

Stock Market Liberalization and International Risk Sharing

Shigeru Iwata and Shu Wu*

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Abstract

This paper empirically examines what macroeconomic risks are shared (or not shared) internationally after stock market liberalization by developing countries. To address this issue, we incorporate an international asset pricing model into a non-linear structural Vector Autoregression (VAR) system that identifies various sources of macroeconomic risk. We find that most of the risk corresponding to exogenous financial market shocks are very well shared, even though other macroeconomic risks associated with exogenous shocks to output, inflation and monetary policies are not fully shared across countries. Our results imply that one of the main benefits from stock market liberalization is to allow a country to better hedge against exogenous and idiosyncratic financial market shocks, and stock market liberalization should be accompanied by other macroeconomic reforms in order to achieve the full benefits of international risk sharing.

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*Address: The University of Kansas, Department of Economics, 213 Summerfield Hall, Lawrence, KS 66045. E-mail: iwata@ku.edu (Shigeru Iwata) and shuwu@ku.edu (Shu Wu). We would like to thank seminar participants at the Chinese University of Hong Kong for helpful comments. Part of the research of the second author was done while he was visiting the Department of Economics and Finance at the City University of Hong Kong. He would like to express gratitude for their hospitality.

1 Introduction

One of the greatest benefits for developing countries from liberalizing their stock markets is to achieve better international risk sharing, which lowers the equity premium, and hence, cuts the cost of capital in the liberalizing country (see, for example, Stulz, 1999). While there is much empirical evidence reported that the cost of capital has indeed declined after opening the domestic stock market to foreign investors,¹ we yet have a clear picture about the nature and degree of the increased risk sharing after market liberalization. This is because that there are two components to a country's cost of capital: the risk-free rate and the equity premium. Economic theory suggests that both will fall when a developing country liberalizes. In principle, the lower cost of capital could result entirely from changes in the risk-free rate.

In a recent study with firm level data, Chari and Henry (2004) provided microeconomic evidence on the effect of increased risk sharing on the cost of capital. They found that roughly two fifth of the total stock price revaluation following a liberalization is due to the reduction in the systematic risk of investible firms in the liberalizing country. They measured the reduction in the systematic risk by the difference between the historical covariance of a firm's stock return with the local market index and that with the world market index. The covariance with the local market index is roughly 200 times larger than the covariance with the world market index for an average investible firm in the liberalizing country. The increase in the degree of risk sharing appears substantial after liberalizations and contributes significantly to the reduction of the cost of capital.

In recent years, however, economists have increasingly expressed doubts about the benefits of capital account liberalization in the wake of financial crises in Asia, Russia and Latin America (see, for example, Bhagwhati, 1998 and Stiglitz, 2002). They argue that a liberalization does not necessarily generate greater efficiency with nontrivial welfare gains. Instead, it invites international speculations that increase the likelihood of financial crises.

¹For example, Bekaert and Harvey (2000) found that the impact of stock market liberalizations on the cost of capital varies from 5 to 75 basis points. Henry (2000) showed that, in anticipation to a forthcoming stock market liberalization, the liberalizing country's aggregate equity price index experiences an abnormal return of 3.3% per month in real dollar term on average. Henry (2003) reported that liberalizations also lead to sharp increases in aggregate investment as cost of capital, measured by dividend yield, falls.

Therefore, if the increased risk sharing is substantial after market liberalizations, it is important to ask what is the nature of the risks that are shared. If the risks are ultimately linked to macroeconomic fundamentals, *what* macroeconomic risks are shared (or not shared) internationally after stock market liberalization? If the liberalization helps developing countries insure against different shocks to the economy, how well is it done using the existing international asset markets?

To address the above issues, we incorporate an international asset pricing model into a non-linear structural Vector Autoregression (VAR) system that identifies various sources of macroeconomic risk, including exogenous shocks to the domestic and foreign industrial output growth, inflation and monetary policies as well as exogenous shocks to asset markets. We then examine the dynamic effect of these shocks on the domestic-foreign marginal utility growth differential, which is approximated by the change in the real exchange rate. If one macroeconomic risk is fully shared across countries, then domestic and foreign marginal utility growth rates would move together in response to the shock. Hence, their *difference* would not be affected by this shock, and the shock should not account for much of the volatility of the change in the real exchange rate. Consistent with the results in Chari and Henry (2004), we find that most of the risks of exogenous financial market shocks are very well shared between developed and developing countries after the latter liberalize their stock markets. On the other hand, other macroeconomic risks such as those associated with exogenous shocks to output, inflation and monetary policies are not fully shared across countries. Therefore, our results imply that a major part of the benefits from stock market liberalization in developing countries is to allow them to better hedge against exogenous and idiosyncratic financial shocks.

One of the difficulty of studying international risk sharing is that the marginal utility of consumers is not observable. The common approach to studying the gains from trade of risky assets is to make explicit assumptions about the functional form of, as well as the inputs into, consumers' utility function, and calibrate the hypothetical welfare losses from the lack of international risk sharing. See, for example, Obstfeld (1994), Lewis (1996), Van Wincoop (1999) and Davis, Nalewaik and Willen (2000) among others. In this study we take a different approach based on recent papers by Brandt, Cochrane and Santa-Clara (2001) and Iwata and Wu (2004). Recognizing that the real exchange rate moves by the domestic-foreign marginal utility

growth differential, Brandt, Cochrane and Santa-Clara (2001) derive the information about the marginal utility growth directly from asset returns and analyze the degree of risk sharing among major developed countries. Using a similar method, Iwata and Wu (2004) study, for those countries, how well international investors insure against different kinds of macroeconomic risk using the existing asset markets.

The rest of the paper is organized as follows. Section 2 lays out the empirical model used in the paper. Section 3 discusses the main results and section 4 provides a summary.

2 The Framework

In this section, we describe the theoretical background that motivates our empirical specification. We then introduce a structural VAR analysis with various sources of macroeconomic risk and incorporate the asset-market view of international risk sharing. The main advantage of this asset-market view of international risk sharing is that one does not have to make assumptions about the functional form of (or the inputs into) the utility function of international investors.

2.1 The background

We begin by a key economic relationship underlying our empirical analysis. Under the assumption of absence of arbitrage in international financial markets, variations in the real exchange rate are directly linked to the difference between the growth rates of marginal utility of domestic and foreign investors. That is,

$$\log S_{t+1} - \log S_t = -(\log M_{t+1} - \log M_{t+1}^*), \quad (1)$$

where S_t is the real exchange rate (in units of domestic goods/foreign goods) between the two countries, and $\log M_{t+1}$ and $\log M_{t+1}^*$ are the growth rates of the domestic and foreign marginal utility, respectively (see Appendix A for details of derivation of this and other relations in this section).²

²See Brandt et al. (2001). Other studies that also exploit this relationship include Iwata and Wu (2004), Backus, Foresi and Telmer (2001), Brandt and Santa-Clara (2001), Hollifield and Yaron (2000) among others.

Suppose that there were perfect risk sharing across countries after stock market liberalization. Then, M_{t+1} and M_{t+1}^* would be equal and hence, the real exchange rate would stay constant. On the other hand, if risk sharing remains poor after stock market liberalization, $\Delta \log S_{t+1}$ will fluctuate as $\log M_{t+1}$ and $\log M_{t+1}^*$ move differently in response to various economic shocks.

To get a convenient empirical specification, we assume that M_{t+1} and M_{t+1}^* both follow the log-normal distribution. More specifically, let ε_t represent a vector of fundamental economic shocks (to be described below) distributed as $\mathcal{N}(\mathbf{0}, \mathbf{I})$ and assume that

$$\log M_{t+1} = \mu_t - \lambda_t' \varepsilon_{t+1} \quad \text{and} \quad \log M_{t+1}^* = \mu_t^* - \lambda_t^{*'} \varepsilon_{t+1} \quad (2)$$

where $\mu_t = E_t(\log M_{t+1})$, $\mu_t^* = E_t(\log M_{t+1}^*)$. Parameters λ_t and λ_t^* are often referred to as the market prices of risk. Let i_t and i_t^* be the one-period risk-free real interest rates in the home and foreign country, respectively. Then (1) may be expressed as

$$\Delta \log S_{t+1} - (i_t - i_t^*) = \frac{1}{2}(\lambda_t' \lambda_t - \lambda_t^{*'} \lambda_t^*) + (\lambda_t - \lambda_t^*)' \varepsilon_{t+1}. \quad (3)$$

Furthermore, if r_t and r_t^* are the domestic and foreign real stock returns, then under the no-arbitrage condition and log-normal assumption, we find

$$r_{t+1} - i_t = -\frac{1}{2}\sigma' \sigma + \sigma' \lambda_t + \sigma' \varepsilon_{t+1} \quad (4)$$

$$r_{t+1}^* - i_t^* = -\frac{1}{2}\sigma^{*'} \sigma^* + \sigma^{*'} \lambda_t^* + \sigma^{*'} \varepsilon_{t+1}. \quad (5)$$

where σ and σ^* are the volatilities of the stock returns.

Equation (3) together with (4) and (5) can be used to find a link between the excess returns and macroeconomic shocks. In what follows, we specify λ_t and λ_t^* as functions of observable macroeconomic variables,³ which are in turn driven by the fundamental macroeconomic shocks.

³In the finance literature, there have been several trials to use observable economic variables as priced risk factors in asset pricing models, including classic studies by Chen, Roll and Ross (1986), Chan, Chen and Hsieh (1985) and Ferson and Harvey (1991). More recently, Ang and Piazzesi (2003) incorporates macroeconomic variables in a VAR analysis of the term structure of interest rates.

2.2 Building a VAR model

In our analysis, we postulate two types of shocks. The first type of shocks includes exogenous innovations to output growth, inflation and monetary policies in the home and foreign countries. The second type of shocks is the exogenous financial market shocks orthogonal to those macroeconomic shocks. In particular, the ε_t is assumed to have nine elements: $\varepsilon_t = (\varepsilon'_{Y,t}, \varepsilon'_{\Pi,t}, \varepsilon'_{M,t}, \varepsilon'_{S,t})'$. The first two vectors of shocks $\varepsilon_{Y,t} = (\varepsilon_{y,t}, \varepsilon_{y,t}^*)'$ and $\varepsilon_{\Pi,t} = (\varepsilon_{\pi,t}, \varepsilon_{\pi,t}^*)'$ can be thought of as the home and foreign countries' aggregate supply and demand shocks, respectively. The third vector of shocks $\varepsilon_{M,t} = (\varepsilon_{m,t}, \varepsilon_{m,t}^*)'$ includes exogenous shocks to monetary policies in the two countries and the last vector of shocks $\varepsilon_{S,t} = (\varepsilon_{s,t}, \varepsilon_{r,t}, \varepsilon_{r,t}^*)'$ represents exogenous shocks to international financial markets, including the exogenous innovations to the foreign exchange rate as well as the domestic and foreign stock market returns.

We assume that the current state of the economy may be summarized by \mathbf{z}_t : a 9×1 vector of macroeconomic variables. We include in \mathbf{z}_t the home and foreign output growth rates (y_t and y_t^*), inflation rates (π_t and π_t^*) as well as real short-term interest rates (i_t and i_t^*) in the two countries. The last three elements of \mathbf{z}_t are the change in the real exchange rate ($\Delta \log S_t$) and the domestic and foreign real stock returns (r_t and r_t^*).

The market prices of risk are assumed to be linear functions of \mathbf{z}_t ⁴

$$\lambda_t = \mathbf{\Gamma} \mathbf{z}_t \quad \text{and} \quad \lambda_t^* = \mathbf{\Gamma}^* \mathbf{z}_t, \quad (6)$$

where $\mathbf{\Gamma}$ and $\mathbf{\Gamma}^*$ are 9×9 matrices. We assume that the dynamics of the first six elements of \mathbf{z}_t (denoted by \mathbf{z}_t^+) is described by the reduced-form equation

$$\mathbf{z}_t^+ = \mu + \mathbf{B}_1^+ \mathbf{z}_{t-1} + \dots + \mathbf{B}_p^+ \mathbf{z}_{t-p} + \mathbf{u}_t^+ \quad (7)$$

where $\mathbf{z}_t = (\mathbf{z}_t^+, \Delta \log S_t, r_t, r_t^*)'$, $\mathbf{B}_1^+, \dots, \mathbf{B}_p^+$ are 6×9 matrices and μ are a 6×1 vector of constants. The \mathbf{u}_t^+ stands for a vector of one-step-ahead forecast errors and we assume that $\mathbf{u}_t^+ \sim \mathcal{N}(\mathbf{0}, \mathbf{\Sigma})$, where $\mathbf{\Sigma}$ is a symmetric positive definite matrix. The error term \mathbf{u}_t^+ is related to the structural shocks according to $\mathbf{u}_t^+ = \mathbf{C} \varepsilon_t$, where \mathbf{C} is a 6×9 matrix. Using (3), (4)

⁴Similar parameterizations of the market price of risk have been widely used in the literature where \mathbf{z}_t is treated as a latent state variable, including Constantinides (1992), Ahn et al. (2000) and Dai and Singleton (2002) among many others.

and (5) together with (6), the last three elements of \mathbf{z}_t may be written as

$$\Delta \log S_t = (i_{t-1} - i_{t-1}^*) + \frac{1}{2} \mathbf{z}'_{t-1} (\mathbf{\Gamma}' \mathbf{\Gamma} - \mathbf{\Gamma}^{*'} \mathbf{\Gamma}^*) \mathbf{z}_{t-1} + \mathbf{z}'_{t-1} (\mathbf{\Gamma} - \mathbf{\Gamma}^*)' \varepsilon_t \quad (8)$$

$$r_t = -\frac{1}{2} \sigma' \sigma + \sigma' \mathbf{\Gamma} \mathbf{z}_{t-1} + i_{t-1} + \sigma' \varepsilon_t \quad (9)$$

$$r_t^* = -\frac{1}{2} \sigma^{*'} \sigma^* + \sigma^{*'} \mathbf{\Gamma}^* \mathbf{z}_{t-1} + i_{t-1}^* + \sigma^{*'} \varepsilon_t \quad (10)$$

Equations (7), (8), (9) and (10) therefore constitute a constrained non-linear VAR, on which our empirical analysis will be based.

2.3 Identification issue

In conventional VAR models, the identification problem reduces to the restrictions on matrix \mathbf{C} . In the nonlinear VAR model above, we need identification of the pricing matrices $\mathbf{\Gamma}$ and $\mathbf{\Gamma}^*$ as well. We briefly discuss these two identification conditions in turn.

We impose the following restrictions to identify the macroeconomic shocks. First, since it is widely believed that monetary policy actions only affect the real economy as well as inflation with a delay, we assume that output growth and inflation do not respond contemporaneously to shocks to monetary policies in both countries. We also assume that exogenous shocks to the exchange rate and the stock returns have no immediate impact on output growth and inflation. This kind of recursive identification assumption is commonly used in the monetary VAR literature (Christiano et al. 1999), especially when monthly data (such as in this study) are used to estimate the model. Second, we assume that, when setting its policy instrument, the monetary authority in one country does not respond contemporaneously to the other country's aggregate supply and demand shocks as well as the monetary policy shocks. The main reason is that the exact information about a foreign country's output, price and monetary policy stance may not be available immediately to the domestic central bank. Finally, we allow monetary policies to respond to the exogenous shocks to the exchange rate, but not to the innovations to the stock returns in our model. This is mainly because monetary authorities very rarely respond to the development in stock markets while stability in foreign exchange markets has been a more common policy goal of many central banks.

To identify Γ and Γ^* , we make the following three assumptions (see Appendix B for more details). First, we assume that while the foreign stock returns respond contemporaneously to an exogenous shock to the US stock returns, the domestic (U.S.) stock returns do not respond in such a fashion to the shock to the foreign stock returns. Second, consistent with the usual representative-agent approach in macroeconomics, we assume that home investors and foreign investors price the risk factors in a symmetrical fashion as detailed in Appendix B. Third, we assume that the contribution of y_t^* to the market price of home consumption risk is equal in size to the contribution of y_t to the market price of foreign consumption risk. This assumption of symmetry is obviously a simplification.

3 Empirical Results

3.1 The Data and the Risk Sharing Index

In this study, we focus on major Latin American countries that liberalized their stock markets in the late 1980s. The data used in this study are monthly observations on industrial output, consumer price indices (CPI), short-term interest rates, stock market returns, foreign exchange rates for Mexico, Brazil, Chile and the United States over the period from 1988 through 2002. The data on international stock market returns are from Morgan Stanley Capital International Indices. The data on other macroeconomic variables are from International Financial Statistics of the IMF.⁵ All variables are in nominal terms. We obtain the real output growth rate by subtracting the CPI inflation rate from the industrial output rate. Similarly, the real interest rates, real stock returns and the change in the real exchange rates are obtained by adjusting for the CPI inflation rates.⁶ Table 1 presents the summary statistics of our financial market data.

⁵We did not include Argentina in this study because Argentina kept fixed exchange rate regime after financial liberalization in the late 1980s, which makes our model inapplicable. We did not include other Latin America countries mainly because of lack of sufficient time series data required for implementing our model.

⁶Note that we use the ex-post real interest rates, instead of the ex-ante real interest rate, in our empirical exercise. The impact of this choice on the estimation results is minimal because the dependent variables in the regression are the excess rates of returns, see equations (3), (4) and (5).

Brandt, Cochrane, and Santa-Clara (2003) proposed an index of international risk sharing given by

$$1 - \frac{Var_t(\Delta \log S_{t+1})}{Var_t(\ln M_t^*) + Var_t(\ln M_t)}$$

From the observed data, we can calculate the lower bound of the above index, using the Hansen-Jogannathan (1991) bounds. Brandt et al (2003) calculate this index for US/UK, US/Germany, and US/Japan, using monthly stock returns and exchange rates data from January 1975 through June 1998. The results for three pairs of countries are all around 0.980 – 0.986 with standard errors 0.005 – 0.016. These high numbers suggest that international risk sharing is better than one usually thinks.

Using Latin American data, we calculate the above index for US/Mexico, US/Brazil, and US/Chile pairs. The result is shown in Table 2. Taking the figures at their face values, risk sharing between US and Chile is as good as US and the developed countries such as UK, Germany, and Japan. The index between US and Mexico is about 8% lower than that of US and Chile. Risk sharing between US and Brazil appears significantly poorer than those two cases; the index is more than 20% lower than that of in the US and developed countries cases. The standard error for the US/Brazil is, however, quite large because of its short sample.⁷

Using the maximum likelihood method, we next estimate the 9-variable VAR described in section 2.2 separately for three pairs of countries: US/Mexico, US/Brazil and US/Chile. In each case, the variables included in \mathbf{z}_t are the growth rates of the U.S. and foreign real output growth rates, (y_t, y_t^*) , the U.S. and foreign rates of inflation, (π_t, π_t^*) , the U.S. and foreign real interest rates, (i_t, i_t^*) , the depreciation of the real exchange rate $(\Delta \log S_t)$ and the U.S. and foreign real stock market returns, (r_t, r_t^*) .⁸

In the following discussion, we focus on the exchange rate equation given in (8). The equation can be written as

$$\Delta \log S_t = \mathbf{z}'_{t-1} \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_{t-1} + \mathbf{b}' \mathbf{z}_{t-1} + (\mathbf{C}_S \mathbf{z}_{t-1})' \varepsilon_t \quad (11)$$

where \mathbf{B}_S and \mathbf{C}_S are, respectively, 5×4 and 9×9 matrices whose elements are to be estimated, $\mathbf{b} = (0, 0, 0, 0, 1, -1, 0, 0, 0)'$, and \mathbf{A}_1 and \mathbf{A}_2 are some

⁷The sample used for US/Brazil is from January 1996 through December 2002.

⁸Given the large dimension of the model and limited data, we only allow one lag in our VAR model in the current paper.

constant matrices (see Appendix B). Under the symmetry assumption, the matrix \mathbf{C}_S has a simple structure with only 7 unknown coefficients. The estimates of the parameters in \mathbf{B}_S and \mathbf{C}_S are reported, respectively, in Tables 3 and 4.

There are two points worth mentioning. First, stochastic volatility appears to be an important characteristic of the exchange rate movements. This is because the conditional variance of $\Delta \log S_t$ given \mathbf{z}_{t-1} is determined by $\mathbf{z}'_{t-1} \mathbf{C}'_S \mathbf{C}_S \mathbf{z}_{t-1}$, and most estimates of the parameters in \mathbf{C}_S reported in Table 3 are highly significant. Many previous VAR studies on the exchange rate usually assume homoskedasticity. Our result suggests the importance of the stochastic volatility in order to understand the dynamic behavior of the exchange rate. It is interesting to note that the most significant estimates of the elements in matrix \mathbf{C}_S are C_{55} , C_{37} , C_{57} and C_{99} (see Table 3), where C_{55} is the coefficient on the interest rates i_{t-1} and i_{t-1}^* , both C_{37} and C_{57} are coefficients on the exchange rate $\Delta \log S_{t-1}$, and C_{99} is the coefficient on the lagged stock returns r_{t-1} and r_{t-1}^* . (see Appendix B for the definition of matrix \mathbf{C}_S and Section 2.2 for definition of \mathbf{z}_t). This suggests that the lagged financial market variables are the most important empirical factors determining the stochastic volatility of the exchange rate.

Second, we find that there is a substantial deviation from the uncovered interest parity and that the time-varying risk premiums are an important feature of the exchange rate movements. These findings are consistent with those in the large literature on the forward premium puzzle. In particular, given the exchange rate equation (11), the ex-ante UIP deviation can be expressed as

$$\begin{aligned} \phi_t &\equiv E_t [\Delta \log S_{t+1} - (i_t - i_t^*)] \\ &= \mathbf{z}'_t \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_t \end{aligned}$$

and the foreign exchange risk premium as defined in (A14) of Appendix A is given by

$$\begin{aligned} u_t &\equiv E_t [\Delta \log S_{t+1} - (i_t - i_t^*)] + \frac{1}{2} Var_t(\Delta \log S_{t+1}) \\ &= \mathbf{z}'_t \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_t + \frac{1}{2} \mathbf{z}'_t \mathbf{C}'_S \mathbf{C}_S \mathbf{z}_t. \end{aligned}$$

The significant estimates of \mathbf{B}_S and \mathbf{C}_S reported in Tables 3 and 4 are evidence against uncovered interest parity and constant foreign exchange risk premiums for all the countries we considered.

3.2 What macroeconomic risks are shared after stock market liberalization?

As argued in section 2.1 above, equation (1) implies that we may use the movement of the real exchange rate, $-\Delta \log S_t$, to measure the difference of the marginal utility growth rates, $\log M_t - \log M_t^*$, between the domestic and foreign investors. If there were perfect risk sharing across countries after stock market liberalization, $\log M_t$ would be equal to $\log M_t^*$ across *every* state of nature, and hence, we would obtain $\Delta \log S_t = 0$, i.e. the constant real exchange rate over time. In reality, even after stock market liberalization, the lack of complete markets and the existence of non-traded goods as well as transport costs most likely prevents full risk sharing across countries. When investors face a domestic shock that is difficult to diversify across national borders, the marginal utility growth ($\log M_t$) of the domestic investors would be driven apart from that of foreign investors ($\log M_t^*$), leading to fluctuations in $\Delta \log S_t$. On the other hand, if a particular economic shock is perfectly shared by the domestic and foreign investors, $\log M_t$ and $\log M_t^*$ would move in lockstep in response to the shock, leaving $\Delta \log S_t$ unchanged. As a result, this shock would account for only small portion of the volatility of the real exchange rate. Our structural VAR enables us to examine what macroeconomic risks, each corresponding to a specific fundamental shock ε_t , are shared across countries after stock market liberalization, and what macroeconomic risks are not.

For the above purpose, we calculate a variance decomposition for $\Delta \log S_t$, which is treated as a proxy for $\log M_t^* - \log M_t$. For conventional linear VAR models, the variance decomposition can be obtained as a transformation of the model parameters. For nonlinear models, however, no such a simple relation exists. Therefore, the variance decomposition is computed based on Monte Carlo simulations. Specifically, random shocks (ε_{t+j} , $j = 1, \dots, 4$) are drawn and the forecasting errors for $\Delta \log S_t$ are calculated from the estimated exchange rate equation. This process is repeated 500 times. The sample variances of the forecast errors due to each component of ε_{t+j} (namely $\varepsilon_{Y,t+j}$, $\varepsilon_{\Pi,t+j}$, $\varepsilon_{M,t+j}$ and $\varepsilon_{S,t+j}$) are then computed. Since the variances are state-dependent due to the nonlinearity of the exchange rate movement, we first compute the variance decomposition conditional on each observation of \mathbf{z}_t in our sample period (1988 to 2002). We then take the average of the variances across different states. Table 5 reports those variances as percentages of the overall volatility of the forecast errors of $\Delta \log S_t$, or $\log M_t^* - \log M_t$,

at each time horizon.⁹

The most striking finding from the above exercise is that exogenous financial shocks (the shocks to the exchange rate and stock market returns) that are orthogonal to other macroeconomic shocks, jointly account for only a very small fraction of the volatility of the domestic/foreign marginal utility growth differential ($\log M_t - \log M_t^*$). In the US/Mexico case, the financial shocks jointly account for about 3.78% of the volatility of $\log M_t^* - \log M_t$ four months ahead. Similar results are found for other countries as well, namely 10.5% in the US/Brazil case and 4.1% in the US/Chile case. Most of the variances of the marginal utility growth differentials are due to exogenous shocks to output, inflation and monetary policies.

To show the identified financial shocks are indeed important sources of asset market volatilities, we also compute the variance-decomposition for stock returns, using the US/Mexico, US/Brazil and US/Chile data. The results are reported in Tables 6, 7 and 8. Across all countries, we find that the financial shocks in fact account for the largest share (32% and up) of the variance of stock market returns, a result in sharp contrast with the finding reported in Table 5 that the financial shocks appear to have little impact on the marginal utility growth differentials across countries. In the US/Mexico case, the financial shocks jointly account for 78.8% and 57.2% of the volatility of the U.S. and Mexico stock market returns, respectively (Table 6). In the other two cases, the percentages are 33% and 32%, respectively, for the U.S. and Brazil stock market returns (Table 7), and 35.6% and 45.5%, respectively, for the U.S. and Chile stock market returns (Table 8).

In summary, the results imply that, after stock market liberalization, even though investors in the developing countries face a significant amount of financial market risks, the underlying financial shocks have little impact on the domestic/foreign marginal utility growth differentials. At the same time, however, international investors do not appear to fully share other macroeconomic risks such as exogenous shocks to output growth, inflation and monetary policies, as each of these shocks contributes significantly to the volatility of the marginal utility growth differential across countries. If stock market liberalization indeed contributes to the reduction of the cost of capital through increased risk sharing, the primary source of such benefit is

⁹The variance decompositions are reported up to 4 months in Table 5. A longer horizon beyond 4 months yields no change in the main result.

that the liberalization allows investors in developing countries better hedge against the risk of exogenous and idiosyncratic financial market shocks.

It should be emphasized that our results do not suggest that exogenous financial market shocks are less important in the post liberalization era. Those shocks could still account for a large fraction of the volatility of the marginal utility growth, $\log M_t$ or $\log M_t^*$, in each individual country, and hence present a major risk to investors across the world. Our results, however, imply that those risks are nonetheless shared very well between the developing countries and the U.S.

The above results are consistent with our knowledge about the incompleteness of financial markets. Full risk sharing through the existing financial markets requires that asset returns span the space of the underlying economic shocks, a proposition that is strongly rejected by empirical evidence (Davis et al., 2000). For example, while labor earnings account for a major portion of national income, they are not securitized because of the non-marketable nature of human capital. Labor income in turn is shown to have near-zero correlation with aggregate equity returns (see Fama and Schwert, 1977 and Botazzi et al. 1996). On the other hand, aggregate supply and demand shocks as well as monetary policy shocks are probably most responsible for the uncertainties in labor income.

This lack of perfect risk sharing may also reflect the impact of non-traded consumption goods and transport cost, which can be affected significantly by shocks to aggregate consumption, inflation and monetary policies. Suppose that a typical investor's utility function is characterized by non-separability between traded and non-traded consumption goods. The macroeconomic shocks would then drive the marginal utility growth of domestic investors away from that of foreign investors. As a result, there would be large fluctuations in the movement of the real exchange rate.

One major caveat of this study is the reliance on the movement of the real exchange rate ($\Delta \log S_t$) as a proxy of the marginal utility growth differential ($\log M_t - \log M_t^*$) across countries (the key equation in Section 2.1). This relation is based on the assumption of no-arbitrage in international asset markets. However, equation (1) will not hold for arbitrary pairs of domestic and foreign stochastic discount factors (M_t and M_t^*) if markets are incomplete.¹⁰ In such a case, $\log M_t$ and $\log M_t^*$ in equation (1) should

¹⁰ Absence of arbitrage, however, is sufficient for the existence of the stochastic discount

be interpreted as the linear projection of the marginal utility growth onto the space of asset returns as pointed out by Brandt et al. (2001). Therefore, the movements of the real exchange rate may not be exactly linked to the fluctuations in the marginal utility growth differential. The movement of investors' marginal utility growth orthogonal to asset returns is not captured by our empirical model. For example, risk sharing can also be achieved through international government transfers and aids, rather than through financial markets. Our empirical exercise therefore only addresses the issue of what macroeconomic risks are or are not shared by international investors through the *existing* asset markets.

4 Concluding Remarks

Whether or not developing countries should liberalize their capital markets has been a highly controversial issue in recent years. Standard economic theory suggests that one of the main benefits of financial liberalization is to achieve better international risk sharing, and hence to reduce the equity premium and the cost of capital of the liberalizing country. However, the volatility in the financial markets of developing countries under speculative money flows seems to question such benefits of stock market liberalization.

In this paper, we find empirical evidence, using a sample of developing countries that had liberalized their stock markets, that most of the risk of exogenous financial market shocks are in fact very well shared internationally, even though other macroeconomic risks such as those associated with exogenous shocks to output growth, inflation and monetary policies are not fully shared across countries. On one hand, our results imply that stock market liberalization indeed helps developing countries better insure against exogenous and idiosyncratic financial shocks, therefore contributes to the reduction of the cost of capital in those countries as documented by other studies. On the other hand, our results also suggest that stock market liberalization should be accompanied by other macroeconomic reforms in order to achieve the full benefits of international risk sharing.

factors. See, for example, Harrison and Kreps (1979).

Appendix A: derivation of the key relations in Section 2.1

Absence of arbitrage in asset markets implies that there exists a positive stochastic discount factor M_{t+1} such that for any domestic asset (Harrison and Kreps, 1979),

$$1 = E_t(M_{t+1}R_{t+1}) \quad (\text{A1})$$

where R_{t+1} is the real gross rate of return on the domestic asset between time t and $t + 1$, and expectation is taken with respect to the investors' information set at time t . In various versions of the consumption-based asset pricing model, M_{t+1} is equal to $u'(c_{t+1})/u'(c_t)$ where $u'(c_t)$ is the marginal utility of consumption at time t . Hence, $\log M_{t+1}$ is simply the growth rate of marginal utility of domestic investors.

Let S_t be the real exchange rate (in units of domestic goods/foreign goods) between the two countries. Then for any foreign asset that can be purchased by domestic investors, (A1) implies

$$1 = E_t \left[M_{t+1} \left(\frac{S_{t+1}}{S_t} \right) R_{t+1}^* \right] \quad (\text{A2})$$

where R_{t+1}^* is the real gross rate of return in terms of foreign goods. But for foreign investors, absence of arbitrage implies that there must also exist a foreign stochastic discount factor satisfying

$$1 = E_t(M_{t+1}^*R_{t+1}^*). \quad (\text{A3})$$

where $\log M_{t+1}^*$ may be similarly interpreted as the growth rate of marginal utility of foreign investors.

If markets are complete, M_{t+1} and M_{t+1}^* are unique. Therefore, (A2) and (A3) imply that

$$\frac{S_{t+1}}{S_t} = \frac{M_{t+1}^*}{M_{t+1}} \quad (\text{A4})$$

or, in terms of logarithms,

$$\log S_{t+1} - \log S_t = -(\log M_{t+1} - \log M_{t+1}^*). \quad (\text{A5})$$

If markets are incomplete, there would be multiple stochastic discount factors that satisfy (A1) and (A3), and equation (A5) would not hold for arbitrary pairs of domestic and foreign discount factors. However, as pointed

out by Brandt et al. (2001), the projection of the discount factor (or the marginal utility growth) on the space of asset returns is unique. Therefore (A5) still holds for the so-called minimum variance discount factors.

We assume that M_{t+1} and M_{t+1}^* follow the log-normal distribution according to

$$\log M_{t+1} = \mu_t - \lambda_t' \varepsilon_{t+1} \quad (\text{A6})$$

$$\log M_{t+1}^* = \mu_t^* - \lambda_t^{*'} \varepsilon_{t+1} \quad (\text{A7})$$

where ε_t is distributed as $\mathcal{N}(\mathbf{0}, \mathbf{I})$. Now consider a one-period risk-free pure discount bond with continuously compounded real interest rates i_t and i_t^* in the home and foreign country, respectively. (A1) and (A3) implies that

$$i_t = -\log(E_t M_{t+1}) \quad (\text{A8})$$

$$i_t^* = -\log(E_t M_{t+1}^*) \quad (\text{A9})$$

Using (A6) - (A9), we can express μ_t and μ_t^* as

$$\mu_t = -(i_t + \frac{1}{2} \lambda_t' \lambda_t) \quad (\text{A10})$$

$$\mu_t^* = -(i_t^* + \frac{1}{2} \lambda_t^{*'} \lambda_t^*). \quad (\text{A11})$$

Note that if M_{t+1} and M_{t+1}^* are not distributed as log-normal, (A10) and (A11) still hold as the second order approximations to (A8) and (A9) respectively, as shown in Backus et al (2001). Equations (A10) and (A11) together with (A5) imply

$$\Delta \log S_{t+1} - (i_t - i_t^*) = \frac{1}{2} (\lambda_t' \lambda_t - \lambda_t^{*'} \lambda_t^*) + (\lambda_t - \lambda_t^*)' \varepsilon_{t+1}. \quad (\text{A12})$$

It is easy to see from (A12) that the conventional uncovered interest rate parity (UIP) does not hold in general. And the UIP deviation ϕ_t can be expressed as a quadratic function of the home and foreign country's market prices of risk

$$\phi_t = \frac{1}{2} (\lambda_t' \lambda_t - \lambda_t^{*'} \lambda_t^*). \quad (\text{A13})$$

Decompose ϕ_t as $\phi_t = u_t + v_t$, where $u_t = (\lambda_t - \lambda_t^*)' \lambda_t$ and $v_t = -\frac{1}{2} (\lambda_t - \lambda_t^*)' (\lambda_t - \lambda_t^*)$. Note that, using equation (A6) and (A12), u_t can be expressed as

$$u_t = Cov_t[\Delta \log S_{t+1} - (i_t - i_t^*), -\log M_{t+1}], \quad (\text{A14})$$

In other words, u_t is the conditional covariance between the excess return on the foreign exchange and the log of the stochastic discount factor, therefore the risk premium from investing in the foreign exchange. By (A6), we can write u_t as $u_t = \sum_{i=1}^N \lambda_{i,t} \text{Cov}_t[\Delta \log S_{t+1} - (i_t - i_t^*), \varepsilon_{i,t+1}]$, which explains why λ_t is called the market price of risk. The i th component λ_{it} of λ_t prices the covariance between the foreign exchange return and the i th fundamental economic shock. For example, if $\varepsilon_{i,t+1}$ is an exogenous shock to monetary policy in the home country, then the risk associated with the policy when investing in the foreign exchange is characterized by the conditional covariance between the foreign exchange return and the policy shock, and $\lambda_{i,t}$ is the expected excess rate of return per unit of such covariance. Note that similar results hold for the foreign country as well. The foreign exchange risk premium for foreign investors can be expressed as $u_t^* = \sum_{i=1}^{N+1} \lambda_{i,t}^* \text{Cov}_t[-\Delta \log S_{t+1} - (i_t^* - i_t), \varepsilon_{i,t+1}]$, and the similar interpretation applies to $\lambda_{i,t}^*$.

The second term v_t in the decomposition of ϕ_t is simply the Jensen's inequality term when taking logarithm of the foreign exchange return, or

$$v_t = -\frac{1}{2} \text{Var}_t[\Delta \ln S_{t+1} - (i_t - i_t^*)] \quad (\text{A15})$$

This term does not have any economic significance and disappears in a continuous time setting.¹¹

Next, we assume that the domestic and foreign real gross stock returns R_t and R_t^* also have log-normal distribution

$$\log R_{t+1} = E_t(r_{t+1}) + \sigma' \varepsilon_{t+1} \quad (\text{A16})$$

$$\log R_{t+1}^* = E_t(r_{t+1}^*) + \sigma^{*'} \varepsilon_{t+1} \quad (\text{A17})$$

where $r_{t+1} = \log R_{t+1}$ and $r_{t+1}^* = \log R_{t+1}^*$. Using (A1) and (A3) again, we have

$$r_{t+1} - i_t = -\frac{1}{2} \sigma' \sigma + \sigma' \lambda_t + \sigma' \varepsilon_{t+1} \quad (\text{A18})$$

$$r_{t+1}^* - i_t^* = -\frac{1}{2} \sigma^{*'} \sigma^* + \sigma^{*'} \lambda_t^* + \sigma^{*'} \varepsilon_{t+1}. \quad (\text{A19})$$

¹¹It is, however, interesting to note that both the conditional volatility of the exchange rate and the risk premium are determined by the home and foreign country's market price of risk. Since in the finance literature the market price of risk is routinely treated as time-varying, it is not surprising that movements of the exchange rate are characterized by stochastic volatilities and time-varying risk premiums.

It is assumed for simplicity that the stock returns have constant volatilities (i.e. σ and σ^* are independent of time). Note that the risk premiums are still time-varying under this specification because of the presence of the market prices of risk λ_t and λ_t^* .

Similar interpretations can be made for the excess stock returns in (A18) and (A19). The risk premiums on the domestic and foreign stocks can be expressed as, after adjusting for the Jensen's inequality term,

$$\begin{aligned} E_t(r_{t+1} - i_t) + \frac{1}{2}\sigma'\sigma &= \text{Cov}_t(r_{t+1} - i_t, -\log M_{t+1}) \\ &= \sum_{i=1}^N \lambda_{i,t} \text{Cov}_t[r_{t+1} - i_t, \varepsilon_{i,t+1}] \end{aligned} \quad (\text{A20})$$

$$\begin{aligned} E_t(r_{t+1}^* - i_t^*) + \frac{1}{2}\sigma^{*'}\sigma^* &= \text{Cov}_t(r_{t+1}^* - i_t^*, -\log M_{t+1}^*) \\ &= \sum_{i=1}^N \lambda_{i,t}^* \text{Cov}_t[r_{t+1}^* - i_t^*, \varepsilon_{i,t+1}]. \end{aligned} \quad (\text{A21})$$

Appendix B: Identification of Γ and Γ^*

We rewrite (9) and (10) here for convenience

$$r_t = -\frac{1}{2}\sigma'\sigma + \beta'\mathbf{z}_{t-1} + i_{t-1} + \sigma'\varepsilon_t \quad (\text{B1})$$

$$r_t^* = -\frac{1}{2}\sigma^{*'}\sigma^* + \beta^{*'}\mathbf{z}_{t-1} + i_{t-1}^* + \sigma^{*'}\varepsilon_t \quad (\text{B2})$$

The first identifying restriction is based on the assumption that home and foreign investors price risk factors in a symmetrical fashion as described below. For example, let us consider the first two elements of ε_t : the shocks to the home and foreign country's consumption growth ($\varepsilon_{y,t}$ and $\varepsilon_{y,t}^*$). To investors in the home country, the foreign exchange risk associated with the shock to home consumption is $\text{Cov}_{t-1}[\Delta \log S_t, \varepsilon_{y,t}]$,¹² while to investors in the foreign country the foreign exchange risk associated with foreign consumption is $\text{Cov}_{t-1}[-\Delta \log S_t, \varepsilon_{y,t}^*]$. We assume that if the market price of the risk (or the expected excess rate of return per unit of the covariance) in the home country is given by

$$\lambda_{1,t} = \Gamma_{11}y_t + \Gamma_{12}y_t^* + \Gamma_{13}\pi_t + \Gamma_{14}\pi_t^* + \Gamma_{15}i_t + \Gamma_{16}i_t^* + \Gamma_{17}\Delta \log S_t + \Gamma_{18}r_t + \Gamma_{19}r_t^*$$

then the foreign counterpart is given by

$$\lambda_{2,t}^* = \Gamma_{12}y_t + \Gamma_{11}y_t^* + \Gamma_{14}\pi_t + \Gamma_{13}\pi_t^* + \Gamma_{16}i_t + \Gamma_{15}i_t^* - \Gamma_{17}\Delta \log S_t + \Gamma_{19}r_t + \Gamma_{18}r_t^*$$

where Γ_{ij} refers to the element on the i th row and j th column of matrix Γ . And similar parameterizations apply to $\lambda_{i,t}$ and $\lambda_{i,t}^*$ for $i = 2, \dots, 6$ and $i = 8, 9$. For $i = 7$ we need a little different condition. Note that if the foreign exchange risk due to the exogenous shock $\varepsilon_{S,t}$ is expressed as $\text{Cov}_{t-1}[\Delta \log S_t, \varepsilon_{S,t}]$ for home investors, its foreign analogy would be $\text{Cov}_{t-1}[-\Delta \log S_t, -\varepsilon_{S,t}]$. Hence, the symmetric assumption implies that if $\varepsilon_{S,t}$ has a market price of risk in the home country given by

$$\lambda_{7,t} = \Gamma_{71}y_t + \Gamma_{72}y_t^* + \Gamma_{73}\pi_t + \Gamma_{74}\pi_t^* + \Gamma_{75}i_t + \Gamma_{76}i_t^* + \Gamma_{77}\Delta \log S_t + \Gamma_{78}r_t + \Gamma_{79}r_t^*$$

then in the foreign country its market price of risk must be

$$\lambda_{7,t}^* = -\Gamma_{72}y_t - \Gamma_{71}y_t^* - \Gamma_{74}\pi_t - \Gamma_{73}\pi_t^* - \Gamma_{76}i_t - \Gamma_{75}i_t^* + \Gamma_{77}\Delta \log S_t - \Gamma_{79}r_t - \Gamma_{78}r_t^*$$

In summary, the symmetric treatment of the market prices of risk across countries implies that $\Gamma^* = \mathbf{A}\Gamma\mathbf{A}$ where $\mathbf{A} = \text{diag}\{\mathbf{J}, \mathbf{J}, \mathbf{J}, -1, \mathbf{J}\}$

¹²We have ignored the term $(i_{t-1} - i_{t-1}^*)$ here since it does not affect the conditional covariance.

with $\mathbf{J} = \begin{bmatrix} 0 & 1 \\ 1 & 0 \end{bmatrix}$. Pre and post multiplication of matrix \mathbf{A} has an effect on matrix $\mathbf{\Gamma}$ in the following manner: First it changes the position of the first and second rows, the third and fourth rows, the fifth and sixth rows, and the eighth and ninth rows of matrix $\mathbf{\Gamma}$ and then changes the sign of the seventh row. Second, it changes the position of the first and second columns, the third and fourth columns, the fifth and sixth columns, and eighth and ninth columns of matrix $\mathbf{\Gamma}$ and then changes the sign of the seventh column.

With this restriction, equation (8) can be expressed as

$$\Delta \log S_t = \mathbf{z}'_{t-1} \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_{t-1} + \mathbf{b}' \mathbf{z}_{t-1} + (\mathbf{C}_S \mathbf{z}_{t-1})' \varepsilon_t \quad (\text{B3})$$

where \mathbf{B}_S and \mathbf{C}_S are respectively 5×4 and 9×9 matrices whose elements are to be estimated, $\mathbf{b} = (0 \ 0 \ 0 \ 0 \ 1 \ -1 \ 0 \ 0 \ 0)'$, and the matrices \mathbf{A}_1 and \mathbf{A}_2 are given by

$$\mathbf{A}_1 = \begin{pmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & -1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & -1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & -1 & 1 \end{pmatrix}$$

and

$$\mathbf{A}_2 = \begin{pmatrix} 1 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 1 \end{pmatrix}$$

To see this, note that $\mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^*\mathbf{\Gamma}^*)\mathbf{z} = \mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{A}\mathbf{\Gamma}'\mathbf{\Gamma}\mathbf{A})\mathbf{z} = \mathbf{z}'\mathbf{\Gamma}'\mathbf{\Gamma}\mathbf{z} - \tilde{\mathbf{z}}'\mathbf{\Gamma}'\mathbf{\Gamma}\tilde{\mathbf{z}} = (\mathbf{z} - \tilde{\mathbf{z}})'\mathbf{\Gamma}'\mathbf{\Gamma}(\mathbf{z} + \tilde{\mathbf{z}})$, where $\tilde{\mathbf{z}} = \mathbf{A}\mathbf{z}$. Now

$$\mathbf{z} - \tilde{\mathbf{z}} = \begin{bmatrix} z_1 - z_2 \\ z_2 - z_1 \\ z_3 - z_4 \\ z_4 - z_3 \\ z_5 - z_6 \\ z_6 - z_5 \\ 2z_7 \\ z_8 - z_9 \\ z_9 - z_8 \end{bmatrix} \quad \text{and} \quad \mathbf{z} + \tilde{\mathbf{z}} = \begin{bmatrix} z_1 + z_2 \\ z_2 + z_1 \\ z_3 + z_4 \\ z_4 + z_3 \\ z_5 + z_6 \\ z_6 + z_5 \\ 0 \\ z_8 + z_9 \\ z_9 + z_8 \end{bmatrix}$$

Note also that

$$\begin{bmatrix} z_1 - z_2 \\ z_3 - z_4 \\ z_5 - z_6 \\ z_7 \\ z_8 - z_9 \end{bmatrix} = \mathbf{A}_1 \mathbf{z} \quad \text{and} \quad \begin{bmatrix} z_1 + z_2 \\ z_3 + z_4 \\ z_5 + z_6 \\ z_8 + z_9 \end{bmatrix} = \mathbf{A}_2 \mathbf{z}.$$

Therefore, if there is no restriction on $\mathbf{\Gamma}$, we can express the original quadratic form as $\mathbf{z}'(\mathbf{\Gamma}'\mathbf{\Gamma} - \mathbf{\Gamma}^*\mathbf{\Gamma}^*)\mathbf{z} = \mathbf{z}'\mathbf{A}'_1\mathbf{B}_S\mathbf{A}_2\mathbf{z}$ as claimed.

The second set of restrictions is based on another type of symmetric assumption to simplify the expression of matrix \mathbf{C}_S . We assume, for example, the contribution of y^* to the market price of home consumption risk is assumed to be equal in size to the contribution of y to the market price of foreign consumption risk. This type of symmetric assumption implies restrictions on matrix $\mathbf{\Gamma}$ in the form of $\Gamma_{1+2i,1+2j} = \Gamma_{2+2i,2+2j}$ and $\Gamma_{1+2k,2+2l} = \Gamma_{2+2k,1+2l}$ for $i, j, k, l = 0, 1, 2$ and $i \neq j$. It makes all off-diagonal elements of \mathbf{C}_S except the seventh row and the seventh column equal to zero. The resulting matrix \mathbf{C}_S becomes

$$\mathbf{C}_S = \begin{pmatrix} C_{11} & 0 & 0 & 0 & 0 & 0 & C_{17} & 0 & 0 \\ 0 & -C_{11} & 0 & 0 & 0 & 0 & C_{17} & 0 & 0 \\ 0 & 0 & C_{33} & 0 & 0 & 0 & C_{37} & 0 & 0 \\ 0 & 0 & 0 & -C_{33} & 0 & 0 & C_{37} & 0 & 0 \\ 0 & 0 & 0 & 0 & C_{55} & 0 & C_{57} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & -C_{55} & C_{57} & 0 & 0 \\ C_{17} & C_{17} & C_{37} & C_{37} & C_{57} & C_{57} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & C_{99} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & -C_{99} \end{pmatrix}$$

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Table 1: Summary Statistics

	Mexico		Brazil		Chile		
	US Stock	Stock	X-Rate	Stock	X-Rate	Stock	X-Rate
	Returns(%)						
Mean	1.09	-5.75	19.52	-12.53	24.80	-2.16	11.36
Std Dev	22.01	45.25	26.20	54.12	38.96	35.79	13.90
	Return Correlations						
US Stock		0.48	-0.02	0.58	-0.18	0.36	-0.10
Foreign Stock			-0.13		-0.07		-0.16

Note: This table shows annualized summary statistics for real excess returns on stock indices and exchange rates for the US, Mexico, Brazil, and Chile. The stock indices are total market returns from Datastream, the interest rates are for one-month Eurocurrency deposits from Datastream, and the CPI is from the International Monetary Fund's IFS database. The stock returns (Stock) are excess returns over the same country's one-month interest rates. The exchange rate returns (X-Rate) are excess returns for borrowing in dollars, converting to the foreign currency, lending at the foreign interest rate, and converting the proceeds back to dollars. Monthly data from January 1988 through December 2002 for Mexico and Chile, from January 1996 through December 2002 for Brazil.

Table 2: Risk Sharing Index			
	US vs Mexico	US vs Brazil	US vs Chile
	Risk Sharing Index		
	0.917	0.791	0.984
	(0.066)	(0.219)	(0.007)
	Std Dev of Marginal Utility Growth		
Domestic	0.76	0.73	0.83
	(0.11)	(0.17)	(0.10)
Foreign	0.50	0.43	0.69
	(0.15)	(0.22)	(0.11)

Note: This table shows the risk sharing index and annualized standard deviations of the discount factors recovered from asset markets. The domestic country is the US and the foreign country is Mexico, Brazil, and Chile. The investable assets are the domestic interest rate, the domestic stock market, foreign interest rate, and the foreign stock market. The risk sharing index is defined in Cochrane et al (2003) Figures in parentheses are standard errors calculated with the bootstrap method.

Table 3: Estimates of the matrix \mathbf{C}_S

	US/Mexico Ex-rate	US/Brazil Ex-rate	US/Chile Ex-rate
C_{11}	0.2181 (0.2571)	-0.8104 (1.7437)	-0.4839 (0.2266)
C_{33}	0.6018 (0.3344)	1.1989 (0.8436)	-0.0125 (0.1861)
C_{55}	-1.4411 (0.1906)	0.2822 (0.6825)	0.2804 (0.1028)
C_{17}	-0.2655 (0.1443)	-0.8597 (0.5150)	0.2660 (0.1279)
C_{37}	0.8462 (0.1148)	1.3972 (0.4414)	0.6721 (0.0757)
C_{57}	0.9304 (0.0602)	1.4593 (0.3194)	1.1905 (0.0547)
C_{99}	-0.1203 (0.0359)	-0.2845 (0.1218)	0.1088 (0.0188)

Note: This table reports the estimates of the elements of the 9×9 matrix \mathbf{C}_S , whose definition can be found in Appendix B. The figures in parentheses are the robust standard errors. Under the symmetry assumption, \mathbf{C}_S has 7 unknown parameters. C_{ij} represents the element on the i th row and j th column of the matrix. The exchange rate equation is given by $\Delta \ln S_t = \mathbf{z}'_{t-1} \mathbf{A}_1' \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_{t-1} + b' \mathbf{z}_{t-1} + (\mathbf{C}_S \mathbf{z}_{t-1})' \varepsilon_t$. Hence \mathbf{C}_S determines the conditional variance of $\Delta \ln S_t$, which can be obtained as: $Var_{t-1}(\Delta S_t) = \mathbf{z}'_{t-1} \mathbf{C}'_S \mathbf{C}_S \mathbf{z}_{t-1}$.

Table 4: Estimates of the matrix \mathbf{B}_S

	US/Mexico Ex-rate	US/Brazil Ex-rate	US/Chile Ex-rate
B_{11}	-0.0256 (0.0223)	-0.3223 (0.3898)	0.0860 (0.0583)
B_{12}	-0.0196 (0.0773)	-0.3620 (0.3468)	0.1230 (0.0673)
B_{13}	-0.0894 (0.0413)	0.0731 (0.1038)	0.0143 (0.0325)
B_{14}	-0.0304 (0.0129)	0.0010 (0.0518)	0.0529 (0.0250)
B_{21}	-0.0043 (0.0094)	-0.0651 (0.0327)	0.0101 (0.0061)
B_{22}	-0.0298 (0.0740)	0.5710 (0.5441)	-0.0439 (0.0538)
B_{23}	-0.0515 (0.037)	-0.1571 (0.2002)	-0.0968 (0.0268)
B_{24}	-0.0702 (0.0252)	0.1787 (0.1206)	-0.0215 (0.0179)
B_{31}	-0.0183 (0.0061)	-0.2116 (0.0944)	0.0327 (0.0137)
B_{32}	0.0090 (0.0063)	-0.0793 (0.0441)	0.0022 (0.0051)
B_{33}	0.0429 (0.0331)	-0.4318 (0.2156)	0.0653 (0.0394)
B_{34}	0.0639 (0.0314)	-0.1911 (0.1298)	0.0299 (0.0249)
B_{41}	-0.0214 (0.0067)	-0.0498 (0.0270)	0.0023 (0.0059)
B_{42}	0.0071 (0.0054)	0.0458 (0.0308)	-0.0042 (0.0110)
B_{43}	-0.0030 (0.0048)	0.0331 (0.0142)	0.0012 (0.0025)
B_{44}	0.0070 (0.0069)	0.0226 (0.0317)	-0.0012 (0.0058)
B_{51}	0.0113 (0.0089)	0.0833 (0.0382)	0.0013 (0.0064)
B_{52}	-0.0011 (0.0027)	-0.0088 (0.0069)	0.0012 (0.0021)
B_{53}	-0.0020 (0.0014)	0.0001 (0.0028)	0.0014 (0.0022)
B_{54}	0.0001 (0.0006)	-0.0009 (0.0010)	0.0002 (0.0003)

Note: This table reports the estimates of the elements of the 5×4 matrix \mathbf{B}_S . B_{ij} represent the element on the i th row and j th column of the matrix. The figures in parentheses are the robust standard errors. The exchange rate equation is given by $\Delta \ln S_t = \mathbf{z}'_{t-1} \mathbf{A}_1' \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_{t-1} + (i_t - i_t^*) + (\mathbf{C}_S \mathbf{z}_{t-1})' \varepsilon_t$. Hence \mathbf{B}_S determines the ex-ante UIP deviation, which can be obtained as $E_t[\Delta \log S_{t+1} - (i_t - i_t^*)] = \mathbf{z}'_t \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_t$, where the matrixes \mathbf{A}_1 and \mathbf{A}_2 are given in Appendix B. Moreover, the foreign exchange risk premium as defined in (5) can be obtained as $E_t(\Delta \log S_{t+1} - (i_t - i_t^*)) + \frac{1}{2} \text{Var}_t(\Delta \log S_{t+1}) = \mathbf{z}'_t \mathbf{A}'_1 \mathbf{B}_S \mathbf{A}_2 \mathbf{z}_t + \frac{1}{2} \mathbf{z}'_t \mathbf{C}'_S \mathbf{C}_S \mathbf{z}_t$

Table 5: Variance decomposition of $\Delta \log S_t$

US vs Mexico				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0151	0.1415	0.4482	0.3952
2 month	0.0580	0.3296	0.5055	0.1068
3 month	0.0619	0.3863	0.5045	0.0473
4 month	0.0730	0.4046	0.4845	0.0378
US vs Brazil				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0735	0.1604	0.1816	0.5845
2 month	0.1848	0.3419	0.3553	0.1180
3 month	0.1869	0.3452	0.3713	0.0966
4 month	0.2076	0.3509	0.3363	0.1052
US vs Chile				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0218	0.1176	0.3645	0.4962
2 month	0.0624	0.2149	0.5913	0.1314
3 month	0.0540	0.2327	0.6618	0.0515
4 month	0.0634	0.2483	0.6472	0.0410

Note: This table reports the results of variance decomposition for the US/foreign marginal utility growth differential $\log M_t - \log M_t^*$, which is approximated by the movement of the real exchange rate $-\Delta \log S_t$.

Table 6: Variance decomposition of the US and Mexico stock returns

The U.S. stock market returns				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0180	0.0247	0.0197	0.9376
2 month	0.0522	0.0577	0.0533	0.8369
3 month	0.0572	0.0628	0.0580	0.8219
4 month	0.0684	0.0742	0.0699	0.7875
Mexico stock market returns				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0006	0.0005	0.0107	0.9882
2 month	0.0419	0.0419	0.0502	0.8659
3 month	0.0854	0.0853	0.0921	0.7371
4 month	0.1410	0.1409	0.1456	0.5724

Table 7: Variance decomposition of the US and Brazil stock returns

The U.S. stock market returns				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0842	0.1825	0.0071	0.7261
2 month	0.1367	0.2048	0.0821	0.5764
3 month	0.1701	0.2167	0.1318	0.4814
4 month	0.2198	0.2383	0.2056	0.3363
Brazil stock market returns				
	Output shock	inflation shock	monetary shock	financial shock
1 month	0.0577	0.0665	0.0072	0.8687
2 month	0.1528	0.1578	0.1299	0.5596
3 month	0.1805	0.1864		