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

Stefanos Delikouras†
Robert F. Dittmar‡
Haitao Li§

November 15, 2012

*All errors are the responsibility of the authors.

†Department of Finance, Stephen Ross School of Business, University of Michigan, Ann Arbor, MI 48109

‡Department of Finance, Stephen Ross School of Business, University of Michigan, Ann Arbor, MI 48109

§Department of Finance, Stephen Ross School of Business, University of Michigan, Ann Arbor, MI 48109 and Cheung Kong Graduate School of Business, Beijing, PR China 100738

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. 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]

Thus, from the perspective a U.S.-domiciled investor, an emerging market bond denominated in U.S. dollars might be viewed as free of foreign exchange risk.

While buying debt denominated in dollars may directly hedge an investor against currency risk, issuing debt in dollars exposes a company to foreign exchange risk. In particular, if the company's revenue is denominated in local currency, and this risk is unhedged in derivatives markets, the borrower's cash flow available for debt service will be sensitive to the dollar exchange rate. Consequently, the borrower is potentially exposed to greater default risk.[4] An investor in these bonds should recognize this increased risk, and demand compensation for the default risk induced by the dollarization through a higher yield. We investigate the degree to which foreign exchange

[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.

[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.

risk impacts implied default rates, and thus the yield that investors demand on dollar-denominated emerging market bonds.

Our approach to investigating this issue is through the lens of an affine reduced-form model of defaultable debt, as in Duffie and Singleton (1997, 1999). Specifically, we estimate the parameters of the reduced-form model and recover the associated default intensities. We then investigate whether the recovered default intensities are affected by the level and the volatility of exchange rates between bonds' home countries and the U.S. Dollar. We find strong evidence to suggest that the default intensity is increasing in exchange rate volatility. Across all bonds in our sample, volatility in foreign exchange spot rates can explain 35% of the variation in default intensities on average. This finding is consistent with the hypothesis that dollarization of debt results in increased default risk, as firms are no longer naturally hedged against adverse shocks that affect their home country's economy.

Given the evidence of links in default intensity and volatility of exchange rates, we extend the Duffie and Singleton (1999) model to include sensitivity of yields to exchange rate volatility. We re-estimate the model to investigate the effect of foreign exchange volatility on yields and any potential improvements in pricing. The parameter estimates suggest that, at the median, a 1% increase in exchange rate volatility results in a 0.16% increase in the yield on the dollar-denominated bond. Moreover, the results suggest that the model reduces the pricing error of bonds relative to the case in which exchange rate volatility sensitivity is not included. In particular, the maximum pricing errors in most countries in our sample are substantially reduced. Our results suggest that although purchasing dollar-denominated bonds may eliminate direct risk of exposure to exchange rates, this risk is traded at least partially for increased default risk, and that prices of emerging market dollar-denominated bonds reflect this indirect exposure.

The question that is at the center of this paper, whether foreign exchange risk affects default rates, is related to a large literature investigating "excessive dollar debt." It is well-recognized that, for a domestic company with domestic revenues, issuing debt in local currency provides a hedge against currency depreciation. In contrast, issuing in foreign currency results in balance sheet exposure that may lead to default in currency crises. Consequently, researchers observe emerging market companies issuing levels of dollar debt that appear to be excessive given the exposure to currency crisis risk. While a number of explanations have been advanced for this apparently sub-optimal behavior, we are not aware of any that examine the pricing of these bonds from an investor perspective. In particular, if an investor in a dollar-denominated emerging market bond is aware that she is shifting foreign exchange risk onto the borrower, resulting in increased default risk, does she require compensation for this increased risk? Our goal is to quantify the impact of this tradeoff in terms of the price paid on emerging market corporate debt.

We focus on corporate rather than sovereign debt because sovereign entities have the ability to directly affect their own currencies' values. Eichengreen and Hausmann (2004) refer to the difficulty of sovereign entities' ability to issue debt in local currencies to "original sin," and an extensive literature has been devoted to explaining the phenomenon. Calvo (2001) suggests that dollar-denominated debt is a self-disciplining mechanism for sovereign entities who cannot commit credibly to sound monetary policy. The issuance of debt in dollars prevents the entity from eliminating its debt through a currency devaluation. Consequently, understanding the pricing of dollar-denominated sovereign debt requires modeling the behavior of a monetary authority that is outside the scope of this paper.

In contrast to sovereign entities, corporations cannot inflate away their debt; Calvo (2001) suggests that the commitment mechanism cannot apply to private borrowers since they must take their domestic monetary policy as given. Jeanne (2004) argues, however, that decreasing monetary policy credibility can induce firms to issue more dollar debt. He considers a model in which monetary policy credibility affects domestic real interest rates. As the probability that a sovereign entity will depreciate its currency increases, the domestic real interest rate will rise, resulting in increased borrowing costs for firms issuing in domestic currency if the currency is not depreciated. Dollar-denominated debt hedges against this cost of borrowing.

Alternative explanations are presented in Caballero and Krishnamurthy (2003) and Korinek (2010). Caballero and Krishnamurthy (2003), solve a model in which domestic agents undervalue insuring against an exchange rate depreciation and foreign lenders have limited participation in emerging markets due to limited development, resulting in firms borrowing in dollars. Investors in their model are risk neutral; thus the focus is on the actions of the corporate agents rather than the risk premia required by investors to invest in dollar- versus local currency-denominated debt. Korinek (2010) develops an equilibrium model of demand for foreign and locally-denominated currency with risk averse agents and a social planner. Agents choose an amount of local currency debt such that the insurance benefits of this debt offset the cost of obtaining the insurance, resulting in the equilibrium risk premium for local currency debt. When decision making is decentralized, borrowing constraints result in a higher equilibrium risk premium for local currency debt than in the presence of a social planner. This risk premium arises due to the financial accelerator effect; if a negative economic shock causes borrowing constraints to bind, then investors reduce consumption, resulting in exchange rate depreciation, resulting in greater binding of the borrowing constraint. Kedia and Mozumadar (2003) empirically investigate the decision to issue foreign currency debt and find support for hedging motives and segmented capital markets affect the choice of foreign currency debt.

Our paper differs from those discussed above in that it is not concerned with the reasons that emerging market *firms* issue excessive amounts of dollar denominated debt. This existing literature on the denomination of emerging market debt approaches the issue from the standpoint of the issuer and implications for monetary policy and development. In contrast, we ask whether investors trade off the risks induced by local currency debt (such as the financial accelerator effect discussed above) relative to the increased default risk induced by dollarization. Thus, our focus is on whether dollar denominated debt represents an externality to creditors, rather than borrowers, or whether these risks are accurately compensated in dollar-denominated debt.

The remainder of this paper is organized as follows. In Section 2, we discuss the baseline and extend model for dollar denominated debt. In Section 3, we outline our strategy for estimating the model and the data employed in the model. Section 4 presents the results of our empirical investigation. Section 5 contains concluding remarks and directions for further research.

2 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. corporate debenture and assume that both types of bonds are the same instrument. 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

$$P\left(t,T\right)=E_t^Q\left[e^{-\int_t^T R_s ds}\right], \tag{1}$$

with R_s representing the instantaneous default-adjusted discount rate,

$$R_t = r_t + \left(1-\delta\right)\lambda_{d,t} \tag{2}$$

where r_t is the instantaneous risk free rate, δ is the rate of recovery on the debt, and $\lambda_{d,t}\left(1-\delta\right)$ is the spread in excess of the risk-free rate.

Our goal in pursuing this approach is to ask whether the hazard rates implied by the estimates of the model exhibit sensitivity to fluctuations in foreign exchange spot rates. Under the hypothesis that dollar-denominated bonds completely hedge investors against foreign exchange risk, currency spot rates should have no incremental power to explain hazard rates and yields. As a result, the reduced-form model of equations (1) and (2) should adequately describe the pricing of dollar-denominated emerging market corporate bonds.

2.1 Risk Free Bond Prices

Equation (2) shows that the default-adjusted discount rate is a function of the risk-free term structure and hazard rates. We specify the risk free term structure following Duffie and Kan (1996), and assume an affine framework for the pricing of risk-free debt. Specifically, we assume that the risk free rate is affine in two state variables,

$$r_t = a_f + s_{1,t} + s_{2,t}, \tag{3}$$

where the state variables $s_{1,t}$ and $s_{2,t}$ follow Cox, Ingersoll, and Ross (1985) dynamics

$$d\mathbf{s}_t = \boldsymbol{K}\left(\boldsymbol{\Theta} - \mathbf{s}_t\right) dt + \boldsymbol{\Sigma}\sqrt{\mathbf{S}_t} d\mathbf{W}_t^P. \tag{4}$$

The matrix $\mathbf{K}$ is a 2×2 diagonal matrix of mean reversion coefficients, $\boldsymbol{\Theta}$ is a 2×1 vector of long-tun means, $\boldsymbol{\Sigma}$ is a 2×2 diagonal matrix, and $d\mathbf{W}_t^P$ is a 2×1 vector of independent Browian motions under the physical probability measure, P. The matrix $\mathbf{S}_t$ is a 2×2 diagonal matrix with the state variables on the diagonal.

Under an equivalent risk neutral measure, Q, the dynamics of the state variables are given by

$$d\mathbf{s}_t = \left(\mathbf{K}\boldsymbol{\Theta} - \left(\mathbf{K} + \boldsymbol{\Lambda}\right)\mathbf{s}_t\right) dt + \boldsymbol{\Sigma}\sqrt{\mathbf{S}_t} d\mathbf{W}_t^Q. \tag{5}$$

where $\boldsymbol{\Lambda}$ is a 2×2 diagonal matrix with prices of risk η_1 and η_2 on the diagonal. Since this is a standard Cox, Ingersoll, and Ross (1985), we know that risk-free bond yields are given by

$$Y\left(s_t, \tau\right) = a_f - \frac{A\left(\tau\right) + \mathbf{B}'\left(\tau\right)\mathbf{s}_t}{\tau} \tag{6}$$

where $\tau = T - t$ is the time to maturity in years and the coefficients $A\left(\tau\right)$ and $\mathbf{B}\left(\tau\right)$ are given by

$$B_i\left(\tau\right) = -\frac{2\left(e^{\gamma_i \tau} - 1\right)}{2\gamma_i + \left(\kappa_i + \eta_i + \gamma_i\right)\left(e^{\gamma_i \tau} - 1\right)} \tag{7}$$

$$A\left(\tau\right) = \sum_{i=1}^{2} \frac{2\kappa_i\theta_i}{\sigma_i^2} \ln\left[\frac{2\gamma_i e^{\frac{1}{2}\left(\kappa_i + \eta_i + \gamma_i\right)\tau}}{2\gamma_i + \left(\kappa_i + \eta_i + \gamma_i\right)\left(e^{\gamma_i \tau} - 1\right)}\right], \tag{8}$$

where $\gamma_i = \sqrt{\left(\kappa_i + \eta_i\right)^2 + 2\sigma_i^2}$.

One issue that deserves mention is our specification of a two-factor default free term structure. Litterman and Scheinkmann (1991) document three term structure factors, representing a level, slope, and curvature factor, in the term structure of U.S. interest rates. We follow Duffee (1999) in using a two-factor default-free term structure. Our motivation is simply parsimony, since the first two factors empirically dominate the determination of the U.S. term structure.

2.2 Risky Bond Prices

Following Duffie and Singleton (1999), we augment the risk-free term structure by a term structure of default risk. Our modeling approach represents a special case of the Duffie and Singleton (1999) framework examined in Duffee (1999). The Brownian motion driving the evolution of the default rate is independent of the Brownian motions governing the riskless rate, and the default rate does not depend explicitly on the state variables that determine the risk-free rate. However, the spread itself depends on the risk-free rate state variables as

$$R_t - r_t = (1-\delta)\,\lambda_{d,t} = a_d + h_{d,t} + \boldsymbol{\beta}'\left(\mathbf{s}_t - \bar{\mathbf{s}}\right). \tag{9}$$

The parameter β allows for correlation between the default-free term structure and the spread on the bond above the risk free rate. Longstaff and Schwartz (1995) argue theoretically for a negative relation between the credit spread and the risk-free rate, since the risk-neutral drift of the value of the firm's assets, and consequently the distance to default, is increasing in the risk-free rate. This relation is docmented empirically in Duffee (1998).

Terming the default factor $h_{d,t}$, we also specify its dynamics as a square root process under the physical and risk neutral measures:

$$dh_{d,t} = \kappa_d\left(\theta_d - h_{d,t}\right)dt + \sigma_d\sqrt{h_{d,t}}dW^P_{d,t} \tag{10}$$

$$dh_{d,t} = \left(\kappa_d\theta_d - \left(\kappa_d + \eta_d\right)h_{d,t}\right)dt + \sigma_d\sqrt{h_{d,t}}dW^Q_{d,t}. \tag{11}$$

As discussed in Duffie and Singleton (1999), the default factor is a combination of the hazard rate and fractional loss given default process for the bond. Given simply the information in bond prices, we cannot separately identify these components of the default intensity. The authors also note that one can view this factor as the arrival intensity of a jump that first occurs as default. Thus, although default is a discrete event, this intensity follows a diffusion.

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.

The solution for the contribution of default risk to bond prices is isomorphic to the solution for risk-free bond prices above. First define the processes for state variables $i = 1, 2$,

$$ds^*_{i,t} = \left(\kappa_i \theta^*_i - (\kappa_i + \eta_i)\, s^*_{i,t}\right) dt + \sigma_i \sqrt{(1 + \beta_{d,i})\, s^*_{i,t}} dW^Q_{i,t} \tag{12}$$

where $s^*_{i,t} = (1 + \beta_{d,i})\, s_{i,t}$ and $\theta^*_i = (1 + \beta_{d,i})\, \theta_i$. The price of a zero-coupon bond with maturity T is given by

$$P_d(\tau) = e^{A^*(\tau) + \mathbf{B}^{*\prime}(\tau)\mathbf{s}^*_t + A_d + B_d(\tau) h_{d,t}} \tag{13}$$

where

$$B^*_i(\tau) = -\frac{2\left(e^{\gamma^*_i \tau} - 1\right)}{2\gamma^*_i + (\kappa_i + \eta_i + \gamma^*_i)\left(e^{\gamma^*_i \tau} - 1\right)} \tag{14}$$

$$A^*(\tau) = \sum_{i=1}^{2} \frac{2\kappa_i \theta_i}{\sigma_i^2} \ln\left[\frac{2\gamma^*_i e^{\frac{1}{2}\left(\kappa_i + \eta_i + \gamma^*_i\right)\tau}}{2\gamma^*_i + (\kappa_i + \eta_i + \gamma^*_i)\left(e^{\gamma^*_i \tau} - 1\right)}\right] \tag{15}$$

$$B_d(\tau) = -\frac{2\left(e^{\gamma_d \tau} - 1\right)}{2\gamma_d + (\kappa_d + \eta_d + \gamma_d)\left(e^{\gamma_d \tau} - 1\right)} \tag{16}$$

$$A(\tau) = \frac{2\kappa_d \theta_d}{\sigma_d^2} \ln\left[\frac{2\gamma_d e^{\frac{1}{2}(\kappa_d + \eta_d + \gamma_d)\tau}}{2\gamma_d + (\kappa_d + \eta_d + \gamma_d)\left(e^{\gamma_d \tau} - 1\right)}\right], \tag{17}$$

with $\gamma^*_i = \sqrt{(\kappa_i + \eta_i)^2 + 2\sigma^{*2}_i}$ and $\gamma_d = \sqrt{(\kappa_d + \eta_d)^2 + 2\sigma_d^2}$.

Finally, we note that the risky bonds that we will examine are coupon bonds. We treat these coupon bonds as a series of zero coupon bonds with face value c plus a final payment 1. Mathematically, the price of the coupon bond with maturity T is given by

$$P_{d,t}(\tau, c) = E^Q_t\left[c \sum_{m=1}^{T-t} e^{-\int_t^{t+m} R_s ds} + e^{-\int_t^T R_s ds}\right]. \tag{18}$$

where m indexes the periodic coupon payments.

2.3 Estimation Procedure

The state variables of the model, s_1 and s_2, as well as the hazard rate h_d, 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

$$\mathbf{Y}_t(\boldsymbol{\tau}) = a_f \boldsymbol{\iota} - \frac{1}{\tau}\left(\mathbf{A}(\boldsymbol{\tau}) + \mathbf{B}'(\boldsymbol{\tau})\,\mathbf{s}_t\right) + \mathbf{u}_t \tag{19}$$

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

Transition equations for the state variables are given by:

$$\begin{pmatrix} s_{1,t} \\ s_{2,t} \end{pmatrix} = \begin{pmatrix} \theta_1 (1 - e^{-\kappa_1}) \\ \theta_2 (1 - e^{-\kappa_2}) \end{pmatrix} + \begin{pmatrix} e^{-\kappa_1} & 0 \\ 0 & e^{-\kappa_2} \end{pmatrix} \begin{pmatrix} s_{1,t-1} \\ s_{2,t-1} \end{pmatrix} + \begin{pmatrix} w_{1,t} \\ w_{2,t} \end{pmatrix}, \tag{20}$$

where

$$\mathbf{w}_t \sim \mathcal{N}\left(\mathbf{0}, \begin{pmatrix} Q_{1,t} & 0 \\ 0 & Q_{2,t} \end{pmatrix}\right) \tag{21}$$

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

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 the defaultable bond price equation, measured with error:

$$P_{d,t}(\tau, c) = c \sum_{m=1}^{\tau} P_d(m) + P_d(\tau) + u_{d,t}, \tag{23}$$

where $P_d(\tau)$ is the zero-coupon risky bond price in equation (13). 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:

$$h_{d,t} = \frac{\theta_d \kappa_d}{\kappa_d + \eta_d} \left(1 - e^{-\kappa_d}\right) + e^{-\kappa_d} h_{d,t-1} + w_{d,t}, \tag{24}$$

where

$$w_{d,t} \sim \mathcal{N}(0, Q_{d,t}), \tag{25}$$

$$Q_{d,t} = h_{d,t-1} \frac{\sigma^2}{\kappa_d} \left(e^{-\kappa_d} - e^{-2\kappa_d}\right) + \theta_d \frac{\sigma_d^2}{2\kappa_d} \left(1 - e^{-\kappa_d}\right)^2. \tag{26}$$

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.

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.

3 Data and Estimation Results

3.1 Data

Our estimation of the parameters of the default-free term structure utilizes constant maturity Treasury yields from the Federal Reserve H.15 report. We sample data at the daily frequency for a set of eight yields; 3 month, 6 month, 1 year, 2 year, 5 year, 7 year, 10 year, and 20 year maturities. Data are sampled at the daily frequency over the period January 3, 1994 through September 28, 2010, for 4,190 observations per maturity. The time series of the 10-Year and 3-Month yields are depicted in Figure 3. As shown in the figure, yields have exhibited considerable variation over this period, with the yield curve moving through upward sloping, flat, and inverted shapes. Summary statistics for Treasury yields are presented in Table 1. As shown in the table, the average yield curve is upward sloping over this period, with an average term premium of 20 year maturity bonds in excess of 3 month maturity bonds of 2.12%.

Data on emerging market corporate bonds are taken from Datastream. We compile a set of noncallable dollar-denominated bonds issued by companies of non-U.S. origin. We exclude bonds that are not fixed-coupon debentures, ruling out perpetuities, floating rate debt, and other non-standard structures. Additionally, in order to identify parameters and filter state variables, we focus on bonds that are actively traded. Since we have only price information, we use as a measure of whether the bond is actively traded the fraction of non-zero price change days, as in Lesmond, Ogden, and Trzcinka (1999). We restrict attention to those bonds that have ratio of non-zero price change days to total observations in excess of 0.75. Observations with prices that imply negative yields are also eliminated. Finally, we eliminate bonds with fewer than 250 trading days of data available. This procedure yields a total of 86 bonds in six countries.

Summary statistics for these bonds are presented in Table 2. The number of bonds are fairly evenly distributed across countries; the fewest bonds available are in Mexico (11) and the most in South Korea (19). Company representation is also fairly evenly distributed, with the fewest companies in Mexico (3) and Chile, Russia, and South Korea having the most (8). For most countries the minimum initial maturity is five years and the maximum is 30. Exceptions are found in Chile (minimum initial maturity of 9.5 years), Russia (maximum initial maturity of 10.0 years), and South Korea (maximum initial maturity of 20.0 years). Average coupons range from 5.68% in Mexico to 8.28% in Brazil and Russia. For the bonds in our sample, coupon rates range from a minimum of 4.25% (SK Telecom 7-year debentures maturing 1/4/11) to a maximum of 10.50% (JBS 10-year debentures maturing 4/8/16).

Because each country has a different number of bonds and different length time series, it is difficult to summarize information on bond yields across countries. Data in Chile start in December, 2000, while data in Mexico do not start until January, 2005. To provide some information about the bond yields in the different countries, we take an average of yields across the bonds in each country on each date. We report the mean of this averaged series, its standard deviation, the minimum, and maximum. As shown in the table, the highest mean average yield is 8.86% in Russia, and the lowest is 5.53%, in Singapore. Russia also appears to have the most volatile yields, with a standard deviation of 3.69%. The highest maximum average yield is also in Russia, at 28.00%, whereas the lowest maximum average yield is in Chile, at 8.68%. Finally, the lowest minimum average yield is in Singapore, at 2.59%, with the highest minimum in Brazil at 5.55%.

To get some sense of the time series of these yields, we plot the time series of the cross-sectional average yield in Figure 2. Some patterns are fairly common across countries. For example, in most countries, the general trend in yields appears to be declining across the sample, which is commensurate with the yields in developed countries, such as the United States. Yields exhibit a sharp upward spike at the end of 2008, lasting until mid- to late-2009 for each country. The average level of yields is relatively high in Brazil and Russia, and relatively low in South Korea and Singapore.

3.2 Model Estimation

3.2.1 Default Free Term Structure

As discussed in Section 2.1, we first estimate parameters of a two state variable square root process for risk free bond prices using yields on U.S. Treasury securities. Parameter estimates and standard errors are presented in Table 3. The first state variable is characterized by low mean reversion ($\kappa_1 = 0.083$), a high long-run mean ($\theta_1 = 0.966$), and relatively low volatility and prices of risk ($\sigma_1 = 0.029$, $\eta_1 = -0.006$). The second state variable has quite high mean reversion ($\kappa_2 = 0.943$), a lower long-run mean ($\theta_2 = 0.068$), and higher volatility and price of risk ($\sigma_2 = 0.083$, $\eta_2 = -0.053$).

,All parameters are statistically distinguishable from zero at conventional significance levels. The parameter a_f is set to -1 as in Duffee (1999). The rationale is largely empirical; as noted in Duffee (1999), α_f is not precisely estimable since the gradient of the likelihood function is too flat. However, it is clearly negative, and a similar issue is faced in Pearson and Sun (1994).

Despite little overlap in the time series over which the parameters are estimated, our estimates are quite similar to those presented in Duffee (1999). In contrast to his results however, our first state variable has low mean reversion and large mean, while in his study the first state variable has both high mean reversion and high long-run mean. The mean reversion in the second state variable is considerably higher than those that he estimates. Nonetheless, the state variables appear strongly linked to the level and slope of the term structure, as suggested by principal components analysis in Litterman and Scheinkmann (1991). The first state variable is 99% correlated with the 10-year constant maturity Treasury yield, and the second is approximately 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. Duffee (1999) also reports that his state variables are related to the negative of the slope of the term structure and long-term bond yields. The state variables are plotted in Figure 3, where the high correlation in the state variables and the observable proxies for the state variables is readily apparent.

The model performs reasonably well in pricing the default free securities. The patterns in pricing errors are generally similar to those reported in Duffee (1999). Pricing of short-term securities, particularly the 90-month Treasury Bill is relatively poor, with a mean (root mean square) pricing error of -9.42 (38.88) basis points. The model fares best at pricing Treasuries of intermediate maturities; mean and root mean square errors are lowest for the one-, two, five-, and seven-year Treasury yields. Since most of the bonds in our sample are intermediate- to long-term, the lack of fit on the short end of the yield curve is not of extraordinary concern for our purposes. While more precise fitting of the risk-free term structure is an important question, it is beyond the scope of this paper.

3.2.2 Hazard Rates

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 86 bonds in our sample we estimate the mean reversion coefficient, $\kappa_{i,d}$, the long-run mean, $\theta_{i,d}$, the price of risk, $\eta_{i,d}$, and the diffusion parameter, $\sigma_{i,d}$. We also estimate the parameters of the spread, α_i, $\beta_{i,1}$, and $\beta_{i,2}$.

In Table 4, 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_d (7.451 across all bonds) and θ_d (0.069 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, $\eta_d = -1.362$, is negative and suggests that investors demand compensation for default risk. The magnitude of this median parameter is considerably larger than that estimated by Duffee (1999), who finds the median price of default risk in his sample of U.S. firms is -0.235. 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 -0.545 and -0.657, 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 Brazil ($\kappa_d = 12.510$), with high median estimates also in Mexico ($\kappa_d = 8.455$), South Korea ($\kappa_d = 7.137$), and Russia ($\kappa_d = 6.847$). Again, 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 $\theta_d = 0.048$ (Singapore) to $\theta_d = 0.085$ (Brazil); Singapore exhibits the lowest 25th percentile ($\theta_d = 0.016$), and the Russian Federation the highest 75th percentile ($\theta_d = 0.118$).

Compensation for default risk appears to vary widely across the countries in the sample. For all six countries, the interquartile range of the price of risk, η_d is negative, suggesting positive compensation for default risk. The medians, however range from $\eta_d = -1.879$ in the Russian Federation and $\eta_d = -1.645$ in Brazil to $\eta_d = -1.147$ in South Korea. Variation is quite large in most countries as well; the interquartile ranges in the six countries are 1.195, 1.420, 1.606, 1.421, 1.488, and 1.124 for Brazil, Chile, Mexico, Russia, Singapore, and South Korea, 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 basis points, with a 25th percentile of 12 basis points and a 75th percentile of 38 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 basis points, with an interquartile range of 38 to 67 basis points. The model also has difficulty in pricing Brazilian bonds with a median RMSE of 32 basis points and an interquartile range of 16 to 39 basis points. In contrast, the remaining countries are better represented by the overall estimates. One issue that we can investigate in this paper is whether these pricing issues are particularly exacerbated by currency issues. However, it is quite possible that there are other country-specific sources of pricing that we are not accounting for in this model that are particularly pertinent to Russia and Brazil.

The time series of default intensities are plotted in Figure 4. In each subfigure, we plot the average across firms of the filtered default intensity for subsets of the data; averages across all firms are in Panel A and results by country are depicted in Panels B-G. The time series patterns appear reasonable; both the overall default intensity and the intensity in individual countries exhibit a sharp spike during the financial crisis of 2008-2009. There is some interesting cross-sectional variation in these increases in default intensity. Mexican default intensities increase in excess of 2000% of their pre-crisis average. Russian Federation default intensities also increase in excess of 1500% relative to pre-crisis averages; South Korean default intensities also increase in excess of 1100% of pre-crisis average. In contrast, increases in Brazil, Chile, and Singapore are somewhat more muted, at 298%, 324%, and 774% of pre-crisis averages, respectively.

In the case of the Latin American countries the smaller increase during the financial crisis relative to pre-crisis averages is partly due to relatively high default intensities in these countries prior to the crisis. Brazil and Chile have pre-crisis default intensities that are in some subperiods as high as the financial crisis default intensities. These high default rates correspond to periods in the early 2000s, as Argentina defaulted on its debt in 2002. The effects of this crisis are apparent in the overall sample in Panel A as well. In general, the default intensities appear to behave as would be expected, spiking during times of crisis, generally increasing through global economic downturns, and decreasing through global economic booms.

3.2.3 Default and Foreign Exchange

The preceding section presents estimates of default intensities that assume that the default intensity is exposed only to systematic risk in the default-free term structure and not to exchange rate risk. That is, default intensity results from default-free term structure factors and firm-specific default risk. As discussed earlier, due to financial stress caused by the need to repay debt in dollars that have potentially appreciated relative to home currency cash flows, we speculate that default intensities may also be affected by exchange rate risk. In this section, we examine the hypothesis that default intensities are related to exchange rate risks.

Our data on exchange rates are taken from Datastream, and cover the period January 3, 1994 through September 28, 2010, sampled at the daily frequency. We plot the log first difference of the exchange rates in Figure 5. The first difference plots are scaled such that they are comparable across currencies. As shown in the plots, log currency returns exhibit volatility clustering across countries. Brazilian Real, Chilean Peso, Mexican Peso, and South Korean Won currency returns exhibit particularly pronounced volatility. The graphs also clearly show the effects of the 1998 Asian currency crisis and the 2008 financial crisis. Perhaps not surprisingly, the effects of the Asian crisis are stronger in the Asian than the Latin American currencies in our sample. The plots also exhibit large currency movements associated with the Mexican Peso crisis in 1995.

To investigate the link between default intensity and exchange rate risk, we consider the possibility that default intensities are related to changes in the level of exchange rates and the conditional volatility of the exchange rate. Specifically, for each bond's estimated default intensity, we estimate the following regression,

$$h_{d,t} = a_d + b_{1,d}\Delta fx_t + b_{2,d}v_{fx,t} + e_{d,t}, \tag{27}$$

where $h_{d,t}$ is the estimated hazard rate, fx_t is log foreign exchange rate between the home currency and US Dollars, and $v_{fx,t}$ is the estimated volatility (standard deviation) of the exchange rate. We hypothesize that the coefficient $b_{1,d} < 0$ assuming that the exchange rates are specified in local currency per US Dollar. A positive innovation in the exchange rate implies that the home currency has appreciated against the dollar, making repayment in local cash flows easier, and thus reducing default intensities. In contrast, we hypothesize that $b_{2,d} > 0$; an increase in exchange rate risk (volatility) increases probabilities of crossing default thresholds.

Modeling the dynamics of exchange rates, in particular volatility, is 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 intraday data available, and instead use daily data on exchange rates from Datastream. 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. We follow their lead, but use an EGARCH volatility model, which seems to behave better in our estimation than standard GARCH or asymmetric volatility modeled in Glosten, Jagannathan, and Runkle (1993). For brevity, the results of the exchange rate estimation are not reported, but are available from the authors upon request. Volatilities implied by the model are plotted in Figure 6. Again, volatilities spike around crisis events such as the Peso crisis, Asian crisis, Argentinian default, and the global financial crisis.

Results of estimating equation (27) are reported in Table 5. We report mean coefficient estimates and associated cross-sectional t-statistics, and averages of time series regression R^2 in the table. Means and t-statistics are calculated across all countries in Panel A and within each individual country in Panels B-G. In addition to the cross-sectional means, we present a count of the number of bonds overall and within each country that have t-statistics for regression coefficients that are statistically significant. In the case of the coefficient $b_{1,d}$, we consider a coefficient statistically significant if it is smaller than -1.96, as we hypothesize that $b_{1,d} < 0$. In the case of the coefficient $b_{2,d}$, we consider a coefficient statistically significant if it is larger than 1.96, testing the hypothesis $b_{2,d} > 0$. In addition to the count of the significant t-statistics, we present the average of the time-series t-statistics across all bonds and within each country.

The main message of Table 5 is that hazard rates are strongly related to exchange rate volatility, but not to innovations in levels. Panel A presents summary results across all six countries and 86 bonds in our sample, and shows that on average, an innovation in exchange rates has a positive impact on hazard rates when not controlling for volatility. Only one bond has a statistically significant negative t-statistic in the time series. Controlling for volatility, the average coefficient remains positive, but the average t-statistic becomes negative, but insignificant. In this case, six bonds exhibit statistically significant negative coefficients in the time series. Finally, on average, innovations in exchange rates are able to explain less than 1% ($R^2 = 0.002$) of the time series of default intensities. Taken together, these results suggest little role for innovations in levels of exchange rates in determining default intensity.

In contrast to the level, the volatility of the exchange rate has generally strong, positive and statistically significant effects on hazard rates. In panel A, the results suggest that volatility in exchange rates has a positive effect on hazard rates. Point estimates are positive and statistically significant for 80 of the 86 bonds in our sample, or 93%. Finally, on average volatility explains a bit over one third (35.1%) of variation in hazard rates. Adding the innovation in the level has very little impact on the volatility coefficient, the average explanatory power, or the statistical significance of the point estimates. Both the average t-statistic (23.910 in single and 23.886 in multiple regression) and the cross-sectional t-statistic (5.331 in single and 5.345 in multiple regression) suggest that the impact of volatility on hazard rates is statistically, in addition to economically, significant.

Within-country results present much the same picture. The only example of a statistically significant estimate of level exposure in the cross-section is in the Russian Federation, where in single regression the cross-sectional t-statistic is 2.977 and the average time series t-statistic is 2.480. However, the coefficient is positive on average, and in fact none of the bonds in the Russian Federation subsample have a significantly negative t-statistic. After controlling for volatility, the average time-series t-statistic for the level exposure falls to 1.514, suggesting that on average default intensities are not exposed in the time series to level innovations in exchange rates. Finally, the R^2 of 0.7% in the single regression suggests that even if exposures are statistically significant, the economic significance of default intensity exposure to exchange rate innovations is limited. For the remaining countries, exposures to exchange rate levels explain no more than 0.2% of variation in the time series of default intensities.

Volatility exposures, in contrast, present a picture that is quite similar to that in the overall data. Average exposures are strongest in the Russian Federation, with a mean point estimate in regressions of default intensities on exchange rate volatility of 199.732 (cross-sectional t-statistic of 2.852) and lowest in Chile with a mean point estimate of 33.724 (cross-sectional t-statistic of 2.370). On average, exchange rate volatility explains between 17.9% (Chile) and 54.0% (Singapore) of time series variation in default intensities. In every country, nearly all bonds have positive and statistically significant exposures of default intensity on exchange rate volatility, and the average time series t-statistic is greater than 10 in every country.

The conclusion that we draw from these results is that dollar-denominated foreign bonds are not immune from risks in exchange rate fluctuations. An increase in the uncertainty (volatility) in exchange rates between the home country and the United States leads to an increase in the default risk embedded in the bond's price. Thus, the results suggest that investors in these bonds, although perhaps directly immunized from exchange rate risk, are bearing increased default risk induced by exchange rate fluctuations. The economic significance is quite important; as noted above, on average 35% of the variation in hazard rates, and up to 54% of the variation in hazard rates (Singapore) can be related to foreign exchange volatility.

4 Pricing Dollar-Denominated Bonds with Exchange Rate Risk

4.1 Model and Estimation Procedure

Given the evidence in the preceding section that a substantial portion of the default risk in dollar-denominated bonds is related to exchange rates, we model and test a framework for the pricing of dollar-denominated bonds in the presence of 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 physical (P) and risk-neutral measures (Q),

$$dv_{fx,t} = \kappa_v \left(\theta_v - v_{fx,t}\right) dt + \sigma_v \sqrt{v_{fx,t}} dW_{v,t}^P \tag{28}$$

$$dv_{fx,t} = \left[\kappa_v \theta_v - \left(\kappa_v + \eta_v\right) v_{fx,t}\right] dt + \sigma_v \sqrt{v_{fx,t}} dW_{v,t}^Q. \tag{29}$$

We use the volatility series estimated using an EGARCH(1,1) model above for each exchange rate as observations 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. Specifically, equation (2) becomes

$$R_{d,t} = r_t + (1-\delta)\,\lambda_{d,t} + \beta_{d,3} v_{fx,t}, \tag{30}$$

where r_t represents the risk free rate, δ is the rate of recovery given default, and $v_{fx,t}$ is the exchange rate volatility defined above. In our expression for the risky yield, $\beta_{d,3}$ represents an issue-specific sensitivity to exchange rate volatility. We model this sensitivity, rather than the sensitivity multiplied by a price of foreign exchange risk, in order to simplify our estimation of the parameters of the risky yield and exchange rates.[5]

Under this specification, the price of a zero-coupon security with face value c becomes

$$P_d(\tau) = c \exp\left(\bar{A}(\tau) + B_1^*(\tau)\, s_{1,t}^* + B_2^*(\tau)\, s_{2,t}^* + B_d(\tau)\, h_{d,t} + B_v^*(\tau)\, v_{fx,t}^*\right), \tag{31}$$

where $v_{fx,t}^*$ satisfies the stochastic differential equation

$$dv_{fx,t}^* = \left(\kappa_v \beta_{d,3} \theta_v - \left(\kappa_v + \eta_v\right) v_{fx,t}^*\right) dt + \sigma_v \sqrt{\beta_{d,3} v_{fx,t}^*} dW_{fx,t}^Q. \tag{32}$$

The constant in the bond price expression, $\bar{A}(\tau)$ is a composite of the constants for each element's price. Specifically, we can express

$$\bar{A}(\tau) = A^*(\tau) + A_d(\tau) + A_v(\tau) - \left(a_f \tau + a_d \tau + \beta_{d,1} \bar{s}_1 + \beta_{d,2} \bar{s}_2\right) \tau, \tag{33}$$

where $A^*(\tau)$ and $A_d(\tau)$ are given in expressions (??) and (??) above. Given independence of the Brownian motions, the solutions for parameters related to the risk-free rate and default probabilities are the same as before. Solutions for the coefficients related to exchange rate volatility are given

[5] Specifically, we would have to impose a constant price of risk for exchange rate risk across securities in our estimation in order to separate sensitivity from the price of risk. Imposing this cross-sectional constraint is beyond the scope of the estimation pursued in this paper.

by

$$B_v^*(\tau) = -\frac{2\left(e^{\gamma_v^*\tau}-1\right)}{2\gamma_v^*+(\kappa_v+\eta_v+\gamma_v^*)\left(e^{\gamma_v^*\tau}-1\right)} \tag{34}$$

$$A_v^*(\tau) = \frac{2\kappa_v\theta_v^*}{(\sigma_v^*)^2}\ln\left[\frac{2\gamma_v^* e^{\frac{1}{2}(\kappa_v+\eta_v+\gamma_v^*)\tau}}{2\gamma_v^*+(\kappa_v+\eta_v+\gamma_v^*)\left(e^{\gamma_v^*\tau}-1\right)}\right], \tag{35}$$

with $\gamma_v^* = \sqrt{(\kappa_v+\eta_v)^2+2(\sigma_v^*)^2}$.

Our estimation procedure is mostly unchanged from that discussed in Section 2.3. Since the default-free term structure is modeled using Brownian motions that are independent of the defaultable term structure, the parameter estimates and state variables for the risk-free term structure are unchanged from those reported in section 3.2.1, and the transition equation for the latent hazard rate variable is given in equation (24). In order to identify parameters, we assume that the latent hazard variable, $h_{d,t}$ is independent of the exchange rate volatility. The difference in estimation lies in the measurement equation, equation (23), which becomes

$$P_{d,t}(\tau,c) = c\sum_{m=1}^{\tau} P_d^*(m) + P_d^*(\tau) + u_{d,t}^*, \tag{36}$$

where $P_d^*(\tau)$ is the price of a τ-maturity zero coupon bond given the price equation (31).

4.2 Estimation with Exchange Rate Volatility

We take the default-free term structure estimates as given, and re-estimate the model incorporating exchange rate volatility as discussed in Section 4.1. Parameter estimates are presented in Table 6. As in Table 4, when we do not account for foreign exchange rate volatility, we present the median and the interquartile range (first and third quartiles of the parameter estimates). Results are reported for each country individually and across all countries collectively.

Our particular interest in the table is the parameter β_3, which represents the sensitivity of a bond issue's yield to exchange rate volatility. As shown in the table, the median parameter estimate is positive, with a point estimate of 0.163. This estimate suggests that increases in exchange rate volatiltiy lead to increases in yield spreads; a 1% increase in volatility implies a 0.163% increase in the yield. The median sensitivity varies considerably across countries; bonds in Chile and South Korea have median sensitivities that are slightly negative and close to zero (-0.022 and -0.002, respectively), while Brazil and Singapore have relatively high median sensitivities (0.746 and 0.546, respectively). These estimates indicate that there are important cross-sectional differences in the importance of foreign exchange risk for dollar-denominated bonds. For the full sample of bonds, most bonds within the interquartile range exhibit positive sensitivities to exchange rate volatiltiy; the sensitivities range from -0.020 (25th percentile) to 0.632 (75th percentile). The medians by country reported above suggest that the median Russian Federation bond is in the top quartile of sensitivity, whereas the median Chilean bond is in the first quartile of sensitivity.

We report median, 25th, and 75th percentiles of root mean square error across bonds in Panel H. Comparing these results to those in Table 4, we observe that for the full sample, there is moderate improvement in the pricing of the bonds in our sample. The median estimate of RMSE drops to 16 basis points from 18 basis points, and the interquartile range narrows from a range of 12 to 38 basis points to a range of 11 to 32 basis points. Thus, the results suggest that the impact of incorporating foreign exchange volatility exposure is greatest for bonds in the extremes of the distribution of pricing errors. In general, the results produce a leftward shift in the distribution of pricing errors. Particularly notable shifts include a leftward shift in the median from 32 basis points to 19 basis points in Brazil.

While the improvement in pricing errors is modest, it is not especially surprising. Although we allow bond yields to exhibit exposure to currency volatility, we do not explicitly include a price of currency volatility risk in the bond pricing model. Since this risk is common to bonds at the country level at a minimum, the model would require a cross-sectionally constant compensation for exchange rate volatility risk. This would require cross-equation restrictions in estimation that are difficult to impose. However, we acknowledge that a richer model of dollar-denominated emerging market bonds would include compensation for this risk, and we conjecture that doing so may improve the pricing performance of the model considerably.

The other noteworthy parameter in this estimation is the estimated price of default risk, η_d. In Table 4, we report a median estimate of -1.362 across all bonds. Accounting for exchange rate volatility exposures, the median price of risk is now -1.676. The range of estimated prices of risk also shifts substantially; the 25th percentile falls from -2.135 to -2.630 and the 75th percentile falls from -0.846 to -0.670. These estimates suggest that failing to account for exchange rate volatility exposure in pricing dollar-denominated emerging market bonds results in underestimation of the price that investors demand to compensate for default risk.

4.3 Discussion of the Results and Implications

The goal of this paper is to understand whether foreign exchange risk affects the pricing of emerging-market dollar-denominated corporate debt. As stated in the introduction, a primary motive for purchasing dollar-denominated debt is to reduce exchange rate risk exposure. Our empirical analysis yields the following pieces of evidence relative to this question:

1. The pricing errors in a standard model of default risk are larger than those estimated in a context in which U.S. firms issue U.S. dollar-denominated corporate debt. This result suggests that the standard model fails to capture some critical variables that investors use to price dollar-denominated emerging market debt.

2. Hazard rates implied by the standard model of default risk exhibit strong exposure to exchange rate volatility. In general, hazard rates appear to be positively affected by exchange rate volatility. Hence, issuing in dollars is not removing exchange rate exposure; exchange rate exposure is being realized through increased hazard rates.

3. A model accounting for volatility in exchange rates suggests that emerging market dollar-denominated corporate yields have positive sensitivities to exchange rate volatility. Thus, the price of these bonds are not immune to exchange rate risk.

4. Incorporating foreign exchange rate risk somewhat improves the pricing of dollar-denominated emerging market bonds. This results suggests that investors are demanding compensation for risks associated with foreign exchange that are induced by the dollarization of the bond payments.

In sum, the results suggest that dollarization of emerging market bond prices does not eliminate foreign exchange risk.

The question that naturally emerges from these results is why do investors demand dollar-denominated emerging market corporate bonds? Our results suggest that by investing in these bonds, investors are attempting to trade foreign exchange risk for default risk and, since default risk is impacted by foreign exchange risk, remaining exposed to risk inherent in foreign exchange. It is difficult from our analysis to assess the relative magnitudes of these risks. It is possible that investors receive a net benefit; that the increase in default risk is more than offset by the reduction in exchange rate risk. However, the reduction is clearly not complete; some risk remains. As a result, it is natural to ask whether these types of bonds are the best mechanism for hedging exchange rate risk induced by holding an emerging market corporate bond. Development in financial markets has resulted in the availability of other products for hedging currency risk, including options, futures, and swaps. An open question is whether an investor would be better off purchasing a local currency-denominated bond in combination with one of these hedging contracts than purchasing the dollar-denominated bond.

Our results also suggest that there is significant room for improvement in the modeling of prices of emerging market corporate bonds. While our simple approach to modeling exchange rate risk generally reduces pricing errors, the errors remain quite large. Root mean squared errors of our model are larger than those implied by the parameter estimates in Duffee (1999) for U.S. corporate bonds. Some of this magnitude may simply be due to the larger credit risk inherent in these bonds; Duffee (1999) shows that pricing errors increase in credit risk rating. However, a limitation of our approach is that our model does not directly incorporate a price of foreign exchange volatility risk, which may inhibit our ability to improve pricing further. As discussed above, a difficulty in imposing such a price of risk is that it must be common across all bonds. Nonetheless, incorporating such a restriction may provide further insights into the pricing of dollar-denominated emerging market securities.

5 Conclusion

The rationale frequently given for the prevalence of emerging market corporate bonds issued in major currencies is that investors demand these bonds to hedge against exchange rate risk. However, as noted in theoretical work, a company issuing major currency-denominated debt when operations are denominated in local currencies faces increased default risk. The reason for this increase in default probability is that debt denomination in dollars poses a dual problem for issuers. Revenues denominated in local currencies are likely to fall due to weakened economic conditions at the times when the local currency depreciates against the dollar. However, liabilities continue to be dominated in the relatively strong dollar, resulting in an increase in default risk. By demanding dollar-denominated debt, investors remove a natural hedge from companies and trade risk of currency fluctuations for increased risk of default.

We fit a model of risky bond prices to the yields of corporate bonds in seven emerging market countries and retrieve implied hazard rates. Our results document that the hazard rates implied by the model are sensitive to foreign exchange volatility, suggesting that while investors may have directly hedged their foreign currency exposure, their claims are indirectly exposed to foreign currency risk through default risk. A one percent increase in volatility of exchange rates between the dollar and the home country results in, at the median, a 0.16% increase in the hazard rate and thus a 16 basis point increase in the required yield on emerging market dollar-denominated corporate bonds. When we incorporate compensation for foreign exchange volatility risk into a reduced-form risky bond pricing framework as in Duffie and Singleton (1999), we find that the risk is incorporated in bond yields and that the fit of the model improves modestly.

Our results have important implications for companies and investors. Perhaps the most important is that dollarization does not immunize investors from foreign exchange risk. The foreign exchange risk is borne in an alternative form, through increased default risk. This implication raises significant questions for investors in dollar-denominated emerging currency bonds. Are the yields that the investor receives adequately compensating the investor for the increased default risk? Could welfare be improved through greater issuance of local currency debt and alternative foreign exchange hedging mechanisms? Answering these questions is important in furthering our understanding of both asset prices and implications of capital market development for emerging nation economies.

References

Andersen, Torben G, and Tim Bollerslev, 1998, Dm-dollar volatility: intraday activity patterns, macroeconomic announcements and longer-run dependencies, *Journal of Finance* 53, 219–265.

Baillie, Richard T, and Tim Bollerslev, 1989, The message in daily exchange rates: A conditional variance tale, *Journal of Business and Economic Statistics* 7, 297–305.

Caballero, Ricardo, and Arvind Krishnamurthy, 2003, Excessive dollar debt: Financial development and underinsurance, *Journal of Finance* 58, 867–894.

Calvo, Guillermo, 2001, Capital markets and the exchange rate: with special reference to the dollarization debate in latin america, *Journal of Money, Credit, and Banking* 33, 312–334.

Cox, John C, Jonathan Ingersoll, and Stephen A Ross, 1985, A theory of the term structure of interest rates, *Econometrica* 53, 385–408.

Dai, Qiang, and Kenneth J Singleton, 2000, Specification analysis of affine term structure models, *Journal of Finance* 55, 1943–1978.

Duffee, Gregory R, 1998, The relation between treasury yields and corporate bond spreads, *Journal of Finance* 53, 2225–2241.

——— , 1999, Estimating the price of default risk, *Review of Financial Studies* 12, 197–226.

Duffie, Darrell, and Rui Kan, 1996, A yield-factor model of interest rates, *Mathematical Finance* 6, 379–406.

Duffie, Darrell, and Kenneth J Singleton, 1997, An econometric model of the term structure of interest-rate swap yields, *Journal of Finance* 52, 1287–1321.

——— , 1999, Modeling term structures of defaultable bonds, *Review of Financial Studies* 12, 687–720.

Eichengreen, Barry, and Ricardo Hausmann, 2004, The pain of original sin, in *Other People's Money: Debt Denomination and Financial Instability in Emerging Market Economics* (University of Chicago Press).

Fisher, Irving, 1933, The debt-deflation theory of great depressions, *Econometrica* 1, 337–357.

Glosten, Lawrence R, Ravi Jagannathan, and David E Runkle, 1993, On the relation between the expected value and the volatility of the nominal excess return on stocks, *Journal of Finance* 48, 1779–1801.

Jeanne, Olivier, 2004, Why do emerging economies borrow in foreign currency, in *Other People's Money: Debt Denomination and Financial Instability in Emerging Market Economics* (University of Chicago Press).

Kang, Johnny, and Carolin E Pflueger, 2011, Inflation risk in corporate bonds, unpublished manuscript, Harvard University.

Kedia, Simi, and Abon Mozumadar, 2003, Foreign currency-denominated debt: An empirical examination, *Journal of Business* 76, 521–546.

Korinek, Anton, 2010, Excessive dollar borrowing in emerging markets: Balance sheet effects and macroeconomic externalities, unpublished manuscript, University of Maryland.

Lesmond, David, Joseph Ogden, and Charles Trzcinka, 1999, A new estimate of transaction costs, *Review of Financial Studies* 12.

Litterman, Robert, and José Scheinkmann, 1991, Common factors affecting bond returns, *Journal of Fixed Income* pp. 54–61.

Longstaff, Francis A, and Eduardo S Schwartz, 1995, A simple approach to valuing risky fixed and floating rate debt, *Journal of Finance* 50, 789–819.

Pearson, Neal, and T S Sun, 1994, Exploiting the conditional density in estimating the term structure: An application to the cox, ingersoll, and ross model, *Journal of Finance* 49, 1279–1304.

Table 1: Summary Statistics: U.S. Treasury Constant Maturity Yields

Table 1 presents summary statistics for the yields on constant maturity Treasuries for 3 months, 6 months, 1 year, 2 years, 5 years, 7 years, 10 years, and 20 years to maturity. Data are sampled on a daily frequency over the period January 3, 1994 through September 28, 2010. Constant maturity yields are obtained from the Federal Reserve Report H.15. The table reports the mean, standard deviation, minimum and maximum of each yield over the sample period.

Maturity:	3 Mos	6 Mos	1 Yr	2 Yr	5 Yr	7 Yr	10 Yr	20 Yr
Mean	3.45	3.62	3.76	4.05	4.59	4.86	5.04	5.57
Std	1.98	2.00	1.97	1.90	1.52	1.37	1.21	1.06
Min	0.00	0.13	0.24	0.37	1.25	1.59	2.08	2.86
Max	6.42	6.67	7.32	7.74	7.90	7.92	8.05	8.30

Table 2: Summary Statistics: Emerging Market Corporate Bonds

Table 2 presents summary statistics for data on emerging market corporate bonds. We sample corporate bonds with payments denominated in dollars from Datastream over the period January 21, 1997 through September 28, 2010 at the daily frequency. We limit attention to bonds that have price changes on 75% of their trading days to capture more liquid bonds. We also include only straight bonds with fixed maturity and coupon. This results in a sample of 86 bonds across six countries; Brazil (BR), Chile (CL), Mexico (MX), Russia (RS), Singapore (SG), and South Korea (SK). For each country, we report the number of bonds sampled, the number of companies represented, and the average, minimum, and maximum of both the coupon rate and maturity of the bonds in years when issued. In addition, we calculate the time series of average yields across bonds in each country, and report the mean, standard deviation, minimum, and maximum of the time series averages by country . Finally, we report the first observation available for each country.

Country	BR	CL	MX	RS	SG	SK
Bonds	12	14	11	17	13	19
Companies	7	8	3	8	5	8
Avg. Coupon	8.28	7.00	5.68	8.28	6.09	5.84
Min. Coupon	6.25	5.13	4.75	5.67	5.00	4.25
Max. Coupon	10.50	8.63	6.63	10.00	7.38	8.75
Avg. Maturity	12.96	15.64	13.47	7.00	11.70	8.47
Min. Maturity	5.00	9.50	5.00	5.00	5.00	5.00
Max. Maturity	30.00	30.00	30.00	10.00	30.00	20.00
Mean Average Yield	8.16	6.26	5.64	8.86	5.53	5.71
Std. Average Yield	1.73	0.73	0.71	3.69	1.06	1.26
Min. Average Yield	5.55	4.95	3.67	4.93	2.59	3.03
Max. Average Yield	15.32	8.68	8.30	28.00	9.78	10.61
First Observation	1/14/04	12/28/00	1/7/05	4/16/04	11/19/01	7/24/01

Table 3: Parameter Estimates for Risk-Free Term Structure

Table 3 presents parameter estimates for a two-factor square root process latent term structure model. Parameters are estimated via the Kalman Filter, utilizing the following measurement and transition equations:

$$
\begin{aligned}
\mathbf{Y}_t(\boldsymbol{\tau}) &= a_{f}\iota - \frac{1}{\boldsymbol{\tau}}\left(\mathbf{A}(\boldsymbol{\tau}) + \mathbf{B}'(\boldsymbol{\tau})\,\mathbf{s}_t\right) + \mathbf{u}_t \\
\begin{pmatrix} s_{1,t} \\ s_{2,t} \end{pmatrix} &= \begin{pmatrix} \theta_1\left(1-e^{-\kappa_1}\right) \\ \theta_2\left(1-e^{-\kappa_2}\right) \end{pmatrix} + \begin{pmatrix} e^{-\kappa_1} & 0 \\ 0 & e^{-\kappa_2} \end{pmatrix} \begin{pmatrix} s_{1,t-1} \\ s_{2,t-1} \end{pmatrix} + \begin{pmatrix} w_{1,t} \\ w_{2,t} \end{pmatrix},
\end{aligned}
$$

and

$$
\begin{aligned}
\mathbf{w}_t &\sim \mathcal{N}\left(\mathbf{0}, \begin{pmatrix} Q_{1,t} & 0 \\ 0 & Q_{2,t} \end{pmatrix}\right) \\
Q_{i,t} &= s_{i,t}\frac{\sigma_i^2}{\kappa_i}\left(e^{-\kappa_i} - e^{-2\kappa_i}\right) + \theta_i \frac{\sigma_i^2}{2\kappa_i}\left(1-e^{-\kappa_i}\right)^2.
\end{aligned}
$$

$\mathbf{Y}_t(\boldsymbol{\tau})$ is a vector of yields observed at time t, $\mathbf{A}(\boldsymbol{\tau})$ and $\mathbf{B}'(\boldsymbol{\tau})$ are a vector and matrix, respectively, of coefficients as a function of the parameters of the model, and $\mathbf{u}_t$ is a vector of mean-zero error terms. The parameters of the model, $\boldsymbol{\theta}$, $\boldsymbol{\kappa}$, $\boldsymbol{\eta}$ and $\boldsymbol{\sigma}$ govern the dynamics of the state variables, $\mathbf{s}_t$. We utilize constant maturity Treasury yields with maturities $\boldsymbol{\tau}' = \{0.25, 0.50, 1.00, 2.00, 5.00, 7.00, 10.00, 20.00\}$ obtained from the H.15 report of the Federal reserve to estimate the parameters. Panel A reports parameter estimates and standard errors; Panel B reports pricing errors in basis points. The first row is the mean error and the second row is the root mean square error (RMSE). Data are sampled at the daily frequency over the period January 3, 1994 through September 28, 2010, for 4,190 time series observations.

Panel A: Parameter Estimates

Parameter	Estimate	Parameter	Estimate
κ_1	0.083	κ_2	0.943
SE	(0.012)	SE	(0.073)
θ_1	0.966	θ_2	0.068
SE	(0.024)	SE	(0.002)
η_1	-0.006	η_2	-0.053
SE	(0.002)	SE	(0.070)
σ_1	0.029	σ_2	0.083
SE	(0.003)	SE	(0.007)

Panel B: Pricing Errors

	3 Mo	6 Mo	1 Yr	2 Yr	5 Yr	7 Yr	10 Yr	20 Yr
Mean Error	-9.42	-0.44	0.01	4.44	0.41	-0.11	-10.79	2.85
RMSE	38.88	18.52	5.11	19.11	10.83	6.88	17.05	27.03

Table 4: Parameter Estimates of Risky Bond Prices

Table 4 presents parameter estimates of the defaultable component of bond prices as modeled in equation (18),

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

where the risk-neutral defaultable yield, R_s, is specified as

$$R_s = a_d + h_{d,t} + \boldsymbol{\beta}' (\mathbf{s}_t - \bar{\mathbf{s}}),$$

$\mathbf{s}_t = \{s_{1,t}, s_{2,t}\}$ are the state variables implied by parameter estimates from Table 3, and $h_{d,t}$ is the hazard rate, which follows the stochastic differential equation

$$dh_{d,t} = (\kappa_d \theta_d - (\kappa_d + \eta_d))\, dt + \sigma_d \sqrt{h_{d,t}} dW_d^Q, t.$$

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_d	2.643	7.451	11.972
θ_d	0.037	0.069	0.102
η_d	-2.135	-1.362	-0.846
σ_d	0.374	0.921	1.492
β_1	-0.819	-0.545	-0.254
β_2	-1.475	-0.657	0.088
a_d	-0.114	-0.080	-0.044

Panel B: Brazil

Pct	25	50	75
κ_d	7.857	12.510	13.989
θ_d	0.053	0.085	0.115
η_d	-2.477	-1.645	-1.282
σ_d	1.278	1.484	1.636
β_1	-1.330	-0.819	-0.560
β_2	-1.761	-1.491	-0.859
a_d	-0.131	-0.077	-0.058

Panel C: Chile

Pct	25	50	75
κ_d	3.342	5.413	11.643
θ_d	0.029	0.049	0.090
η_d	-1.924	-1.307	-0.504
σ_d	0.396	0.717	1.364
β_1	-0.574	-0.478	-0.024
β_2	-1.955	-0.797	0.263
a_d	-0.111	-0.058	-0.037

Panel D: Mexico

Pct	25	50	75
κ_d	0.692	8.455	9.675
θ_d	0.035	0.055	0.107
η_d	-1.890	-1.270	-0.284
σ_d	0.229	0.871	0.992
β_1	-0.737	-0.538	-0.331
β_2	-1.483	-0.667	-0.360
a_d	-0.139	-0.083	-0.044

Table continued on the next page.

Panel E: Russia

Pct:	25	50	75
κ_d	2.643	6.847	9.962
θ_d	0.048	0.080	0.118
η_d	-2.671	-1.879	-1.250
σ_d	0.450	1.050	2.116
β_1	-1.169	-1.027	-0.611
β_2	-1.430	-0.887	0.498
a_d	-0.145	-0.054	-0.030

Panel F: Singapore

Pct	25	50	75
κ_d	0.678	3.786	10.240
θ_d	0.016	0.048	0.066
η_d	-2.443	-1.241	-0.955
σ_d	0.249	0.467	1.091
β_1	-0.498	-0.394	-0.165
β_2	-0.714	-0.541	-0.253
a_d	-0.092	-0.057	-0.050

Panel G: South Korea

Pct	25	50	75
κ_d	2.379	7.137	14.009
θ_d	0.055	0.072	0.097
η_d	-1.618	-1.147	-0.494
σ_d	0.433	0.667	1.485
β_1	–0.642	-0.291	-0.009
β_2	-1.036	-0.251	0.784
a_d	-0.114	-0.092	-0.070

Panel H: Pricing Errors

Country	25	50	75
All	12.070	18.335	38.020
Brazil	15.710	32.415	38.945
Chile	11.070	14.110	22.510
Mexico	8.510	11.600	15.140
Russia	37.740	50.920	66.530
Singapore	9.780	11.120	19.440
South Korea	13.390	17.760	24.710

Table 5: Regressions of Default Intensities on Exchange Rates Levels and Volatilities

Table 5 presents regressions of the sensitivity of default intensities on the level and volatility of the log spot exchange rate between the the U.S. and the home country. Regressions are specified as

$$h_{d,t} = a_d + b_{1,d}\Delta fx_t + b_{2,d}v_{fx,t} + e_{d,t},$$

where $h_{d,t}$ represents the default intensity as filtered using parameter estimates in Table 4, fx_t represents the log exchange rate, and $v_{fx,t}$ represents the volatility from an EGARCH(1,1) model. Regressions are estimated for 100 bonds. We present the mean estimates across all countries in Panel A, Brazil in Panel B, Chile in Panel C, Mexico in Panel D, Russia in Panel E, Singapore in Panel F, and South Korea in Panel G. Reported standard errors are calculated as the cross-sectional standard deviation of the parameter estimate, scaled by the square root of the number of observations. The proportion of coefficients with 5% critical value statistical positive significance for changes in the log level and negative significance for volatility are shown in parentheses below the standard errors. The column R^2 is the average R^2 across the regressions.

Panel A: All Countries

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.048	0.169		0.002
SE	(0.011)	(0.036)		
Frac Sig.		(0.011)		
Coeff.	0.008		6.593	0.350
SE	(0.003)		(1.528)	
Frac Sig.			(0.955)	
Coeff.	0.008	-0.013	6.591	0.351
SE	(0.003)	(0.044)	(1.530)	
Frac Sig.		(0.159)	(0.955)	

Panel B: Brazil

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.055	0.047		0.002
SE	(0.021)	(0.033)		
Frac Sig.		(0.077)		
Coeff.	0.008		4.863	0.366
SE	(0.002)		(2.146)	
Frac Sig.			(0.923)	
Coeff.	0.008	0.032	4.862	0.367
SE	(0.002)	(0.028)	(2.146)	
Frac Sig.		(0.077)	(0.923)	

Panel C: Chile

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.031	0.002		0.000
SE	(0.010)	(0.023)		
Frac Sig.		(0.000)		
Coeff.	0.017		2.131	0.214
SE	(0.009)		(0.974)	
Frac Sig.			(0.929)	
Coeff.	0.017	-0.084	2.145	0.216
SE	(0.009)	(0.052)	(0.981)	
Frac Sig.		(0.571)	(0.929)	

Panel D: Mexico

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.037	0.073		0.000
SE	(0.013)	(0.040)		
Frac Sig.		(0.000)		
Coeff.	0.010		4.000	0.208
SE	(0.004)		(1.555)	
Frac Sig.			(0.917)	
Coeff.	0.010	-0.011	4.001	0.208
SE	(0.004)	(0.022)	(1.555)	
Frac Sig.		(0.000)	(0.917)	

Table continued on next page.

Panel E: Russia

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.056	0.583		0.007
SE	(0.011)	(0.115)		
Frac Sig.		(0.000)		
Coeff.	0.022		5.708	0.302
SE	(0.006)		(1.165)	
Frac Sig.			(0.941)	
Coeff.	0.0022	0.264	5.664	0.303
SE	(0.004)	(0.072)	(1.555)	
Frac Sig.		(0.000)	(0.941)	

Panel F: Singapore

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.028	0.253		0.001
SE	(0.008)	(0.097)		
Frac Sig.		(0.000)		
Coeff.	-0.025		15.010	0.567
SE	(0.010)		(4.669)	
Frac Sig.			(1.000)	
Coeff. -0.025	-0.017	15.012	0.567	
SE	(0.010)	(0.029)	(4.668)	
Frac Sig.		(0.000)	(1.000)	

Panel G: South Korea

	Intercept	FX	σ_{FX}	R^2
Coeff.	0.068	0.005		0.001
SE	(0.047)	(0.017)		
Frac Sig.		(0.000)		
Coeff.	0.011		7.646	0.425
SE	(0.006)		(5.733)	
Frac Sig.			(1.000)	
Coeff.	0.011	-0.238	7.661	0.427
SE	(0.005)	(0.171)	(5.744)	
Frac Sig.		(0.263)	(1.000)	

Table 6: Parameter Estimates of Risky Bond Prices with Foreign Exchange Risk

Table 6 presents parameter estimates of the defaultable component of bond prices as affected by exchange rate risk,

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

where the risk-neutral defaultable yield, R_s^*, is specified as

$$R_s = a_d + h_{d,t} + \boldsymbol{\beta}'\left(\mathbf{s}_t - \bar{\mathbf{s}}\right) + \beta_3 v_{fx,t}.$$

The variable $v_{fx,t}$ is the volatility of the log foreign exchange rate, $\mathbf{s}_t = \{s_{1,t}, s_{2,t}\}$ are the state variables implied by parameter estimates from Table 3, and $h_{d,t}$ is the hazard rate, which follows the stochastic differential equation

$$dh_{d,t} = \left(\kappa_d\theta_d - \left(\kappa_d + \eta_d\right)\right) dt + \sigma_d\sqrt{h_{d,t}} dW_d^Q, t.$$

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

$$dv_{fx,t}^* = \left(\kappa_3\beta_3\theta_v - \left(\kappa_v + \eta_v\right) v_{fx,t}^*\right) dt + \sigma_v\sqrt{\beta_3 v_{fx,t}^*} dW_{fx,t}^Q$$

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_d	1.498	4.523	9.262
θ_d	0.039	0.071	0.105
η_d	-2.630	-1.676	-0.670
σ_d	0.326	0.708	1.145
β_1	-0.530	-0.334	-0.106
β_2	-0.697	-0.313	0.182
β_3	-0.020	0.163	0.623
a_d	-0.110	-0.078	-0.024

Panel B: Brazil

Pct	25	50	75
κ_d	5.112	7.630	13.759
θ_d	0.056	0.087	0.095
η_d	-2.929	-2.339	-1.576
σ_d	1.010	1.536	2.104
β_1	-0.993	-0.467	-0.252
β_2	-1.309	-0.564	0.087
β_3	0.381	0.746	1.477
a_d	-0.105	-0.088	-0.073

Panel C: Chile

Pct	25	50	75
κ_d	0.799	3.094	8.636
θ_d	0.020	0.049	0.080
η_d	-1.965	-1.321	-0.574
σ_d	0.164	0.403	0.835
β_1	-0.364	-0.100	0.048
β_2	-0.970	-0.099	0.565
β_3	-0.055	-0.022	0.013
a_d	-0.094	-0.068	-0.019

Panel D: Mexico

Pct	25	50	75
κ_d	0.907	2.920	6.259
θ_d	0.017	0.057	0.112
η_d	-2.279	-1.527	-0.434
σ_d	0.226	0.405	0.761
β_1	-0.427	-0.378	-0.190
β_2	–0.882	-0.133	0.343
β_3	-0.025	0.186	0.437
a_d	-0.180	-0.104	-0.011

Table continued on the next page.

Panel E: Russia

Pct:	25	50	75
κ_d	1.498	2.744	6.705
θ_d	0.048	0.097	0.113
η_d	–3.140	-2.473	-1.408
σ_d	0.356	0.745	1.066
β_1	-0.927	-0.743	-0.444
β_2	-0.702	-0.464	0.121
β_3	0.116	0.286	1.109
a_d	-0.139	-0.102	-0.026

Panel F: Singapore

Pct	25	50	75
κ_d	0.456	3.262	9.587
θ_d	0.030	0.053	0.074
η_d	-2.200	-1.290	-0.844
σ_d	0.277	0.383	1.197
β_1	-0.333	-0.234	-0.098
β_2	–0.697	-0.371	-0.197
β_3	0.093	0.546	1.824
a_d	-0.097	-0.025	-0.006

Panel G: South Korea

Pct	25	50	75
κ_d	1.600	4.674	8.594
θ_d	0.046	0.073	0.113
η_d	-2.050	-1.276	1.213
σ_d	0.494	0.753	1.018
β_1	-0.467	-0.177	-0.009
β_2	-0.395	-0.094	0.265
β_3	-0.093	-0.002	0.352
a_d	-0.105	-0.076	-0.033

Panel H: Pricing Errors

Country	25	50	75
All	11.160	16.190	31.780
Brazil	14.835	19.295	40.395
Chile	9.350	10.795	15.980
Mexico	8.350	10.570	12.500
Russia	31.780	47.290	65.440
Singapore	9.160	11.570	16.730
South Korea	12.310	16.470	23.240

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.

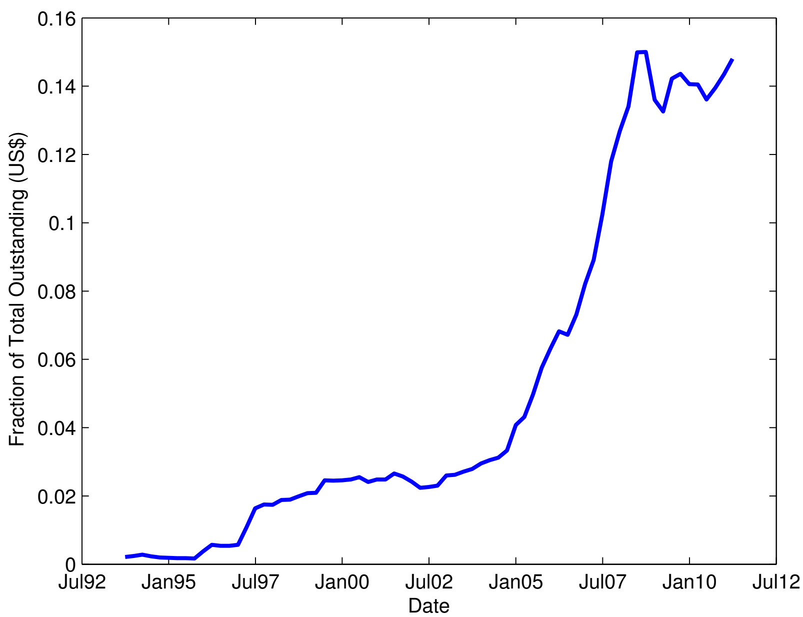
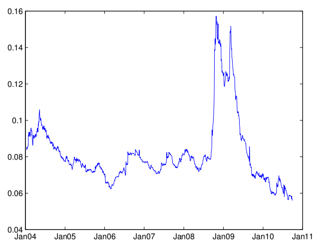
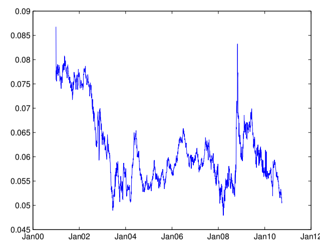


Figure 2: Emerging Market Corporate Average Yields

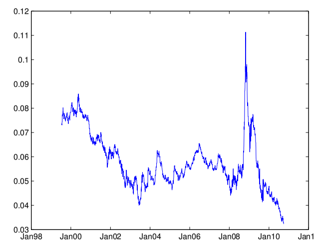
Figure 2 depicts the time series of average bond yields in each country. We sample 99 bonds in seven countries from Datastream at the daily frequency over the period January 21, 1997 through September 28, 2010. We retain bonds that have price changes on 75% of available trading days, at least 250 trading days of data, and contracts that provide for a fixed coupon and maturity. Each day, we calculate the average yield across the bonds in each country, and plot the time series in panels a-g.



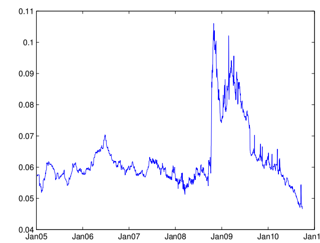
(a) Brazil



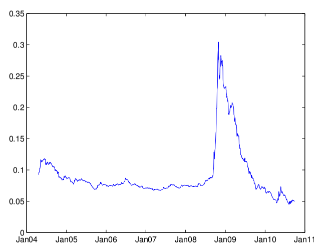
(b) Chile



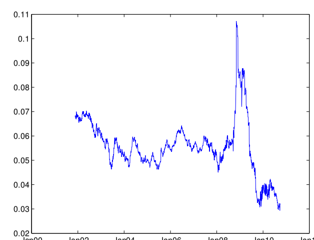
(c) Hong Kong



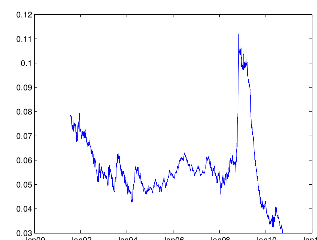
(d) Mexico



(e) Russia



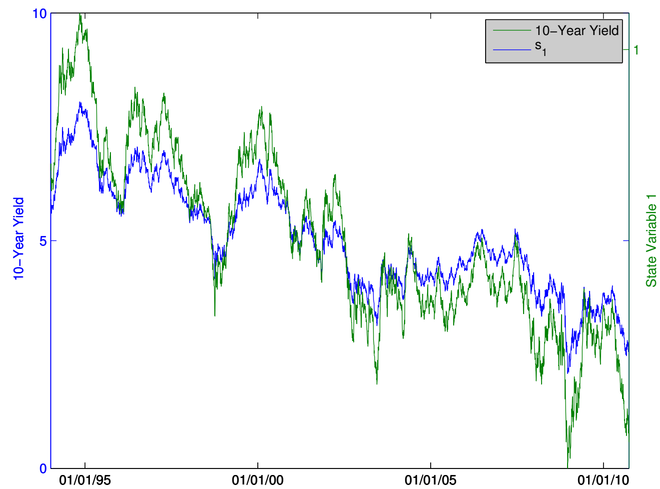
(f) Singapore



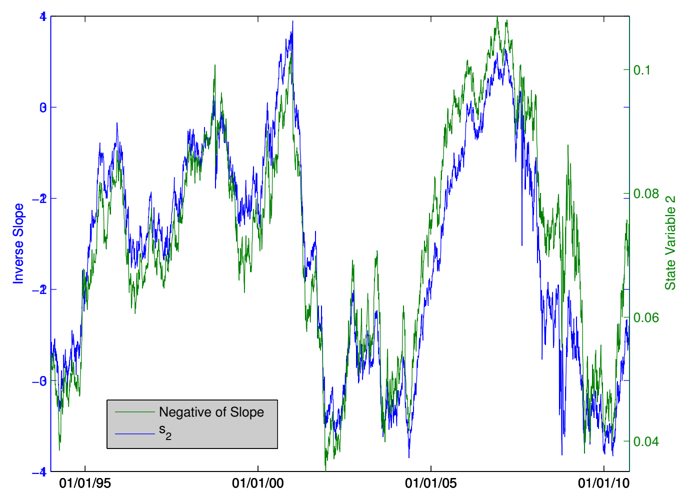
(g) South Korea

Figure 3: State Variables

Figure 3 depicts the filtered state variables retrieved from the Kalman filter estimation. Panel A plots the first state variable and the 10-year constant maturity Treasury yield. Panel B plots the negative of the slope of the yield curve, measured as the difference in a 3-month constant maturity Treasury yield and a 10-year constant maturity Treasury yield. Data are sampled on a daily frequency over the period January 3, 1994 through September 28, 2010. Constant maturity yields are obtained from the Federal Reserve Report H.15.



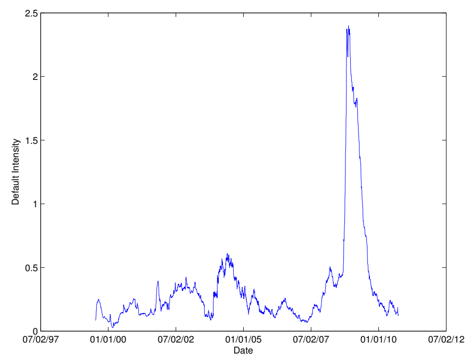
(a) Panel A: State Variable 1



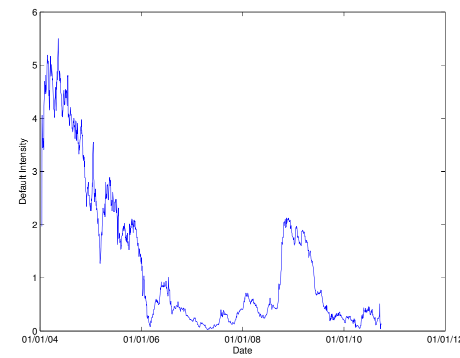
(b) Panel B: State Variable 2

Figure 4: Default Intensities

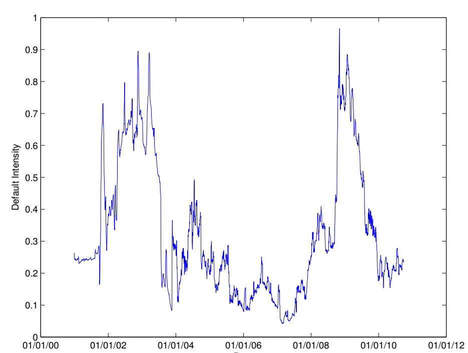
Figure 4 presents the default intensities recovered via Kalman filter from the Duffee (1999) and Duffie and Singleton (1999) models of bond pricing with default risk in Section 3. The panels plot the average of retrieved hazard rates across all countries and within countries across bonds. Panel A depicts the time series for all countries and bonds, Panel B for Brazilian bonds, Panel C for Chilean Bonds, Panel D for Hong Kong Bonds, Panel E for Mexican bonds, Panel F for Russian bonds, Panel G for Singaporean bonds, and Panel H for South Korean bonds. Data are sampled at the daily frequency and cover ranges from January, 1994 through September, 2010.



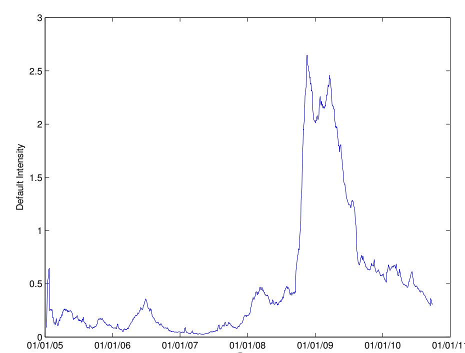
(a) Panel A: Overall



(b) Panel B: Brazil

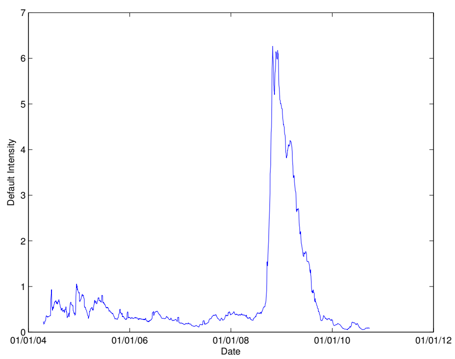


(c) Panel C: Chile

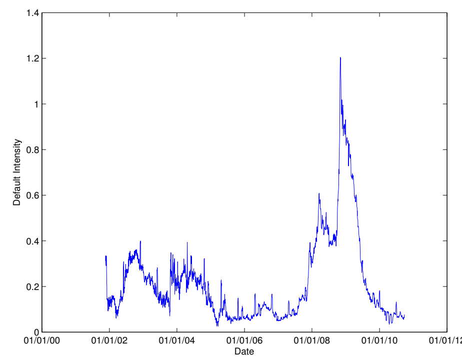


(d) Panel D: Mexico

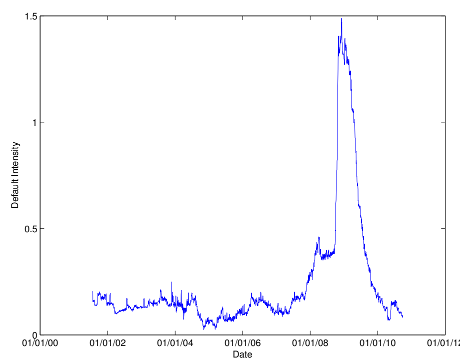
Figure continued on next page.



(e) Panel E: Russian Federation



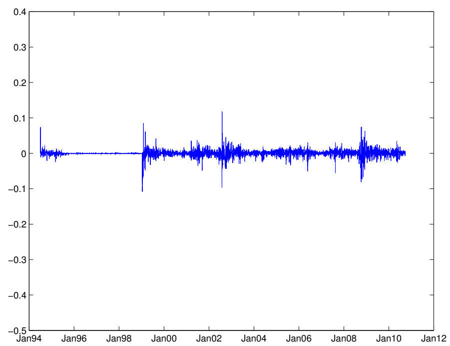
(f) Panel F: Singapore



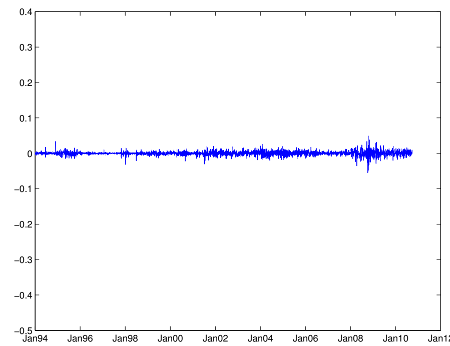
(g) Panel G: South Korea

Figure 5: Changes in Log Exchange Rates

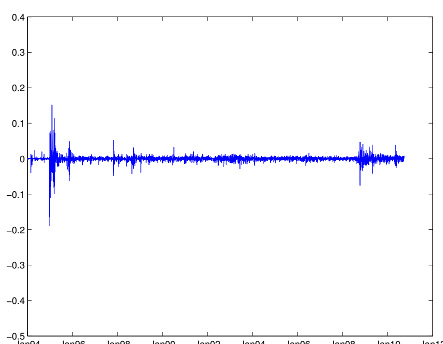
Figure 5 presents the first differences in the log exchange rate between the United States Dollar and foreign currencies. Data are sampled at the daily frequency from January, 1994 through September, 2010 and are obtained from DataStream. Panel A presents plots of the Brazilian Real, Panel B the Chilean Peso, Panel C the Hong Kong Dollar, Panel D the Mexican Peso, Panel E the Russian Ruble, Panel F the Singaporean Dollar, and Panel G the South Korean Won. All plots are scaled to the same levels.



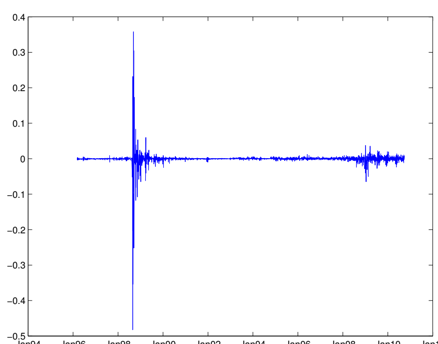
(a) Panel A: Brazil



(b) Panel B: Chile

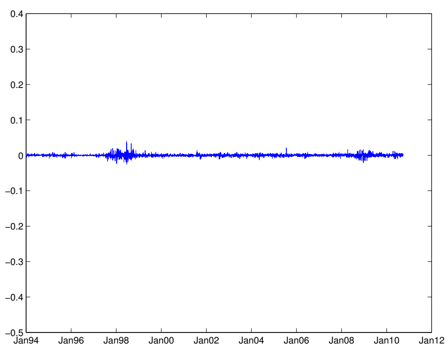


(c) Panel C: Mexico

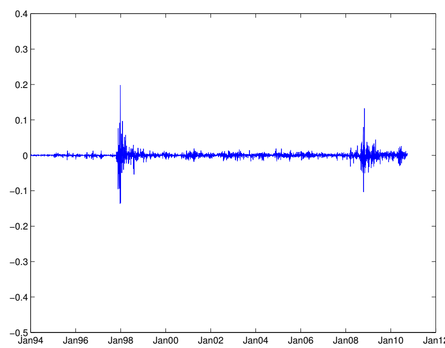


(d) Panel D: Russia

Figure continued on next page.



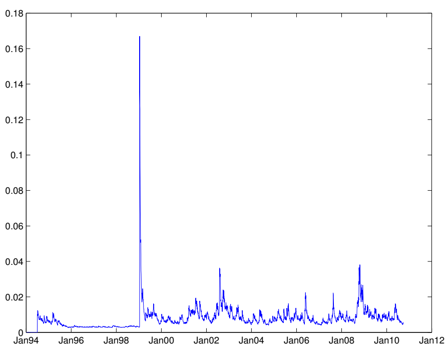
(e) Panel E: Singapore



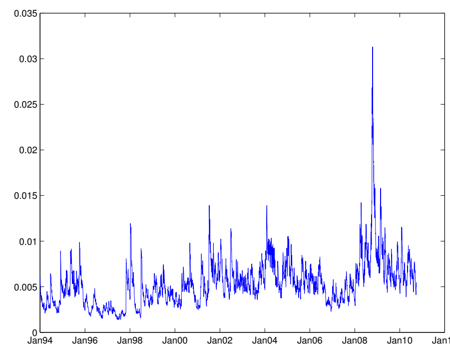
(f) Panel F: South Korea

Figure 6: Volatility in Log Exchange Rates

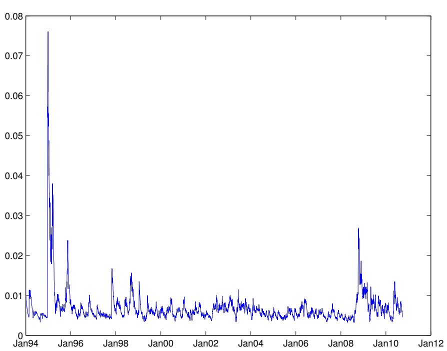
Figure 6 presents the standard deviation in the log exchange rate between the United States Dollar and foreign currencies. Standard deviations are retrieved from the estimates of an EGARCH(1,1) model with MA(1) innovations. Data are sampled at the daily frequency from January, 1994 through September, 2010 and are obtained from DataStream. Panel A presents plots of the volatility in the Brazilian Real, Panel B the Chilean Peso, Panel C the Hong Kong Dollar, Panel D the Mexican Peso, Panel E the Russian Ruble, Panel F the Singaporean Dollar, and Panel G the South Korean Won.



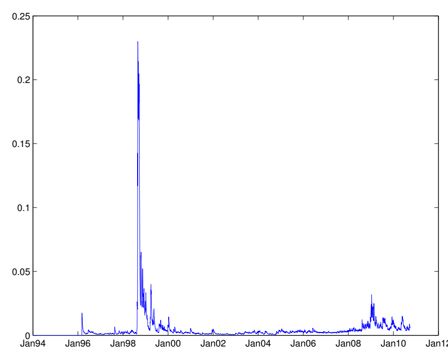
(a) Panel A: Brazil



(b) Panel B: Chile

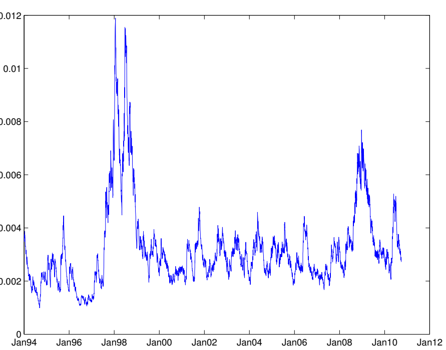


(c) Panel D: Mexico

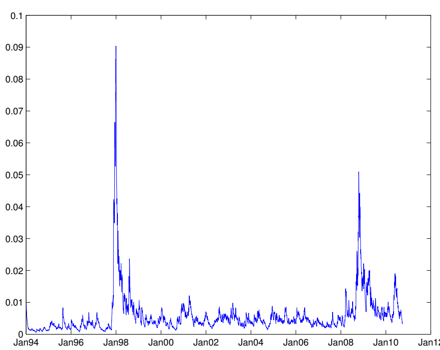


(d) Panel D: Russia

Figure continued on next page.



(e) Panel E: Singapore



(f) Panel F: South Korea