# Equity market comovements and financial contagion: a study of Latin America and the United States


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**Abstract**

This paper examines the extent of short-term comovements and financial contagion for four major Latin American emerging markets and the United States by employing a dynamic conditional correlation (DCC) approach and a structural change analysis. Cross-market linkages are found to be time-varying and subjected to several breaks over the period 1988-2009. The patterns followed by DCC estimates are, however, not consistent with higher comovements between these markets in recent years. Finally, our results do not support the financial contagion hypothesis for the Mexican crisis of 1994, the Asian crisis of 1997-1998, and the global financial crisis of 2008-2009, in almost all cases.

***Keywords**:* Stock market comovements, Latin American emerging markets, DCC-MGARCH models, structural change analysis.

***JEL Classification**:* F37; G15

## 1. Introduction

The issue of equity market comovements and financial contagion has received particular attention from academic researchers in recent years. Their main motivations are related to a desire to measure the potential benefits of international portfolio diversification, and to gauge the intensity of the interdependencies between national stock markets during crisis periods in order to have a reliable method for managing the risks of contagion which may result from the multilateral transmission of volatility and shocks[1]. The ongoing process of globalization, greatly stimulated by market-opening policies in many emerging market countries during the late 1980s, also contributes to an expanding interest in the study of stock market comovements[2]. This is explained by the fact that regional and global financial integration would further reduce the potential gains from diversifying portfolios internationally.

Past empirical studies have employed a wide variety of methods and data frequencies at both firm and country levels to measure equity market comovements and have attempted to look for their determinants. The abundant existing literature can without loss of generality be divided into three main groups: short-term comovements, long-term comovements, and extreme comovements. We typically observe that short-term comovement analysis emphasizes not only the study of correlations, but also the spillovers of stock return and volatility across equity markets (e.g., Hamao et al., 1990; Karolyi, 1995; Forbes and Rigobon, 2002; Chiang et al., 2007; Syriopoulos and Roumpis, 2009). Accordingly, the empirical strategies adopted rely essentially on Granger (1969)'s causality test, Sims (1980)'s vector autoregressive models, and GARCH methodology. For example, Hamao et al. (1990) explore the relationships between three marketplaces (New York, London, and Tokyo) using an ARCH-type model and find significant spillovers of prices and price volatility. Subsequent studies confirm these findings

[1] It should be noted that the recurrence as well as the gravity of recent financial crises, such as the 1994-1995 Mexican debt crisis, the 1997 Asian crisis, the 2000 technological sector collapses, and the current global financial crisis, naturally calls for careful analysis of financial contagion because comovements in international equity markets tend to be particularly strong in the aftermath of a crisis and cannot be explained by the actual degree of market integration.

[2] Recent studies have shifted their focus on emerging stock markets to examine whether the diversification benefits documented by, for example Errunza (1977), have been significantly reduced now that they are open to foreign investors' capital flows and have become fairly well integrated into the world market system (Bekaert and Harvey, 1995; Carrieri *et al.*, 2007).

and further point out that there is evidence of a positive relationship between correlation and volatility, i.e., the correlations between international stock markets tend to be important in periods of high volatility or in the times of financial trouble.

For its part, the analysis of long-term comovements focuses on the joint behavior of international equity markets based on the cointegration concept introduced by Engle and Granger (1987). Following the works of Taylor and Tonks (1989) and Kasa (1992), who provide evidence of cointegration between stock prices in major equity markets within bivariate and multivariate frameworks respectively, many papers have analyzed this issue, and established that equity markets display significant comovements (e.g., Gilmore et al., 2008; Diamandis, 2009). Moreover, it is well documented that stock prices respond strongly to international common factors.

The analysis of extreme comovements among equity markets has recently been the focus of several works (e.g., Longin and Solnik, 2001; Chan-Lau et al., 2004; De Melo Mendes, 2005). They argue that traditional models might fail to describe the true pattern of equity market comovements because stock returns often depart widely from the normal distributions commonly-assumed, and exhibit nonlinear and extreme interdependencies. Using extreme value theory and a copula approach, these studies present strong evidence of extreme dependence structure in international equity market returns.

In this paper we focus our attention on the comovements between four major emerging markets in Latin America and the United States (US). Reasons for doing so include especially the pivotal role of these markets in international portfolio diversification because of their high expected returns and their low correlations with mature markets. In addition, from the American investors' point of view, it is interesting to question the degree of comovements among these markets now that market liberalization has gradually integrated emerging markets into the world financial system. The issue of equity market comovement in Latin America has already been investigated by several studies. Choudry (1997) employs unit root tests, cointegration tests, and error correction models to examine the long-term relationship between six Latin American markets and the US market, and finds evidence of a cointegration relationship and significant causality among these markets. Chen et al. (2002) investigate the interdependencies of six equity markets in Latin America and document that diversification benefits are limited when investing in these markets, particularly owing to their high level of comovement. By combining a vector autoregressive (VAR) model with a multivariate exponential GARCH process, Christofi and Pericli (1999) show evidence of significant cross-market linkages in five Latin American markets. Johnson and Soenen (2003) examine the cross-country comovement for eight Latin American equity markets with the US market using the Geweke (1982) measure of contemporaneous feedback between return series. They find a statistically significant linkage between eight equity markets in the Americas and the US stock market. In a more recent paper, Fujii (2005) finds evidence of the intra-causal linkages among four Latin American markets, and further demonstrates that these causal linkages are stronger during times of major financial crisis.

Our empirical investigation differs from the related literature in several aspects. First, instead of modeling the comovement by VAR and realized correlations which capture the causal linkages but do not allow the comovement to be quantified, we directly infer the cross-market linkage from the stock data using a multivariate Dynamic Conditional Correlation GARCH model (DCC-GARCH)[3]. Second, we are interested in dating the structural breaks in the time-paths of the conditional correlation indices to highlight whether the cross-market comovement encompasses significant changes in its nature or not. Lastly, our methodology enables investigation of the differences in stock market comovements between normal and crisis periods. Using monthly data from February 1988 to April 2009, our findings reveal significant time-varying market linkages between Latin American and US stock markets. In addition, stock returns in the US markets significantly affect the contemporaneous dynamics of stock returns in two of the four emerging markets considered. These results suggest a certain degree of financial interdependence and market integration among the sample markets. The analysis of structural changes, based on Bai and Perron (2003)'s procedure, shows that the comovements of Latin American emerging stock markets with the US stock markets are not stable over time, but are subject to several

[3] Prior to our investigation, some papers had found evidence of the time-varying aspect of international equity market comovements (e.g., Bae et al., 2003 and references therein).

structural breaks. These are generally found to coincide with major market events including market liberalization policies and the Asian financial crisis of 1997. Our findings do not provide support for the existence of contagion effects around the most recent financial crisis.

The remainder of the article is organized as follows. Section 2 presents our empirical methodology. Section 3 presents the data used and reports the empirical results. We provide some concluding remarks in Section 4.

## 2. Econometric method

The financial literature on volatility modeling of emerging stock markets has been extensively developed since the introduction of ARCH/GARCH-type models by Engle (1982) and Bollerslev (1986) respectively. Examples of applications include, among others, Bekaert and Harvey (1997, 2000), and Kim and Singal (2000). As pointed out by previous studies, GARCH models appear to successfully describe the stochastic properties and irregular features of stock returns such as leptokurtic behavior, time-variation, and volatility persistence.

In this paper we rely on time-varying correlation coefficients estimated from a multivariate DCC-GARCH model developed by Engle (2002) to investigate the comovements among and between Latin American emerging stock markets and the US stock markets. Ultimately, the proposed model offers several advantages compared to its alternatives such as BEKK GARCH (Engle and Kroner, 1995) and the Constant Conditional Correlation (CCC) GARCH model (Bollerslev, 1990). First, it allows us to directly infer the cross-market conditional correlations from the data, which is not possible when using the CCC-MGARCH, and also to avoid the proliferation of parameters to be estimated in a full parameterization of a BEKK MGARCH specification. Second, the use of the DCC-GARCH model is entirely justified by the presence of conditional heteroscedasticity in our return data (see Table 1). Moreover, in its multivariate version the model used provides a flexible way of capturing the dynamics of a large correlation matrix over time. This flexibility arises from the fact that we can estimate the model by using a two-step procedure, i.e., in the first stage a set of univariate GARCH models is estimated for each market in the sample, and in the second stage a simple specification is employed to measure the conditional correlations based on the standardized residuals obtained from the first stage. Accordingly, the economic interpretation of a univariate GARCH specification in directly capturing the presence of GARCH effects in stock returns remains meaningful, and we are able to provide consistent estimates of the conditional correlation matrix (Kearney and Poti, 2006). In addition, the model is sufficiently flexible to account for the impact of common factors by adding additional explanatory variables in the mean equations (Chiang et al., 2007). Last but not least, the DCC-MGARCH enables the quantification of the amounts of comovements between stock markets, which is not possible when using a VAR modeling, as noted earlier.

Turning to the econometric specification of the model, and assuming that stock market returns from the *k* series are multivariate normally-distributed with zero mean and a conditional variance-covariance matrix $H_t$, our multivariate DCC-GARCH model can be written as follows:

$$\begin{cases} \tilde{r}_t = \mu_t + \varepsilon_t,\ \varepsilon_t | I_{t-1} \rightarrow N(0, H_t) \\ H_t \equiv D_t R_t D_t \end{cases} \tag{1}$$

In these formulas, $\tilde{r}_t$ is the ($k\times 1$) vector of the returns on stock market indices; and $\varepsilon_t$ is a ($k\times 1$) vector of zero mean return innovations conditional on the information available at time *t-1*. Let $\tilde{r}_i$ and $\tilde{r}_w$ be the rates of return on an individual emerging market and world stock market index respectively; we further assume that stock market returns are generated by an autoregressive process such as:

$\mu_{i,t} = \delta_{i0} + \delta_{i1}\tilde{r}_{i,t-1} + \delta_{i2}\tilde{r}_{w,t} + \delta_{i3}\tilde{r}_{w,t-1}$ for emerging market *i*

and

$\mu_{w,t} = \delta_{wo} + \delta_{w1}\tilde{r}_{w,t-1}$ for the world stock market

$D_t$ refers to a ($k\times k$) diagonal matrix with the elements on its main diagonal being the conditional standard deviations of the returns on each market in the sample, and $R_t$ is the ($k\times k$) conditional correlation matrix. $D_t$ and $R_t$ are defined as follows:

$$D_t = diag(h_{1kt}^{1/2} ... h_{kkt}^{1/2}) \quad (2)$$

where $h_{iit}$ is chosen to be a univariate GARCH(1,1) process;

$$R_t = (diag(Q_t))^{-1/2} Q_t (diag(Q_t))^{-1/2} \quad (3)$$

where $Q_t = (1-\alpha-\beta)\overline{Q} + \alpha u_{t-1}u_{t-1}^{'} + \beta Q_{t-1}$ refers to a ($k\times k$) symmetric positive definite matrix with $u_{it} = \varepsilon_{it} / \sqrt{h_{iit}}$, $\overline{Q}$ is the ($k\times k$) unconditional variance matrix of $u_t$, and $\alpha$ and $\beta$ are non-negative scalar parameters satisfying $\alpha + \beta \prec 1$.

The conditional correlation coefficient $\rho_{ij}$ between two markets *i* and *j* is then expressed by the following equation:

$$\rho_{ij} = \frac{(1-\alpha-\beta)\overline{q}_{ij} + \alpha u_{i,t-1}u_{j,t-1} + \beta q_{ij,t-1}}{\left((1-\alpha-\beta)\overline{q}_{ii} + \alpha u_{i,t-1}^{2} + \beta q_{ii,t-1}\right)^{1/2}\left((1-\alpha-\beta)\overline{q}_{jj} + \alpha u_{j,t-1}^{2} + \beta q_{jj,t-1}\right)^{1/2}} \quad (4)$$

In Equation (4), $q_{ij}$ refers to the element located in the *i*th row and *j*th column of the symmetric positive definite matrix $Q_t$. As discussed above, the estimation of our empirical DCC-MGARCH model is carried out by using a two-stage procedure, i.e., we first estimate a univariate GARCH(1,1) model for each time series, and then employ the resulting residuals standardized by their conditional standard deviations to infer the conditional correlation estimators. The set of unknown parameters is estimated by maximizing the following log-likelihood function based on the *T* observations available and the Berndt-Hall-Hall-Hausman (BHHH) optimization method:

$$L = -\frac{1}{2}\sum_{t=1}^{T}\left(n\log(2\pi) + \log\left|D_t R_t D_t\right| + \varepsilon_t^{'} D_t^{-1} R_t^{-1} D_t^{-1} \varepsilon_t\right)$$

Since $u_t = \varepsilon_t / \sqrt{h_t} = D_t^{-1}\varepsilon_t$, the log-likelihood function can be rewritten as follows:

$$L = -\frac{1}{2}\sum_{t=1}^{T}\left(n\log(2\pi) + 2\log\left|D_t\right| + \log\left|R_t\right| + u_t^{'} R_t^{-1} u_t\right)$$

It should be noted that a conditional variance term can be added into the mean equation to control for the risk-return tradeoffs in previous work. However, the majority of these studies found evidence of an insignificant impact of conditional volatility on stock returns. Lundblad (2007) studied the same issue over the period from 1802 to the present and concluded that the mean-variance tradeoffs are positively significant only over a very long period. Accordingly, we deliberately do not model the in-mean effect of conditional volatility because we focus on a short-run comovement across equity markets.

Once we obtain the dynamic conditional correlations between sample stock markets, we proceed to investigate whether structural changes are present in their time-series properties. Our motivation comes essentially from the complexity of stock market liberalizations in Latin American countries and the recurrence of financial crises that have unhinged world markets in recent years. Indeed, Latin American emerging markets have experienced significant reforms in their capital markets over the last three decades. They occurred in various periods of financial turbulence including, among others, the Mexican crisis in 1994-1995, the Brazilian crisis in 1998, and the Argentinean crisis in 2001. Both Latin American emerging and US markets were also severely affected by the recent global financial crisis of 2007-2009. Thus the time-paths of cross-market comovements might accordingly be subjected to structural changes. The idea is that a higher degree of financial openness can strengthen the relationships between international stock markets, but the depth of serious financial crises may affect a country's economic and financial structure, which in turn leads to changes in the nature of its comovements with other countries. Since the changes in the cross-market comovements may affect the policy coordination of two or more countries as well as the actions of portfolio managers, it is opportune to examine the issue of structural changes. To do so, we attempt to detect structural breaks in the time-varying cross-market conditional correlations based on the Bai and Perron (2003)'s testing procedure, which consists of determining the number and location of breaks in a linear regression framework.

More precisely, suppose there are $m$ breaks $(n_1,\ldots,n_m)$ in the time-path of the dependant variable, the problem of dating structural breaks amounts to finding the breakpoints $(\tilde{n}_1,\ldots,\tilde{n}_m)$ that minimize the objective function:

$$(\tilde{n}_1,\ldots,\tilde{n}_m) = \arg\min_{(n_1,\ldots,n_m)} RSS_n(n_1,\ldots,n_m)$$

where $RSS_n$ is the resulting residual sum of squares based on the *m* regressions as shown by the following equation:

$$y_t = \beta x_t^{\mathrm{T}} + \varepsilon_t \qquad (t = 1,\ldots, n) \tag{5}$$

In Equation (5), $y_t$ plays the role of the estimated conditional correlation series at the time *t*, $x_t = (1, y_{t-1})^{\mathrm{T}}$ is the (2×1) vector of observations of the independent variables with the first component equal to unity, $\beta$ is the (2×1) vector of regression coefficients, i.e., a constant term and an autoregressive coefficient, and $\varepsilon_t$ is assumed to be independent and identically distributed according to a normal distribution with zero mean and variance $\sigma^2$. Our structural stability test is then concerned with testing the null hypothesis of '*no structural break*' against the alternative that the regression coefficients vary over time. The breakpoint selection procedure is based on the Bayesian Information Criteria (BIC). Empirically, we set the maximum number of optimal breaks to be 5 and run the test. Whenever the effective number of breaks is equal to 5, a higher number of breaks will be automatically chosen so that the testing procedure captures all possible breakpoints. The optimal number of breaks corresponds to the one with the lowest BIC score.

## 3. Data and empirical findings

### *3.1 Data and properties*

The data used are collected on a monthly basis and consist of the MSCI (Morgan Stanley Capital International) total return indices for four major Latin American emerging markets (Argentina, Brazil, Chile, and Mexico) and the US stock market, sampled over the period from February 1988 to April 2009. All the indices are obtained from MSCI Barra and expressed in US dollars to preserve homogeneity across markets and also to avoid the effects of currency risks. They are converted to return series by calculating the differences in natural log prices. The basic statistics and stochastic properties of the monthly returns are presented in Table 1.

**Table 1**
**Descriptive statistics and stochastic properties of stock market returns**

*Panel A: Summary statistics*

| | Argentina | Brazil | Chile | Mexico | US |
|---|---|---|---|---|---|
| Mean *(% per month)* | 0. 979 | 1.125 | 1.011 | 1.296 | 0.489 |
| Std-Dev. *(% per month)* | 14.962 | 16.155 | 7.204 | 9.587 | 4.328 |
| Skewness | $0.365^{**}$ | $-1.396^{*}$ | $-0.654^{*}$ | $-1.026^{*}$ | $-0.817^{*}$ |
| Kurtosis | $3.956^{*}$ | $9.790^{*}$ | $2.843^{*}$ | $3.237^{*}$ | $1.844^{*}$ |
| Jarque-Bera | $172.011^{+}$ | $1101.296^{+}$ | $104.112^{+}$ | $156.148^{+}$ | $64.575^{+}$ |
| Q(12) | 6.421 | $19.366^{+++}$ | 13.089 | $20.647^{+++}$ | $18.233^{+++}$ |
| ARCH(12) | $38.731^{+}$ | $18.988^{+++}$ | $20.960^{++}$ | $23.451^{++}$ | $29.447^{+}$ |

*Panel B: Autocorrelations*

| Lag | Argentina | Brazil | Chile | Mexico | US |
|---|---|---|---|---|---|
| 1 | 0.059 | -0.105 | 0.163 | 0.090 | 0.071 |
| 2 | -0.015 | 0.026 | 0.002 | 0.034 | -0.020 |
| 3 | 0.001 | -0.108 | -0.068 | 0.073 | 0.103 |
| 4 | 0.057 | -0.039 | 0.072 | -0.040 | 0.059 |
| 5 | -0.087 | 0.004 | 0.026 | -0.023 | 0.026 |
| 6 | -0.021 | -0.047 | -0.049 | -0.114 | -0.041 |

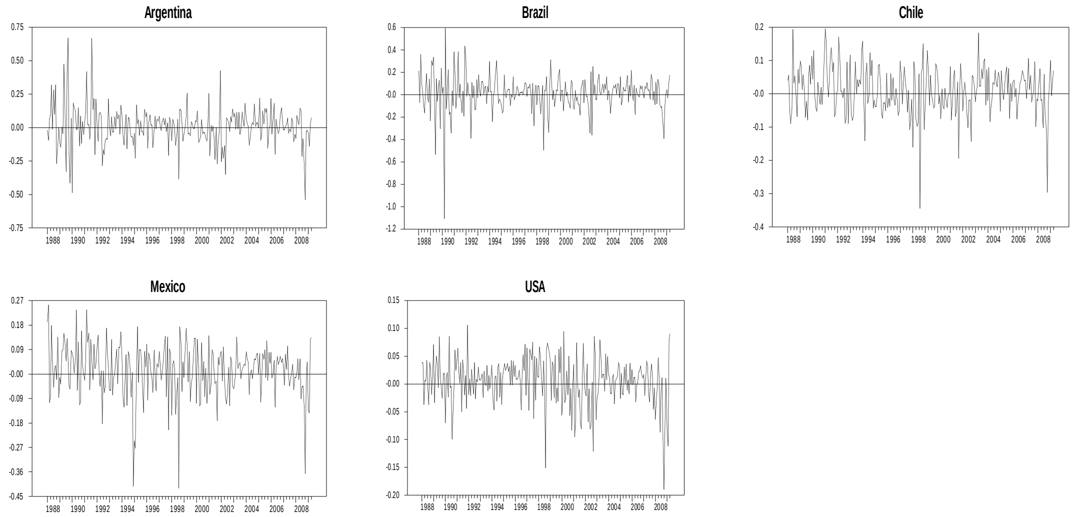
*Panel C: Unconditional correlation matrix of stock market returns*

| | Argentina | Brazil | Chile | Mexico | US |
|---|---|---|---|---|---|
| Argentina | 1.000 | | | | |
| Brazil | 0.293 | 1.000 | | | |
| Chile | 0.348 | 0.413 | 1.000 | | |
| Mexico | 0.479 | 0.378 | 0.468 | 1.000 | |
| US | 0.336 | 0.383 | 0.467 | 0.568 | 1.000 |

Notes: The test for Kurtosis coefficient has been normalized to zero. JB is the Jarque-Bera test for normality based on skewness and kurtosis. Q(12) is the Ljung-Box test for autocorrelation of order 12. ARCH is the Engle (1982)'s test for conditional heteroscedasticity. $^{*}$, $^{**}$ and $^{***}$ indicate significance of coefficients at the 1%, 5% and 10% respectively. $^{+}$, $^{++}$ and $^{+++}$ indicate rejection of the null hypotheses of no autocorrelation, normality and homoscedasticity at the 1%, 5% and 10% levels of significance respectively for statistical tests.

Panel A indicates that the monthly average of Latin American market returns ranges from 0.979% in Argentina to 1.296% in Mexico. Their returns are indeed consistently above those provided by the US stock market (0.489% per month). However, it is important to note that all the sample emerging markets experienced a very high level of unconditional volatility, Brazil being the most volatile market with a standard deviation of 16.155%, followed by Argentina, Mexico, and Chile. Investors should therefore be aware of the fact that some emerging markets may not be attractive in terms of risk-return tradeoff. Skewness and kurtosis coefficients are all significant at the conventional levels. The Jarque-Bera test for normality based on the third and fourth moments strongly rejects the hypothesis of normally distributed returns. These facts support our decision to use the quasi-maximum likelihood (QML) approach of Bollerslev and Wooldridge (1992) to estimate the empirical model. We also performed the Engle (1982) test for conditional heteroscedasticity and find that the null hypothesis of no ARCH effects is rejected for all of the stock markets. The null hypothesis of no autocorrelation of order 12 is rejected for Brazil, Mexico, and the US. The first-order autocorrelation, reported in Panel B, is significant for Chile. Altogether, this suggests that a final auto-regressive correction is needed in the mean equations.

**Figure 1**
**Dynamic patterns of monthly stock market returns**




We depict in Figure 1 the time-variations in return series from the sample stock markets. Unsurprisingly, these series are quite unstable and testify to periods of high volatility, especially during times of crisis. If we look closely at each market, we see that the Mexican stock market was particularly sensitive to the Tequila debt crisis of 1994-1995, and that the stock markets in Argentina and Brazil reacted strongly to their market-opening events during the years from 1989 to 1993 (Bekaert and Harvey, 2000). All the markets experienced sharp declines in returns at the time of the Asian financial crisis of 1997-1998 and the global financial crisis of 2007-2008, except for Brazil.

Panel C reports the unconditional correlations among markets. As expected, they are all positive and range from 0.293 (Brazil-Argentina) to 0.568 (Mexico-US). The important linkages between Mexico and the US can be broadly explained by the fact that both markets are members of the trilateral trade block in North America (NAFTA). Within the Latin American region, the highest unconditional correlation is between Argentina and Mexico (0.479). The values of the unconditional correlations are quite low, suggesting that the diversification benefits from investing in these emerging markets still remain substantial.

*3.2 Results of the DCC-MGARCH model*

Table 3 contains parameter estimates and a number of diagnostic tests for the DCC-MGARCH model. The coefficients relating the return series to the one-lag local and US market returns (Panel A) are insignificant, except for Chile where current returns are predictable from their AR(1) values. The effects of the US current stock returns on the dynamics of emerging stock returns are significant only in the cases of Argentina and Mexico. One can explain this finding by a high degree of stock-market integration, particularly between Mexico and the US. This result is consistent with that of Johnson and Soenen (2003) who use Geweke measures of feedback and find about 91% of contemporaneous association between the United States and Mexico. In their study, the same-day responses of Argentinean stock markets to the US ones are also significant. Our result confirms, to the fullest extent, the evidence of increased integration between Latin American emerging markets and the US markets, as documented by Choudry (1997) and Aggarwal and Kyaw (2005) on the basis of unit root and cointegration tests.

As for the ARCH and GARCH coefficients reported in Panel B, they are significant at the conventional levels for all the countries except Brazil. This is consistent with the time-varying volatility and justifies our choice of a GARCH-type model. More specifically, the small size of ARCH-term coefficients $\varpi_1$ indicates a low rate of change in conditional market volatility over time, while the large magnitude of GARCH-term coefficients $\varpi_2$ testifies to its time-dependence. It is equally important to note that conditional volatility in Argentina, Brazil, and the US tends to be particularly persistent over time in view of the sum $(\varpi_1 + \varpi_2)$, which is very close to unity. The DCC estimates, $\alpha = 0.065$ and $\beta = 0.394$, are significantly different from zero at the 1% level and satisfy the mean-reverting condition $\alpha + \beta \prec 1$. This typically implies that the conditional volatility of the sample stock markets slows to adjust to their "normal" equilibrium level, which is governed by the state of the economy.

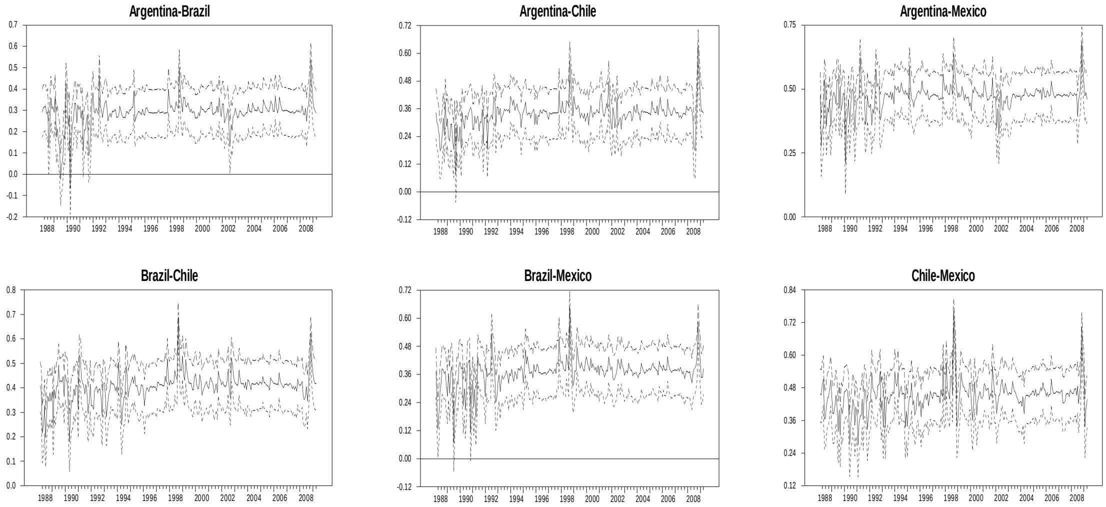
**Table 2**
**Estimation results of the DCC-MGARCH model**

| *Panel A – parameter estimates of the mean equations* | | | | | |
|---|---|---|---|---|---|
| | Argentina | Brazil | Chile | Mexico | US |
| $\delta_0$ | 0.0077 | 0.010 | 0.009** | 0.013*** | 0.006* |
| | (0.007) | (0.011) | (0.004) | (0.013) | (0.002) |
| $\delta_1$ (1st lag of local market returns) | -0.028 | -0.047 | 0.127** | -0.007 | -0.019 |
| | (0.058) | (0.105) | (0.066) | (0.058) | (0.058) |
| $\delta_2$ (current lag of the US market returns) | 0.559** | 0.224 | 0.110 | 0.440* | --- |
| | (0.246) | (0.397) | (0.136) | (0.147) | |
| $\delta_3$ (1st lag of the US market returns) | 0.244 | 0.187 | 0.062 | 0.142 | -0.019 |
| | (0.177) | (0.223) | (0.091) | (0.128) | (0.058) |
| *Panel B - parameter estimates of the variance processes* | | | | | |
| | Argentina | Brazil | Chile | Mexico | US |
| $\varpi_0$ | 0.001 | 0.026** | 0.006* | 0.001 | 0.000 |
| | (0.001) | (0.013) | (0.001) | (0.001) | (0.000) |
| $\varpi_1$ | 0.134* | 0.121 | 0.059 | 0.121* | 0.174* |
| | (0.042) | (0.113) | (0.044) | (0.048) | (0.031) |
| $\varpi_2$ | 0.853* | 0.189 | 0.470*** | 0.875* | 0.826* |
| | (0.053) | (0.327) | (0.278) | (0.055) | (0.031) |
| Variance persistence $[(\varpi_1 + \varpi_2)]$ | 0.987 | 0.310 | 0.529 | 0.996 | 1.000 |
| $\alpha$ | 0.065* | | | | |
| | (0.023) | | | | |
| $\beta$ | 0.394* | | | | |
| | (0.159) | | | | |
| *Panel C - Robust tests for model standardized residuals* | | | | | |
| | Argentina | Brazil | Chile | Mexico | US |
| Mean | -0.017 | -0.019 | -0.021 | -0.036 | -0.020 |
| Std-Dev. | 1.018 | 0.996 | 0.981 | 0.991 | 0.954 |
| Skewness | 0.228 | -1.917* | -0.626** | -1.105* | -0.656 |
| Kurtosis | 2.053* | 12.424* | 2.554* | 4.349* | 1.047* |
| JB | 47.004+ | 1796.291+ | 86.019+ | 252.924+ | 29.991+ |
| Q(12) | 4.107 | 16.059 | 7.224 | 19.164+++ | 13.152 |
| ARCH(12) | 11.288 | 16.084 | 13.433 | 3.071 | 3.277 |

Notes: Bollerslev and Wooldridge (1992)'s robust standard errors are given in parentheses. $\varpi_0$, $\varpi_1$ and $\varpi_2$ refer to the parameters of a GARCH(1,1) process. The test for Kurtosis coefficient has been normalized to zero. JB is the Jarque-Bera test for normality based on excess skewness and Kurtosis. Q(12) is the Ljung-Box test for autocorrelation of order 12. ARCH is the Engle (1982) test for conditional heteroscedasticity. *, ** and *** indicate significance of coefficients at the 1%, 5%, and 10% respectively. +, ++ and +++ indicate rejection of the null hypotheses of no autocorrelation, normality, and homoscedasticity at the 1%, 5%, and 10% levels of significance respectively for statistical tests.

We provide some diagnostics of the model's residuals in Panel C in order to assess the appropriateness of the empirical model. The indices of kurtosis in the filtered return series are lower in most cases than what we found for the raw returns. Unfortunately, the results of the Jacque-Bera test for normality do not support the proposition that the conditionally normal GARCH process is sufficiently fat-tailed to accommodate the excess kurtosis in the data. We also apply the Ljung-Box test for autocorrelation and the Engle (1982)'s test for ARCH effects to the estimated residuals and find that our DCC-MGARCH specification is powerful enough to capture the dynamics of the returns and the conditional covariance matrix. There is, after all, weak evidence of significant autocorrelation of return innovations in Mexico (at the 10% level).

**Figure 2**
**DCCs among Latin American emerging markets**



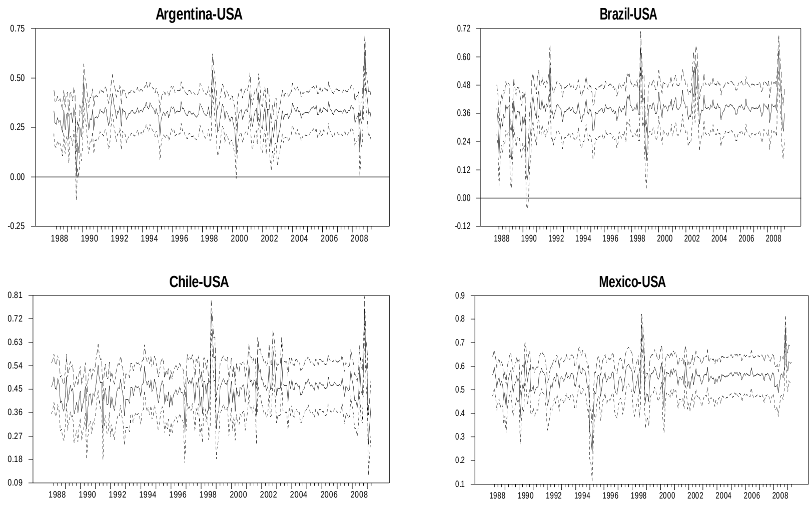

The patterns of dynamic conditional correlations among Latin American emerging markets, and between them and the US markets together with their 95% confidence levels, are shown in Figure 1. Several interesting observations can be made. First, cross-market dynamic correlations are positive throughout the study period, confirming that the sample stock markets exhibit a certain degree of financial interdependence and significant comovements. They are relatively low and average 0.401 for the whole sample. The average correlation of the Latin American markets studied and the US market is 0.422, compared to 0.387 for the emerging universe. These findings are in agreement with evidence in the literature that global integration in Latin American stock markets proceeds faster than regional integration (Barari, 2004).

Second, the dynamic linkages between stock markets vary considerably over time and bear witness to some periods of great instability. The highest conditional correlations were observed in November 2008 during the global financial crisis period. In fact, during that month, the cross-market conditional correlations grew to 0.656 on average, while the average correlation of emerging and US markets increased to 0.700. Other important peaks of market comovement were observed during the Mexican crises (1994), during the Asian crisis (1997-1998), and after the terrorist attacks on the US World Trade Center (2001). For US investors, this should mean low diversification gains from adding Latin American emerging market assets during these crisis periods. Note however that for all the abrupt jumps observed, the correlations did not increase sharply at the start of the crisis, which implies that crisis shock transmission among markets is not immediate but occurs with a certain time lag.

Third, contrary to expectation, there was a notable drop in conditional correlations between Mexico and the US in March 1995, just after the start of the debt crisis marked by the peso's 34% fall. This behavior indicates that the contagion effect was not present, and can be explained by the reimposition of controls on foreign ownership and the increase in the capital limit requirements for banks. For example, the Mexican government decided to reduce the maximum amount of equity in a NAFTA affiliate bank that a NAFTA investor was authorized to hold from 99% to only 51%.

Finally, despite the fact that emerging markets have become more open to foreign direct investments and portfolio flows following the market liberalization policies of the 1990s, dynamic correlations in Figure 3 seem to indicate that their financial links with the US did not strengthen significantly.

**Figure 3**
**DCCs between emerging Latin America and the US**



Summarizing all of the above, the correlations of four emerging markets with the US market vary over time, and reached some peaks, but their levels did not greatly change after crises.

*3.3 Structural changes*

We now investigate whether any structural change has occurred in the conditional correlations of four Latin American markets and the US market over the estimation period. The Bai and Perron (2003)'s testing procedure, described in Section 2, was applied to detect and date the structural breakpoints. The optimal number of breakpoints ($m$) should be the one associated with the minimum BIC. In Table 3 we report the selected optimal breakpoints for each market together with their 95% confidence intervals.

The null hypothesis of stability is rejected for all the markets, since the Bai-Perron test detects at least one breakpoint in the correlations of all four emerging markets with the US. In Argentina, three significant breakpoints are obtained, while in the three remaining countries one breakpoint is detected. This implies a clear change in the nature of the comovement structure, and thus has important implications for market authorities and global investors in managing their coordination policies and portfolios respectively. For instance, an upward trend in inter-market comovement after a structural break would tend to reduce the allocation rate of a US investor's portfolio into Latin American markets.

Another interesting question arising from structural change analysis is whether the break dates estimated from Bai and Perron's procedure coincide with important market events in the various counties under consideration. We first observe that the break dates in the correlations of three emerging markets with the US occurred around the time of the Asian financial crisis of 1997-1998 (Argentina, Chile, and Mexico). Other crisis episodes did not appear to cause structural changes in market comovements. We then attempt to match the remaining break dates with the liberalization events in emerging markets obtained from Bekaert and Harvey (2000). Our motivation comes from the fact that stock market liberalization is part of a broader set of economic and financial reforms undertaken by emerging countries to improve economic growth and market efficiency, and structural changes in their relationship with other markets are therefore expected. Accordingly, we find that the official liberalization date in Brazil (May 1991) falls within the 95% confidence intervals of the correlation break, whereas the 95% confidence intervals of one correlation break in Argentina contain the dates of the first ADR (American Depository Receipt) and country fund introduction (August and October 1991 respectively) as well as the date of a break in US capital flows to Argentina (March 1993).

**Table 3**

**Estimates of structural breakpoints in the conditional correlations with the US markets**

| Countries | Test parameters | | | Optimal breakpoints | | | |
|---|---|---|---|---|---|---|---|
| | Best *m* breakpoints | RSS | BIC | Optimal number of breakpoints | Estimated breakpoint dates | 95% confidence levels for breakpoint dates | |
| | | | | | | *Lower bound* | *Upper bound* |
| Argentina-US | 0 | 0.8961 | -706.2539 | 3 | Mar-91 | Oct-90 | Jun-93 |
| | 1 | 0.8379 | -712.2976 | | | | |
| | 2 | 0.8218 | -706.1763 | | Dec-98 | Oct-96 | Oct-99 |
| | 3 | 0.7677 | **-712.4295** | | | | |
| | 4 | 0.7641 | -702.5746 | | Nov-03 | Mar-02 | Aug-05 |
| | 5 | 0.7806 | -686.0250 | | | | |
| Brazil-US | 0 | 0.9864 | -681.7696 | 1 | Mar-91 | Oct-90 | Mar-93 |
| | 1 | 0.8464 | **-709.7081** | | | | |
| | 2 | 0.8303 | -703.5334 | | | | |
| | 3 | 0.8234 | -694.5772 | | | | |
| | 4 | 0.8217 | -684.0340 | | | | |
| | 5 | 0.8191 | -673.7593 | | | | |
| Chile-US | 0 | 0.8565 | -717.7826 | 1 | Aug-98 | Nov-94 | Mar-00 |
| | 1 | 0.7875 | **-728.1122** | | | | |
| | 2 | 0.7759 | -720.8143 | | | | |
| | 3 | 0.7708 | -711.4178 | | | | |
| | 4 | 0.7695 | -700.7624 | | | | |
| | 5 | 0.7700 | -689.5067 | | | | |
| Mexico-US | 0 | 0.7236 | -760.7628 | 1 | Jun-97 | Oct-95 | Oct-00 |
| | 1 | 0.6599 | **-773.2063** | | | | |
| | 2 | 0.6482 | -766.6577 | | | | |
| | 3 | 0.6434 | -757.5006 | | | | |
| | 4 | 0.6388 | -748.2408 | | | | |
| | 5 | 0.6371 | -737.8267 | | | | |

Notes: The breakpoint selection procedure in the works of Bai and Perron (2003) is based on the Bayesian Information Criteria (BIC). We arbitrarily set the maximum number of breaks to be 5. If the effective number of breaks is equal to 5, a higher number of breaks will be chosen so that the testing procedure captures all possible breakpoints. A model's optimal number of breakpoints should be the one associated with the minimum BIC. For the countries considered in this present study, none of the volatility series has more than 5 breakpoints.

Overall, major stock market events such as financial crises, official liberalizations, and ADR and country fund introduction have significant impacts on the comovements of Latin American emerging markets. Their correlations with the US generally increased in the year following the lifting of international investment barriers.

*3.4 Comovement asymmetries around financial crises*

The majority of past studies claim that the comovement of stock markets is stronger during a crisis period than during normal or tranquil ones. In this paper we test this assertion by investigating the differences in the level of conditional correlations for three crisis periods: the Mexican crisis (1994), the Asian crisis (1997), and the subprime crisis (2007). More specifically, we compare the average correlation of each emerging market with the US market 24 months prior to the crises, to those computed 24 months after the crises[4]. Our method consists of using a simple two-tailed parametric test, called a T-test to compare the means between two subperiods. This test investigates the null hypothesis of no increase in correlations; the empirical *t*-statistic, used to make test decisions, is given by $t^* = (\bar{x} - m)/(s/\sqrt{n})$, where $\bar{x}$ refers to the average correlation of the crisis period, $m$ is the average of the tranquil period, $s$ refers to the unbiased standard deviation of the crisis period, and $n$ is equal to 24. Assume that the conditional correlation series is normally distributed, then under the null hypothesis of no increase in correlations the $t^*$ follow a Student-*t* distribution with ($n$-1) degrees of freedom[5].

Tables 4, 5, and 6 report the T-test results for the three crisis episodes we consider. As far as the Mexican crisis is concerned, the post-crisis correlations are lower than the pre-crisis ones. The T-test rejects the null hypothesis of no increase in correlations for three markets: Argentina (at the 10% level), Chile (5%), and Mexico (1%). Given the symmetry of the Student-*t* distribution and the negative

[4] The event window is shorter for the global financial crisis of 2007-2008 (22 months before and after the crisis) owing to insufficient data points at the time this research was performed.

[5] Note that we also tested for the significance of changes in correlations by using a standard *Z*-test after standardizing the correlation coefficients of the two subperiods according to the Fisher transformation formula. However, the test results are however similar to those of the T-test presented here.

values of empirical *t*-statistics, the test results indicate that cross-market correlations did not increase significantly after the Mexican crisis. The change in correlation for Brazil is not significant.

**Table 4**

**Test of changes in conditional correlations around the 1994 Mexican crisis**

| **Pre-crisis period:** December, 1992 to November 1994 | | | | |
|---|---|---|---|---|
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.336 | 0.367 | 0.448 | 0.559 |
| Std. Dev. | 0.020 | 0.022 | 0.037 | 0.024 |
| Jarque-Bera | 0.043 | 0.424 | 2.306 | 0.109 |
| Probability | 0.979 | 0.809 | 0.316 | 0.947 |
| **Post-crisis period:** December, 1994 to November, 1996 | | | | |
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.325 | 0.365 | 0.435 | 0.496 |
| Std. Dev. | 0.032 | 0.026 | 0.028 | 0.094 |
| Jarque-Bera | 61.297 | 35.247 | 0.192 | 13.604 |
| Probability | 0.000 | 0.000 | 0.908 | 0.001 |
| **T-test for the null hypothesis of mean equality** | | | | |
| Changes in means | -0.011 | -0.002 | -0.014 | -0.063 |
| *t*-statistics | -1.751 | -0.455 | -2.424 | -3.275 |
| *p*-value | 0.093 | 0.654 | 0.024 | 0.003 |

Notes: this table shows the differences in the means of conditional correlation coefficients around the 1994 Mexican crisis (24 months before and 24 months after). *t*-statistics and *p*-values refer respectively to the empirical statistics and the associated probability of the two-tailed Student-*t* test for the null hypothesis of no increase in correlations. The T-test was chosen owing to the non-normality of sample data. Jarque-Bera denotes the empirical statistic of the test for normality of the series studied.

In sharp contrast to the previous crisis, the 24-month average comovement between Latin American emerging and US markets rose in three countries (Brazil, Chile, and Mexico) following the Asian financial crisis of 1997-1998, but according to the T-test results this positive increase is only significant for Mexico (Table 5).

Turning to the global financial crisis of 2007-2008, which is often compared to the Great Depression in 1929 for its harmful impacts on world economy, no significant change in inter-market comovements is observed, as the null hypothesis of the T-test cannot be rejected. This finding is particularly interesting because the US mortgage crisis has spread to almost all areas of the globe, including the financial markets in Latin America, and many economists pleaded in favor of contagion effects.

**Table 5**

**Test of changes in conditional correlations around the 1997 Asian crisis**

| **Pre-crisis period:** July, 1995 to June, 1997 | | | | |
|---|---|---|---|---|
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.327 | 0.367 | 0.440 | 0.511 |
| Std. Dev. | 0.028 | 0.025 | 0.041 | 0.081 |
| Jarque-Bera | 114.469 | 31.329 | 37.506 | 46.335 |
| p-value | 0.000 | 0.000 | 0.000 | 0.000 |
| **Post-crisis period:** July, 1997 to June, 1999 | | | | |
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.317 | 0.386 | 0.454 | 0.557 |
| Std. Dev. | 0.068 | 0.073 | 0.084 | 0.071 |
| Jarque-Bera | 21.015 | 35.378 | 20.328 | 5.753 |
| p-value | 0.000 | 0.000 | 0.000 | 0.056 |
| **T-Test for the null hypothesis of mean equality** | | | | |
| Changes in means | -0.011 | 0.019 | 0.014 | 0.046 |
| *t*-statistics | -0.925 | 1.546 | 0.969 | 3.889 |
| *p*-value | 0.362 | 0.131 | 0.339 | 0.000 |

Notes: see notes to Table 4.

**Table 6**

**Test of changes in conditional correlations around the 2007-2008 global financial crisis**

| **Pre-crisis period:** September, 2005 to June, 2007 | | | | |
|---|---|---|---|---|
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.334 | 0.382 | 0.468 | 0.564 |
| Std. Dev. | 0.015 | 0.013 | 0.011 | 0.011 |
| Jarque-Bera | 11.668 | 5.043 | 3.198 | 5.251 |
| p-value | 0.003 | 0.080 | 0.202 | 0.072 |
| **Post-crisis period:** July, 2007 to April, 2009 | | | | |
| | *ARG-US* | *BRA-US* | *CHI-US* | *MEX-US* |
| Mean | 0.341 | 0.396 | 0.456 | 0.572 |
| Std. Dev. | 0.096 | 0.073 | 0.100 | 0.056 |
| Jarque-Bera | 23.005 | 15.587 | 8.586 | 32.163 |
| p-value | 0.000 | 0.000 | 0.014 | 0.000 |
| **T-Test for the null hypothesis of mean equality** | | | | |
| Changes in means | 0.006 | 0.013 | -0.012 | 0.008 |
| *t*-statistics | 0.308 | 0.873 | -0.579 | 0.650 |
| *p*-value | 0.761 | 0.393 | 0.569 | 0.523 |

Notes: see notes to Table 4.

Taken together, our results do not provide evidence to support the proposition according to which Latin American stock markets tend to comove much more with the US stock markets during times of crisis than during normal times. In other words, US investors would still benefit from investing in these markets, provided that they are aware of significant peaks of comovement in times of crisis and possible changes in the nature of the cross-market comovement structure.

## 4. Conclusion

In this paper we investigate the comovements between four Latin American emerging stock markets (Argentina, Brazil, Chile, and Mexico) and the US stock market, using the recent period from February 1988 to April 2009. Our empirical analysis was conducted within the framework of a DCC-MGARCH model proposed by Engle (2002) and the Bai and Perron (2003)'s structural change test. Our method thus enables us to examine not only the time-varying trends in market comovements (or more broadly the regional and global integration of Latin American emerging markets), but also the changing nature of these comovements.

Several interesting findings emerge from our analysis. First, the estimates of DCC-MGARCH and time-varying correlation graphs show significant market linkages between Latin American and US stock markets. However, these correlations remain below 0.50 almost all of the time. In particular, the US markets allow us to explain the contemporaneous dynamics of stock returns in two emerging markets, Argentina and Mexico, with the impact being bigger for Mexico. Second, we document the presence of at least one structural breakpoint in the dynamic conditional correlations of emerging markets and the US market. These breakpoints are generally found to coincide with some major market events such as a financial crisis and stock market reforms. Third, our results are not consistent with the view that contagion effects exist during times of crisis, based on sample market data. Correlations increased significantly only in the aftermath of the Asian crisis in the case of Mexico and the US.

Overall, our results, in line with the findings of Diamandis (2009), suggest that US investors may still benefit from including assets issued by Latin American stock markets in their portfolios. Mean-while, the presence of structural breaks in the correlation patterns implies that a reassessment of stock market comovements is indeed necessary before implementing any investment strategy in Latin America, and in particular in a follow-up of major stock market events.

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