Characterizing Global Risk in Emerging Markets During Financial Crises: 1998-1999

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Abstract

Movements in the bond risk premia of nine emerging markets during the Russian, LTCM and Brazilian financial crises, are explained in terms of the risk preferences of investors. Three broad characteristics of risk are considered: Global risk factors comprising credit, liquidity and volatility risks; country risk arising from idiosyncratic shocks of countries; and contagion risk caused by additional cross-border linkages arising during financial crises. The empirical results show that all risk components are generally important in explaining the widening of spreads during the Russian and LTCM crises, whereas the Brazilian crisis is better characterized in terms of changes in global credit risk and country risk.

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1. Introduction

During February 1998 to May 1999, emerging markets experienced an enormous widening of bond spreads. Figure 1 highlights this phenomenon, showing the risk premia of nine emerging markets across three regions (Asia, Europe and Latin America) over this period. All markets show a sharp increase in spreads during the Russian crisis (August to September, 1998). Further increases in spreads occur during the Long Term Capital Management (LTCM) crisis (September to October, 1998) and the Brazilian crisis (January to February, 1999). During the interval between the LTCM and Brazilian crises spreads were broadly declining with the exception of Russia. The risk premium for Russia is sustained for most of the period at the level reached during the Russian crisis.[2]

Whilst changes in spreads between sovereign bonds provide some indication of the risk premia between countries, they do not distinguish between the risks associated purely with default (credit risk), or other market factors including liquidity and volatility risks, which are important to international investors. In addition, there are other risks which may explain movements in spreads, including risks that are either idiosyncratic to countries (country risk), or arise from the transmission of shocks across national borders during periods of financial crisis (contagion risk).[3]

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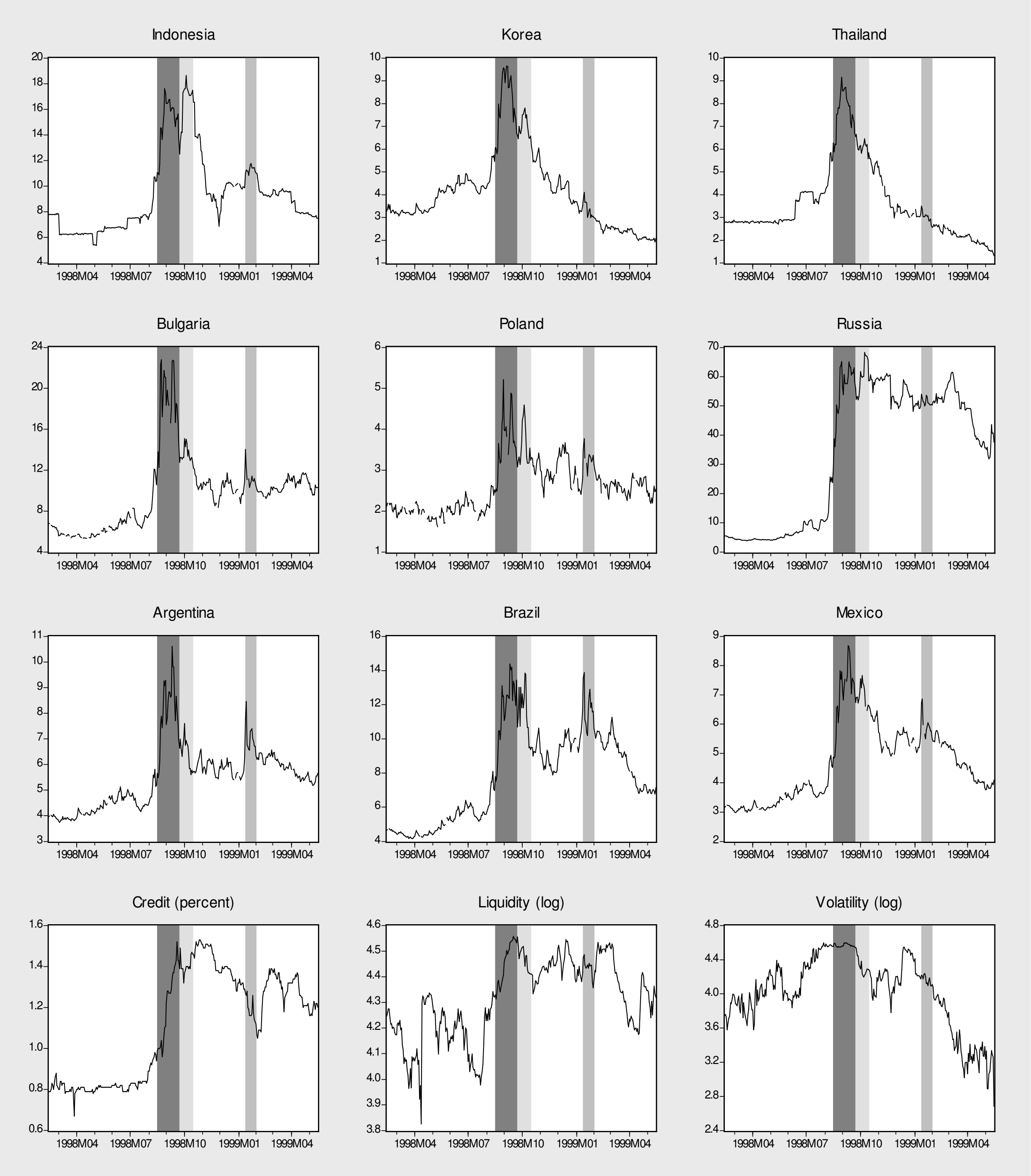
[2] See Appendix A for definitions of the variables. See Jorion (2000), Lowenstein, (2001) and Goldfajn (2003) for an overview of the Russian, LTCM and Brazilian crises, and Appendix B for a discussion of the choice of crisis dates.

[3] See Dornbusch, Park and Claessens (2000) and Pericoli and Sbracia (2003) for general definitions of contagion.

Figure 1. Risk premia and risk indicators, February 1998 to May 1999. Shaded areas from left to right refer to crisis periods: Russian crisis, August 17 to September 23, 1998; LTCM crisis, September 23 to October 15, 1998; Brazilian crisis, January 13 to February 2 1999.



The problem of identifying the underlying components of risk is compounded when the risk factors themselves are correlated. For example, a rise in credit risk can impact upon the liquidity of the market resulting in international investors offloading liquid assets, despite their relatively low default risk (Greenspan, 1999). This suggests that at any point in time more than one risk factor is at play and classifying crises in terms of different types of risk, for example the Russian crisis as a credit crisis and the LTCM crisis as a liquidity crisis, may be too simplistic.

There is a growing literature on the impact of risk preferences of investors and consumers on asset market prices. Burger and Warnock (2003) examine credit and liquidity risk for international bond portfolio allocation and find credit risk to be important. Recently, emphasis has shifted to the effects of changes in risk preferences on asset markets. Kumar and Persaud (2002) and the International Monetary Fund (2003) look at measures of risk appetite and particularly relate it to portfolio behaviour during periods of financial crises. Bekaert, Engstrom and Xing (2006) and Bekaert, Engstrom and Grenadier (2006) find a substantial role for time varying risk aversion in both equities and bonds. Bollerslev, Gibson and Zhou (2004) estimate time varying volatility risk from option prices for equities.

This paper decomposes the risk premia of nine emerging market bonds during periods of financial crises in terms of their three broad risk characteristics: global risk (credit, liquidity, volatility), country risk and contagion risk. The approach is based on developing a model of risk aversion using the stochastic discount factor model, whereby the expected risk premia is expressed as a function of the key characteristics that determine risk; see also Flood and Rose (2004) who adopt a similar modeling framework, and Cochrane and Piazzesi (2005) for an alternative approach using forward rates.

In the empirical application liquidity and volatility risks are measured using indices compiled by JP Morgan, while credit risk is measured as the spread between US industrial BBB1 10 year yields and the 10 year US Treasury bond. Time series plots of the three global risk variables are given in the bottom three panels of Figure 1. The country and contagion variables are treated as latent. Country risk is measured as the idiosyncratic shocks from a dynamic model based on a structural vector autoregression (SVAR). Contagion risk is defined as the transmission of shocks from one country to another country during a financial crisis. In keeping with the contagion literature, contagion is treated as a short-run phenomenon which dissipates in the long-run (Masson, 1999). This leads to a set of restrictions which are imposed on the long-run dynamics of the SVAR to identify contagion risk.

An important feature of the model is that it can be used to track the key risk factors underlying changes in the risk premia of bonds over time during the three crises. The empirical analysis is presented in two ways. First, risk premia are decomposed in terms of the quantity and price of risk. The risk quantities are computed directly from the long-run variance-covariance matrix of the SVAR, whilst the risk prices in the crisis periods are implied from the risk quantity estimates and the observed risk premia in each crisis period. These estimates are then compared with the risk price estimates obtained for a noncrisis period. Second, a historical decomposition is performed which separates risk premia into a benchmark spread and the contributions of shocks of risk factor innovations during each financial crisis.

The rest of the paper proceeds as follows. A theoretical model of risk preferences is developed in Section 2. Estimates of the risk quantities and prices are given in Section 3, while Section 4 contains the dynamic decompositions of risk using the historical decomposition. Concluding comments are presented in Section 5.

2. A Model of Risk Premia

This section develops the key theoretical and empirical models of risk used in the empirical analysis to explain the movements in risk premia of the nine countries presented in Figure 1. The modeling approach consists of using the stochastic discount factor model to price the risks of all assets for the international investor. This has the effect of imposing a set of restrictions on the factor structure of bond spreads which enables the risk quantities and prices to be identified from the panel of countries; see Piazzesi (2006) for a recent review of factor models of bonds. In specifying the model, a nonparametric approach is adopted to circumvent the need to specify the utility function governing the functional form of the stochastic discount factor.

2.1 Model Specification

Consider the following stochastic discount factor model associated with the j^{th} portfolio with return $R_{j,t}$, which represents the Euler equation of an intertemporal portfolio model (Campbell, Lo and McKinley, 1997),

$$E_{t-1}\left[M_t\left(1+R_{j,t}\right)\right]=1, \tag{1}$$

where M_t is the stochastic discount factor or pricing kernel which is assumed to be positive, and $E_{t-1}[.]$ is the conditional expectations operator based on information at time t-1. Expanding the left-hand side gives

$$E_{t-1}\left[M_t\right]+E_{t-1}\left[M_tR_{j,t}\right]=E_{t-1}\left[M_t\right]+Cov_{t-1}\left(M_t,R_{j,t}\right)+E_{t-1}\left[M_t\right]E_{t-1}\left[R_{j,t}\right]. \tag{2}$$

The covariance risk of the portfolio is defined in terms of a set of risk factors ($F_{i,t}$) as follows

$$Cov_{t-1}(M_t, R_{j,t}) = E_{t-1}[M_t]\sum_{i=1}^{K}\gamma_i E_{t-1}\left[\left(R_{j,t} - E_{t-1}\left[R_{j,t}\right]\right)\left(F_{i,t} - E_{t-1}\left[F_{i,t}\right]\right)\right], \qquad (3)$$

where the inclusion of the term $E_{t\text{-}1}[M_t]$ acts as a convenient scalar adjustment in deriving the estimating equation. In the empirical analysis $K = 12$, corresponding to three global risk factors (credit, liquidity and volatility) and nine country risks. Using (2) and (3) in (1) gives

$$E_{t-1}[M_t] + E_{t-1}[M_t]\sum_{i=1}^{K}\gamma_i E_{t-1}\left[\left(R_{j,t} - E_{t-1}\left[R_{j,t}\right]\right)\left(F_{i,t} - E_{t-1}\left[F_{i,t}\right]\right)\right] + E_{t-1}[M_t]\,E_{t-1}\left[R_{j,t}\right] = 1. \qquad (4)$$

For the case where the pricing kernel is defined in terms of the risk free rate ($R_{f,t}$) then (1) becomes

$$E_{t-1}\left[M_t\left(1 + R_{f,t}\right)\right] = \left(1 + R_{f,t}\right)E_{t-1}[M_t] = 1, \qquad (5)$$

in which case

$$E_{t-1}[M_t] = \frac{1}{1 + R_{f,t}}. \qquad (6)$$

Using this expression in (4) gives

$$1 + \sum_{i=1}^{K}\gamma_i E_{t-1}\left[\left(R_{j,t} - E_{t-1}\left[R_{j,t}\right]\right)\left(F_{i,t} - E_{t-1}\left[F_{i,t}\right]\right)\right] + E_{t-1}\left[R_{j,t}\right] = 1 + R_{f,t}, \qquad (7)$$

or

$$E_{t-1}\left[R_{j,t}\right] - R_{f,t} = \sum_{i=1}^{K}\beta_i E_{t-1}\left[\left(R_{j,t} - E_{t-1}\left[R_{j,t}\right]\right)\left(F_{i,t} - E_{t-1}\left[F_{i,t}\right]\right)\right], \qquad (8)$$

where $\beta_i = -\gamma_i$. This equation shows that the expected excess return over a risk free asset is expressed as a weighted average of the covariances between the innovations in the return on the portfolio and the innovations in the risk factors. The covariances represent the quantity of risk whilst the weighting parameters, β_i, $i = 1,2,...,K$, are the prices of risk associated with each risk factor in $F_{i,t}$. The prices of risk are a function of, amongst other things, the risk parameters of the investor's utility function as well as the parameters that summarize the dynamics of the underlying model linking asset returns. In the case of the CCAPM, Campbell (1996) shows that the risk price is equal to the product of the relative risk aversion parameter less unity and the correlation between innovations in asset returns and revisions in expected future market returns. For a risk aversion parameter greater than unity and a positive correlation between the innovations in the return on an asset and revisions in expected future market returns, this results in a positive risk price and implies a positive trade-off between expected excess returns and covariance risk.

To provide a benchmark with which to compare the variation in prices of the risk factors across crises, the unconditional stochastic discount factor model is defined by expressing (8) in terms of unconditional expectations

$$E[R_{j,t}]-R_{f,t}=\sum_{i=1}^{K}\beta_i E\left[\left(R_{j,t}-E[R_{j,t}]\right)\left(F_{i,t}-E[F_{i,t}]\right)\right]. \tag{9}$$

In the empirical analysis, the unconditional covariance matrix $E\,[(R_{j,t}-E[R_{j,t}])\,(F_{i,t}-E[F_{i,t}])\,]$ is identified by imposing long-run restrictions on the dynamics of a SVAR model, which is specified below.

2.2 Identifying Risk Quantities

To estimate the long-run risk quantities in (9), a 12-variate SVAR model is specified containing the three global risk variables (credit, liquidity and volatility), and the nine bond spreads $R_{j,t}$, $j=1,2,...,9$. Let the full set of variables at t be represented by the (12×1) vector

Z_t, and let the long-run variance-covariance matrix between Z_t and the innovations to the risk factors be represented by

$$E\left[\left(Z_t - E[Z_t]\right)\left(F_t - E[F_t]\right)\right] = E\left[\left(Z_t - E[Z_t]\right)e_t\right] = H\,, \tag{10}$$

where

$$e_t = F_t - E[F_t], \tag{11}$$

are the innovations to the risk factors, and *H* contains the unknown long-run parameters. The $e_t = \{\, e_{i,t}, i = 1,2,...,12 \,\}$, is assumed to be an independent normal random vector with zero mean and unit covariance matrix

$$e_t \sim N(0, I). \tag{12}$$

The matrix of long-run parameters in (10) is defined as

$$H = \begin{bmatrix}
\lambda_c &&&&&&&&&&& \\
& \lambda_l &&&&&&&&&& \\
&& \lambda_v &&&&&&&&& \\
\delta_1 & \gamma_1 & \rho_1 & \phi_1 &&&&&&&& \\
\delta_2 & \gamma_2 & \rho_2 && \phi_2 &&&&&&& \\
\delta_3 & \gamma_3 & \rho_3 &&& \phi_3 &&&&&& \\
\delta_4 & \gamma_4 & \rho_4 &&&& \phi_4 &&&&& \\
\delta_5 & \gamma_5 & \rho_5 &&&&& \phi_5 &&&& \\
\delta_6 & \gamma_6 & \rho_6 &&&&&& \phi_6 &&& \\
\delta_7 & \gamma_7 & \rho_7 &&&&&&& \phi_7 && \\
\delta_8 & \gamma_8 & \rho_8 &&&&&&&& \phi_8 & \\
\delta_9 & \gamma_9 & \rho_9 &&&&&&&&& \phi_9
\end{bmatrix}, \tag{13}$$

where blank cells represent zeros. The long run global risk parameters for credit, liquidity and volatility are respectively $\lambda_c, \lambda_l, \lambda_v$. This implies that the three global risk factors are assumed to be independent processes determined entirely by their own dynamics in the long run. The parameters $\delta_1, \delta_2, \ldots \delta_9$, measure the long-run effects of a credit innovation on bond spreads (credit risk), the parameters $\gamma_1, \gamma_2, \ldots \gamma_9$, measure the long-run effects of a liquidity innovation on bond spreads (liquidity risk), and the parameters $\rho_1, \rho_2, \ldots \rho_9$, measure the long-run effects of a volatility innovation on bond spreads (volatility risk). The effects of country innovations on bond spreads (country risk) is controlled by the $\phi_1, \phi_2, \ldots \phi_9$ parameters.

The long-run risk quantities, $E\,[(R_{j,t} - E[R_{j,t}\,])\,(F_{i,t} - E[F_{i,t}\,])\,]$ in (9), correspond to the submatrix represented by the last 9 rows and 12 columns of *H*. The zeros in (13) impose the restriction of no contagion risk in the long-run. In the short-run however, contagion risk can occur whereby shocks in the risk premia of one country can impact on the bond risk premia of another country. In addition, the three risk indicators are also allowed to be interconnected in the short-run as they can be affected by all of the innovations in the model, including idiosyncratic country shocks of the emerging markets as well as shocks related to all risk indicators. To capture these short-run interactions a VAR with *p* lags, is specified as

$$\left(I - \Phi_1 - \Phi_2 L^2 - \ldots - \Phi_p L^p\right)\Delta Z_t = \alpha + v_t, \tag{14}$$

where $L^k Z_t = Z_{t-k}$ defines the lag operator, $\Delta = (I - L)$ is the first difference operator, Φ_k are (12×12) matrices of autoregressive parameters, α is a (12×1) vector of intercept parameters to capture the levels of the variables, and v_t is the VAR disturbance term which is assumed to be distributed as[4]

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[4] An extension of the assumption of a constant covariance matrix is to specify a conditional covariance matrix based on a multivariate GARCH model (Bekaert, Harvey and Ng, 2005), or a factor GARCH model (Dungey, Fry, González-Hermosillo and Martin, 2006).

$$v_t \sim N(0,\Omega). \tag{15}$$

The specification of the VAR in first differences follows Blanchard and Quah (1989) as the approach to impose the long-run restrictions defined in (13).

The short-run dynamics of the variables in the model are identified by the relationship between the VAR disturbance term v_t, and the innovations in the risk factors, e_t in (11),

$$v_t = Ge_t, \tag{16}$$

where from (15) and the definition of e_t in (12), implies that $\Omega = GG'$, while G is related to H as

$$G = \left(I - \Phi_1 - \Phi_2 - \ldots - \Phi_p\right)H. \tag{17}$$

2.2 Estimation

The long-run parameters in (13) are estimated by maximum likelihood. Given the assumption of normality for v_t, the log of the likelihood at the t^{th} observation is

$$\ln L_t = -\frac{N}{2}\ln(2\pi) - \frac{1}{2}\ln\left|G'G\right| - \frac{1}{2}v_t'\left(G'G\right)^{-1}v_t, \tag{18}$$

where v_t is the vector of VAR disturbances in (14), G is defined by (17) and $N = 12$. The log of the likelihood function for a sample of $t=1,2,\ldots,T$ observations, is given by

$$\ln L = \sum_{t=1}^{T} \ln L_t, \tag{19}$$

which is maximized using the procedure MAXLIK in GAUSS, version 5.0. The BFGS iterative gradient algorithm is used with derivatives computed numerically. Estimation is performed in two steps. First, the VAR parameters $\hat{\Phi}$ are estimated. Second, the long run parameters in H are estimated by maximizing the likelihood function in (19), subject to the restriction in (17) with Φ replaced by $\hat{\Phi}$ from the first step.

3. Empirical Estimates of Risk Quantities and Risk Prices in the Long Run

This section presents estimates of the risk quantities and risk prices from daily data on the global risk variables (credit, liquidity and volatility) and the bond spreads of nine emerging markets across three regions: Asia (Indonesia, Korea and Thailand), Eastern Europe (Bulgaria, Poland and Russia), and Latin America (Argentina, Brazil and Mexico). The sample period begins February 12, 1998 and ends May 17, 1999. The 12-variate VAR in (14) is specified with a lag length of $p=5$.[5]

3.1 Risk Quantities

The long run parameter estimates of H are reported in Table 1, with QMLE standard errors given in parentheses. The VAR parameter estimates are not reported. Instead the contribution of these estimates is presented graphically in Section 4.

The estimates of the long-run risk quantities in (9) are given in the second block of Table 1. All risk quantity parameter estimates are positive showing that a positive innovation to each of the risk factors on average widens bond spreads. Most of the risk quantity estimates are statistically significant with the main exceptions being the covariances between liquidity risk innovations and the three European spreads, which are all statistically insignificant.

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[5] This choice is based on the AIC and the likelihood ratio lag structure tests.

Table 1: Long-run parameter estimates, *H* in (13): QMLE standard errors are in brackets. The risk quantity estimates are given in the second part of the table.

Variables	Risk Factors			
	Credit risk	Liquidity risk	Volatility risk	Country risk
Credit	0.031			
	(0.002)			
Liquidity		0.032		
		(0.005)		
Volatility			0.058	
			(0.004)	
Risk premia: Asia				
Indonesia	0.132	0.125	0.182	0.825
	(0.054)	(0.061)	(0.051)	(0.073)
Korea	0.010	0.022	0.023	0.151
	(0.009)	(0.009)	(0.008)	(0.009)
Thailand	1.265	0.364	0.366	2.310
	(0.153)	(0.215)	(0.143)	(0.168)
Risk premia: Eastern Europe				
Bulgaria	0.085	0.012	0.081	0.583
	(0.036)	(0.028)	(0.034)	(0.046)
Poland	0.062	0.037	0.062	0.307
	(0.021)	(0.022)	(0.022)	(0.025)
Russia	0.057	0.039	0.065	0.260
	(0.018)	(0.021)	(0.018)	(0.023)
Risk premia: Latin America				
Argentina	0.085	0.064	0.033	0.301
	(0.020)	(0.021)	(0.019)	(0.022)
Brazil	0.134	0.093	0.056	0.387
	(0.024)	(0.023)	(0.022)	(0.024)
Mexico	0.074	0.040	0.025	0.203
	(0.014)	(0.014)	(0.013)	(0.014)

The country risk quantity estimates dominate the three global risk innovations for all countries. Of the quantity risk estimates associated with the three global risk innovations the relative rankings varies across countries. For the Latin American countries, credit risk dominates liquidity risk which in turn, dominates volatility risk. The credit risk quantity estimate for Thailand dominates the liquidity and volatility estimates, with the results for Indonesia and Thailand mixed. For the Eastern European countries, the credit and volatility quantity estimates are similar in magnitude, whilst the liquidity estimates tend to be at least 50% smaller.

Another way to compare the relative magnitudes of the size of the quantity risk estimates given in Table 1, is to decompose the long-run volatility of the risk premia into the contributions of the global risk factors and the country risk components. Using (13) and the independence assumption of the innovations e_t, the volatility decomposition for the j^{th} risk premium is given by

$$Var(R_{j,t}) = \delta_j^2 + \gamma_j^2 + \rho_j^2 + \phi_j^2, \tag{20}$$

where δ_j^2 is the contribution of credit risk, γ_j^2 is the contribution of liquidity risk, ρ_j^2 is the contribution of volatility risk, and ϕ_j^2 is the contribution of country risk.

The results of the long-run volatility decomposition in (20) are given in Table 2, where all terms are expressed in percentages. The risk quantity estimates are dominated by the country risk factors, where the estimates range from as low as 53.667% for Russia and as high as 76.801% for Indonesia. The contribution of the three global risk factors to long-run volatility is relatively even for most countries. The contribution of liquidity risk to quantity risk is smallest for the Asian countries, while the contribution of credit risk to quantity risk is smallest for the Latin American countries.

Table 2: Long-run volatility decomposition of risk premia: Based on (20), expressed in percent.

Country	Credit risk δ_j^2	Liquidity risk γ_j^2	Volatility risk ρ_j^2	Country risk ϕ_j^2
Asia				
Indonesia	10.596	1.388	11.214	76.801
Korea	13.308	8.022	12.442	66.229
Thailand	15.512	9.315	13.529	61.645
Eastern Europe				
Bulgaria	14.423	9.889	10.459	65.229
Poland	11.210	10.637	4.755	73.398
Russia	8.496	8.452	29.383	53.669
Latin America				
Argentina	6.841	13.278	17.537	62.344
Brazil	8.338	13.834	19.999	57.829
Mexico	7.240	11.728	21.670	59.362

3.2 Risk Prices

The estimates of the risk prices of the three global risks and the country risk, are now presented for the three crisis periods. The estimates of the risk prices are obtained by computing β_i in (9), as (Campbell, 1996)

$$\beta = Q^{-}\mu, \tag{21}$$

where β is the (9×12) vector containing the risk prices, μ is a (9×1) vector representing the average value of risk premia and Q^{-} is the (generalized) inverse of the matrix of quantity risks given by the submatrix represented by the last 9 rows and 12 columns of H in (13). To track the changes in risk prices during each of the three crises, the estimates of the quantity risk are fixed at their long-run estimates, while μ is estimated using the sample mean of bond spreads in each of the three crisis periods. For comparative purposes the long-run estimates of the risk prices are also computed by choosing μ as the sample mean of the bond spreads for the total sample period.

The estimates of the risk prices are given in Table 3. The results show that the prices of all three global risk factors increase above their long-run levels during the Russian crisis. A similar result occurs during the LTCM crisis. In the case of the Brazilian crisis, only the price of credit risk shows any substantial increase above its long-run level, whilst the liquidity and volatility risk prices tend to remain at their respective long-run levels.

A comparison of the country risk prices during the Russian and LTCM crises reveals that both crises are widespread as all countries with the exception of Bulgaria, experience increases in the prices of their country risks. Bulgaria experiences an increase in its country risk price during the Russian crisis, but this falls immediately during the LTCM crisis. In contrast, the Brazilian crisis is more localized, with just the Latin American countries experiencing increases in their country risk prices. During this period all Asian country risk prices return to their long-run levels, while the country risk prices in Poland and Russia maintain the levels achieved during the Russian and LTCM crises.

Table 3: Risk price estimates in the long-run and during crises: Based on (21).

Risk Factor Price		Crisis		
	Long-run	Russia	LTCM	Brazil
Global risk				
Credit	17.452	27.935	26.982	22.792
Liquidity	11.627	18.451	16.521	14.523
Volatility	11.602	20.337	18.055	12.828
Country risk: Asia				
Indonesia	11.768	18.541	21.651	13.960
Korea	6.489	14.849	12.119	2.184
Thailand	5.477	15.582	10.658	1.069
Country risk: Eastern Europe				
Bulgaria	4.613	10.648	5.609	4.827
Poland	12.806	16.929	16.818	16.188
Russia	1.524	2.471	5.693	5.472
Country risk: Latin America				
Argentina	9.663	12.810	8.458	12.301
Brazil	9.870	13.144	13.887	16.955
Mexico	13.089	18.388	18.959	16.344

4. Historical Decomposition of Risk Premia

In the previous section, the calculations of the risk prices for the three crises are computed relative to the long-run risk quantity estimates. An alternative way of identifying the relative importance of the components of risk during financial crises is to decompose the risk premia of each country at each point in time in terms of the innovations associated with each measure of risk: global factors, country and contagion. Formally this is accomplished by using the estimated SVAR to compute a historical decomposition over the crisis periods.

From (14), rewrite the VAR in terms of its vector moving average (VMA) representation and use (16) to express the model in terms of the risk innovations e_t

$$\Delta Z_t = (I - \Phi_1 L - \Phi_2 L^2 - \ldots - \Phi_p L^p)^{-1}(\alpha + Ge_t),$$
$$\Delta Z_t = \psi + (I - \Theta_1 L - \Theta_2 L^2 - \ldots - \Theta_q L^q)^{-1} Ge_t, \tag{22}$$

where Θ_k are (12×12) matrices of moving average parameters which are functions of the autoregressive parameters of the VAR, and $\psi = (I - \Phi_1 - \Phi_2 - \ldots - \Phi_p)^{-1}\alpha,$ is a (12×1) vector of intercept parameters.[6] To identify the contributions of the innovations of all risk factors during the crisis periods, benchmark spreads for each country risk premium are computed which represent counterfactual estimates of the risk premia if a crisis had not historically occurred. The benchmark spreads are computed as the conditional expectations of the risk premia based on a pre-crisis period. Letting T^* representing the end of the pre-crisis period, from (22), the conditional expectation over the crisis period is

$$\Delta Z_{T^*+D|T^*} = \psi + \sum_{i=D}^{\infty} \Theta_i G e_{T^*+D-i}\ . \tag{23}$$

The forecast error over the crisis period corresponding to the deviations between changes in actual and benchmark spreads is

$$\Delta Z_{T^*+D} - \Delta Z_{T^*+D|T^*} = \sum_{i=D}^{D-1} \Theta_i G e_{T^*+D-i}\ , \tag{24}$$

which combined with (23)

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[6] The matrix polynomial inversion used to generate the vector moving average representation is usually computed numerically; see Hamilton (1994, p.260).

$$\Delta Z_{T^*+D} = \left[\psi + \sum_{i=D}^{\infty} \Theta_i G e_{T^*+D-i} \right] + \sum_{i=0}^{D-1} \Theta_i G e_{T^*+D-i} \,, \tag{25}$$

provides the historical decomposition of the change in risk premia over the crisis period in terms of changes in benchmark spreads based on pre-crisis information (first term) and the various innovations to all of the risk variables (second term). The historical decomposition is computed by replacing the unknown parameters and innovations by their estimated values. To express the historical decomposition in terms of the levels of the risk premia, (25) is cumulated over the sample.

The historical decompositions are computed with the pre-crisis period ending at *T**=May 30, 1998. This date is several months before the actual disclosure of Russia's default on August 17, 1998, but is chosen to deal with the potential mounting pressure in financial markets that may have occurred before the actual crisis became newsworthy.

4.1 Benchmark Spread Estimates

Figure 2 compares the observed risk premia with the benchmark spreads for each country over the crisis period. Average estimates over selected periods are also given in Table 4 for both series. With the exception of the three Asian countries, the benchmark spreads steadily increase over the period of the historical decomposition. For the case of Russia, the benchmark spread deteriorates markedly from around 6.5 percent on June 1, 1998 to about 35 percent by May 17, 1999, reflecting the sustained level of risk experienced in the Russian market following the Russian debt default. The benchmark spreads of the Asian countries either tend to be relatively flat over the period (Indonesia), or show a slight fall (Korea and Thailand). This result reflects the recovery from the Asian financial crisis starting in July 1997.

Figure 2. Actual and benchmark risk premia, June 1, 1998 - May 17, 1999.

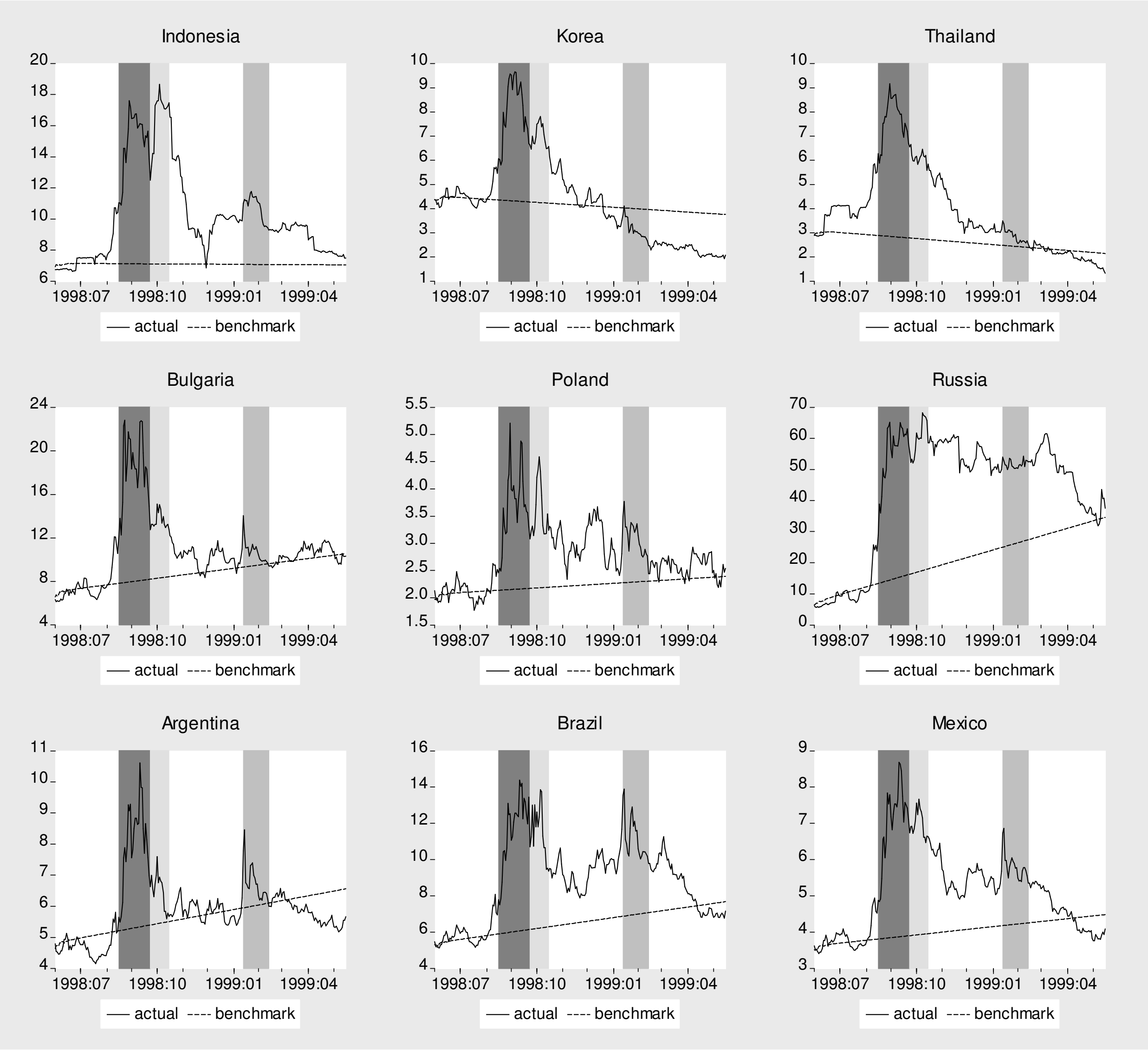


Table 4: Decomposition of risk premia in terms of innovations to the risk factors: percentage per annum.

Country	Actual	Bench.	Credit	Liquidity	Volatility	Country	Asia	Europe	Lat. Am.
Indonesia									
Russian crisis	15.03	7.12	1.21	0.07	1.05	5.71	-0.08	-0.03	-0.02
LTCM crisis	16.56	7.11	1.38	0.10	0.81	7.05	0.09	0.09	-0.08
Brazilian crisis	11.29	7.08	0.32	0.08	0.83	3.07	-0.05	-0.05	0.01
Korea									
Russian crisis	8.12	4.32	0.82	0.30	0.80	1.83	-0.04	-0.01	0.10
LTCM crisis	7.01	4.25	0.93	0.34	0.61	0.66	0.10	0.07	0.04
Brazilian crisis	3.34	4.01	0.21	0.28	0.63	-1.77	-0.03	-0.02	0.04
Thailand									
Russian crisis	7.69	2.84	0.75	0.32	0.85	2.90	-0.04	0.02	0.05
LTCM crisis	6.13	2.76	0.92	0.36	0.64	1.31	0.08	0.02	0.04
Brazilian crisis	3.01	2.46	0.27	0.29	0.67	-0.66	-0.04	0.01	0.01
Bulgaria									
Russian crisis	18.50	8.01	1.85	1.02	2.36	4.87	0.04	-0.04	0.38
LTCM crisis	13.56	8.31	2.12	1.12	1.77	-0.42	0.11	0.21	0.32
Brazilian crisis	11.23	9.41	0.50	0.93	1.84	-1.61	0.14	0.00	0.01
Poland									
Russian crisis	3.69	2.15	0.18	0.18	0.30	0.77	0.02	0.00	0.09
LTCM crisis	3.57	2.18	0.16	0.20	0.22	0.73	0.03	-0.01	0.06
Brazilian crisis	3.29	2.28	-0.01	0.16	0.23	0.66	-0.01	-0.02	-0.01
Russia									
Russian crisis	55.19	14.76	15.35	2.92	4.73	18.28	0.14	-0.66	-0.33
LTCM crisis	59.89	17.15	20.21	3.30	3.60	14.40	1.02	0.02	0.19
Brazilian crisis	51.45	25.71	6.96	2.71	3.76	12.75	0.58	-0.53	-0.50
Argentina									
Russian crisis	8.07	5.30	1.04	0.52	0.43	0.63	0.00	-0.03	0.19
LTCM crisis	6.48	5.45	1.35	0.58	0.34	-1.35	0.13	0.08	-0.09
Brazilian crisis	7.04	6.00	0.44	0.48	0.35	-0.52	0.06	0.00	0.24
Brazil									
Russian crisis	11.67	6.02	1.71	0.76	0.72	2.35	-0.01	-0.02	0.13
LTCM crisis	11.52	6.22	2.13	0.84	0.54	1.53	0.12	0.01	0.12
Brazilian crisis	11.78	6.94	0.64	0.70	0.57	2.98	0.06	-0.04	-0.08
Mexico									
Russian crisis	7.06	3.86	0.95	0.33	0.32	1.65	-0.03	0.00	0.00
LTCM crisis	6.97	3.93	1.19	0.37	0.25	1.11	0.07	0.01	0.04
Brazilian crisis	5.95	4.20	0.36	0.30	0.26	0.88	0.02	-0.03	-0.04

4.2 Global Risk Factor Estimates

The contribution of the innovations of the global risk factors to the bond risk premia of the emerging countries during the financial crisis periods are presented in Figure 3 with summary estimates given in Table 4.

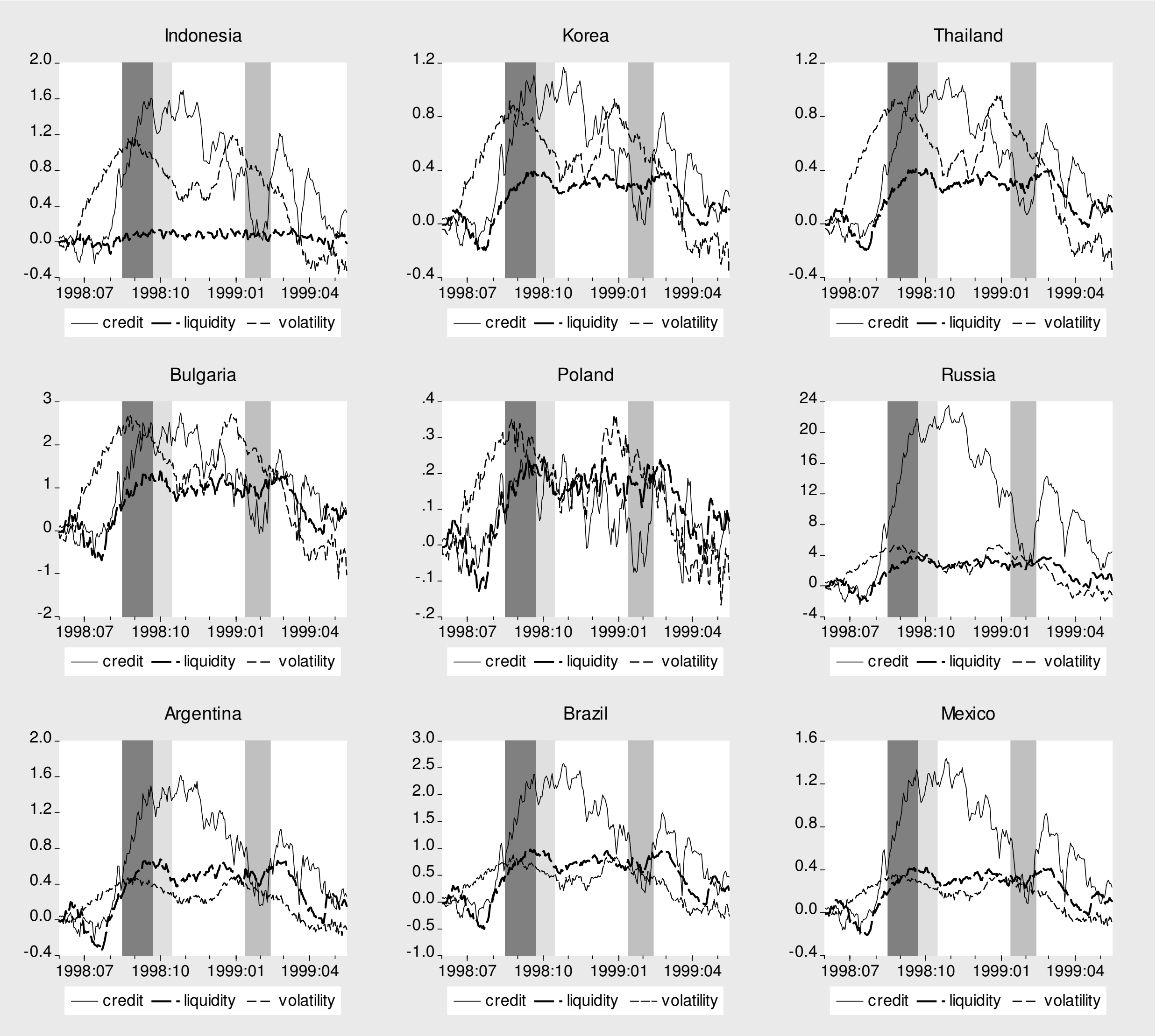
Figure 3 shows that of the three global risk factors, credit risk has the largest contribution to bond risk premia. This is especially true in Russia and the Latin American countries during the period preceding the Brazilian crisis. The importance of credit risk rises sharply during the Russian crisis for all countries, with the largest effect occurring in Russia where the contribution to Russian bond spreads is about 20%. The effect of credit risk is maintained during the LTCM crisis and approximately for another two months after this crisis. The relative importance of credit risk beginning with the Russan crisis, gives credence to the claim that the Russian crisis was a crisis of credit.

The contribution of credit risk to bond risk premia trends downwards for the remainder of the sample period, dipping below trend during the Brazilian crisis. During this crisis all the global risk factors tend to have similar impacts on the bond risk premia. From Table 4, the average contribution of these factors during the Brazilian crisis is between 0.35% and 0.48% for Argentina, between 0.57% and 0.70% for Brazil, and between 0.26% and 0.36% for Mexico.

Figure 3 also shows that the importance of volatility risk to bond risk premia began much earlier than that of the other global risk factors. This rise began closer to the beginning of the historical decomposition period, suggesting early signs of the forthcoming crisis in Russia.

Liquidity risk rose sharply during the Russian crisis for all countries, with the exception of Indonesia. Having obtained a new higher level during the Russian crisis, the liquidity risk

Figure 3. Contribution of global risk innovations to risk premia, June 1, 1998 - May 17, 1999.



Indonesia, Korea, Thailand, Bulgaria, Poland, Russia, Argentina, Brazil, Mexico factor remains relatively high during the historical decomposition period, and does not substantially decline until after the Brazilian crisis. Liquidity risk is the least likely of the identified risk factors to account for a substantial portion of the decomposition of the risk premia for all but the Latin American countries, where it is often marginally more important than volatility risk.

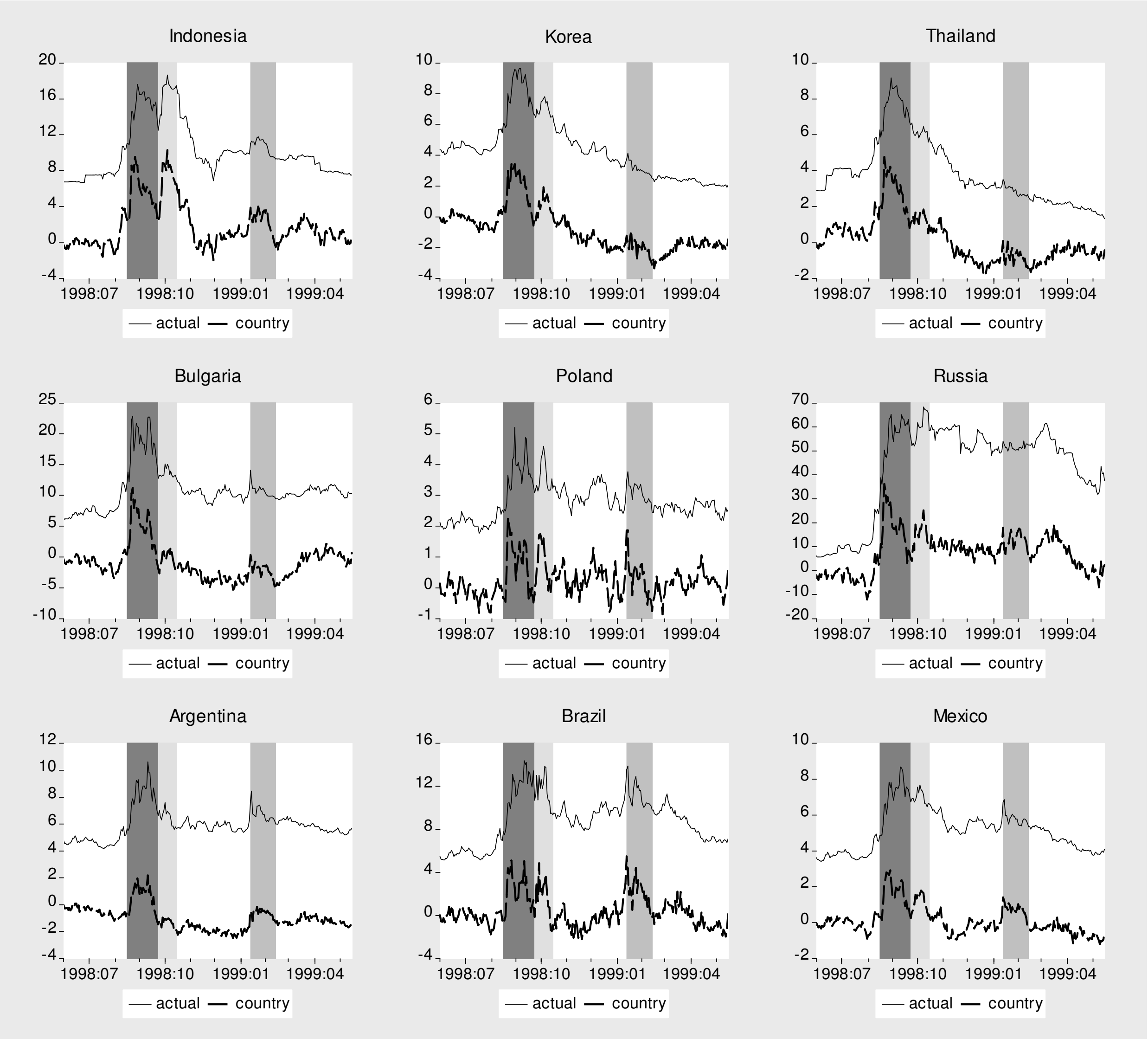
4.3 Country Risk Factor Estimates

The contribution of innovations to country risk to the risk premia of the nine countries during the financial crisis period are presented in Figure 4, with summary estimates given in Table 4. A comparison of the country risk innovations with the actual risk premia in Figure 4, highlights the importance of idiosyncratic country risks in pricing for all assets. It is interesting to note that the fall in country risk after the Russian crisis reduces spreads for Korea and Thailand, indicating that credit concerns of these two countries become less of an issue compared to the pre-crisis period.

4.4 Contagion Risk Factor Estimates

The contributions of contagion risk to the risk premia are summarised in the last three columns of Table 4 which breakdown the contagion effects originating from each region. A comparison of these estimates with the global risk estimates and country risk estimates, shows that contagion has a relatively small effect on risk premia during the three crisis periods. This result is consistent with the early empirical results of Forbes and Rigobon (2002) who found little evidence for the existence of contagion for a broad range of asset markets across a number of financial crises.

Figure 4. Contribution of country risk innovations to risk premia, June 1, 1998 - May 17, 1999.



5. Conclusions

This paper has identified and quantified the effects of changes in global risks on the risk premia of sovereign bonds issued by emerging markets. Three measures of global risk were used consisting of credit, liquidity and volatility risk. Additional risks in the form of country and contagion factors were also identified.

The approach consisted of developing a theoretical framework to price risk using the stochastic discount factor model. This yielded a multifactor asset pricing model which resulted in a set of restrictions being imposed on the long-run dynamics of a SVAR model. The model was applied to analysing the daily movements in the bond spreads of nine emerging markets from Asia, Eastern Europe and Latin America, from 1998 to 1999, a period containing the Russian, LTCM and Brazilian crises. The three global risk variables were based on the JP Morgan indices of risk, whilst the country and contagion risks were identified and incorporated into the SVAR using long run restrictions.

The empirical results suggested that different characterizations of global risk patterns were at play during the recent the financial crises analyzed in this paper. The Russian and LTCM crises were characterized by increases in the price of risk of all three global risk factors. The results of the Brazilian crisis were different however, where the only significant increase in the price of global risk came from credit risk.

All bond markets experienced increases in country risk prices during the Russian crisis which were sustained during the LTCM crisis, whereas it was just the Latin American and Russian bond markets that incurred increases in the price of country risk during the Brazilian crisis. In contrast, contagion was found to have a smaller effect on bond spreads than the global and country measures of risk during all three financial crises.

Within the set of global risk factors investigated, credit risk was on average the most important contributor to bond spreads over the period investigated. Credit risk was dominant particularly during the Russian crisis, consistent with the view that this period could be viewed as a global credit risk shock. However, the relative importance of credit risk during the Brazilian crisis diminished, where its role was comparable in magnitude to liquidity and volatility risk. Liquidity risk was the least important of the three risk factors.

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Appendix A: The Data

The sample period is from February 12, 1998 to May 17, 1999, for a total of 328 observations.

Table A.1: Data Definitions and Sources

Variable	Description	Source
Bonds[1]		
Indonesia	Indonesian Yankee bond	US Federal Reserve
Korea	Government of Korea 8 7/8% 4/2008 bond	US Federal Reserve
Thailand	Kingdom of Thailand Yankee bond	US Federal Reserve
Bulgaria	Bulgarian Discount Stripped Brady bond	US Federal Reserve
Poland	Poland Par Stripped Brady bond	US Federal Reserve
Russia	Government of Russia 9.25% 11/2001 bond	Bloomberg
Argentina	Republic of Argentina bond	US Federal Reserve
Brazil	Republic of Brazil bond	US Federal Reserve
Mexico	JP Morgan Eurobond Index Mexico Sovereign bond	US Federal Reserve
Credit[1]	US Industrial BBB1 corporate 10-year bond spread over the US Treasury bond of comparable maturity	Bloomberg (IN10Y3B1)
Liquidity[2]	JP Morgan Chase Bank's (LCVI) liquidity index. 0= low risk, 100=high risk, log. Components are: (a) US Treasury yield spreads of benchmark and off-the-run bonds for different maturities; (b) 10-year US swap spreads.	JP Morgan Chase Bank
Volatility[2]	JP Morgan Chase Bank's (LCVI) volatility index. 0= low risk, 100=high risk, log. Components are: (a) implied 12-month foreign exchange volatility for six currencies (EUR, JPY, CHF, GBP, CAD, AUD against the USD); (b) implied equity volatility based on option markets on the Chicago Board of Options Exchange; (c) JP Morgan Global Risk Appetite Index (GRAI) based on measures of correlation between the rank of the $ returns of 15 currencies of the past two months, and (d) the rank of risk measured by historical yield.	JP Morgan Chase Bank

1. The risk premia are constructed from a representative long term emerging country sovereign bond (rather than Brady bonds) issued in US dollars less a US treasury bond of comparable maturity. Missing observations are replaced with the previous day's observation.

2. The JP Morgan risk components are expressed as indices. A cumulative distribution function is applied to the raw data, and each observation is then expressed as a percentile of the distribution function such that higher risk is associated with higher percentiles. The indices are the average of these transformed data. See JP Morgan (1999) and JP Morgan Chase Bank (2002).

Appendix B: Crisis Dates

The Russian crisis begins on August 17 corresponding to the announcement of it's bond default and ends on September 23, 1998, with the announcement of the orchestrated rescue plan for the hedge fund LTCM which marks the start of the LTCM crisis. The end of the LTCM crisis corresponds to the inter-FOMC Fed interest rate cut on October 15, 1998 consistent with the findings of the Committee on the Global Financial System, 1999. The Brazilian crisis is dated to begin January 13, 1999 with the effective devaluation of the Real and ends February 2, 1999 to give a crisis duration of two weeks.

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