

Mexico's integration into the North American capital market

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Abstract

We explore a model of time varying regional market integration that includes three factors for the North American equity market, the local Mexican equity market and the peso/dollar exchange rate. We argue that a useful instrument for the degree of integration is the sovereign yield spread. Applying our methodology to Mexico over the 1991–2002 period, we show that the degree of market integration was higher at the end of the period than at the beginning but that it exhibited wide swings that were related to both global as well as local events. We also discover that Mexico's currency risk is priced. Further, the currency returns process reveals strongly significant asymmetric volatility that is strongly related to the asymmetric volatility of the Mexican equity market returns process. A plausible reason for these results is that currency devaluations in emerging markets like Mexico can cause default-risk crises in local banking systems that mismatch local-currency assets and hard currency liabilities, whereas appreciations produce no such problems. Devaluations that destabilize banking systems are, therefore, more likely than appreciations to increase the volatilities of both the currency's and the equity market's returns.

Keywords: Integration; Currency; Equity market

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1. Introduction and summary

The purpose of this paper is to extend what we know about the processes through which emerging capital markets become integrated into world markets. In the spirit of the current literature, we use the word integration loosely. Strictly speaking, partial segmentation exists in any market when a given identifiable subset of investors is barred from access to a given list of securities. We do not test this possibility directly. Instead, we characterize a market as less than fully integrated when the local factor premium is large relative to the regional or world factor premium, a condition that could also have other causes. We investigate the issue through a longitudinal study of a single stock market, Mexico's BOLSA, using weekly data over the eleven-year period from January 1991 to February 2002.

Our paper takes as its point of departure the methodology of Bekaert and Harvey (Bekaert and Harvey, 1995), which itself has roots in Errunza and Losq (1985, 1989) but provides an intertemporal, time varying extension of their static model. We choose not to compare directly the pricing of risk in the Mexican and North American stockmarkets for the following reason. Mexico's equity market is an emerging market: Harvey (1995a), Bekaert and Harvey (1995) and Bekaert and Urias (1996) conclude that there are significant local components in emerging market asset returns. The sources and the number of the risk factors of emerging markets may be dramatically different from those of developed countries. This multiplicity of factors makes it difficult to estimate differences between prices of risk or to compute and compare costs of equity capital in different countries. The basic model of Bekaert and Harvey (1995) combines the domestic and international versions of the CAPM. It tests the power of local factors relative to that of common factors to explain expected returns and empirically infers segmentation when the weight of the local factors is high. We use a simple version of their approach. As we are interested in the integration problem for the Mexican market as a whole, we employ the market index rather than individual stock returns as our dependent variable.

Within the limits of defining integration as the weight of the North American factor premium in Mexican expected returns, we shall argue that integration in general is influenced by global as much as by local events and does not necessarily start immediately after local capital markets are liberalized. We further claim that in Mexico's case the integration process is strongly and negatively related to the market's changing assessment of sovereign default risk.² The result is to relate the process in a very intuitive way to ongoing news that changes investors' risk perceptions but that need not be reflecting only domestic regime changes. Behind our results lies a simple general observation. Emerging capital markets become integrated when world investors notice them and add them to their portfolios. These markets show signs of disassociation that is greater partial segmentation, when

² Credit ratings were used as a factor to explain expected returns by Erb, Harvey and Viskanta (1996). Credit risk is priced also in Bansal and Dahlquist (2002). Credit ratings and the sovereign yield spreads that we use are highly correlated cross sectionally: see Erb et al. (2000).

international investors liquidate their positions because of increasing worries over sovereign default risk and take their money out.^{3,4}

We differ from Bekaert and Harvey (1995) in several additional ways. First, like them, we define our measure of integration as the relative weight of the common factor (in our case the North American market factor) in the Mexican expected returns equation. They make their integration measure a logistic function of the same set of information variables as those to which they relate their market prices of risk. We, however, need an integration measure that depends on variables that can be observed weekly. Therefore, we make our integration measure depend on the excess stripped spread of the Mexican Par bond over a similar duration US Treasury. Changes in sovereign yield spreads over the US Treasury, like ratings changes, generally reflect changes in bond markets' perceptions of an indebted country's creditworthiness and political riskiness. Sudden increases in sovereign risk perceptions cause liquidity to dry up and capital to flow out of debtors' equity markets and to stay out, sometimes for appreciable periods of time. This is observationally similar to what would have happened had the market become suddenly, if temporarily, segmented: some investors behave as if access to the credit-risky market's securities had become suddenly restricted. Sovereign default risk can make emerging markets appear to be segmented just as surely as do barriers to capital flows to the extent that it causes some assets to be systematically excluded from investors' portfolios at least for a time.⁵

Second, following the suggestion in Bekaert and Harvey (1995) followed by Domowitz et al. (1998), we incorporate a pure currency asset and discover that nominal currency risk is priced. We find that the price of currency risk was negative on the average and significantly different from zero. More specifically, it was negative for most of the sample period, but took on sometimes large absolute values during crises. There is a possibility that we examine in the body of the paper that currency risks are related to sovereign credit risks as the effects of the latter on equity and fixed income capital flows simultaneously affect the nominal exchange rate. The processes of the integration measure and the estimated price of currency risk have a small negative correlation of -0.02 , which is not statistically significant. At most, this indicates possibly but plausibly that when integration is higher, the

³ Table 8 in Bekaert et al. (2003) shows that a minimum level of credit rating is necessary for a liberalization to impact investment positively. We have not performed the parallel analysis of whether there are threshold levels of spreads associated with shifts in portfolio investment capital flows.

⁴ While much of the emerging markets literature seems to imply that improved integration is associated with higher correlation among markets, it is useful to reiterate that this is not necessarily the case. Segmentation is about whether risks are priced differently across markets, not about correlation. As Adler and Dumas (1983, p. 967) pointed out markets can be completely integrated and still have zero or even negative correlations depending among other things on the compositions of their outputs and the idiosyncrasies of their politics.

⁵ An obvious case could be made that default risks are diversifiable and should not be priced. However, a mounting body of evidence following Bekaert (1995) and Bekaert et al. (1995, 1997) suggests that political risk on all its dimensions is indeed priced in emerging markets and that credit ratings are strongly related cross sectionally to an integration measure.

price of currency risk is slightly smaller.⁶ The correlation between currency returns and Mexico equity returns (all in US dollars) is 0.549. The correlation between the variances of these returns, measured in moving 10-week windows, is 0.570. These results all suggest that the two processes may be strongly interlinked.

Third, we investigate asymmetric volatility in the Mexican market and currency processes. The nominal returns on Mexico's BOLSA index reveal marked departures from the normal distribution: fat tails; high negative skewness and excessive leptokurtosis. All stock market returns are likely to share these characteristics to the degree that stock prices and volatility are related. This can be inferred from Das and Sundaram (1998) who consider a standard continuous-time stochastic volatility model with a mean-reverting square-root volatility process. They find first that the sign of the skewness is determined by that of the coefficient of correlation between stock prices and volatility and second that the size of the skewness is proportional to that of the coefficient. The result that in the US the correlation coefficient is negative dates back to Black (1976) who ascribed this asymmetric volatility to leverage.⁷ When stock prices fall, corporate leverage and riskiness both rise making stocks more volatile when prices fall than when they rise. We confirm this pattern with strong evidence of asymmetric volatility in Mexican returns although we are unable to determine which of the BOLSA or the currency returns processes is its primary source. We further record that the currency returns process itself reveals highly significant asymmetric volatility and that it is associated via cross effects with the asymmetric volatilities of equity market returns. A different leverage related explanation of this finding appears in the concluding section.

2. A model of market integration

2.1. The GARCH model

Let us emphasize from the outset that our main concern is not with Mexico's integration into the world market. Because of the prevalence in the early part of the sample period of regional events like the negotiations leading up to the signing of NAFTA, we focus on regional integration.⁸ Since the degree of market integration

⁶ That the correlation is small and not significantly different from zero is actually good news. Were it much larger, it could destroy the validity of our model.

⁷ A competing explanation for asymmetric volatility is the so-called 'volatility-feedback hypothesis' (see Campbell and Hentschel (1992)), in which positive shocks to volatility drive down returns. More specifically, if expected stock returns increase when volatility increases, and if expected dividends are unchanged, then stock prices should fall when volatility increases. The fall of stock prices can increase the potential for large crashes, which will generate an even higher correlation between negative returns and conditional volatility. Either of these two conjectures could account for the evidence of asymmetric volatility in the currency process that we report below.

⁸ Interpreted in the context of the International CAPM, North America could be said, however, to be serving as a reasonable proxy for the world market portfolio. According to Erb, Harvey and Viskanta (1996, Exhibit 1) the conditional beta of the US on the world market index was 0.88 in September 1995. In our data, measured over the entire period, the correlation between the North American index and the World index is 0.81 and the beta is 0.90 with $R^2=0.74$

is related both to the kinds of risk faced by investors and to the level of reward for given levels of risk, we proceed from an asset-pricing model to characterize these kinds of relations. If the Mexican stock market is fully separate from the North American stock market, a domestic version of the CAPM, following Sharpe (1964) and Lintner (1965), holds true. The conditional CAPM implies that the expected return of asset j must be linearly related to the covariance of its return with return of the market portfolio. Thus,

$$E_{t-1}(r_{j,t}) = \frac{E_{t-1}(r_{m,t})}{\text{var}_{t-1}(r_{m,t})} \text{cov}_{t-1}(r_{j,t}, r_{m,t})$$

In the case of wholly separate markets, investors face only country-specific risks and only the domestic risk is priced in the asset-pricing model. Taking asset j as the Mexican market index and letting $\lambda_{m,t-1}$ be the unit price of covariance risk, the conditional version of the CAPM in the case of a fully segmented market can be written as follows:

$$E_{t-1}(r_{m,t}) = \lambda_{m,t-1} \text{var}_{t-1}(r_{m,t})$$

where $r_{m,t}$ is the excess return on Mexico's local market index defined as the continuously compounded percentage or log return translated into US dollars, and calculated as 100 times the log difference between the current and previous week's closing prices in US dollars, minus the one-month US risk free rate. E_{t-1} is the conditional expectation operator conditional on the information available to the investors at time $t-1$. $\lambda_{m,t-1}$ is the conditional unit price of local country-specific risk. $\text{var}_{t-1}(r_{m,t})$ is the conditional variance of the local market index conditional on the information at $t-1$. As implied by these notations, the conditional unit price of risk and conditional variance all vary through time.

If, on the contrary, Mexico's stock market were fully integrated with the North American stock market, the opportunity set available to investors would include all the stocks on the North American stock markets and the mean-variance efficient portfolio would be the North American market portfolio. So, in the case of a fully integrated market, the conditional CAPM can be written as follows:

$$E_{t-1}(r_{m,t}) = \lambda_{n,t-1} \text{cov}_{t-1}(r_{m,t}, r_{n,t})$$

where $\lambda_{n,t-1}$ is the conditional unit price of risk of the North American market and $\text{cov}_{t-1}(r_{m,t}, r_{n,t})$ is the conditional covariance between the return on the Mexican and the North American stock market indexes. Again, all these variables change through time.

Most theoretical models of international asset pricing incorporate exchange risk [see Solnik (1974a), Stulz (1981), Adler and Dumas (1983), Anderson and Danthine (1983) and Stulz (1995)]. There are many empirical tests of the unit price of foreign exchange risk. Solnik (1974b), Stehle (1977), Hodrick, 1981, Vassalou

(1995) and Korajczyk and Viallet (1991) tested unconditional versions of the model; and Dumas and Solnik (1995), De Santis and Gerard (1997), Vassalou (2000), Hardouvelis et al. (2000) and De Santis et al. (1998) performed conditional tests. We are forced by the low, monthly frequency of inflation data and the absence of an inflation indexed bond in Mexico to follow this literature and model nominal rather than real exchange risk.⁹ To do so, we employ changes in one-month Cete rates adjusted by nominal exchange rate changes to define currency returns and use the monthly inflation differential only later, as an instrumental variable in connection with calculating the price of currency risk.

Following the approach of Adler and Dumas (1983), we assume that four types of nominal asset exist: besides the North American and the Mexican market indexes, there are 28-day Cetes, denominated in pesos, which we take as the nominally riskless Mexican asset, and a reference currency riskless asset, the 1-month US T-Bill. So, in the case of a fully integrated market, the conditional regional CAPM can be written as follows:

$$E_{t-1}(r_{m,t}) = \lambda_{n,t-1} \text{COV}_{t-1}(r_{m,t}, r_{n,t}) + \lambda_{c,t-1} \text{COV}_{t-1}(r_{m,t}, r_{c,t})$$

where $r_{c,t}$ is the gross return of Mexico's 28-day Cete adjusted by nominal exchange rate changes. $\lambda_{c,t-1}$ is the price of currency risk which is a function of instrumental variables at time $t-1$. We use gross returns rather than excess returns on Cetes in order partially to avoid the multi-collinearity problem that otherwise confounds our estimates.¹⁰

If the market were only partially integrated, then investors would face not only country-specific risks, but also the risks of the North American market. Incomplete integration, proxied by a large local factor premium may be due to several factors. The results of previous research indicate that there are significant local components in emerging market asset returns. Also, if one follows the Bekaert and Harvey (1995) reasoning, the partial-integration model nests the cases of full separation and full integration. A partial market integration model should, therefore, describe Mexican market returns better than either extreme.

⁹ Most of the theoretical models call for real returns. When PPP is violated, investors view different returns differently as the same nominal return is translated into different purchasing powers depending on the investor's currency base. International investors select portfolios hedged against their idiosyncratic inflation risk. In equilibrium, assets that are 'hedged against PPP deviations' are priced by the CAPM (see Adler and Dumas, 1983). For a variety of reasons, most of the cited work follows Dumas and Solnik (1995) and investigates the pricing of nominal rather than real currency risk. Because of the low frequency of inflation data, so do we.

¹⁰ The multi-collinearity problem arises from the relatively high correlations among the excess returns of Mexico's equity market, the excess returns of the North American stock market and changes in the post-1994 value of the Peso. We use the gross returns instead of excess returns on our pure currency variable in order to reduce this problem. In experiments where we subtracted the T-Bill rate from the gross currency return, our model failed to converge.

$$\begin{aligned}
E_{t-1}(r_{m,t}) &= (1 - \alpha_{t-1})[\lambda_{m,t-1}\text{var}_{t-1}(r_{m,t})] + \alpha_{t-1}[\lambda_{n,t-1}\text{cov}(r_{m,t},r_{n,t}) \\
&\quad + \lambda_{c,t-1}\text{cov}_{t-1}(r_{m,t},r_{c,t})] \\
E_{t-1}(r_{n,t}) &= \lambda_{n,t-1}\text{var}_{t-1}(r_{n,t}) + \lambda_{c,t-1}\text{cov}_{t-1}(r_{n,t},r_{c,t}) \\
E_{t-1}(r_{c,t}) &= (1 - \alpha_{t-1})[\lambda_{m,t-1}\text{cov}_{t-1}(r_{m,t},r_{c,t})] + \alpha_{t-1}[\lambda_{n,t-1}\text{cov}(r_{c,t},r_{n,t}) \\
&\quad + \lambda_{c,t-1}\text{cov}_{t-1}(r_{m,t},r_{c,t})] \\
\lambda_{m,t-1} &= \exp(Z_{t-1}^m \kappa_m) \\
\lambda_{n,t-1} &= \exp(Z_{t-1}^n \kappa_n) \\
\alpha_{t-1} &= \exp(\theta Z_{t-1}^\alpha)
\end{aligned}$$

where α_{t-1} is the measurement of the conditional level of market integration based on the information up to time $t-1$. We assume that $0 \leq \alpha_{t-1} \leq 1$. If $\alpha_{t-1} = 1$, this is a model of full regional integration; and if $\alpha_{t-1} = 0$, then this is a fully domestic CAPM model. In the second equation, the expected excess return of the North American market index is equal to the unit price of the market risk times the conditional variance of the North American market index plus the price of Mexico's currency risk times the conditional covariance between returns on the North American index with Mexico's real currency process. In the third equation, as a local class of asset, Mexico's pure currency return follows the same pricing rule as Mexico's market index.

We further assume that the unit price of the Mexican market risk and the North American market risk are exponential functions of the instrumental variables Z_{t-1}^m and Z_{t-1}^n . In this manner the unit price of market risk is guaranteed to be non-negative. This assumption is based on Sercu (1980) and Adler and Dumas (1983), which show that in equilibrium the unit price of market risk is related to a country's aggregate risk aversion coefficient. Recently, some have argued that the empirical results show it is possible to have a negative risk and expected return relationship. However, Adler and Dumas (1983) point out that although apparently negative ex-post unit prices of risk can be obtained from the data, the true ex-ante relation between market risk and expected return is necessarily positive if investors are risk averse. We, therefore, continue to assume $\lambda_{m,t-1}$ and $\lambda_{n,t-1}$ to be non-negative. In the last equation, α_{t-1} is assumed to be an exponential function of the underlying instrument variables Z_{t-1}^α . θ is the unknown to be estimated. Our choice of instrumental variables, discussed below, guarantees that Z_{t-1}^α will always be positive: the larger Z_{t-1}^α , the greater the sovereign credit risk. We, therefore, expect the sign of θ to be negative as we expect the integration measure to rise when Z_{t-1}^α falls. A negative sign for θ guarantees also that the measure of integration will fall between 1 and 0.¹¹

To repeat for ease of reading going forward, $r_{c,t}$ is the one-month Cete rate compounded by variations of the nominal exchange rate. $\lambda_{c,t-1}$ is the unit price of currency risk, which is a function of instrumental variables Z_{t-1}^c with loading

¹¹ Following a suggestion from Geert Bekaert, we experimented with a specification of the market integration measure function, $\alpha_{t-1} = \exp(\theta Z_{t-1}^\alpha) / (1 + \exp(\theta Z_{t-1}^\alpha))$, that would have guaranteed that $0 \leq \alpha_{t-1} \leq 1$. Unfortunately, the non-linearity of the formulation caused our model to fail to converge.

coefficients κ_c . Mexico had a fixed exchange rate until December 19, 1994. Before then, most of the variation in the currency return came from interest rate volatility, as the Mexican authorities used monetary policy to protect the peg. After January 1, 1995, variations in the currency return were due both to interest rate volatility and to changes in the nominal exchange rate itself. The unit price of currency risk can be either positive or negative, depending on whether the inverse of the country's risk aversion coefficient is smaller or larger than 1. In our estimates, as we have already pointed out, the price of currency risk is negative on the average.

We use the Autoregressive Conditional Heteroskedastic (ARCH) class of models to estimate the above model. The ARCH class of models were introduced by Engle (1982 and many other referenced papers)¹², and generalized by Bollerslev (1986 plus other referenced papers). De Santis (1994, 1995), De Santis and Gerard (1996, 1997), Bekaert and Harvey (1997) and others have extended GARCH to international asset pricing settings. The key insight of the ARCH or GARCH class of models lies in the distinction between the conditional and the unconditional second moment. While the unconditional covariance matrix for the economic variable may be time invariant, the conditional variance and covariance often depend on past states of world. We add a disturbance term ε_t , orthogonal to the information available at the end of time $t-1$, and assume the conditional volatility follows a GARCH (1,1) process.

$$\varepsilon_t = [\varepsilon_{m,t}, \varepsilon_{c,t}, \varepsilon_{n,t}] | \Omega_{t-1} \sim T(0, H_t)$$

$$H_t = C'C + B'\varepsilon_{t-1}\varepsilon'_{t-1}B + A'H_{t-1}A + D'\eta_{t-1}\eta'_{t-1}D$$

$$\eta_{t-1} = (\varepsilon_{m,t} \mathbf{1}_{\{\varepsilon_{m,t} < 0\}}, \varepsilon_{n,t} \mathbf{1}_{\{\varepsilon_{n,t} < 0\}}, \varepsilon_{c,t} \mathbf{1}_{\{\varepsilon_{c,t} < 0\}})'$$

$$D = \begin{pmatrix} d_{11} & d_{12} & d_{13} \\ d_{21} & d_{22} & d_{23} \\ d_{31} & d_{32} & d_{33} \end{pmatrix}$$

This is essentially a Glosten et al. (1993) model. These authors propose a specification to capture both the volatility clustering and the leverage effect by adding a term for the asymmetry to a GARCH model. ε_t is the vector of unexpected excess returns given the information set Ω_{t-1} available at time $t-1$. ε_t is a three-dimensional random variable with mean 0 and variance H_t ; H_t is the conditional variance and covariance matrix of asset returns at time t . The second equation indicates that the conditional covariance matrix follows a GARCH (1,1) process, and the last term captures the asymmetric volatility effect: see also Kroner and Ng (1998) and, particularly, Bekaert and Wu (2000) and Wu (2000). In the third equation, η_{t-1} is equal to the asymmetrically distributed error term. We capture asymmetric information via an indicator function: if $\varepsilon_t < 0$ it takes the value 1 and is zero otherwise.

¹² Bollerslev et al., 1991, Bollerslev and Engle, 1993, Engle and Kozicki, 1993, Engle and Kroner, 1989, Engle and Ng, 1993, Engle et al., 1987, Engle et al., 1990.

A, B, C, D are the unknown parameters for the variance and covariance matrix H_t . Each element of A measures the autoregressive coefficient of the conditional variance; each element of B measures the coefficient on last period's innovation to the second moment of asset returns; and C is the triangular matrix of constants in the variance equation. The A, B and D matrices necessarily have the same dimensions. D determines the coefficients of the asymmetric volatility effect. Each diagonal element $d_{i,i}$ of D measures the asymmetric effect of each asset return on its own volatility, while off-diagonal elements measure the cross-asset asymmetric volatility effects. For example, $d_{i,j}$ measures the asymmetric effect of last period's shock from market j on the conditional volatility of market i . In order to limit the number of parameters in our model, we assume that $d_{2,1}$ and $d_{2,3}$ are zero, as Mexico's equity and currency markets can be expected to have but a small, marginal effect on the North American stock market. Again because of potential multicollinearity we also assume that $d_{3,2}$, which reflects the effects of North American market shocks on Mexico's currency volatility, is also zero. A justification is that before 1994, Mexico had a fixed exchange rate: the potential impact of North American stock market shocks on the exchange rate was suppressed during this time period so that the asymmetric volatility effect cannot be estimated with adequate precision.¹³

The above model captures the time-varying, clustering and leverage effect of the conditional covariance process. As Bollerslev et al. (1988) and Bollerslev and Wooldrige (1992) point out, these are the important empirical regularities for the asset return volatility to follow. Another is the existence of a thick tail, which means the return process tends to be leptokurtic. In other words, stock returns are not drawn from a normal distribution, but from some other thick-tailed distribution. As argued by Bekaert and Harvey (1995, 1997, 2000), this phenomenon is much more pronounced when dealing with emerging market returns. In this paper, we consequently test the degree of market integration by assuming that the conditional joint distribution of the error term is Student's t .

The structure above is the basic setting for the conditional time-varying model of market integration using a GARCH model. The parameter α_t is our measure of the time-varying degree of market integration. Our model extends the literature in that we simultaneously take account of currency risks and asymmetries in the volatility processes to accommodate the so-called leverage effect. We estimate GARCH all in one step in order to ensure the asymptotic efficiency of the estimators.

3. Data and empirical results

3.1. Summary statistics

Our sample is limited to the period starting shortly after Mexico's Brady bonds began to trade actively until we started writing. We use weekly closing price data

¹³ Not making the assumption that $d_{3,2}=0$ causes our model to produce three $d_{i,2}$ coefficients that are all negative, an absurd result. It is unreasonable to think that shocks transmitted from the North American market could reduce the volatility of the Mexican markets.

Table 1
Breakdown of North American market capitalizations

Year	Mexico (%)	Canada (%)	United States (%)
1990	0.646	6.803	92.556
1991	1.492	6.062	92.447
1992	1.924	5.158	92.918
1993	3.058	5.869	91.073
1994	2.085	5.865	92.051
1995	1.180	5.238	93.582
1996	1.198	5.518	93.284
1997	1.396	5.113	93.491
1998	0.661	3.884	95.455
1999	0.920	4.537	94.543
2000	1.010	4.234	94.756
2001	1.274	4.578	94.148

from the first week of January 1991 to the second week of February 2002, totalling 582 observations. We calculate the continuously compounded return or log return as the natural logarithm of the current period price divided by the last period price. The excess return is defined as the difference of the continuously compounded asset return minus the continuously compounded interest rate, which is the middle rate of the one-month US Treasury bill. Mexican equity index and US Treasury data were obtained from Datastream International Corporation while Cetes and Brady bond data were provided by ABN(AMRO) and J.P. Morgan.

Table 1 displays the annual breakdown of the North American total market capitalizations from 1990 to 2002. The values of market capitalization are converted to US dollars. The United States represents an average of 93% of the total North American market capitalization in the last decade. Canada accounts for an average of 5.5%. It is not surprising that Mexico takes the smallest share of market capitalization; its 10-year average is only approximately 1.5%. This suggests that Mexico's market capitalization is small enough not to introduce collinearity when the Mexican and North American market indices, which are not orthogonal to each other, are used together in a single regression. Our model is not, by this token, misspecified.

There is an informative pattern to the change in the percentage of Mexico's market capitalization in the total of North American financial market from 1990 to 2002. Mexico's market capitalization increased steadily from 1990 to 1994, the percentage gain being largest from 1992 to 1993, up by 58.8% to its peak of 3.1%. This period coincides with President Clinton's signing of NAFTA, NAFTA's ratification by the US Congress and the Mexican government's passage of a new law to liberalize foreign direct investment. In the following years, the percentage of Mexican market capitalization fell continuously, through 1999. Two particularly sharp decreases occurred. One is from 1994 to 1995, a 43% decrease, coincident with Mexico's crisis and the start of the depreciation trend for the Mexican peso. The second sharp drop was from 1997 to 1998 with decreases of up to 54% in the

Table 2
Summary statistics of asset returns on market index portfolios

Panel A:			
Country	Mexico(\$)	North American index (\$)	Cete 28(\$)
Mean(%)	0.121	0.131	-0.087
Std	4.369	2.170	0.999
Minimum	-27.805	-14.418	-12.809
Maximum	19.937	8.798	8.021
Skewness	-0.599 [0.000]	-0.045[0.657]	-4.227 [0.000]
Ex.Kurtosis	3.781 [0.000]	1.537[0.000]	63.896 [0.000]
Berra-Jarque	381.174 [0.000]	20.393[0.000]	2371.00 [0.000]
$\rho_1(r)$	0.132	-0.129	0.074
$\rho_2(r)$	0.145	0.049	0.346
$\rho_3(r)$	0.004	0.051	-0.061
$\rho_4(r)$	0.140	-0.076	0.101
$\rho_5(r)$	0.014	-0.006	-0.165
$\rho_1(r^2)$	0.221	0.271	0.071
$\rho_2(r^2)$	0.265	0.167	0.223
$\rho_3(r^2)$	0.192	0.204	0.323
$\rho_4(r^2)$	0.291	0.169	0.086
$\rho_5(r^2)$	0.221	0.169	0.165

Panel B:			
Correlation	Mexico(\$)	North American	Cete 28
Mexico (\$)	1	0.456	0.549
NorthAmerican	0.456	1	0.081
Cete28	0.549	0.081	1

For each country, the weekly returns on value-weighted market index are calculated based on the closing price index provided by Datastream International Corporations. We report the summary statistics of weekly stock returns from the first week of January, 1991 to the second week of February, 2002 for a total of 582 observations. Weekly returns are not annualized and are measured in US dollars. In the second row, we report the weekly mean returns in percentage points. Row 3 is the weekly standard deviation of the percentage asset return. Row 4 is the minimum observation of the whole sample. Row 5 is the maximum observation of the sample. Row 6 is the skewness of the asset return. Numbers in the brackets are the significant level at which H_0 of skewness equal to zero can be rejected. Row 7 is the excess kurtosis, which is defined to be sample kurtosis less 3. Numbers in brackets are the significant level at which H_0 : no excess kurtosis, can be rejected. The estimators of skewness and kurtosis are normally distributed with mean 0 and 3 and variance $6/T$ and $24/T$, respectively, where T is the number of observations. Row 8 reports the Bera-Jarque test for the normality of the asset return. Numbers in brackets are the significant level at which H_0 of normality can be rejected. Row 9 to row 18 are the autocorrelations of the weekly returns and the squares of the weekly returns.

market capitalization as a percentage of the North American total. This was the period of economic crises first in South East Asia and later in Russia. The Asian crises started with the collapse of the Thai baht (June 1997) and ended after the Korean crisis calmed down at the end of January 1998. The Russian crisis began with rumors of default in June 1998, continued with the collapse in the bond market

at the beginning of August 1998 and lasted until the end of that month. The percentage shares of the Canadian and US stock markets in the total of North American market capitalization were fairly stable across the sample period.

Table 2 provides summary statistics for excess returns. Mexico's stock market has a lower mean weekly return and higher volatility than the North American market index. Our results confirm Harvey (1995a,b) who rejected normality at the 5% level: like him we find the skewness of the Mexican market index to be negative and significantly different from zero. The excess kurtosis and Bera-Jarque tests all reject the hypothesis that Mexico's index return is normally distributed.

Table 2 also reports several lags of autocorrelation for the asset returns and the square of asset returns. Mexico has a higher first order autocorrelation than the North American index return, which may reflect the effects of thin trading, low liquidity and, possibly, low speeds of information transmission in the emerging markets.¹⁴ The autocorrelations of the square of the asset returns are significantly higher than those of the returns themselves. This confirms the fact that the higher moments of the asset returns are themselves autocorrelated, which motivates our GARCH specification of the second moment return as an autoregressive process.

3.2. Instrumental variables

Before we discuss the instrument variables in detail, we give a brief introduction of the several fixed income securities that are issued by the federal government of the United Mexican States. There are several kinds of peso-denominated securities that are issued by the Mexican Government. One of the most important is Certificados de la Tesoreria (Cetes). Cetes are Mexican Treasury bills issued domestically at a discount to face value through a weekly auction managed by the Banco de Mexico. Cetes are peso-denominated and usually have maturities of 28-days, 91-days, 182-days, 364-days and very occasionally 728-days.¹⁵ For present purposes, Cetes are considered to be the short-term interest rate benchmark in Mexico and with rare exceptions, are auctioned on a weekly basis. The Cetes are comparable with US Treasury bills, because the two instruments have several significant characteristics in common. For example, both Cetes and T-bills are (1) direct obligations of their respective governments; (2) discount securities; and (3) fully fungible, i.e. new and existing securities with the same maturity are economically equivalent.

Mexico issued bonds under the Brady Plan in 1990. Mexico's Brady bonds are

¹⁴ Low apparent speeds of information diffusion in emerging markets may be due to insider trading: see Bhattacharya et al. (2002)

¹⁵ Cetes, with yields between 12 and 25% were a favorite means of yield pick-up for US investors during the fixed exchange rate period that ended with the Mexican crisis. The destabilization of Mexico can be dated back to February 1994 when the Fed tightened and precipitated a bond market crisis in the US that cascaded southwards. Throughout that spring and into the summer the Mexican government, at the behest of US investors, began partially to replace Cetes with US dollar-linked Tesobonos. It was on the Tesobonos that the Mexicans subsequently defaulted in January 1995.

issued by the United Mexican States, are denominated in US dollars and have 30-year zero coupon US Treasury bonds held at the US Federal Reserve as the principal collateral. The par bond has a 6.25% fixed coupon rate and the coupon is paid semi-annually. The discount bond has a floating coupon rate reset at 6-month LIBOR plus 13/16% and also pays its coupon semi-annually. Par bonds count 30 days per month and 360 days per year, and the discount bond counts actual days per month and 360 days per year. The coupon dates for the discount bond are April 15 and October 15 of each year, and the par bonds pay on March 30 and September 30.

In our model of market integration, the unit price of market risk is a function of several instrumental variables. In order to control the number of unknown parameters, we use three instrument variables for each of the unit prices of risk. The instruments we use for Mexico's unit price of market risk are: a constant (MCON); lagged excess return on the market index (MLMR) and the lagged dividend yield on the market index (MLDY) in excess of US 30-day Treasury rate. The instruments for the North American market risk are: a constant (NCON); the lagged return on the North American market portfolio (NLMR) and the lagged level of the dividend yield on the market index (NLDY) in excess of the US 30-day Treasury rate. The instruments for Mexican currency risk are a constant (CCON); the change in term spread between yields of 364-day and 28-day Cetes (ΔTS); the change in the interest rate differential between the 28-day Cete rate and 1-month US T-bills (ΔIR) and the inflation rate differential between US and Mexico (INF).¹⁶ Cumby and Obstfeld (1984), Kaminsky and Peruga (1990), Bekaert and Hodrick (1992) and others show that the short-term interest-rate differential (equivalent to the forward premium) helps predict time variation in real currency returns. The Mexican peso exchange rate is affected by movements of both US dollar and Mexican peso interest rates. The term spread of Cetes partially captures the local economic environment, which affects inflation and currency risk. These instruments are widely used in the existing empirical finance literature.

We estimate $\lambda_{m,t-1}$ and $\lambda_{n,t-1}$, respectively the unit market prices of risk for Mexico and North America, by assuming them to be exponential functions of the associated instrumental variables, which restricts them to be positive, as noted above. We do not restrict the sign of the market price of currency risk, $\lambda_{c,t-1}$, as it can be either positive or negative depending on whether investors' aggregate risk tolerance is below or above unity. The actual estimating equations appear in the margin of Table 4. The parameter α_t , the relative weight of the North American market factor in our model measures the degree of market integration as a time-varying variable. Bekaert and Harvey (1995) measure α_t through a regime-switching model, from which they interpret α_t as the conditional probability that a country is

¹⁶ We convert yields into equivalent bond yields first and then take the yield difference between the 364-day Cete and the 28-day Cete to get the term spread. We take the difference between the 28-day Cete and the 1-month US T-Bill rate to get the interest rate differential.

Table 3
Summary statistics of instrument variables

Panel A: instruments				
Instrument	Mean	Std	Min	Max
Brady	6.056	2.107	3.057	16.198
r_m	0.126	4.367	-27.805	19.937
DY_m	1.872	0.794	0.760	5.220
r_n	0.135	2.169	-14.418	8.798
DY_n	2.037	0.717	0.960	3.790
ΔTS	0.000	3.062	-31.030	22.000
ΔIR	0.004	0.129	-0.770	0.910
INF	0.247	0.260	-0.116	1.909

Panel B: correlations								
Instrument	Brady	r_m	DY_m	r_n	DY_n	ΔTS	ΔIR	INF
Brady	1	-0.055	-0.099	0.082	0.023	-0.020	-0.056	0.611
r_m		1	0.017	0.456	0.019	0.123	0.035	0.052
DY_m			1	-0.016	0.769	-0.010	-0.005	-0.010
r_n				1	0.021	0.038	-0.029	0.106
DY_n					1	-0.005	0.011	0.186
ΔTS						1	-0.096	0.081
ΔIR							1	-0.053
INF								1

in the integrated regime. We use the difference between the Mexican Par Brady bond's stripped yield and the US Treasury yield as a proxy for the degree of market integration. α_t is then:¹⁷

$$\alpha_{t-1} = \exp[(\text{stripped yield}_{t-1} - \text{US yield}_{t-1}) * \theta]$$

in which θ is an unknown parameter to be estimated. When the above equation is estimated, θ is always less than zero, implying that α_t is a decreasing function of the difference between the Brady stripped yield and the US Treasury yield. When $\theta \rightarrow 0$, $\alpha_t \rightarrow 1$; and if $\theta \rightarrow -\infty$, then $\alpha_t \rightarrow 0$. Overall, $\alpha_t \in [0, 1]$. When the stripped yield spread increases, the perceived credit risk of the debtor country is becoming larger. As a consequence, the degree of observed market integration should decrease. We, therefore, expect θ to be negative.

¹⁷ We wanted to see whether the integration measure based on the stripped yield spread could be improved. Prompted by Bekaert and Harvey (1997) we experimented with incorporating an additional variable in the alpha function, either Total Trade/GDP, which was used also by Bhattacharya and Daouk (2002), or Market Capitalization/GDP. In all cases, the model failed to converge. The reasons may be the large number of unknowns in this model, which requires 37 initial values; the highly non-linear nature of the model; or the fact that GDP is published at most monthly but not weekly.

The valuation of the Brady bond is a combination of the value of the US Treasury principal guarantee and the emerging market sovereign risks. As the degree of market integration ultimately reflects the sovereign risk of each country, the stripped yield of Brady bond is used as the proxy measure of the creditworthiness of a debtor government. The stripped yield of the Mexican par Brady bond is calculated in three steps: first, we find the price of the principal component of the Brady bond in terms of the value of a US zero coupon with a similar maturity; secondly, we subtract that price from the observed total price of the Brady bond; lastly, we find the yield that makes the future cash flow of the Brady bond equal to that price. The yield obtained is the stripped yield of the Brady bond.^{18,19} Table 3 lists the summary statistics of the instrumental variables.

3.3. Estimation results

We obtain the Maximum Likelihood Estimators of all the unknowns using the Broyden, Fletcher, Goldfarb and Shanno (BFGS) algorithm. Each price of risk is the function of a constant, the lagged return and the lagged dividend yield. We report the coefficients in Table 4 along with our main results. The heteroskedasticity-consistent covariance matrix is used to compute the covariance matrix of the parameters (Bollerslev and Wooldrige, 1992). The estimated value of θ is equal to -0.175 , making the average degree of market integration 0.386 with a standard deviation of 0.111 . The final λ s are functions of each instrument variable times its loading coefficient, κ . The estimated results of the loadings on each instrumental variable for the prices of risk are also presented in Table 4, Panel A. Both North American and Mexico's prices of market risk are positively related to the last period dividend yield. The estimated coefficient of $\lambda_{c,t-1}$ on the change in the Cetes term spread (κ_c^1) is positive. The larger the increase in the term spread of Mexico's interest rates, the higher the unit price of Mexico's nominal currency risk. The estimated coefficient of $\lambda_{c,t-1}$ on the change of interest rate differential (κ_c^2) is positive and the inflation differential coefficient is negative. The higher the interest rate differential between Mexico and US, the larger the prices of currency risk.

Table 4, Panel B reports the results of the estimation from the variance equations. It also presents striking evidence of asymmetric volatility in both equity market and currency returns. The estimated value of the coefficient of asymmetric volatility in North American market returns, d_{22} is statistically significant at 1% level. The coefficient of asymmetric volatility in the BOLSA returns, d_{11} is less significant. Perhaps it should not be considered alone but rather in conjunction with d_{13} : North American investors in the BOLSA bear Mexican market risk and currency risk

¹⁸ To derive the sovereign risk and stripped yield we must make the critical assumption that the current price is equal to the theoretical price. This would be the case in an efficient market.

¹⁹ This method omits the value of the rolling guarantee of three interest payments, which no one has priced properly and can be shown to be a small part of a Brady bond's value.

Table 4

Market integration with inflation and currency risk. Estimating equations and definitions

$$\begin{aligned}
r_{m,t} &= (1 - \alpha_{t-1})(\lambda_{m,t-1}h_{m,t}) + \alpha_{t-1}(\lambda_{n,t-1}h_{mn,t} + \lambda_{c,t-1}h_{mc,t}) + \varepsilon_{m,t} \\
r_{n,t} &= \hat{\lambda}_{n,t-1}h_{n,t} + \lambda_{c,t-1}h_{nc,t} + \varepsilon_{n,t} \\
r_{c,t} &= (1 - \alpha_{t-1})(\lambda_{m,t-1}h_{mc,t}) + \alpha_{t-1}(\lambda_{n,t-1}h_{nc,t} + \lambda_{c,t-1}h_{c,t}) + \varepsilon_{c,t} \\
\varepsilon_t &= [\varepsilon_{m,t}, \varepsilon_{n,t}, \varepsilon_{c,t}] | \Omega_{t-1} \sim t(0, H_t, \delta) \\
H_t &= C'C + B'\varepsilon_{t-1}\varepsilon'_{t-1}B + A'H_{t-1}A + D'\eta_{t-1}\eta'_{t-1}D \\
\eta_{t-1} &= (\text{mdummy}, \text{ndummy}, \text{cdummy})' = (\varepsilon_{m,t} | \{\varepsilon_{m,t} < 0\}, \varepsilon_{n,t} | \{\varepsilon_{m,t} < 0\}, \varepsilon_{c,t} | \{\varepsilon_{c,t} < 0\})' \\
D &= \begin{pmatrix} d_{11} & d_{12} & d_{13} \\ 0 & d_{22} & 0 \\ d_{31} & 0 & d_{33} \end{pmatrix} & H_t &= \begin{pmatrix} h_{m,t} & h_{mn,t} & h_{mc,t} \\ h_{mn,t} & h_{n,t} & h_{nc,t} \\ h_{mc,t} & h_{nc,t} & h_{c,t} \end{pmatrix} \\
C &= \begin{pmatrix} c_{11} & c_{12} & c_{13} \\ 0 & c_{22} & c_{23} \\ 0 & 0 & c_{33} \end{pmatrix} & B &= \begin{pmatrix} b_{11} & b_{12} & b_{13} \\ 0 & b_{22} & 0 \\ b_{31} & 0 & b_{33} \end{pmatrix} \\
A &= \begin{pmatrix} a_{11} & a_{12} & a_{13} \\ 0 & a_{22} & 0 \\ a_{31} & 0 & a_{33} \end{pmatrix} \\
\lambda_{m,t-1} &= \exp(\text{MCON} * \kappa_m^0 + \text{MLMR} * \kappa_m^1 + \text{MLDY} * \kappa_m^2) \\
\lambda_{n,t-1} &= \exp(\text{NCON} * \kappa_n^0 + \text{NLMR} * \kappa_n^1 + \text{NLDY} * \kappa_n^2) \\
\lambda_{c,t-1} &= (\text{CCON} * \kappa_c^0 + \Delta \text{TS}_{t-1} * \kappa_c^1 + \Delta \text{IR}_{t-1} * \kappa_c^2 + \text{INF} * \kappa_c^3) \\
\alpha_t &= \exp((\text{stripped yield} - \text{US yield}) * \theta)
\end{aligned}$$

where $r_{m,t}$ is the excess return on the Mexico market index portfolio. $r_{n,t}$ is the excess return on the North American market portfolio. $r_{c,t}$ is the gross return of 28-day Cetes adjusted by nominal exchange rate changes. α_t is the degree of market integration, which is an exponential function of the difference between the stripped yield of the Brady Par bond and the US Treasury yield. θ is the unknown parameter to be estimated. $\lambda_{m,t-1}$ is the price of Mexico market risk. We assume that $\lambda_{m,t-1}$ is an exponential function of the Mexican instrumental variables $Z_{m,t-1}$, which include a constant, the lagged market return and the lagged dividend yield. The loading coefficients, $\kappa_m = (\kappa_m^0, \kappa_m^1, \kappa_m^2)$, are each associated with each of the parameters. $\lambda_{n,t-1}$ is the price of the North American market risk.

We assume that $\lambda_{n,t-1}$ is an exponential function of the North American instrument variables $Z_{n,t-1}$. The instruments include a constant, the lagged return on market index and the lagged dividend yield. The parameters $\kappa_n = (\kappa_n^0, \kappa_n^1, \kappa_n^2)$ are the coefficients associated with each of the instruments. $\lambda_{c,t-1}$ is the unit price of currency risk, which is a function of instrument variables including a constant, changes in the term spread between Cete 364-day, Cete 28-day and changes in the interest rate differential between US and Mexico.

$h_{m,t}$, $h_{c,t}$, $h_{c,t}$ are the conditional variances and $h_{mn,t}$, $h_{mc,t}$, $h_{nc,t}$ are the conditional covariances. H_t is the 3×3 conditional variance and covariance matrix. The C , B and A matrixes are the constant, moving average and auto-regressive coefficients, respectively, for the conditional variance covariance matrix H_t . The error term, ε_t , is a three-dimensional Student t distribution with degrees of freedom δ . The unknown parameters are estimated by maximizing the log-likelihood function under the assumption of a Student's t distribution for the error term:

$$\begin{aligned}
\ln(L(\phi)) &= T \left(\ln \Gamma \left(\frac{\delta + p}{2} \right) - \ln \Gamma \left(\frac{\delta}{2} \right) - \frac{p}{2} \ln(\delta \pi) \right) \\
&\quad - \frac{1}{2} \sum_{t=1}^T \ln \left| \frac{\delta - 2}{\delta} H_t \right| - \left(\frac{\delta + p}{2} \right) \sum_{t=1}^T \ln \left(1 + \frac{1}{\delta} \varepsilon_t' \left(\frac{\delta - 2}{\delta} H_t \right)^{-1} \varepsilon_t \right)
\end{aligned}$$

For our model, we have a three dimensional error vector, so $p=3$. ϕ is the vector of unknown parameters to be estimated. For the Student's t distribution, we have one more parameter to be estimated—the degree of freedom δ .

Table 4 (Continued)

Panel A: estimation of the mean equations								
κ_m^0 [MCON]	κ_m^1 [MLMR]	κ_m^2 [MLDY]	κ_n^0 (NCON)	κ_n^1 (NLMR)	κ_n^2 (NLDY)			
-4.534 (1.601)	0.213 (0.067)	0.903 (0.240)	-4.035 (1.210)	-0.321 (0.053)	0.709 (0.408)			
κ_c^0 (CCON)	κ_c^1 (Δ TS)	κ_c^2 (Δ IR)	κ_c^3 (INF)	δ	θ			
-0.078 (0.275)	0.061 (0.020)	1.446 (1.029)	-0.371 (0.275)	4.513 (0.460)	-0.175 (0.042)			
Panel B: estimation of the variance equations								
a_{11}	a_{22}	a_{33}	a_{12}	a_{13}	a_{31}	c_{11}	c_{22}	c_{33}
0.940 (0.012)	0.979 (0.005)	0.868 (0.020)	0.190 (0.043)	0.535 (0.129)	0.041 (0.055)	0.817 (0.157)	-0.186 (0.056)	0.023 (0.009)
c_{12}	c_{13}	c_{23}	d_{11}	d_{22}	d_{33}	d_{12}	d_{13}	d_{31}
0.187 (0.100)	0.075 (0.036)	0.000 (0.058)	0.075 (0.101)	0.408 (0.146)	0.554 (0.134)	0.751 (0.211)	2.267 (0.773)	0.035 (0.007)
b_{11}	b_{22}	b_{33}	b_{12}	b_{13}	b_{31}			
-0.235 (0.037)	-0.204 (0.027)	-0.460 (0.047)	-0.094 (0.057)	-0.225 (0.124)	0.003 (0.002)			

together.²⁰ The coefficient of asymmetric volatility in currency returns, d_{33} , is also highly significant. As noted below, this implies that devaluations increase currency volatility more than appreciations, which makes intuitive sense in an emerging market setting. The estimation results further indicate that an asymmetric variance effect exists not only for equity and currency markets considered separately, but also in the cross effect of these two markets. This can be seen from the cross effect between the Mexican market and the currency processes, $d_{1,3}$ which is highly significant. All these coefficients are positive, indicating that a negative shock from either the equity or currency markets increases the conditional volatility. At the level of the second moments, the equity market and currency returns processes in Mexico are strongly linked.

3.4. The evolution of risk premia and integration

The Unit Price of North American Market Risk. In the model of Adler and Dumas (1983), the unit price of market risk is related to each country's representative investor's risk tolerance coefficient and the ratio of each country's capitalization to

²⁰ In experiments with our model in which the currency variables were removed, d_{11} was positive and highly significant. The asymmetric volatility in the currency returns seems to dominate that of the Mexican market's returns when the currency returns are included. We are, however, unable further to decompose the two effects.

Price of North American Stock Market Risk

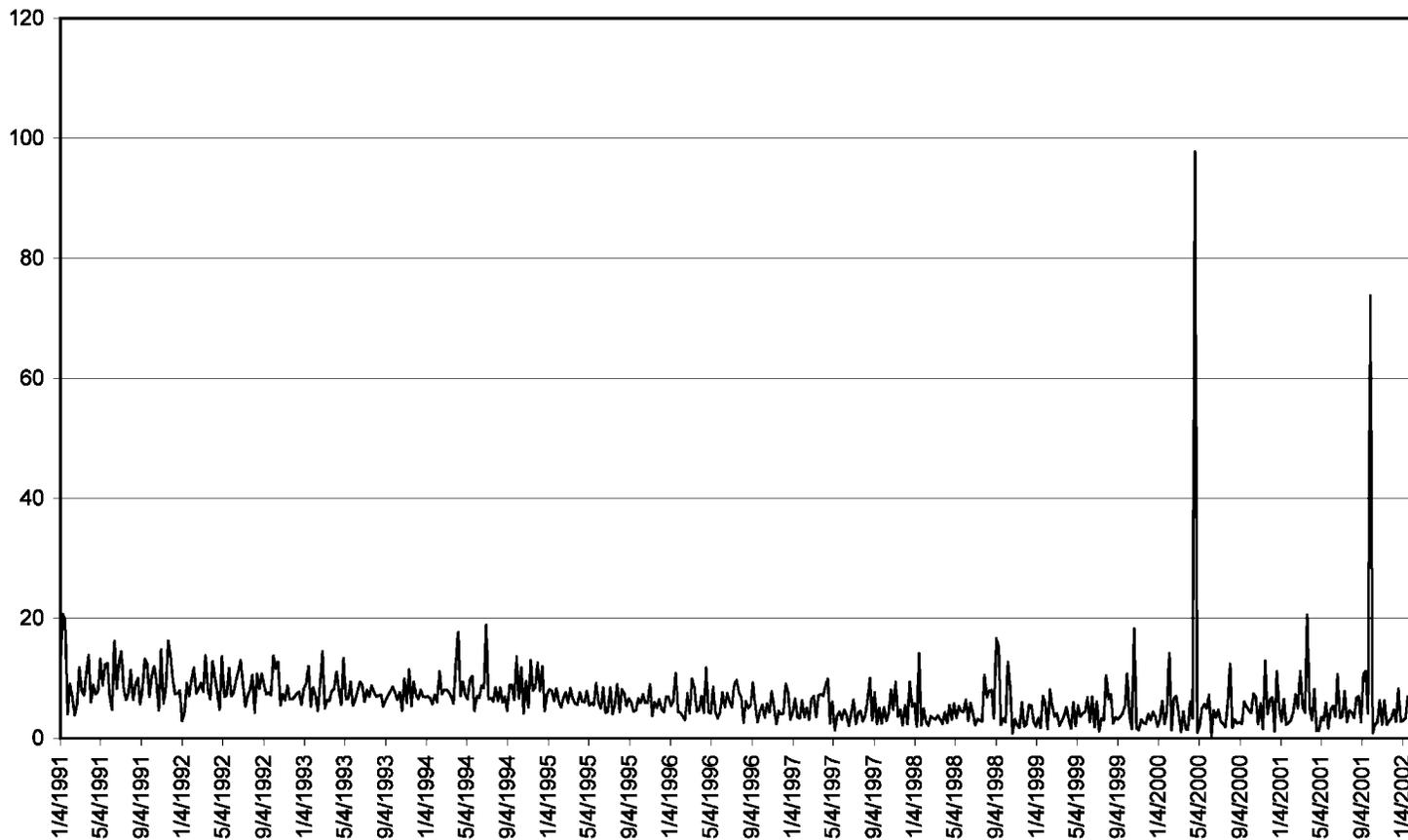


Fig. 1. Price of North American Stock Market Risk.

Price of Mexico's Stock Market Risk

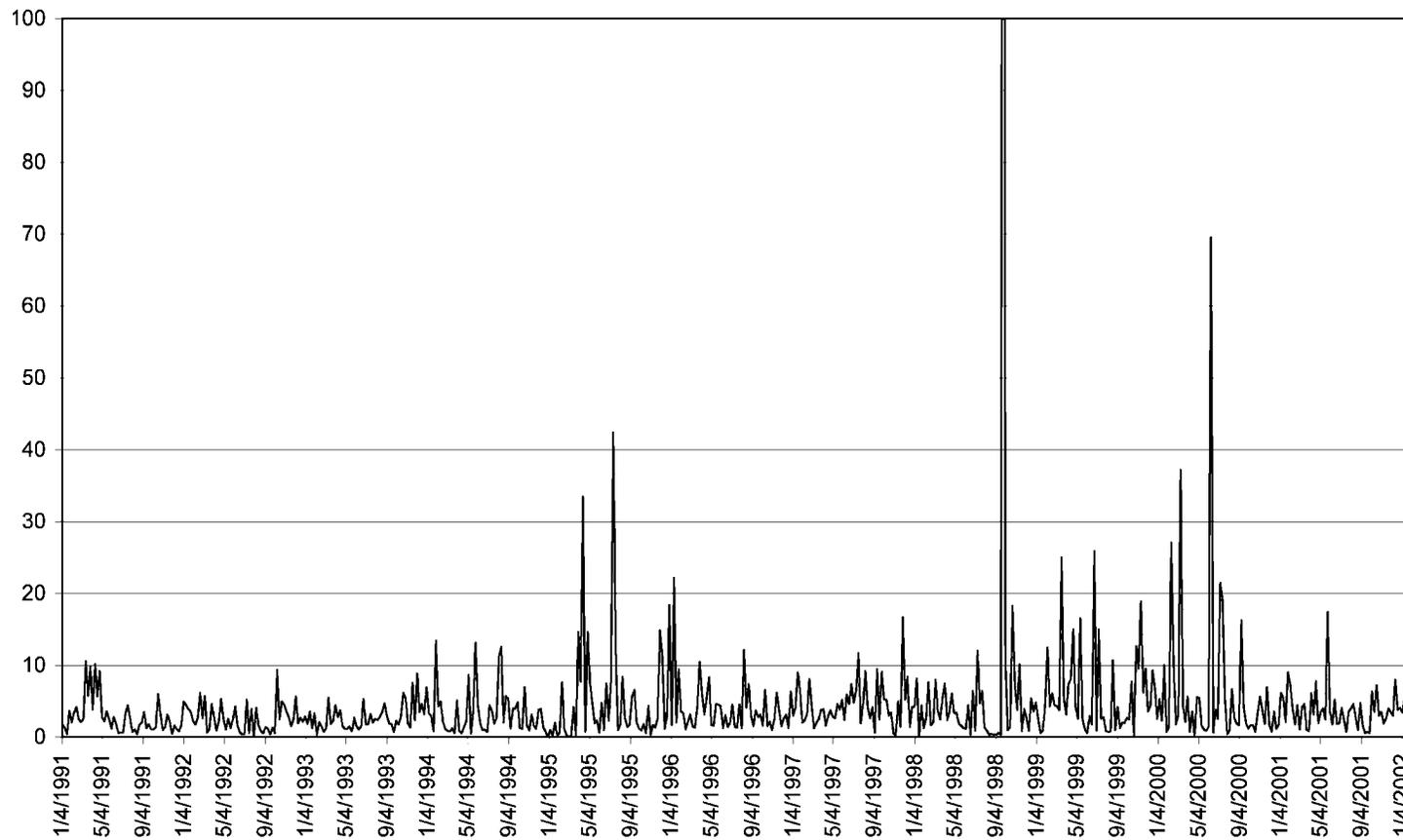


Fig. 2. Price of Mexico's Stock Market Risk.

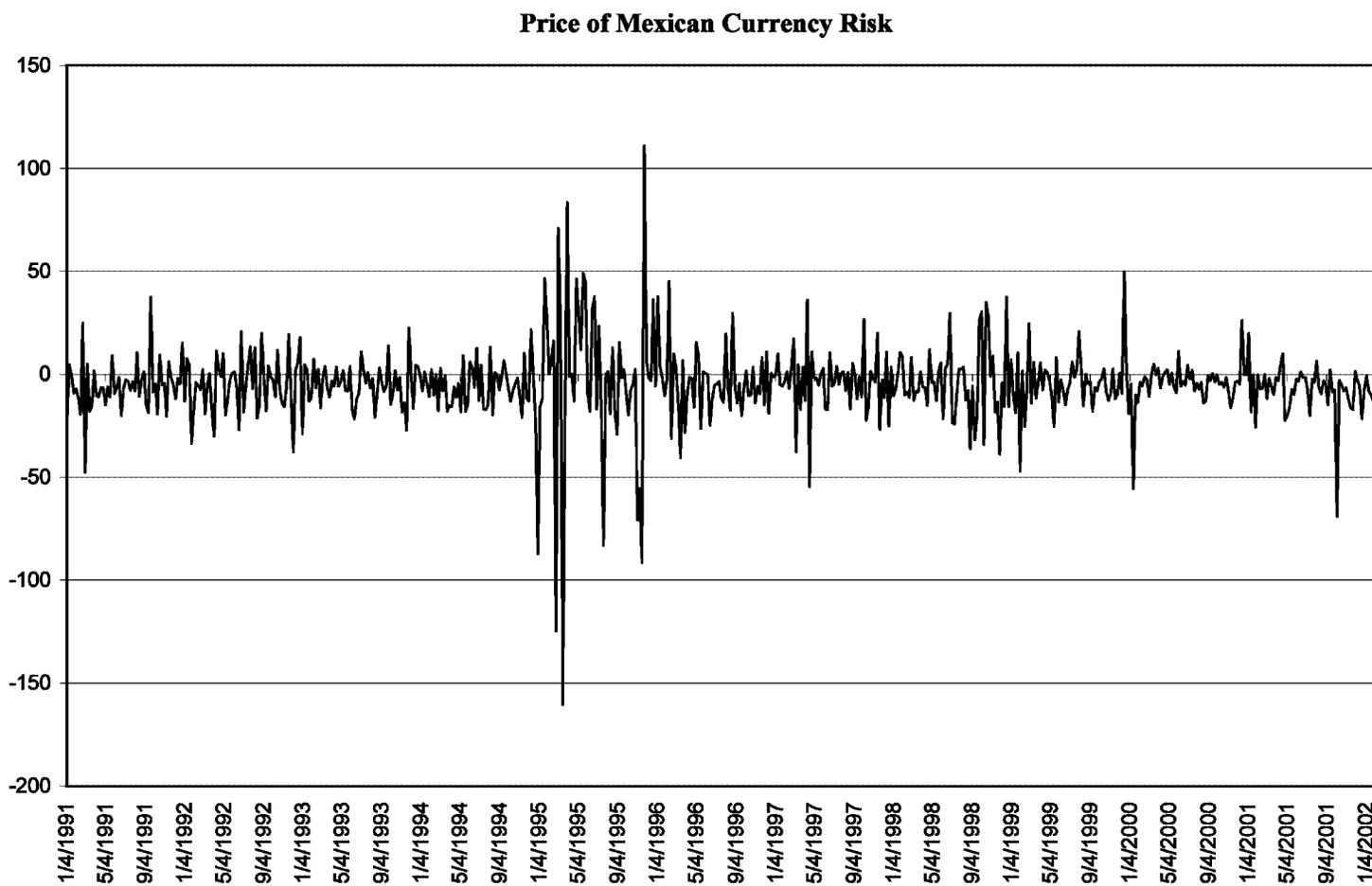


Fig. 3. Price of Mexican Currency Risk.

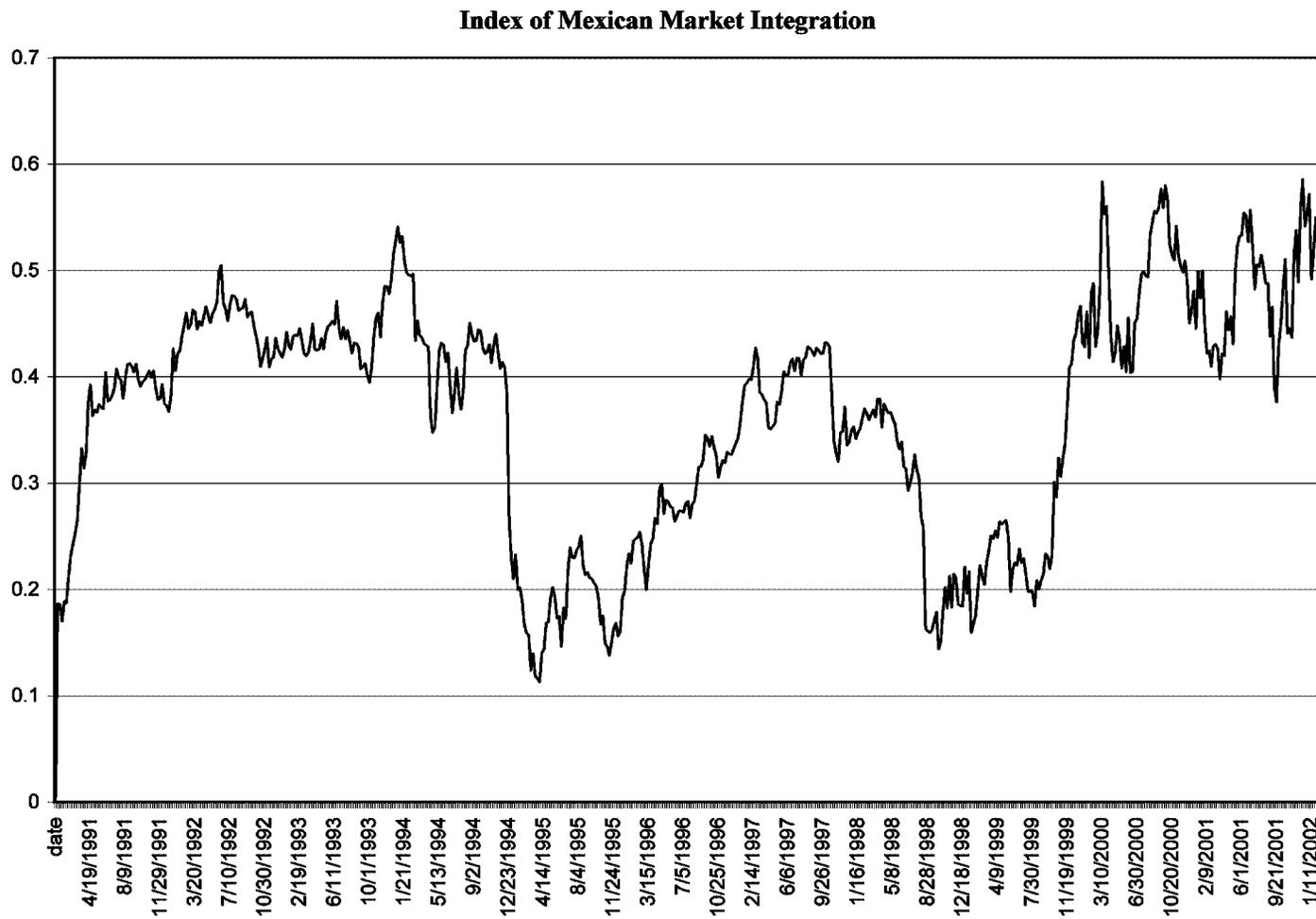


Fig. 4. Index of Mexican Market Integration.

the market total capitalization. In Fig. 1, the average of the unit price of North American market risk, that is, the market risk premium per unit of market variance was 6.59%. From 1991 through the first quarter of 2000 it varied quite tightly around its slowly declining mean. There were two major spikes during the 2000 to 2002 period, when the US stock market slumped following the bursting of the US high-tech bubble.

The Unit Price of Mexican Market Risk. The average range of Mexican market price risks are similar to that of the North American market. In Fig. 2, the average price of Mexican market risk was 4.60%. Before December 1994 it fluctuated quite tightly around its mean. Its volatility rose after January 1995. There were spikes corresponding to the 1995 Mexico crisis, the 1998 Asian and Russian crises and lastly the US stock market slump in 2000.

The Unit Price of Nominal Currency Risk of the Mexican Peso. Perhaps surprisingly, Fig. 3 shows that the price of currency risk for Mexico was significantly negative for most of the sample period.²¹ The average price of peso risk defined as the expected excess return on a unitary US dollar position in pesos per unit of currency variance was -16.36% . The variation of this process showed a plausible pattern. The volatility of currency risk rose substantially during the Peso crisis period from late 1994 until mid-1995. There were also some spikes later corresponding to the 1998 Asian and Russian crises.

Integration. Fig. 4 pictures our main results. It portrays the time-varying process of the degree of market integration. From January 1991 to June 1991 the measure of integration increased rapidly from 0.18 to 0.50. This is the period of the Fed's monetary easing and the beginning of trading in the Mexican Brady bonds. It fluctuated around that level for two years. Between October and December 1993 it moved from 0.40 to 0.53: this is the period of the NAFTA negotiation and Mexico's formal signing of the agreement. It is encouraging that the integration measure fell dramatically when we expect it to during the Mexican crisis and peso devaluation period, reaching its lowest level ever at 0.11 at the end of March 1995. Its rapid recovery after that, to 0.19 by mid-May, was most probably linked to the announcement of the US bailout in late March and may conceivably have had something to do with the beginning of trading in the Peso contract in Chicago in April. Afterwards, along with the recovery of Mexico's economy from the crisis, the integration process recovered to 0.42 by early 1997. The effect of the South East Asian crisis was relatively minor, a decline from 0.42 in October 1997 to 0.32 by November. The effect of the Russian crisis that began in early 1998 was much larger, dragging the integration measure from 0.37 in late March 1998 back to 0.17 in September 1998. As noted by Rigobon (2000), this is a 'contagion' effect. After that, there was a quick recovery from the contagion, which took the integration measure back to 0.58 in March 2000. It has fluctuated around the 0.50 mark ever since. Overall, the Mexican market, on this evidence, seems never to have been

²¹ These findings echo the results of Dumas and Solnik (1995) who explain that following the basic Adler and Dumas (1983) model, the price of exchange risk should be negative if investors' risk aversions are greater than unity. When the price of Mexican currency risk is negative, hedgers require speculators to pay them a premium, rather than the other way around to take a long position in pesos.

Table 5
Structural break tests

Variable	Estimate	S.E.	<i>t</i>	<i>P</i> -value
α_0	0.341	0.005	67.69	0.000
α_1	0.077	0.008	8.87	0.000
α_2	-0.169	0.021	-7.94	0.000
α_3	-0.174	0.041	-4.20	0.000
α_4	0.139	0.014	10.03	0.000

$$\iota_t = \alpha_0 + \alpha_1 D_{(8/1/92 < t < 6/4/94)}^a + \alpha_2 D_{(12/23/94 < t < 3/4/95)}^b + \alpha_3 D_{(8/28/98 < t < 12/28/98)}^c + \alpha_4 D_{(1/21/01 < t < 12/28/01)}^d + \varepsilon_t$$

where ι_t is the degree of market integration at time t . D^a, D^b, D^c and D^d are three dummy variables which corresponding to three economic events: Mexico joins the NAFTA, 1993–1994; the emerging markets crises of 1998, and the collapse of the US high-tech bubble in 2001.

completely integrated with North America. Our integration measure is generally higher than that of Bekaert and Harvey (1995), which was obtained from a different data set.

To confirm the evidence in the figures more formally, we test the effects of special economic events such as NAFTA and financial crises on the degree of market integration. To do so, we split the total sample into several parts: before NAFTA, after NAFTA, before the 1994 crisis, after the 1994 crisis, before the 1998 Asian and Russian crises, after the 1998 crises and the collapse of the US high-tech bubble in 2001. We call this a simple structure break test. The results are shown in Table 5. Dummy variables D^a, D^b, D^c and D^d correspond to each of the economic events. For the process of market integration, the results show that the estimated coefficient for the dummy variable for Mexico joining NAFTA is positive, indicating that joining NAFTA increases the degree of market integration. The estimated coefficients for the two crisis dummy variables are negative, which reflects the fact that during the financial crises, the degree of market integration decreased.

Table 6 lists the results of Wald tests of the robustness of the hypotheses that the unit prices of country and currency risks are zero, respectively. The statistics have a χ^2 distribution with degrees of freedom equal to the number of instruments used. These results indicate that the hypotheses that the unit prices of country risks are zero can be strongly rejected. The hypothesis that the price of currency risk is zero can also be rejected at the 1% significance level.

Table 6
Wald tests for unit price of risks

Null hypothesis	Degrees of freedom	χ^2	<i>P</i> -value
$H_0: \kappa_m^i = 0 (i = 1, 2, 3)$	3	74.67	0.000
$H_0: \kappa_c^i = 0 (i = 1, 2, 3)$	3	50.29	0.000
$H_0: \kappa_c^i = 0 (i = 1, 2, 3, 4)$	4	18.22	0.001

3.5. Robustness tests of the model

In this section, we employ several robustness tests for the model. First, if the model adequately captures the dynamic structure of the expected returns of the Mexican and North American market indexes, the residuals from the expected return equations should be orthogonal to the information variables used to estimate the unknown parameters. Specifically, we regress the residuals on the instrumental variables and perform Wald tests for the orthogonality conditions. Table 7 reports the results of these Wald tests. The values of χ^2 and the associated P -values show that at the 5% significant level, none of the orthogonality assumptions is rejected for the Mexico and North American return residuals. But for the currency risk, the orthogonality condition can be rejected.

The second set of robustness tests is based on the conditional moment of the residuals. Following Bekaert and Harvey (1997), we test the conditional moments of the normalized residuals. We perform the Cholesky decomposition of H_t , and denote the matrix as C_t . The estimated standardized residual is $z_t = C_t^{-1} \hat{\varepsilon}_t$. If the model is correctly specified, the following conditional moments conditions should be valid:

Table 7
Robustness tests

	Orthogonal to Z_m		Orthogonal to Z_n		Orthogonal to Z_c	
	χ^2	P -value	χ^2	P -value	χ^2	P -value
ε_m	3.90	(0.140)	0.214	(0.898)	1.112	(0.573)
ε_n			1.632	(0.442)	2.608	(0.271)
ε_c	13.45	(0.001)	20.35	(0.000)	9.523	(0.008)

Table 8
Conditional moment tests

	Condition a		Condition b		Condition c	
	χ^2	P -value	χ^2	P -value	χ^2	P -value
z_m	0.128	(0.751)	8.955	(0.011)	2.704	(0.109)
Z_n	1.163	(0.201)	1.642	(0.200)	22.16	(0.000)
z_c	6.803	(0.033)	0.939	(0.421)	1.177	(0.185)
Condition d	Conditions f and g				Condition e	
	χ^2	P -value	χ^2	P -value	χ^2	P -value
9.48	(0.090)	9.001	(0.077)	$z_m z_n$	2.703	(0.645)
3.330	(0.501)	10.210	(0.006)	$z_m z_c$	1.750	(0.783)
13.403	(0.001)	5.772	(0.056)	$z_n z_c$	1.253	(0.790)

$$\begin{aligned}
E(\hat{z}_{i,t}) &= 0 \quad (a) \\
E\left(\hat{z}_{i,t}^2 - \frac{v}{v-2}\right) &= 0 \quad (b) \\
E\left[\left(\hat{z}_{i,t}^2 - \frac{v}{v-2}\right)\left(\hat{z}_{i,t-s}^2 - \frac{v}{v-2}\right)\right] &= 0 \quad (c) \\
E(\hat{z}_{i,t}\hat{z}_{i,t-s}) &= 0 \quad (d) \\
E[(\hat{z}_{i,t}\hat{z}_{j,t})(\hat{z}_{j,t-s}\hat{z}_{j,t-s})] &= 0 \quad (e) \\
E(\hat{z}_{i,t}^3) &= 0 \quad (f) \\
E\left[\hat{z}_{i,t}^4 - \frac{3(v-2)}{v-4}\right] &= 0 \quad (g)
\end{aligned}$$

Where z is the estimated standard residual, and v is the degree of freedom of the Student- t distribution. Conditions (a) and (b) are tested separately and have one degree of freedom, respectively. The test statistics for condition (c) and (e) have 4 degrees of freedom if we choose $s = 1, 2, 3, 4$. Conditions (f) and (g) are tested jointly and have two degrees of freedom for each of the standardized residuals. Table 8 reports the results. There are several rejections for the conditional second moment, i.e. equation (b) and the joint tests of (f) and (g). Overall, the robustness tests do not provide much evidence against the model.

4. Conclusions and implications for practice

This paper has investigated the evolution of the process of integration between the Mexican and North American equity markets between 1991 and 2002. In the spirit of the current literature following Bekaert and Harvey (1995), we used a model that combined the domestic and international versions of the CAPM. It tested the power of local factors relative to that of common factors to explain expected returns and empirically inferred segmentation when the weight of the local factors was high. We have acknowledged from the outset that this model is problematic. Tests of multi-factor models of expected returns have routinely found significant local factor premiums even in advanced-country stock markets that cannot meaningfully be said to be segmented by any physical barriers to capital flows.²² These results remain a puzzle. We, therefore, consider our results not necessarily as evidence of true integration or segmentation, but rather as being merely consistent with integration account. Overall, we have shown that Mexico's BOLSA became

²² Following Bekaert (1995) significant local factor premia may have additional explanations. First, there may be investment barriers, which may restrict foreign investment into the domestic market as well as capital outflows to foreign countries [see Black (1974)]. Secondly, following Stulz (1981) asymmetric taxation and differing transaction costs can also affect the behavior of investors when they choose whether to invest domestically or internationally. Thirdly, Lewis (1999) argues that home bias can account for a large local factor premium even without true segmentation. Large local factor loadings can result from other causes as well: currency risk, the local behavior of output or a large non-traded goods sector.

relatively highly, but never completely integrated in the 1990s: at its peak on February 22, 2002 our integration measure never exceeded 0.59.

We have shown that the integration of emerging capital markets into global markets is as likely to be related to changes in international market conditions as to events within the emerging markets themselves. Liberalizing its capital accounts enables, but does not automatically cause a nation's emerging capital market to become integrated with those of other countries. Mexico, for example, made its stock market almost fully investible in May 1989 and issued its first post-debt-crisis bond in June 1989 almost two years before we detected major signs of integration in the first half of 1991. The 1991 integration episode probably had less to do with developments inside Mexico than it had with the Fed's easing program after the Gulf War recession, which expanded US liquidity and forced US investors to look to Mexico's Brady bonds and Cetes for increased returns and risks. More generally, our results suggest that integration is not necessarily a matter of regime shifts within individual emerging markets economies. Instead, it also expands and contracts in association with conditions in the global markets and depends only partly on local developments. Our measure of integration did get a boost from the negotiations leading up to the signing of NAFTA in November 1993 and collapsed in the period leading up to Mexico's default on its Tesobonos in January 1995: these, arguably, might be considered local effects. However, following the bailout in March 1995, the integration measure recovered with ups and downs that are not easily attributable to any particular cause, until contagion from the Asian crisis beginning in 1997 and the Russian crisis in June to August 1998 caused it to collapse once again. It has since recovered. These external crises were not local events: they had nothing to do with Mexico.

Following the strand in the recent literature that showed that credit ratings influence the effects of market liberalizations and subsequent integration, we have argued that in Mexico's case the integration process was related to the market's changing assessment of sovereign default risk as proxied by the excess stripped yield of the Mexican Par bond.^{23,24} In particular the degree of integration was determined by ongoing news that changed investors' credit risk perceptions and that affected market liquidity, but that was not necessarily reflecting only domestic

²³ Spreads and ratings capture much the same information as noted in the introduction. Erb et al. (2000) report in their Exhibit 9 an R^2 of 0.78 in a cross sectional regression of spreads on credit ratings. We could not use ratings in our analysis, however, as they are reported only semi-annually. Besides, ratings reveal much greater cross sectional than intertemporal variation. Mexico's ratings did not fluctuate very much between 1991 and 2002 and varied much less than spreads. The time series correlation between ratings and spreads in our sample is -0.55 : the higher the rating, the lower the spread.

²⁴ This is not to gainsay the success of other measures of integration. Following its introduction by Bekaert (1995), other authors including Bae et al. (2002), De Jong and de Roon (2001) and Edison and Warnock (2003) all employed fruitfully the ratio of the IFC's investible indices to its global indices. From our perspective, the trouble with these and other integration variables that were mentioned before, like trade or market capitalization as fractions of GDP, is that they were observed too infrequently to be used in a time series analysis that requires weekly data.

regime changes. The dependence on credit risk probably also helps account for the strong evidence we find of asymmetric volatility.

We discovered that Mexico's currency risk is priced and that the price of this currency risk was negative on the average. In addition, we found that Mexico's currency returns process revealed strongly significant asymmetric volatility that was strongly related to the asymmetric volatility of the Mexican equity market returns process. The reason for this latter relationship is that the capital inflows and outflows that either accompany or are occasioned by sharp changes in sovereign default risk perceptions are likely to affect equity market and currency returns simultaneously. Devaluations and depreciations of the peso are associated with greater increases in the volatilities of both the currency and stock market returns than are currency appreciations. A different kind of leverage effect offers a plausible explanation for this result. Currency devaluations in emerging markets like Mexico's can cause default-risk crises in local banking systems that mismatch local-currency assets with hard currency liabilities, whereas appreciations produce no such problems. Devaluations are, therefore, more likely than appreciations to destabilize emerging markets' banking systems and increase the volatilities of both the currency and the equity markets' returns. We conjecture that this pattern is more likely to be associated with emerging market currencies than with the exchange rates among the major economies whose banking systems are perceived to be much less vulnerable.

Why should the study of integration interest practitioners? The answer is partly that when markets are fully integrated, identical risks command identical rewards everywhere. When markets are segmented, partially or totally, risks may be priced differently in different locations. In their authoritative survey, Bekaert and Harvey (2003) summarize extensive research showing further that integration with the world market generally leads to lower expected returns and, consequently, lower costs of capital for emerging markets. The obverse of this proposition is that expected returns in emerging markets rise with the degree of segmentation, something that we confirm. A high degree of segmentation, identified from significant, priced local factor and currency premia, implies that firms will typically find project costs of capital in segmented emerging markets to be high. Different degrees of segmentation also imply that project costs of capital will generally differ from market to market. It generally behooves firms to search for the cheapest equity capital for projects in emerging market economies.

Finally, our results have direct implications for estimating the cost of capital in an emerging economy. The US dollar cost of a country's external debt is determined directly by its (stripped) spread over the same duration US Treasury. The higher (lower) the spread, the higher (lower) the cost of US dollar debt. The US dollar cost of emerging market equity is another matter altogether. In principle, it should be higher than the cost of debt as, unprotected by promises to pay and other covenants, equity is riskier and its cost should also rise when the cost of debt does. It does not, however, directly key off the cost of debt and is hard to measure. In a single-factor ICAPM with a single representative investor, it would theoretically be measured as the world's riskless rate plus the product of the world's market risk premium times the country's equity market's beta with the world market portfolio.

In the context of a similar single-factor regional CAPM model, Mexico's cost of equity would be estimated as the North American riskless rate plus the North American market's risk premium times Mexico's beta with the North American market index. However, both the model of this paper and all of those employed in the literature on emerging markets' finance contain multiple factors some of which differ among locations.²⁵ It then becomes difficult, if not impossible, to come up with simple numerical estimates of the cost of equity or to compare them among countries and we have not attempted to do so here.

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²⁵ In Adler and Dumas (1983), the ICAPM even in a frictionless world had multiple priced factors, one for each investor's reference currency. Readers are referred to Dumas and Solnik (1995) for an econometric treatment of this problem. In emerging markets additional special country factors that differ from nation to nation appear also to be priced.

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