}$$

The parameter vector $\boldsymbol{\beta}_i$ allows for correlation between the default-free term structure and the spread on the bond above the risk free rate; as referenced above, Longstaff and Schwartz (1995) argue that structural models in the line of Merton (1974) result in a negative relation between the credit spread and the risk-free rate. The default risk factor, $h_{i,t}$, is referred to as the hazard rate and follows a stochastic process under the risk-neutral probability measure Q defined as

$$dh_{i,t} = (\kappa_i\theta_i - (\kappa_i+\eta_i)\,h_{i,t})\,dt + \sigma_i\sqrt{h_{i,t}}dW_{i,t}^Q. \tag{10}$$

We assume that the Brownian motion driving the evolution of the hazard rate is independent of the Brownian motions governing the riskless rate.[9] Duffie and Singleton (1999) note that one can view the hazard rate as the arrival intensity of a jump that first occurs as default. Thus, although default is a discrete event, the intensity follows a diffusion.

Given the dynamics of the risk free term structure and hazard rates, log zero-coupon bond prices are affine in the state variables and hazard rates,

$$\ln P_{i,t}(\tau) = A_i(\tau) + \mathbf{B}_i'(\tau)\mathbf{x}_{i,t}^* + B_i(\tau)h_{i,t},$$

where $x_{i,k,t}^* = (1+\beta_{i,k})x_{k,t}$ for $k = 1, 2$, and the coefficients $A_i(\tau)$, $\mathbf{B}(\tau)$, and $B_i(\tau)$ are solutions to ordinary differential equations as in Duffie and Singleton (1999) and Cox, Ingersoll, and Ross (1985). The precise form of the coefficients are provided in the Appendix. These solutions are for zero-coupon bond prices, whereas the bonds in our sample are coupon bonds. We treat these coupon bonds as a portfolio of zero coupon bonds with face value c plus a zero coupon bond with face value of 1. Mathematically, the price of the coupon bond with maturity T is given by

$$P_{i,t}(\tau, c) = E_t^Q\left[c\sum_{m=1}^{T-t} e^{-\int_t^{t+m} R_{i,s}ds} + e^{-\int_t^T R_{i,s}ds}\right], \tag{11}$$

where m indexes the periodic coupon payments.

3.2 Incorporating Foreign Exchange Risk

In order to augment the basic reduced form model to incorporate sensitivity to exchange rate risk, we assume that exchange rate risk derives not from the level of exchange rates, which we assume to be tied to differences in risk-free rates across countries, but to its volatility. Specifically, we assume that exchange rate volatility follows dynamics under the risk-neutral measure Q,

$$dv_t = \left[\kappa_v\theta_v - (\kappa_v + \eta_v)\, v_t\right] dt + \sigma_v\sqrt{v_t}dW_{v,t}^Q, \tag{12}$$

where v_t is the foreign exchange volatility. The foreign exchange volatility represents an additional state variable augmenting the original two-dimensional state variable $\mathbf{x}_t$. We use the volatility series estimated using the EGARCH(1,1) model above for each exchange rate volatility as observations

[9] An alternative approach is to use a three-factor model in which the correlation among the state variables is explicit. Dai and Singleton (2000) provide conditions for which affine term structure models are identified. The principal cost of doing so, as the authors note, is that the correlation structure and the stochastic volatility in the hazard rate process are constrained. In order to allow negative correlation between the hazard rate process and the risk-free term structure, one would have to model the hazard process as a Gaussian state variable. This would allow the spread to potentially take on negative values, which is undesirable in the context of a positive premium for default risk.

in estimating the parameters of exchange rate volatility dynamics.

The presence of priced exchange rate risk leads to an alternative specification of the default-adjusted discount rate, accounting now for exchange rate risk. The yield spread becomes

$$R_{i,t} - r_t = a_i + h_{i,t} + \boldsymbol{\beta}_i' \left(\mathbf{x}_t - \bar{\mathbf{x}}\right) + \beta_{i,v} v_t. \tag{13}$$

The credit spread depends on foreign exchange rate volatility in a manner similar to that of the risk-free term structure state variables. That is, $\beta_{i,v}$ allows correlation in the credit spread and the volatility of exchange rates. However, the hazard rate, $h_{i,t}$ is assumed independent of this volatility, similar to the independence of the hazard rate and the risk-free term structure variables. Thus, in this context, hazard rates can be interpreted as the default risk independent of default risk induced by risk-free term structure or foreign exchange volatility.

We define $v_{i,t}^* = \beta_{i,v} v_t$, resulting in an additional term in the log risky bond price,

$$\ln P_{i,t}\left(\tau\right) = A_i\left(\tau\right) + \mathbf{B}'(\tau)\mathbf{x}_{i,t}^* + B_i\left(\tau\right) h_{i,t} + B_v\left(\tau\right) v_{i,t}^*, \tag{14}$$

where expressions for $B_v(\tau)$ and $A_i(\tau)$ are again provided in the appendix. Coupon bond prices are constructed as portfolios of zero coupon bonds as in equation (11). The only modification is that the risky discount rate now incorporates terms representing sensitivity to foreign exchange volatility.

3.3 Estimation Procedure

The state variables of the default-free term structure, x_1 and x_2, as well as the hazard rate h_i, are unobservable. We estimate model parameters and identify the variables using the extended Kalman filter. Our Kalman filtering process first estimates parameters of the risk-free term structure using the measurement equation

$$\mathbf{Y}_t\left(\boldsymbol{\tau}\right) = a_f \boldsymbol{\iota} - \frac{1}{\tau}\left(\mathbf{A}\left(\boldsymbol{\tau}\right) + \mathbf{B}'\left(\boldsymbol{\tau}\right)\mathbf{x}_t\right) + \mathbf{u}_t \tag{15}$$

where $\mathbf{Y}_t\left(\tau\right)$ is a vector of risk-free zero coupon bond yields observed at time t with maturities $\boldsymbol{\tau}$, $\mathbf{A}\left(\boldsymbol{\tau}\right)$ is a vector of coefficients as in equation (8), and $\mathbf{B}\left(\boldsymbol{\tau}\right)$ is a matrix of coefficients as in equation (7). The vector of pricing errors $\mathbf{u}_t$ is assumed to by i.i.d. $\mathcal{N}\left(\mathbf{0}, \boldsymbol{\Sigma}_u\right)$, where $\boldsymbol{\Sigma}_u$ is a diagonal covariance matrix.

Transition equations for the state variables are given by:

$$\begin{pmatrix} x_{1,t} \\ x_{2,t} \end{pmatrix} = \begin{pmatrix} \theta_1 \left(1 - e^{-\kappa_1}\right) \\ \theta_2 \left(1 - e^{-\kappa_2}\right) \end{pmatrix} + \begin{pmatrix} e^{-\kappa_1} & 0 \\ 0 & e^{-\kappa_2} \end{pmatrix} \begin{pmatrix} x_{1,t-1} \\ x_{2,t-1} \end{pmatrix} + \begin{pmatrix} w_{1,t} \\ w_{2,t} \end{pmatrix}, \tag{16}$$

where

$$\mathbf{w}_t \sim \mathcal{N}\left(\mathbf{0}, \begin{pmatrix} Q_{1,t} & 0 \\ 0 & Q_{2,t} \end{pmatrix}\right) \tag{17}$$

$$Q_{k,t} = x_{k,t}\frac{\sigma_k^2}{\kappa_k}\left(e^{-\kappa_k} - e^{-2\kappa_k}\right) + \theta_k \frac{\sigma_k^2}{2\kappa_k}\left(1 - e^{-\kappa_k}\right)^2, \quad k = 1, 2. \tag{18}$$

These transition dynamics represent the conditional means and volatilities of the state variables of square root processes as shown in Cox, Ingersoll, and Ross (1985), where the innovation terms are assumed Gaussian. We use the measurement and transition errors to find parameter estimates and filter state variables by maximizing the log likelihood function of the measurement errors.

Given the estimates of the risk-free term structure parameters and the state variables, we estimate the parameters of the risky term structure and filter hazard rates. Our measurement equation is a discretized version of the risky coupon bond price equation (11), measured with error:

$$P_{i,t}(\tau, c) = c \sum_{m=1}^{\tau} P_i(m) + P_i(\tau) + u_{i,t}, \tag{19}$$

Since we take the latent risk-free variables as given from the estimation of the risk-free term structure, our transition equation applies to the hazard rate:

$$h_{i,t} = \frac{\theta_i \kappa_i}{\kappa_i + \eta_i}\left(1 - e^{-\kappa_i}\right) + e^{-\kappa_i} h_{i,t-1} + w_{i,t}, \tag{20}$$

where

$$w_{i,t} \sim \mathcal{N}\left(0, Q_{i,t}\right), \tag{21}$$

$$Q_{i,t} = h_{i,t-1}\frac{\sigma_i^2}{\kappa_i}\left(e^{-\kappa_i} - e^{-2\kappa_i}\right) + \theta_i \frac{\sigma_i^2}{2\kappa_i}\left(1 - e^{-\kappa_i}\right)^2. \tag{22}$$

As with the risk-free estimation, we estimate parameters and filter hazard rates by maximizing the log likelihood function of the measurement errors for each bond in our sample.

When we incorporate foreign exchange volatility into the estimation, very little changes in the procedure. We add another transition equation for the exchange rate volatility, where we assume that the measured EGARCH(1,1) volatility is the true latent volatility, measured with error. The only additional difference in estimation is that the measurement equation (19) calculates the price of zero coupon bonds using the augmented risky yield process, equation (13).

The standard errors of parameter estimates are constructed according to the quasi-maximum likelihood error approach. The approach uses both the Hessian of the log likelihood function and the outer product estimate for the information matrix. The conditional normality assumption for the log likelihood function is an approximation to the true data generating process which, under the assumption of a square-root process for the state variables, is a non-central χ^2 distribution. In tabulating our results, we do not report the standard errors for the point estimates of the hazard rate process; instead, we report quantiles of the estimates.

Our estimation approach mirrors Duffee (1999). As in his investigation, we estimate parameters of the risk free term structure separately from estimation for individual bonds. Doing so ensures that that common risk free term structure factors and parameters are common to all bonds. In principle, it would be desirable to jointly estimate the parameters of the risky and risk free term structures. However, the technical complications of a joint estimation over a large cross-section of assets renders joint estimation infeasible.

4 Model Estimation Results

In this section, we present and discuss the results of estimating models for risky bond prices with and without foreign exchange risk. Our estimation procedure also provides estimates of the default-free term structure. However, since this estimation is not the focus of our paper, we do not tabulate these results. Our estimates are similar to those presented in Duffee (1999) despite little overlap in our time series. Further, our first state variable is 99% correlated with the 10-year constant maturity Treasury yield and the second is 92% correlated with the negative of the slope of the term structure, measured as the difference in the ten year and three month constant maturity Treasury yield. Complete results of the estimation of the default-free term structure are available from the authors upon request.

4.1 Pricing Without Foreign Exchange Risk

Given the estimates of the parameters of the default free term structure, we next turn to the estimation of parameters of risky bond prices using the reduced form model in Section 2.2. For each of the 68 bonds in our sample we estimate the mean reversion coefficient, κ_i, the long-run mean, θ_i, the price of risk, η_i, and the diffusion parameter, σ_i. We also estimate the parameters of default-free term structure sensitivity, $\beta_{i,1}$, $\beta_{i,2}$, and α_i. In Table 5, we present 25th percentile, median, and 75th percentiles of the parameter estimates for the cross-section of firms in Panel A, and for individual countries in Panels B through H. Median point estimates of κ_i (3.68 across all bonds) and θ_i (0.08 across all bonds) suggest that default intensities are strongly mean-reverting and on average have relatively high long-run means. The point estimates suggest that long run means in hazard rates are an order of magnitude higher than the estimates for domestic bonds in Duffee (1999). As shown by the 25th and 75th percentiles, the point estimates exhibit considerable variation across bonds in both mean reversion and long-run means.

The median price of default risk across all bonds, $\eta_i = -0.34$, is negative and suggests that investors demand compensation for default risk. The magnitude of this median parameter is larger than, but similar to that that estimated by Duffee (1999), who finds the median price of default risk in his sample of U.S. firms is -0.24. Like Duffee (1999), the median sensitivity of the default intensity to the default-free term structure is negative, with median estimates of $\beta_{1,d}$ and $\beta_{2,d}$ of -0.31 and -0.26, respectively. These estimates indicate a somewhat stronger reaction of default intensities to the level and slope of the term structure in emerging markets, such that an increase in the overall level and slope of yields translates into reduced default intensity. As mentioned above, this result may obtain from the effect of the risk-free term structure on the drift of firm asset value. The interquartile ranges of estimates suggest that sensitivity to the risk free level and the price of risk are more tightly clustered than sensitivity to the slope of the yield curve.

Across countries, there are a few notable differences in median parameter estimates. The median mean reversion coefficient is particularly high in Mexico ($\kappa_i = 13.72$). Interquartile ranges suggest considerable variation within each country in the estimation of mean reversion of hazard rates, but estimates are reliably positive. Median long term means of hazard rates range from $\theta_i = 0.06$ (Mexico) to $\theta_i = 0.12$ (Brazil); Singapore and South Korea exhibit the lowest 25th percentile ($\theta_i = 0.04$), and Singapore also exhibits the highest 75th percentile ($\theta_i = 0.16$).

Compensation for default risk appears to vary widely across the countries in the sample. While median prices of risk, η_i, are negative in each country, the 75th percentile estimate is positive in Chile and Singapore. Positive values for η_i are puzzling, as they suggest negative compensation for default risk. Median compensations are especially large in magnitude in Brazil ($\eta_i = -0.55$) and Russia ($\eta_i = -0.58$). Variation is quite large in most countries as well; the interquartile ranges in the six countries are 0.97, 0.69, 0.56, 0.47, 0.41, and 0.30 for Singapore, South Korea, Chile, Brazil, Mexico, and Russia, respectively. These ranges suggest that even on a country-by-country basis, the price of default risk is difficult to pin down.

In Panel H, we present pricing errors for the overall sample and by country. For each set of bonds, we report interquartile ranges and medians (25th, 50th, and 75th percentiles) of the root mean squared error (RMSE) of bond yields. Median estimates of root mean square errors are larger than those in studies of U.S. bonds, such as Duffee (1999). The median RMSE is 18.34 basis points, with a 25th percentile of 12.28 basis points and a 75th percentile of 35.09 basis points. In contrast, Duffee (1999) reports a median estimate of approximately 10 basis points, a 25th percentile of 7 basis points, and a 75th percentile of 11 basis points. Thus, in our estimates, pricing errors are both larger at the median and exhibit greater variation across bonds. The table also shows that pricing difficulties are particularly severe in the Russian Federation, compared to the remaining countries. The median pricing error in Russia is 51.76 basis points, with an interquartile range of 26.30 to 66.53 basis points. The model also has difficulty in pricing Brazilian bonds with a median RMSE of 27.74 basis points and an interquartile range of 13.29 to 38.42 basis points. In contrast, the remaining countries are better represented by the overall estimates.

In order to try to gauge some explanation for the magnitude of pricing errors across these bonds, we examine the relation between pricing errors and both firm- and bond-specific determinants of exchange rate sensitivity of spreads, as in Section 2.4. We regress root mean square errors for the bonds on the maturity and coupon of the bonds, to control for duration-related effects, as well as a hedging dummy, the percentage of sales from the U.S. and the percentage of U.S. dollar debt. Results are tabulated in Table 6. The results suggest that the model prices shorter-term and higher coupon bonds more poorly, as indicated by the negative coefficient on maturity and the positive coefficient on the coupon rate. These results suggest that duration-related factors have an indeterminate effect on pricing error; shorter maturity (low duration) and lower coupon (high duration) bonds are both more poorly priced.

The firm-specific variables also provide somewhat conflicting results. Firms that hedge have lower pricing errors, as indicated by the negative coefficient on the indicator variable for hedging. This result is consistent with the notion that mispricing due to currency-related default risk is less important in firms that hedge currency risk. However, mispricing is higher in firms with a higher percentage of U.S. sales. Since U.S. sales represent a natural hedge for these firms, this evidence is harder to reconcile with the idea that mispricing is related to omitted currency-related default risk. One possible reason for this result is that, as mentioned above, firms with U.S. sales but local currency-denominated costs continue to have a currency mismatch in terms of their natural hedges.

4.2 Pricing With Foreign Exchange Risk

We present estimates of the model with foreign exchange risk, equation (14), in Table 7. As in Table 5, we present 25th percentile, median, and 75th percentile estimates of parameters across all countries and within each country in our sample. Results for all countries are presented in Panel A. The table shows that the new parameter to the model, β_3, is positive at the median, ranging from -0.02 at the 25th percentile to 0.59 at the 75th percentile. The positive median point estimate suggests that the spread of the median firm has positive exposure to exchange rate volatility; an increase in the volatility of the domestic exchange rate results in an increase in spreads. At the 25th percentile, however, the sensitivity is negative. This negative sign may arise because of the natural or explicit hedges of some firms against exchange rate risk as discussed in Section 2.

Panel A of Table 7 also suggests that incorporating foreign exchange volatility into the spread affects some of the other parameter estimates. The long-run mean of the hazard rate, θ_i falls by approximately 15% when foreign exchange volatility is included, from 0.080 to 0.068. The estimates also suggest that the hazard rate reverts more quickly to the mean in this estimation, with a median point estimate of 5.36 as compared to 3.68 without incorporating foreign exchange volatility. The interquartile range of the estimate is reduced as well, suggesting less cross-sectional variation in mean reversion. Finally, the point estimates suggest a much larger in magnitude price of risk. The median point estimate of η_i across all countries is -1.56 compared to -0.34 in the case where foreign exchange volatility is not included. The range of these estimates is also wider; the difference in the 25th and 75th percentile estimates is 2.01 as compared to 0.58 when foreign exchange volatility is not included.

The table also reports significant differences across countries in parameter estimates. Reductions in the median long-run mean of the hazard rate, θ_i, are particularly pronounced in Brazil, Chile, and Singapore, where the median parameter estimates fall by 0.032, 0.031, and 0.022, or 27%, 35%, and 29%, respectively. these countries also exhibit the largest increases in mean reversion estimates, with κ_i rising by 4.06, 2.78 and 3.98 at the median in Brazil, Chile, and Singapore, respectively. The 25th percentile foreign exchange volatility sensitivity coefficient, $\beta_{i,v}$, is positive in four of the six countries, with exceptions in Chile and South Korea. The median sensitivity in Chile is also negative. The Chilean companies in our sample are dominated by mining firms, whose revenues are denominated in U.S. dollars, which may explain this negative sign.

In Panel H, we present the root mean square pricing errors from the model. The median pricing error of 16 basis points is approximately 2 basis points lower than in the case without foreign exchange volatility; the 25th percentile is approximately 1 basis point lower and the 75th percentile is approximately 6 basis points lower. Thus, incorporating foreign exchange volatility appears to reduce both median mispricing and cross-sectional variation in mispricing. Reductions are particularly stark in countries with large pricing errors; the median pricing error falls by 9 basis points to 18.80 basis points in Brazil and by 12 basis points to 39.54 basis points in Russia. Median pricing errors are reduced in all six countries. There are also substantial reductions in the 75th percentile pricing errors in Chile, by 8 basis points to 19.37 basis points, and in Russia, by 13.5 basis points to 53 basis points. Oddly, the 75th percentile pricing error actually increases slightly in Brazil, by 1.68 basis points to 40.10 basis points.

As a last point of comparison between the two models, we plot the default intensities implied by the two models in Figure 3. The figures depict the average default intensities across bonds by country over the period January, 2005 through September, 2010. The plots are depicted on a common y-axis scale to facilitate cross-country comparisons. There are several notable features of the plots. First, like spreads in general, there is a significant spike in the average default intensity corresponding to the onset of the financial crisis in 2007-2008. This spike is most pronounced in Russia, but is also substantially more pronounced in Singapore, Korea, and Brazil than in Mexico or Chile. Second, across all countries except South Korea, the magnitude of the increase in default intensity is more muted accounting for foreign exchange volatility exposure. This effect persists through the remainder of the sample. However, the effect is not solely limited to the financial crisis. In Brazil, default intensities implied by the model incorporating foreign exchange volatility exceed those implied by the model without foreign exchange prior to the crisis. In contrast, the foreign exchange volatility absorbs a considerable portion of the default intensity in Chile, Mexico, and Russia long before the onset of the crisis. Thus, the importance of foreign exchange volatility does not appear to simply be a crisis effect.

The conclusion that we draw from comparing the results of the estimation in this section with the estimation without foreign exchange volatility in the previous section is that incorporating sensitivity to foreign exchange volatility in the default spread is important both quantitatively and qualitatively for pricing of dollar-denominated securities. Sensitivity to foreign exchange volatility is largely positive, and pricing errors fall fairly uniformly across countries. Qualitatively, incorporating foreign exchange volatility results in different estimates of some of the key parameters of the model. Our results suggest that, in order to better understand the pricing of dollar-denominated bonds, accounting for foreign exchange risk exposure, in this case foreign exchange volatility, is a key modeling component.

5 Conclusion

The rationale among practitioners for purchasing dollar-denominated emerging market debt is that holding this debt allows the purchaser to participate in the emerging debt markets without subjecting the purchaser to exchange rate risk. The literatures on excessive dollar-denominated debt and balance sheet effects in currency crises, however, call this rationale into question. Specifically, these literatures suggest that dollarization of debt may result in increased financial distress and probability of default due to exchange rate devaluations. Since credit spreads on risky bonds are primarily driven by default risk, this argument suggests that investors in dollar-denominated bonds may be indirectly exposed to foreign exchange risk. We investigate whether this is the case in this paper.

Our evidence suggests that there is substantial evidence that credit spreads on dollar-denominated corporate bonds are in fact sensitive to variation in the volatility of foreign exchange rates. This sensitivity differs substantially across countries in sign, magnitude, and the proportion of bonds affected. We find that financial and natural hedges reduce the sensitivity to this risk, although a substantial portion of cross-sectional variation in exposures remains unexplained. When we estimate a standard risky term structure model of emerging market dollar-denominated debt, we find that pricing errors across bonds are higher than in comparable U.S. markets, and that the mispricing is significantly related to financial and natural corporate hedging.

In response to this evidence, we suggest a Duffie and Singleton (1997, 1999) term structure model of risky debt that incorporates exposure to foreign exchange risk in a manner analogous to the incorporation of exposure of risky debt to default-free term structure variables. We find that the model reduces pricing errors for many firms, and reduces the dispersion of pricing errors within several countries. Thus, we conclude that accounting for foreign currency exposure can help standard pricing models in pricing dollar-denominated emerging market debt.

An outstanding issue is the relative pricing of dollar-denominated and local currency-denominated debt. The excess dollar debt literature suggests something of a welfare cost to emerging market firms due to issuance of too much emerging market debt. Examining the relative pricing of dollar-denominated and local currency-denominated debt would permit a more thorough analysis of this question. Absent liquidity and legal issues, it is possible that one could create an arbitrage portfolio of local currency-denominated debt and a currency hedge that replicates the payoffs of the dollar-denominated bond. Any difference might be due to mispricing or an unobserved risk premium attached to dollar vs. local currency borrowing. Unfortunately, readily available data on local currency-denominated corporate debt is not available; the question nevertheless remains interesting for further research.

A Appendix

In this appendix, we present the explicit form of bond pricing coefficients for the models estimated in the paper. In our fully specified model with default and foreign exchange risk, a system of four variables follows risk neutral dynamics

$$\begin{pmatrix} dx_{1,t} \\ dx_{2,t} \\ dv_t \\ dh_{i,t} \end{pmatrix} = \left[\begin{pmatrix} \kappa_1 & 0 & 0 & 0 \\ 0 & \kappa_2 & 0 & 0 \\ 0 & 0 & \kappa_v & 0 \\ 0 & 0 & 0 & \kappa_i \end{pmatrix} \begin{pmatrix} \theta_1 \\ \theta_2 \\ \theta_v \\ \theta_i \end{pmatrix} - \begin{pmatrix} \kappa_1+\eta_1 & 0 & 0 & 0 \\ 0 & \kappa_2+\eta_2 & 0 & 0 \\ 0 & 0 & \kappa_v+\eta_v & 0 \\ 0 & 0 & 0 & \kappa_i+\eta_i \end{pmatrix} \begin{pmatrix} x_{1,t} \\ x_{2,t} \\ v_t \\ h_{i,t} \end{pmatrix} \right] dt$$
$$+ \begin{pmatrix} \sigma_1 & 0 & 0 & 0 \\ 0 & \sigma_2 & 0 & 0 \\ 0 & 0 & \sigma_v & 0 \\ 0 & 0 & 0 & \sigma_i \end{pmatrix} \begin{pmatrix} \sqrt{x_{1,t}} & 0 & 0 & 0 \\ 0 & \sqrt{x_{2,t}} & 0 & 0 \\ 0 & 0 & \sqrt{v_t} & 0 \\ 0 & 0 & 0 & \sqrt{h_{i,t}} \end{pmatrix} \begin{pmatrix} dW_{1,t}^Q \\ dW_{2,t}^Q \\ dW_{v,t}^Q \\ dW_{i,t}^Q \end{pmatrix}, \tag{A.1}$$

where $x_{1,t}$ and $x_{2,t}$ are state variables governing the default-free term structure, v_t is the foreign exchange variance, and $h_{i,t}$ is the default intensity for bond i. The instantaneous risk free rate is a linear function of the state variables,

$$r_t = a_f + x_{1,t} + x_{2,t},$$

and the credit spread as

$$R_{i,t} - r_f = (1-\delta_i)\,\lambda_{i,t} = a_i + \beta_{i,1}\,(x_{1,t} - \bar{x}_1) + \beta_{i,2}\,(x_{2,t} - \bar{x}_2) + \beta_{i,v} v_t + h_{i,t},$$

where $R_{i,t}$ is the instantaneous zero-coupon yield on a risky bond.

Log risky zero coupon bond prices are affine in the state variables in the form

$$\ln P_{i,t}(\tau) = A_i(\tau) + B_{i,1}(\tau) x_{i,1,t}^* + B_{i,2}(\tau) x_{i,2,t}^* + B_{i,v}(\tau) v_{i,t}^* + B_{i,h}(\tau) h_{i,t},$$

where τ is the time to maturity in years till the expiration of the zero coupon bond, and $x_{i,1,t}^* = (1+\beta_{i,1})\,x_{1,t}$, $x_{i,2,t}^* = (1+\beta_{i,2})\,x_{2,t}$, and $v_{i,t}^* = \beta_{i,v} v_t$. Collecting the variables into a four-dimensional vector $\mathbf{y}_t = \{x_{i,1,t}^*, x_{i,2,t}^*, v_{i,t}^*, h_{i,t}\}$,

$$B_{i,j}(\tau) = -\frac{2\left(e_j^{\gamma}\tau - 1\right)}{2\gamma_j + (\kappa_j + \eta_j + \gamma_g)\,(e^{\gamma_j \tau} - 1)} \tag{A.2}$$

$$A_i(\tau) = \sum_{j=1}^{4} \frac{2\kappa_j \theta_j}{\sigma_j^2} \ln\left[\frac{2\gamma_j e^{\frac{1}{2}(\kappa_j+\eta_j+\gamma_j)\tau}}{2\gamma_j + (\kappa_j+\eta_j+\gamma_j)\,(e^{\gamma_j\tau} - 1)}\right], \tag{A.3}$$

where $\gamma_j = \sqrt{(\kappa_j + \eta_j)^2 + 2\sigma_j^2}$, and j indexes the parameters associated with the j^{th} element of $\mathbf{y}_t$.

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Table 1: Summary Statistics for Emerging Market Dollar-Denominated Bonds

Table 1 presents summary statistics for emerging market dollar-denominated bonds in our sample. Bonds are sampled from Datastream and represent fixed coupon semi-annual debentures issued by corporations with no call provisions and fixed maturity. All bonds have payments denominated in U.S. Dollars and are issued by companies in countries considered emerging markets as of January, 2001. Bonds must have at least 250 days of price information and 75% of price changes non-zero. The table presents, by country, median, minimum, and maximum coupon rates and years to maturity of the bonds. The countries in our sample are Brazil (BR), Chile (CL), Mexico (MX), Russia (RS), Singapore (SG), and South Korea (SK). Additionally, we report the number of bonds, number of companies issuing bonds, and first observation by country. Data are sampled over the period 12/28/2000 through 9/28/2010 at the daily frequency.

Country:	BR	CL	MX	RS	SG	SK
Number of Bonds	11	11	11	10	7	18
Number of Companies	6	4	3	4	2	5
Minimum Coupon	6.25	5.13	4.75	5.67	5.75	4.25
Median Coupon	8.00	7.38	5.63	8.48	6.38	5.88
Maximum Coupon	10.50	8.63	6.63	9.75	7.38	8.75
Minimum Life at Issue	5.00	9.50	5.00	5.00	5.00	5.00
Median Life at Issue	10.00	10.00	10.00	6.00	10.00	8.50
Maximum Life at Issue	30.00	30.00	30.00	10.00	30.00	20.00
First Observation Year	2004	2000	2005	2004	2001	2001

Table 2: Sensitivity of Spreads to Foreign Exchange Innovations and Volatility

Table 2 presents results of the following regression specification,

$$s_{i,k,t} = a_i + b_{fx,i,k}\Delta fx_{k,t} + b_{v,i,k}v_{k,t} + \epsilon_{i,t},$$

where $s_{i,k,t}$ is the spread on bond i in country k at time t relative to a comparable maturity Treasury security, $\Delta fx_{k,t}$ is the first difference in the log rate of exchange of U.S. Dollar for the home currency of bond i, and $v_{k,t}$ is the volatility of this first difference modeled via an MA(1), EGARCH (1,1) time series specification. Data on emerging market corporate bonds are obtained from Datastream and represent 68 issues from 24 companies across six countries. Treasury yield data are constant maturity yields obtained from the FRED database at the Federal Reserve. Data are sampled at the daily frequency over various horizons with the first observation in December, 2000 and the final observation in September, 2010. The table presents the median of point estimates, t-statistics in parentheses, and R^2 across all countries and within each country. Last, we report the proportion of point estimates for which the null hypothesis that the coefficient is positive and significantly different than zero at the 5% level can be rejected.

Country	a_i	$b_{fx,i,k}$	$b_{v,i,k}$	R^2
All Countries	1.44	−2.04	192.60	0.37
	(17.74)	(−0.46)	(27.14)	
		0.00	*0.94*	
Brazil	2.35	3.57	115.08	0.48
	(28.62)	(0.63)	(29.05)	
		0.00	*0.91*	
Chile	1.06	−6.96	164.24	0.30
	(22.13)	(−2.49)	(28.47)	
		0.00	*1.00*	
Mexico	1.23	0.20	149.31	0.31
	(20.34)	(0.10)	(18.22)	
		0.00	*0.82*	
Russia	2.82	19.66	737.28	0.27
	(10.84)	(1.11)	(15.39)	
		0.00	*1.00*	
Singapore	−0.21	−1.72	490.61	0.62
	(−5.13)	(−0.42)	(43.33)	
		0.00	*0.86*	
South Korea	1.14	−5.74	195.85	0.48
	(17.22)	(−1.59)	(33.55)	
		0.00	*1.00*	

Table 3: Determinants of Emerging Market Bond Spreads

Table 3 presents results of the following regression specification,

$$s_{i,k,t} = a_i + b_{v,i,k} v_{k,t} + b_{y,i} y_t + b_{ts,i} ts_t + b_{vx,i} vx_t + b_{r_m,i} r_{m,t} + b_{r_l,i} r_{l,t} + \xi_{i,t},$$

where $s_{i,k,t}$ is the spread on bond i in country k at time t relative to a comparable maturity Treasury security and $v_{k,t}$ is the volatility of the first difference of the log exchange rate modeled via an MA(1), EGARCH (1,1) time series specification. The covariates are y_t, the yield on a 10-year constant maturity Treasury Bond, ts_t, the difference in the yield on a 10-year and 2-year constant maturity Treasury Bond, vx_t, the level of the VIX index, $r_{m,t}$, the return on the CRSP value-weighted index, and $r_{l,t}$, the return on the local equity market index. Data on emerging market corporate bonds are obtained from Datastream and represent 68 issues from 24 companies across six countries. Treasury yield data are obtained from the FRED database at the Federal Reserve. Data on the VIX and CRSP value-weighted indices are from CRSP, and data for local market yields and equity returns are from Datastream. Data are sampled at the daily frequency over various horizons with the first observation in December, 2000 and the final observation in September, 2010. The table presents the median of point estimates, t-statistics in parentheses, and R^2 across all countries and within each country. Below the t-statistics, we report the proportion of point estimates for which the null hypothesis that the coefficient is positive and significantly different than zero at the 2.5% level can be rejected.

Country	a_i	$b_{v,i,k}$	$b_{y,i}$	$b_{ts,i}$	$b_{vx,i}$	$b_{r_m,i}$	$b_{r_l,i}$	R^2
All Countries	2.86	13.39	−0.55	−24.50	0.09	6.16	6.02	0.81
	(7.02)	(1.77)	(−7.87)	(−6.40)	(19.57)	(2.92)	(2.20)	
		0.53						
Brazil	2.71	−7.24	−0.48	−21.90	0.08	3.57	2.26	0.86
	(9.40)	(−1.56)	(−7.16)	(−8.32)	(19.28)	(2.01)	(1.29)	
		0.46						
Chile	3.08	−6.35	−0.38	6.08	0.07	3.37	6.58	0.78
	(13.76)	(0.82)	(−9.26)	(2.39)	(19.46)	(2.01)	(3.42)	
		0.36						
Mexico	1.67	−19.44	−0.33	−27.17	0.08	4.96	1.97	0.84
	(7.92)	(−2.21)	(−5.61)	(−9.23)	(22.07)	(2.70)	(1.13)	
		0.18						
Russia	1.24	184.85	−0.65	−78.36	0.32	24.92	6.59	0.91
	(1.03)	(7.04)	(−2.35)	(05.36)	(38.14)	(4.77)	(2.03)	
		1.00						
Singapore	0.85	198.55	−0.13	−13.84	0.04	3.05	3.08	0.91
	(3.84)	(7.83)	(−3.68)	(−5.67)	(11.55)	(2.00)	(2.41)	
		1.00						
S. Korea	4.28	7.25	−0.92	−31.81	0.10	7.86	7.77	0.75
	(7.56)	(0.59)	(−10.15)	(−6.80)	(16.26)	(3.49)	(3.96)	
		0.44						

Table 4: Cross-Sectional Determinants of Foreign Exchange Sensitivity

In Table 4, we present estimates of coefficients in the regressions

$$\hat{b}_{v,i} = d_{02} + d_{12} sales_{j,t} + d_{22} debt_{j,t} + d_{32} deriv_{j,t} + d_{42} coup_i + d_{52} mat_i + u_{2i},$$

where $\hat{b}_{v,i}$ is the point estimate of sensitivity of bond i's credit spread to volatility of exchange rates as reported in Table 3. The variable $sales_{j,t}$ is the proportion of firm j's sales derived from U.S. dollars, $debt_{j,t}$ is the proportion of firm j's total long term debt composed of U.S. dollar debentures, $deriv_{j,t}$ is an indicator variable that takes the value 1 if the firm hedges foreign currency risk and 0 otherwise, $coup_i$ is the coupon rate on the bond, and mat_i is the initial maturity of the bond. The index $t = 2006, 2007, 2008, 2009$ reflects fiscal year ends for which accounting data are available. Data are obtained from 20-F filings with the SEC on the EDGAR database, if available, and directly from company financial statements if not. We report point estimates with t-statistics in parentheses, as well as the regression adjusted R^2.

Dep. Var.	Int.	*sales*	*debt*	*deriv*	*coup*	*mat*	$\bar{R}^2$
$\hat{b}_{i,v}$	238.32	−251.97	422.65	−122.53	−46.92	18.31	0.11
	(0.98)	(−2.93)	(2.40)	(−1.13)	(−1.57)	(3.91)	

Table 5: Parameter Estimates of Risky Bond Prices

Table 5 presents parameter estimates of the defaultable component of bond prices as modeled in equation (18),

$$P_{i,t}(\tau, c) = E_t^Q \left[c \sum_{m=1}^{T-t} e^{-\int_t^{t+m}} R_{i,s} ds + e^{-\int_t^T R_{i,s} ds} \right]$$

where the risk-neutral defaultable yield, $R_{i,s}$, is specified as

$$R_{i,s} = a_i + h_{i,t} + \boldsymbol{\beta}' (\mathbf{x}_t - \bar{\mathbf{x}}),$$

$\mathbf{x}_t = \{x_{1,t}, x_{2,t}\}$ are the state variables filtered from estimation of the default-free term structure and $h_{i,t}$ is the hazard rate, which follows the stochastic differential equation

$$dh_{i,t} = (\kappa_i \theta_i - (\kappa_i + \eta_i))\, dt + \sigma_i \sqrt{h_{i,t}} dW_{i,t}^Q.$$

Parameters are estimated via the extended Kalman filter using discrete time Euler approximations to continuous time dynamics. Parameters are estimated for 68 bonds across six countries using daily observations on bond yields. We report 25th percentile, median, and 75th percentile estimates for the full sample and within each country in Panels A-G. In Panel H, we present 25th percentile, median, and 75th percentiles of root mean square pricing errors for the full sample and within each country.

Panel A: All Countries

Pct:	25	50	75
κ_i	1.43	3.68	13.50
θ_i	0.05	0.08	0.12
η_i	-0.60	-0.34	-0.02
σ_i	0.48	0.77	1.60
β_1	-0.47	-0.31	-0.09
β_2	-0.61	-0.26	0.27
a_i	-0.10	-0.06	-0.03

Panel B: Brazil

Pct	25	50	75
κ_i	2.16	3.69	14.00
θ_i	0.06	0.12	0.14
η_i	-0.65	-0.55	-0.18
σ_i	0.53	0.86	2.49
β_1	-0.94	-0.54	-0.30
β_2	-1.44	-0.59	0.49
a_i	-0.13	-0.09	-0.04

Panel C: Chile

Pct	25	50	75
κ_i	0.00	0.28	10.72
θ_i	0.08	0.09	0.14
η_i	-0.32	-0.05	0.24
σ_i	0.10	0.35	1.35
β_1	-0.35	-0.14	0.03
β_2	-0.64	-0.22	0.58
a_i	-0.11	-0.07	-0.02

Panel D: Mexico

Pct	25	50	75
κ_i	1.77	13.72	14.32
θ_i	0.05	0.06	0.09
η_i	-0.52	-0.17	-0.11
σ_i	0.86	1.31	1.58
β_1	-0.40	-0.30	-0.15
β_2	-0.43	-0.10	0.28
a_i	-0.07	-0.05	-0.02

Table continued on the next page.

Panel E: Russia

Pct:	25	50	75
κ_i	1.76	3.85	13.28
θ_i	0.07	0.09	0.12
η_i	-0.76	-0.58	-0.46
σ_i	0.42	0.69	1.69
β_1	-0.78	-0.33	-0.07
β_2	-0.70	0.11	0.55
a_i	-0.12	-0.08	-0.06

Panel F: Singapore

Pct	25	50	75
κ_i	1.63	4.64	19.67
θ_i	0.04	0.07	0.16
η_i	-0.52	-0.16	0.45
σ_i	0.49	0.85	2.27
β_1	-0.43	-0.11	-0.00
β_2	-0.72	-0.15	-0.12
a_i	-0.15	-0.07	-0.03

Panel G: South Korea

Pct	25	50	75
κ_i	1.40	3.47	7.74
θ_i	0.04	0.07	0.14
η_i	-0.71	-0.31	-0.02
σ_i	0.49	0.63	1.29
β_1	-0.42	-0.25	-0.08
β_2	-0.52	-0.33	0.00
a_i	-0.10	-0.05	-0.01

Panel H: Pricing Errors

Country	25	50	75
All	12.28	18.34	35.09
Brazil	13.29	27.74	38.42
Chile	11.007	14.46	27.13
Mexico	8.51	11.60	15.14
Russia	26.30	51.76	66.53
Singapore	10.66	19.44	33.09
South Korea	13.39	17.96	24.71

Table 6: Cross-Sectional Determinants of Pricing Errors

In Table 6, we present estimates of coefficients in the regression

$$rmse_i = d_0 + d_1 sales_{j,t} + d_2 debt_{j,t} + d_3 deriv_{j,t} + d_4 coup_i + d_5 mat_i + u_i$$

where $rmse_i$ is the root mean square pricing error from the estimation of equation (18). The variable $sales_{j,t}$ is the proportion of firm j's sales derived from U.S. dollars, $debt_{j,t}$ is the proportion of firm j's total long term debt composed of U.S. dollar debentures, $deriv_{j,t}$ is an indicator variable that takes the value 1 if the firm hedges foreign currency risk and 0 otherwise, $coup_i$ is the coupon rate on the bond, and mat_i is the initial maturity of the bond. The index $t = 2006, 2007, 2008, 2009$ reflects fiscal year ends for which accounting data are available. Data are obtained from 20-F filings with the SEC on the EDGAR database, if available, and directly from company financial statements if not. We report point estimates with t-statistics in parentheses, as well as the regression adjusted R^2.

Dep. Var.	Int.	*sales*	*debt*	*deriv*	*coup*	*mat*	$\bar{R}^2$
$\hat{b}_{i,fx}$	12.92	−32.99	56.96	−37.61	2.00	−0.87	0.17
	(0.54)	(−3.88)	(3.27)	(−3.49)	(−0.29)	(4.32)	
$\hat{b}_{i,v}$	238.32	−251.97	422.65	−122.53	−46.92	18.31	0.11
	(0.98)	(−2.93)	(2.40)	(−1.13)	(−1.57)	(3.91)	

Table 7: Parameter Estimates of Risky Bond Prices with Foreign Exchange Risk

Table 7 presents parameter estimates of the defaultable component of bond prices as affected by exchange rate risk,

$$P_{i,t}(\tau,c) = E_t^Q\left[c\sum_{m=1}^{T-t} e^{-\int_t^{t+m} R^*_{i,k,s}ds} + e^{-\int_t^T R^*_{i,k,s}ds}\right]$$

where the risk-neutral defaultable yield, $R^*_{i,k,s}$, in country k is specified as

$$R_{i,k,s} = a_i + h_{i,t} + \boldsymbol{\beta}_i'(\mathbf{x}_t - \bar{\mathbf{x}}) + \beta_v v_{k,t}.$$

The variable $v_{k,t}$ is the volatility of the log foreign exchange rate in country k, $\mathbf{x}_t = \{x_{1,t}, x_{2,t}\}$ are the state variables implied by parameter estimates from the risk free term structure, and $h_{i,t}$ is the hazard rate, which follows the stochastic differential equation

$$dh_{i,t} = (\kappa_i\theta_i - (\kappa_i + \eta_i))\,dt + \sigma_i\sqrt{h_{i,t}}dW^Q_{i,t}.$$

The foreign exchange rate volatility is also assumed to follow a stochastic differential equation,

$$dv^*_{k,t} = \left(\kappa_v\beta_v\theta_v - (\kappa_v + \eta_v)\,v^*_{l,t}\right)dt + \sigma_v\sqrt{\beta_v v^*_{k,t}}dW^Q_{k,t}$$

Parameters are estimated via the extended Kalman filter using discrete time Euler approximations to continuous time dynamics. Parameters are estimated for 86 bonds across six countries using daily observations on bond yields. We report 25th percentile, median, and 75th percentile estimates for the full sample and within each country in Panels A-G. In Panel H, we present 25th percentile, median, and 75th percentiles of root mean square pricing errors for the full sample and within each country.

Panel A: All Countries			
Pct:	25	50	75
κ_i	1.73	5.36	9.96
θ_i	0.04	0.07	0.10
η_i	-2.52	-1.56	-0.51
σ_i	0.40	0.76	1.32
β_1	-0.55	-0.33	-0.15
β_2	-0.70	-0.34	0.14
β_v	-0.02	0.18	0.59
a_i	-0.11	-0.08	-0.03

Panel B: Brazil			
Pct	25	50	75
κ_i	5.50	7.75	14.19
θ_i	0.05	0.09	0.10
η_i	-3.00	-2.42	-1.48
σ_i	0.98	1.57	2.18
β_1	-0.89	-0.44	-0.17
β_2	-1.42	-0.53	0.12
β_v	0.36	0.62	1.41
a_i	-0.11	-0.09	-0.07

Panel C: Chile			
Pct	25	50	75
κ_i	0.80	3.06	9.59
θ_i	0.03	0.06	0.36
η_i	-1.87	-1.04	-0.57
σ_i	0.36	0.48	0.92
β_1	-0.36	-0.15	0.05
β_2	-0.97	-0.36	0.31
β_v	-0.08	-0.02	0.01
a_i	-0.10	-0.07	-0.02

Panel D: Mexico			
Pct	25	50	75
κ_i	0.91	2.92	6.26
θ_i	0.02	0.06	0.11
η_i	-2.28	-1.53	-0.43
σ_i	0.23	0.41	0.76
β_1	-0.43	-0.38	-0.19
β_2	-0.88	-0.13	0.34
β_v	-0.02	0.19	0.44
a_i	-0.18	-0.10	-0.01

Table continued on the next page.

Panel E: Russia

Pct:	25	50	75
κ_i	1.02	4.42	10.33
θ_i	0.04	0.08	0.10
η_i	-3.70	-2.43	-0.92
σ_i	0.45	0.81	1.64
β_1	-0.93	-0.83	-0.58
β_2	-0.70	-0.53	0.01
β_3	0.16	0.29	1.11
a_i	-0.11	-0.06	0.00

Panel F: Singapore

Pct	25	50	75
κ_i	0.82	8.62	12.54
θ_i	0.03	0.05	0.07
η_i	-2.20	-1.29	-0.44
σ_i	0.30	0.79	2.06
β_1	-0.39	-0.30	-0.18
β_2	-1.03	-0.70	-0.32
β_3	0.49	1.33	1.92
a_i	-0.10	-0.04	-0.02

Panel G: South Korea

Pct	25	50	75
κ_i	1.60	4.95	8.59
θ_i	0.05	0.07	0.11
η_i	-2.05	-1.17	1.21
σ_i	0.49	0.75	1.02
β_1	-0.47	-0.18	-0.02
β_2	-0.39	-0.16	0.21
β_3	-0.09	0.02	0.35
a_i	-0.10	-0.07	-0.03

Panel H: Pricing Errors

Country	25	50	75
All	11.39	16.43	29.08
Brazil	14.74	18.80	40.10
Chile	10.23	12.07	19.37
Mexico	8.35	10.57	12.50
Russia	26.45	39.54	53.00
Singapore	9.82	16.73	31.85
South Korea	12.31	16.73	23.24

Figure 1: Percent of Emerging Market Debt Denominated in Emerging Currencies

Figure 1 depicts the fraction of total outstanding debt issued in international markets by emerging market issuers denominated in emerging currencies. Emerging markets and currencies follow the definitions of the Bank for International Settlements (BIS). Data are obtained from the BIS Quarterly review. Percentages are calculated by summing the dollar amount outstanding of international bonds and notes denominated in emerging market currencies (as designated by the BIS) from the BIS Quarterly Review Table 13B, and dividing by the total dollar amount outstanding issued by emerging markets issuers in BIS Quarterly Review Table 15B. The data cover the period September, 1993 through December, 2010.

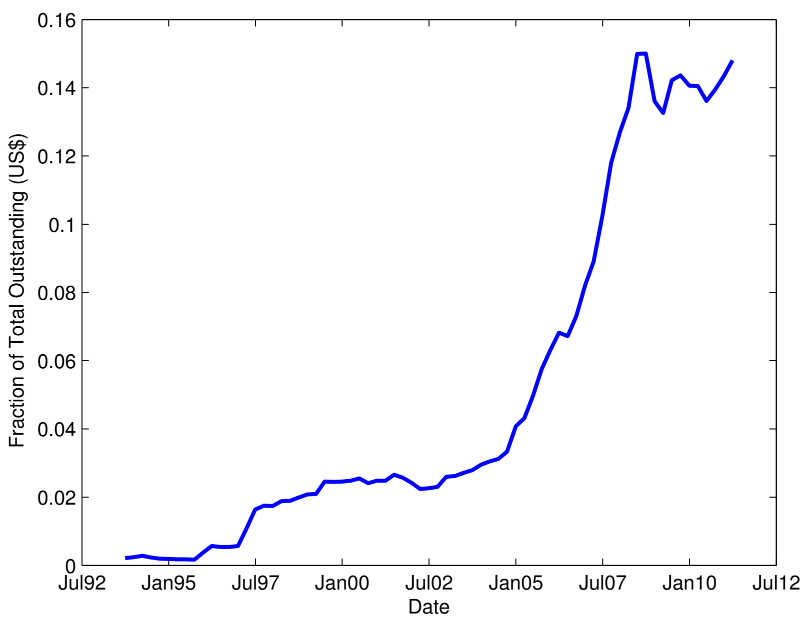
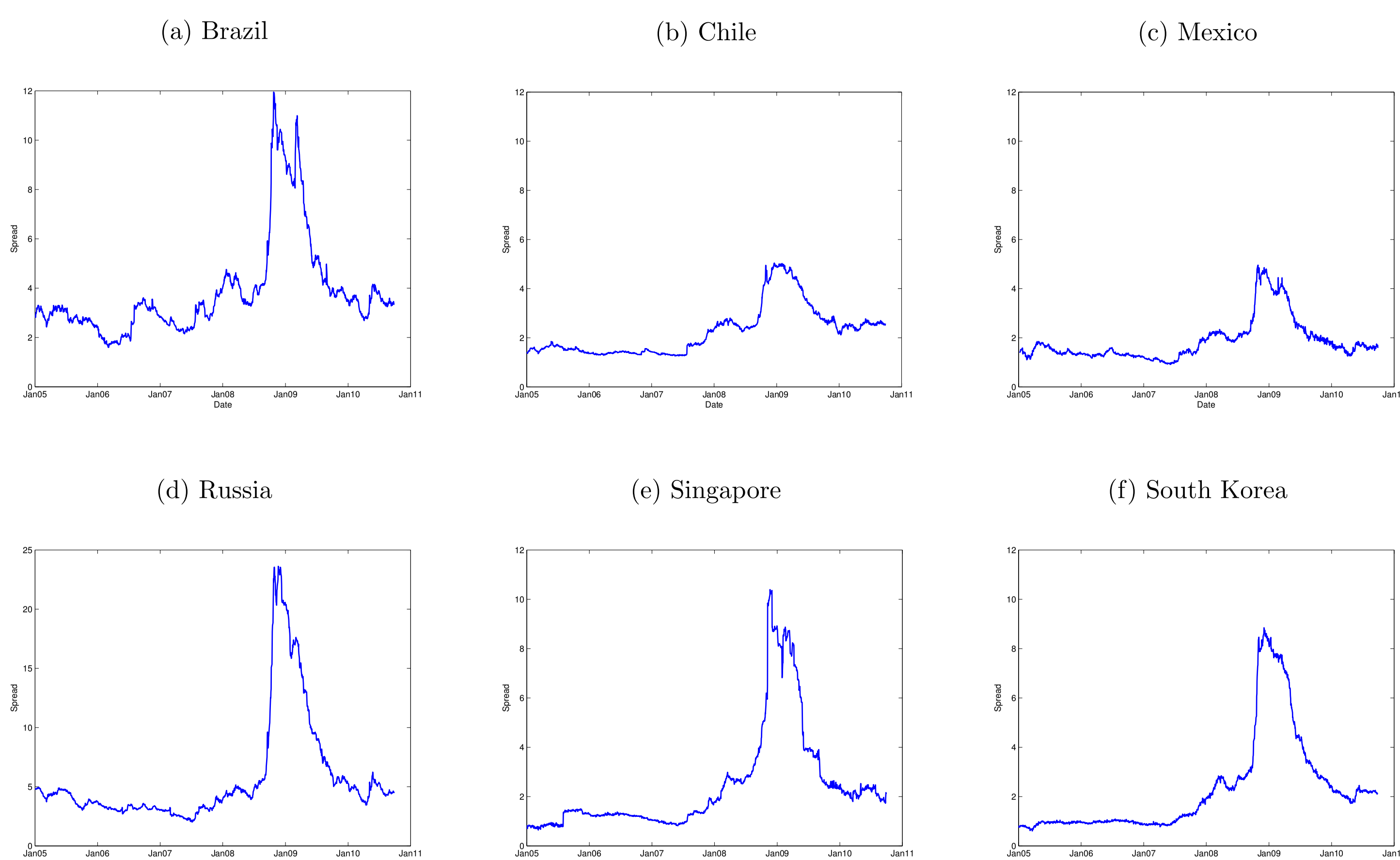


Figure 2: Time Series of Average Bond Yield Spreads

Figure 2 presents the time series of average yield spreads of emerging market corporate bonds plotted for each country in our sample. Yield spreads for individual bonds are calculated as the difference in the yield to maturity on the issue and the yield on a Treasury security with the closest maturity. Yield spreads are then averaged across the bonds within each country on each date to produce a single time series observation for each country. Data on individual bond yields are obtained from DataStream and Treasury yields are constant maturity yields from the Federal Reserve Board of Governors (FRED). Data plotted cover the period January, 2005 through September, 2010, and are sampled at the daily frequency. Panel a) depicts average spreads for Brazilian bonds, b) for Chilean bonds, c) for Mexican bonds, d) for Russian bonds, e) for Singaporean bonds, and f) for South Korean bonds. Panels are depicted on a common y-axis scale with the exception of Russia.



(a) Brazil

(b) Chile

(c) Mexico

(d) Russia

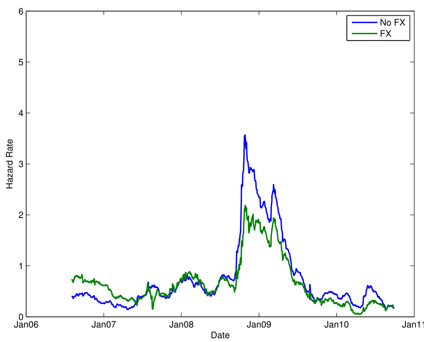
(e) Singapore

(f) South Korea

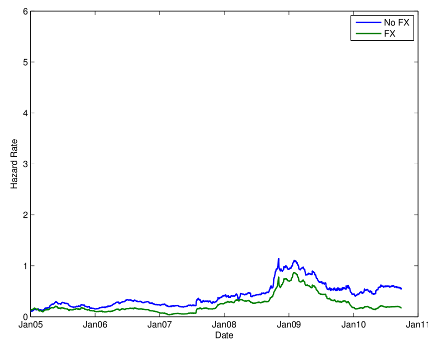
Figure 3: Time Series of Average Default Intensities

Figure 3 presents the time series of average default intensities of emerging market corporate bonds plotted for each country in our sample. Data on individual bond yields are obtained from DataStream. We depict default intensities filtered from a model without foreign exchange sensitivity in blue and with foreign exchange sensitivity in green. Data plotted cover the period January, 2005 through September, 2010, and are sampled at the daily frequency. Panel a) depicts average spreads for Brazilian bonds, b) for Chilean bonds, c) for Mexican bonds, d) for Russian bonds, e) for Singaporean bonds, and f) for South Korean bonds. Panels are depicted on a common y-axis scale.

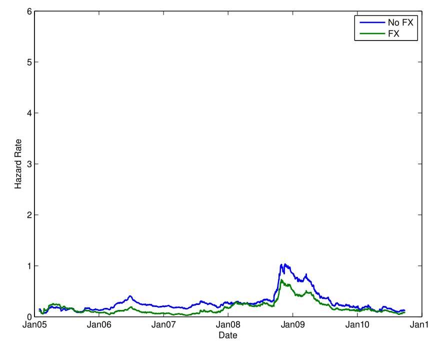
(a) Brazil



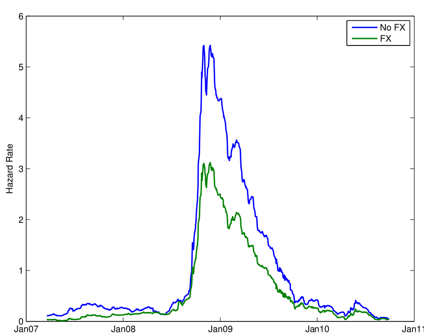
(b) Chile



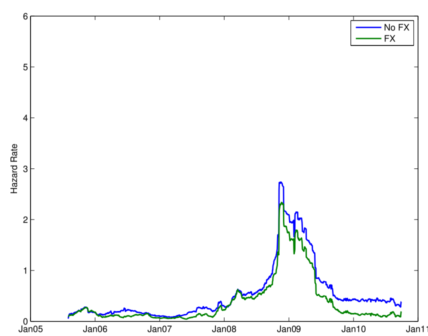
(c) Mexico



(d) Russia



(e) Singapore



(f) South Korea