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Angelo Marsiglia Fasolo

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Edited by Research Department (Depep) – E-mail: workingpaper@bcb.gov.br

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Interdependence and Contagion: an Analysis of Information Transmission in Latin America's Stock Markets

Angelo Marsiglia Fasolo[*]

Abstract

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This paper brings evidences about the hypotheses of financial crisis contagion over Latin American stock markets in the 90's using a multivariate GARCH model. We added to the volatility structure a leverage term in order to avoid problems due to the use of conditional correlation as a measure of relationship between stock markets. Results point to the existence of contagion only during the Asian (1997) and the Russian (1998) crises. The consequences of the change in the exchange rate regime in Brazil (1999) can be identified as a result of interdependence among Latin American markets, while the Mexican (1994) and Argentinean (2001) crises show a specific mechanism of propagation. This result raises questions about the adequacy of the "contagion" and "interdependence" concepts for the information transmission analysis among stock markets.

[*] Research Department, Central Bank of Brazil. The author would like to thank Fernando Alberto S. Rocha, Christiane Rocha Albuquerque, Carlos Hamilton Vasconcelos Araújo e José Álvaro Rodrigues Neto for comments on previous versions of this paper. The views expressed here are solely those of the author and do not necessarily reflect those of the Institution or its staff. E-mail: angelo.fasolo@bcb.gov.br.

1. Introduction

The liberalization of financial markets and capital flows, deepened during the 1990's, raised questions concerning the conduct of national economic policies. In this context, studies on information transmission between financial markets and its effects over asset prices grew in importance. The joint movement of asset markets in different countries, especially in crisis events – the so called "contagion" – became an actual topic of discussion in financial literature.

Latin American stock markets, in particular after last decade's economic reforms, were affected by a sequence of shocks originated from different sources. Their performance was conditioned not only by internal crises – as that of Mexico in 1994 or the devaluation Argentinean currency in 2001 – but also by external crises, which, in general, had origins in countries with little in common with the Latin America's economies. It's important to highlight the collapse of Southern Asian economies in 1997 and the devaluation of the Russian Rublo in 1998.

Amongst the available techniques to study markets' relations, the multivariate ARCH/GARCH models represent the classical framework of analysis. However, the excessive number of parameters in those formulations conditions their use. The approach proposed by ENGLE (2002) not only reduces this problem, but also offers an immediate hypothesis test about information transmission among markets. In this sense, it is possible to test the existence and intensity of the linkages among markets through conditional correlations.

Despite its econometric advantages, the basic model in ENGLE (2002) does not incorporate important developments from traditional univariate models, such as leverage effects of volatility under negative news. Thus, we added a GJR formulation (see GLOSTEN *et alli*, 1993) applied to the variance structure proposed by ENGLE (2002). By doing so, we can expect to detect occasional asymmetry and leverage phenomena over stock markets' returns.

Concerning the results, it is possible to say, in advance, that only the effects from Asian and Russian crises can be seen in the classical definition of contagion. Other shocks, especially those originated from Latin American countries, are cases of interdependence or represent an alternative pattern of information transmission that we call "isolation effect".

Section 2 brings a brief literature review, focusing mainly in the measure of contagion among stock markets, which is the objective of this article. Section 3 describes the methodology used, emphasizing ENGLE (2002)'s multivariate model. Results are presented in section 4, highlighting the relations among markets, while section 5 discusses, based on some results presented in the previous section, the sufficiency of the concepts proposed in the literature about markets' interdependence. Finally, section 6 concludes.

1. Review of the Literature: Contagion and Spillover Effects

The three classical approaches to test contagion among markets are[1]: the assets returns' correlation, cointegration analysis and volatility models. The first methodology consists of calculation the returns'correlation in the so-called normal periods of time and compares results to those obtained in crisis' periods. Cointegration tests try to establish a long-run relationship between markets, leaving behind components such as the markets' short-run volatilities. Volatility models try to conciliate the long run and short-run analysis, considering not only the first but also higher statistical moments of assets' returns. Other hypothesis, such as non-normality in the returns' distribution or the influence of markets' volatility over the mean of returns, may be tested with no harm to results, especially due to large samples from high frequency data.

Evidences about the existence of a relevant channel of information transmission among markets are mixed. LIN, ENGLE and ITO (1994) find little evidence of causality between the US and Japanese markets. However, BAE and KAROLYI (1994) state that not considering the asymmetric responses of shocks may bias this kind of causality analysis. FLEMING and LOPEZ (1999) use GARCH models to find evidences of the interdependency among New York, London and Tokyo's Treasury markets. Their conclusions points to the influence of the US market over the others, but not the other way round.

Concerning Latin America's stock market, the literature is still incipient. CHRISTOFI and PERICLI (1999) estimate an EGARCH, with a VAR structure in the mean equations, while the estimated covariances follow a constant conditional correlation (CCC) pattern. Authors find asymmetric effects in the markets' responses, especially in the presence of bad news. However, the US market is not incorporated in the analysis, what may cause distortions in results, due to absence of an influential factor to the average return.

CHOEIRI (1999) estimates a common factor model to the foreign exchange reserves' stock in Argentina, Brazil, Chile and Mexico, based on the similar pattern in the movements of capital flows to Asian and Latin American markets throughout the 70's and in the early 90's. Her study confirms a common dynamics among the variables in times of high international liquidity. The same does not occur in crisis period, when the estimated factor is not significant. The author points out that, maybe, these movements are determined by financial markets' linkages, instead of economies' fundamentals.

EDWARDS (1998) tests the volatility of the interest rates of Chile, Argentina and Mexico. The author finds contagion effects from Mexico to Argentina during 1994 crisis, but

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[1] See FORBES e RIGOBON (2000) about the first three methods.

do not find the same result from Mexico to Chile. The absence of a relationship in more than one direction is confirmed by DÍEZ DE LOS RIOS and HERRERO (2003), when testing the causality relations among external debt bonds' returns from emerging markets economies. Using the Par Bonds from Mexico, Argentina, Brazil and Venezuela, MORAIS and PORTUGAL (2001) detect common movements among the returns' volatilities, especially between the first two countries. LOPES and MIGON (2002) also apply common factors models to Latin America and point out that, during every crisis periods in the sample, the four major stock markets in the region had suffered contagion.

FORBES and RIGOBON (2000) show some necessary corrections in the contagion measures to analyze financial markets. After those corrections, the changes in bonds and stock markets' returns in Latin America show only the interdependence among the countries, and not contagion, as it could be expected. However, as a long-run concept, the acceptance of the interdependence hypotheses must be seen with some careful, once TABAK and LIMA (2002) finds only a short-run causality among those markets, rejecting the existence of a cointegration vector[2]. The corrections proposed by FORBES and RIGOBON (2000) are also criticized by BAIG and GOLDFAJN (2000), who point that volatility increases are an inherent characteristic of crisis. Thus, correcting the estimated correlation by eliminating the change in volatility is the same as estimating a new value without considering relevant features of those periods.

2. Applied Volatility Models: Methodology

This section consists of two parts. The first one presents the concepts that are the basis to the applications from which the results were derived. The second part presents the econometric methodology and the consequences of the hypothesis that were set according to the chosen model.

By contagion among stock markets we mean a significant increase in their relations after a shock in one of the markets, as defined in FORBES and RIGOBON (2000, p.13). It is important to notice that, according to the authors, high values of correlation (or of any other measure of relation between assets) mean interdependence among markets, but not contagion. What is important to identify contagion is a significant change in level of the relation among markets during crisis periods.

Here, the relation among stock markets is measured by the conditional correlation that is estimated in the volatility models of section 4. This assumption turns the tests into simpler ones, since contagion can be verified only by the analysis of the median correlation between normal and crisis periods[3]. The VAR form of the mean equation allows us to correct the bias

[2] These results confirm the conclusions reached by SINHA e CASTAÑEDA (1999).

[3] About the validity of the concept, its implications and the transmition mechanisms, see FORBES and RIGOBON (2000, 2002). In section 5, it is discussed only the sufficiency of the concept, based in one of the

in the analysis of contagion based on ARCH/GARCH models, as pointed out in FORBES and RIGOBON (2000, p. 23), concerning the omission of relevant variables. Hence, occasional shocks in one of the countries involved would propagate the impact through all over the system.

The use of ARCH/GARCH models is highly spread in the literature, with various applications in economics and finance. Recently, there has been an increase in the number of studies with multivariate formulations. The growth in the number of parameters to be estimated is a restriction to these models[4], bringing some problems to guarantee the convergence of the estimation's algorithms.

Amongst the possible restrictions proposed to the covariance matrix, ENGLE (2002)'s Dynamic Conditional Correlation (DCC) seems to be the most interesting. Consider that the conditional variance of the returns of an asset *i* is given by $h_{i,t} = E_{t-1}(r_{i,t}^2)$. Then the returns may be written as:

$$r_{i,t} = \sqrt{h_{i,t}}.\varepsilon_{i,t}, where: \varepsilon_{i,t} \sim N(0, R_t) \qquad (1)$$
$$i = 1,2,...,n$$

In this case, ε_t is, in matrix notation, the vector of standardized residuals, R_t is the returns' correlation matrix and h_t is the variance-covariance matrix. Using these definitions in the estimator of the conditional correlation between two assets, *i* and *j*, we have that:

$$\rho_{i,j,t} = \frac{E_{t-1}(\varepsilon_{i,t}\varepsilon_{j,t})}{\sqrt{E_{t-1}(\varepsilon_{i,t}^2)E_{t-1}(\varepsilon_{j,t}^2)}} = E_{t-1}(\varepsilon_{i,t}\varepsilon_{j,t}) \qquad (2)$$

Defining the variance-covariance structure as:

$$H_t = D_t R_t D_t, \qquad (3)$$

where $D_t = diag\{\sqrt{h_{i,t}}\}$, the equations representing the individual variances of the assets follow a GARCH process similar to the univariated models, while the correlation matrix is estimated in each moment of time[5]. The restriction applied to the variance matrix asserts a unique GARCH process to the conditional correlation matrix. Here, the correlation estimator has the following form:

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obtained results.

[4] In a GARCH (1,1) formulation, with two dependent variables, it is necessary to compute the values of 21 parameters to get the estimation completed. The same formulation using three dependent variables needs the estimation of 78 parameters.

[5] Note that, leaving D_t on the left side of the equation, we obtain the conditional correlation estimator. See equation 4.

$$Q_t = S\left(1 - \sum_{r=1}^{r} \alpha_r - \sum_{s=1}^{s} \beta_s\right) + \sum_{r=1}^{r} \alpha_r (\varepsilon_{t-r}\varepsilon_{t-r}^{|}) + \sum_{s=1}^{s} \beta_s Q_{t-s} \qquad (4)$$
$$R_t = diag\{Q_t\}^{-1} Q_t diag\{Q_t\}^{-1}$$

where α_i and β_i are scalars.

The linkage between the concepts of contagion and interdependence is a natural consequence. Following LOPES and MIGON (2002), the relation between two markets is given by the conditional correlation between the assets' returns, estimated at each period. This scheme also corrects problems caused by omitted variables and heteroskedasticity, pointed out by FORBES and RIGOBON (2000) as potential sources of distortion in the correlations´ results.

Finally, considering the asymmetry effects of the responses to shocks, we adopt the volatility model first presented in GLOSTEN *et al* (1993), denoted by GJR. In this formulation, the variance equations are given by:

$$h_{i,t} = w_i + \sum_{p=1}^{p} A_{i,p} h_{i,t-p} + \sum_{q=1}^{q} B_{i,q} \mu_{i,t-q}^2 + \sum_{d=1}^{d} \chi_{i,d} B_{i,d} \mu_{t-d}^2 \qquad (5)$$
$$where \qquad \chi_i = \begin{cases} 1, if \mu_{i,t-1} < 0 \\ 0, if \mu_{i,t-1} \geq 0 \end{cases}$$

In equation 5, the dummy χ_i is set endogenously throughout the estimation. It reflects the additional leverage effect caused by a negative surprise in the market. Thus, for each market, the dummy variable will be estimated in the equation of its own variance process.

In order to simplify the computational process, we assumed that the mean equations have no more than one lag in the VAR structure, the GARCH-GJR(p,q,d) structure must respect the conditions Max(p)=Max(q)=2 and Max(d)=1, while the correlation equation D(r,s) obeys that Max(r)=Max(s)=2. A higher number of lags would generate problems similar to those present in the estimations with smaller restrictions in the variance equations. The models are chosen according to the likelihood ratio tests, in the case of nested models, and out-of-sample forecasts statistics to the other comparisons. All tested models are bounded to present a positive semidefinite variance matrix. This necessary condition excluded almost every model with an additional lag on the μ^2 term.

3. Application: Analysis to Latin-American Market

Our data set consists of daily information about the end of the day returns, in US dollars, of the following Latin-American stock markets indexes: Brazil (IBOVESPA), Mexico (IPC), Argentina (MERVAL) and Chile (IGPA). The returns of US markets are represented by the Standard & Poors 500 (S&P 500), because this index has the largest coverage, in terms of number of companies and role in the economy. All these indexes represent benchmarks in their own countries. The sample ranges from August 1st, 1994 to May 5th, 2003, with a total of 2263 observations, after the necessary adjustments. Absent information, due to national holidays, were replaced by the data of the previous day.[6]

The choice of representative assets for Latin America was based on two aspects: the distortion caused by the model's restrictions and the daily average value, in US dollars, traded by each stock market. The restrictions upon the DCC structures, mentioned in ENGLE and SHEPPARD (2001), end up causing a natural restriction over the number of variables. In simulations using Dow Jones and S&P 500 stocks, the portfolio's choice was not always optimal because of the unique dynamic structure of correlations: as the number of assets in the portfolio increases, the DCC model tends to underestimate the assets' variances. Thus, in order to avoid estimation problems, a total of five assets seems to be a reasonable representation for the stock market in the region. Table A shows the average volume (in US dollars) traded in each market. Following this criterion, a natural choice is the four largest Latin American stock markets, considering the discrepancy when compared to the ones in other countries.

TABLE A – Average Traded Volume – Latin-American Stock Market – August, 94 to June, 03

Stock Market	Average Volume (in US$)
1. Brazil – IBOVESPA	323,529,584
2. Mexico – IPC	158,076,318
3. Chile – IGPA	25,410,683
4. Argentina – MERVAL	24,871,015
5. Venezuela – IBC	9,598,514
6. Peru – IGBVL	5,431,841

Source: Bloomberg.

Table B shows the descriptive statistics of the returns. One can notice that only the US market had a positive average return in that period. Concerning risk, the markets with smaller variance and standard errors of the sample are those of Chile and the USA. The kurtosis values, indicating a platykurtic distribution and the low asymmetry of returns also draw attention. Hence, the rejection of the normal-distribution hypothesis by the Jarque-Bera test was expected.

[6] Common holidays (Eastern, Christmas and January 1st) were excluded from the sample.

TABLE B – Descriptive Statistics of the Returns

	S&P 500 – US	*IBOVESPA - BRA*	*IPC - MEX*	*MERVAL - ARG*	*IGPA – CHI*
Average	0.000312	-0.000029	-0.000049	-0.000391	-0.000091
Standard Error	0.011738	0.028401	0.021291	0.027398	0.009062
Sample Variance	0.000138	0.000807	0.000453	0.000751	0.000082
Kurtosis	3.116009	6.177906	13.331070	62.180004	3.734255
Asymmetry	-0.105220	0.099108	-1.172025	-3.379893	0.081100
Jarque-Bera	914.1046	3583.530	17194.11	367228.6	1309.681
Minimum	-0.071127	-0.172462	-0.227132	-0.518372	-0.048967
Maximum	0.055697	0.237176	0.117112	0.161165	0.059337

Table C presents the sample correlation between assets. The highest values are found between the IPC and the S&P500 (0.4426), the IBOVESPA and the MERVAL index (0.4772) and the IBOVESPA and the IGPA (0.4822). It is obvious that these values must be seen carefully, because they reveal the “average” correlation in the analyzed period: if the variance changes across time, these correlations are no longer valid. Table D, for instance, shows, in the lower diagonal, the correlations during the Mexico’s crisis (October, 1994 to December, 1995) and, in the upper diagonal, those during the change in the exchange rate regime in Brazil (January to December, 1999). The differences between correlations during crises and with the whole sample correlation show the importance of a careful analysis of this phenomenon.

TABLE C – Returns’ Correlation – August, 94 to June, 03

	S&P 500 USA	*IBOVESPA BRA*	*IPC MEX*	*MERVAL ARG*	*IGPA CHI*
S&P 500 – USA	**1**				
IBOVESPA – BRA	0.413124	**1**			
IPC – MEX	0.442572	0.388471	**1**		
MERVAL – ARG	0.294935	0.477175	0.341101	**1**	
IGPA – CHI	0.312176	0.482160	0.304869	0.338996	**1**

Before any time-series analysis, one needs to consider the stationarity of the series involved. Table E reports the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests and graph 1 shows the evolution of the series in levels. Indexes’ returns, calculated by the difference between the natural logarithm of the series in level, have the expected stationary behavior. The PP test, in financial series, is fundamental due to the correction of the test statistics for the heteroskedaticity problem.

TABLE D – Returns' Correlation –Crises' Periods

	S&P 500 USA	*IBOVESPA BRA*	*IPC MEX*	*MERVAL ARG*	*IGPA CHI*	
S&P 500 – USA	**1**	0.469918	0.575103	0.487571	0.229126	Exchange Rate Regime – Brazil – Jan/99 to Mai/99
IBOVESPA – BRA	-0.478585	**1**	0.414804	0.777306	0.259236	
IPC – MEX	-0.364736	0.021291	**1**	0.412716	0.257158	
MERVAL – ARG	-0.437293	0.816172	0.207634	**1**	0.169838	
IGPA – CHI	-0.513253	0.805349	-0.141970	0.567627	**1**	
	Mexican Crisis– 12/19/94 to 12/31/94					

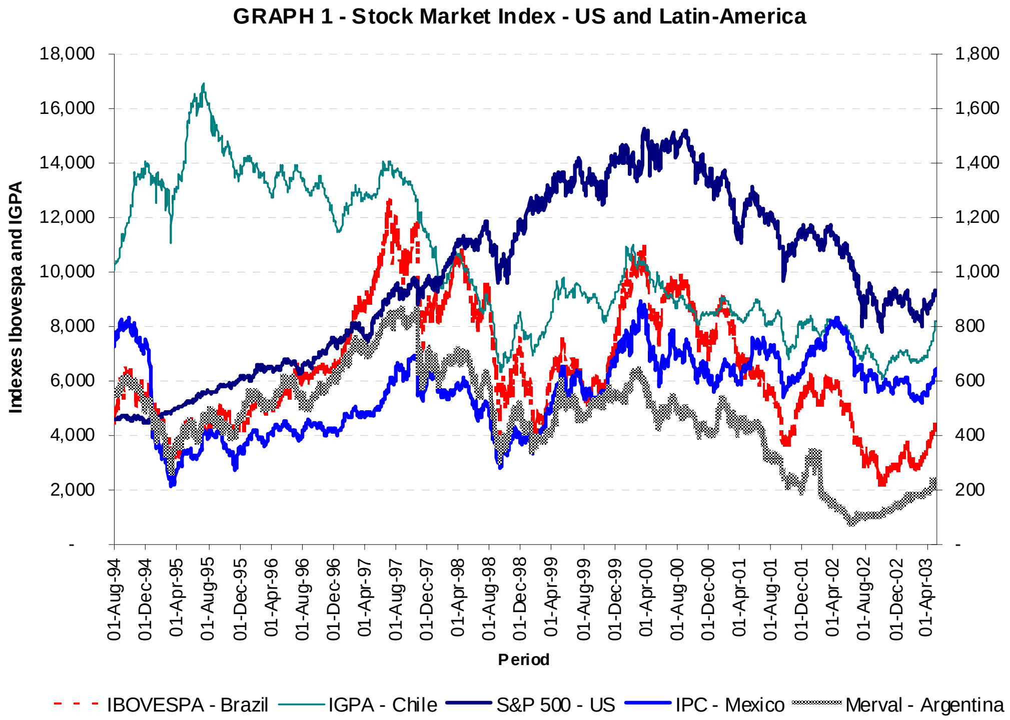


TABLE E – Unit Root Tests

Series in Levels			Series in Differences (Returns)		
	ADF	PP		ADF	PP
S&P 500 (C)	-1.6978 (3)	-1.6832	SP (C/T)	-29.262* (2)	-48.434*
IBOVESPA (C)	-1.8750 (10)	-1.9406	BOV (C)	-13.284* (10)	-42.933*
IPC (C/T)	-3.2986 (2)	-3.2263	IPC (C)	-15.059* (7)	-46.114*
MERVAL (C)	-1.3569 (10)	-1.2210	MER	-34.027* (1)	-44.025*
IGPA (C/T)	-3.0156 (1)	-3.1364	IGPA	-18.134* (4)	-36.937*

Note: (*) Indicates rejection of the null hypothesis of a unit root at 5%. The numbers among brackets show the lag used in the test, chosen by the AIC. (C/T) Indicate that the test used as deterministic terms a Constant and a Trend.

Table F reports some likelihood ratio tests (LR), comparing models with different mean processes. It also compares the VAR structure in the mean, testing against a simple AR(1) formulation, trying to verify if the VAR significance is a strict consequence of autocorrelation. The rejection of the AR(1) formulation implies the existence of an information channel through the mean of the returns, a characteristic of spillover effects. Table G compares the variance structure, also testing the GJR formulation against the traditional GARCH formulation[7].

TABLE F – LR Tests – Nested Models – Mean Structure

Null Hypothesis	Alternative Hypothesis	Test	Degree of Freedom	Choice
C-G(1,1)D(1,1)	AR(1)-G(1,1)D(1,1)	204.716*	5	AR(1)-G(1,1)D(1,1)
AR(1)-G(1,1)D(1,1)	V(1)-G(1,1)D(1,1)	94.024*	20	V(1)-G(1,1)D(1,1)
C-G(1,1)D(1,1)	V(1)-G(1,1)D(1,1)	298.739*	25	V(1)-G(1,1)D(1,1)
C-G(1,1)D(2,2)	AR(1)-G(1,1)D(2,2)	206.930*	5	AR(1)-G(1,1)D(2,2)
AR(1)-G(1,1)D(2,2)	V(1)-G(1,1)D(2,2)	97.221*	20	V(1)-G(1,1)D(2,2)
C-G(1,1)D(2,2)	V(1)-G(1,1)D(2,2)	304.151*	25	V(1)-G(1,1)D(2,2)
C-G(2,1)D(1,1)	V(1)-G(2,1)D(1,1)	291.936*	25	V(1)-G(2,1)D(1,1)
C-GJR-G(1,1)D(1,1)	V(1)-GJR-G(1,1)D(1,1)	301.642*	25	V(1)-GJR-G(1,1)D(1,1)
C-GJR-G(2,1)D(2,1)	V(1)-GJR-G(2,1)D(2,1)	294.612*	25	V(1)-GJR-G(2,1)D(2,1)
C-GJR-G(2,1)D(1,2)	V(1)-GJR-G(2,1)D(1,2)	297.673*	25	V(1)-GJR-G(2,1)D(1,2)

Note: (*) Indicate rejection of the null hypothesis at 5%.

As in the case of the mean process, a large number of parameters are accepted when the variance equations are compared. However, structures that already have a large number of parameters are not always inferior to those with the maximum number allowed. Table G shows that the GJR term always provides a better performance than the simple GARCH structure, an evidence of the presence of asymmetries in the markets responses caused by surprises.

[7] Additional tests are available with the author upon request.

TABLE G – LR Tests – Nested Models – Variance Structure – GARCH e GJR

Null Hypothesis	Alternative Hypothesis	Test	Degree of Freedom	Choice
		Variance Structure		
C-G(1,1)D(1,1)	C-G(2,1)D(1,1)	12.632*	5	C-G(2,1)D(1,1)
C-G(1,1)D(1,2)	C-G(2,1)D(1,2)	14.276*	5	C-G(2,1)D(1,2)
V(1)-G(1,1)D(1,1)	V(1)-G(2,1)D(1,1)	6.044	5	V(1)-G(1,1)D(1,1)
V(1)-G(1,1)D(2,1)	V(1)-G(2,1)D(2,1)	6.089	5	V(1)-G(1,1)D(2,1)
C-GJR-G(1,1)D(1,1)	C-GJR-G(2,1)D(1,1)	12.632*	5	C-GJR-G(2,1)D(1,1)
C-GJR-G(1,1)D(2,2)	C-GJR-G(2,1)D(2,2)	16.807*	5	C-GJR-G(2,1)D(2,2)
V(1)-GJR-G(1,1)D(2,1)	V(1)-GJR-G(2,1)D(2,1)	4.733	5	V(1)-GJR-G(1,1)D(2,1)
V(1)-GJR-G(1,1)D(1,2)	V(1)-GJR-G(2,1)D(1,2)	4.986	5	V(1)-GJR-G(1,1)D(1,2)
		GARCH × GJR		
C-G(1,1)D(1,1)	C-GJR-G(1,1)D(1,1)	193.778*	5	C-GJR-G(1,1)D(1,1)
C-G(2,1)D(2,1)	C-GJR-G(2,1)D(2,1)	219.812*	5	C-GJR-G(2,1)D(2,1)
V(1)-G(1,1)D(1,1)	V(1)-GJR-G(1,1)D(1,1)	196.681*	5	V(1)-GJR-G(1,1)D(1,1)
V(1)-G(1,1)D(2,2)	V(1)-GJR-G(1,1)D(2,2)	193.527*	5	V(1)-GJR-G(1,1)D(2,2)
V(1)-G(2,1)D(1,1)	V(1)-GJR-G(2,1)D(1,1)	195.092*	5	V(1)-GJR-G(2,1)D(1,1)
V(1)-G(2,1)D(2,1)	V(1)-GJR-G(2,1)D(2,1)	195.145*	5	V(1)-GJR-G(2,1)D(2,1)
		Conditional Correlation Structure		
C-G(1,1)D(1,1)	C-G(1,1)D(1,2)	5.668*	1	C-G(1,1)D(1,2)
V(1)-G(1,1)D(1,1)	V(1)-G(1,1)D(2,2)	11.141*	2	V(1)-G(1,1)D(2,2)
V(1)-G(2,1)D(1,1)	V(1)-G(2,1)D(2,1)	5.399*	1	V(1)-G(2,1)D(2,1)
C-GJR-G(1,1)D(1,1)	C-GJR-G(1,1)D(2,2)	3.251	2	C-GJR-G(1,1)D(1,1)
V(1)-GJR-G(1,1)D(2,1)	V(1)-GJR-G(1,1)D(2,2)	2.813	1	V(1)-GJR-G(1,1)D(2,1)
V(1)-GJR-G(1,1)D(1,1)	V(1)-GJR-G(1,1)D(2,2)	7.987*	2	V(1)-GJR-G(1,1)D(2,2)

Note: (*) Indicate rejection of the null hypothesis at 5%.

Therefore, tables F and G present three important results on model selection and the market's information transmission mechanism:

- Models with a VAR structure perform better than those with only a constant in the mean equation or with a simple autoregressive structure, as well as a variance structures with small number of parameters.

- Models with the GJR formulation in the variance equation perform better than those with a simple GARCH structure.

- The VAR structure confirms the existence of a transmission mechanism among markets that may affect their returns. However, it is not possible, yet, to make inferences about the relationship among volatilities, since it is necessary additional information that can be brought by the GJR structure and through the observation of the estimated volatilities.

Since the LR tests do not offer a unique solution to the conditional correlation structure selection problem, table H reports the forecasting evaluation statistics[8] applied to the estimated conditional volatilities. The tests were performed in one, five and ten-steps-ahead forecasts, trying to identify the best model among non-nested formulations, i.e. those that cannot be directly compared by a LR test. As an example, the model with two autoregressive terms in the correlation equation (V1-GJR-G11-D21) is not a restricted case of the formulation with two shock terms in the same equation (V1-GJR-G11-D12). In this sense, a LR test cannot be implemented to solve the model selection problem.

TABLE H –Model Selection – Out-of-Sample Statistics

Models	Mean Squared Root Error			Mean Absolute Percentage Error		
	1 Day	5 Days	10 Days	1 Day	5 Days	10 Days
V1-GJR-G11-D11	0.00059597	0.00064567	0.00118115	732.843%	1967.775%	2163.287%
V1-GJR-G11-D12	0.00059682	0.00064714	0.00118330	736.885%	1973.490%	2171.871%
V1-GJR-G11-D21	0.00059621	0.00064613	0.00118228	734.307%	1968.129%	2165.833%
V1-GJR-G11-D22	0.00071439	0.00101382	0.00198188	1072.170%	2660.042%	4389.019%
Choice:	V1-GJR-G11-D11	V1-GJR-G11-D11	V1-GJR-G11-D11	V1-GJR-G11-D11	V1-GJR-G11-D11	V1-GJR-G11-D11

The results show that the large number of parameters is not an important feature to describe data, since both statistics selected the same model on every horizon. Based on these outome, the V1-GJR-G11-D11 model is chosen as the best selection. Table I presents the results, emphasizing the estimated conditional correlation equation. The parameters' standard errors are estimated with heteroskedaticity correction. Initial values were picked up from univariate GJR models. The estimated structure follows the process:

[8] Values on table H report the simple mean of the statistic for each of the five markets. The use of weighted averages does not change results qualitatively.

$$R_{i,t} = C_i + \alpha_i R_{i,t-1} + \mu_{i,t}$$
$$\mu_{i,t} \sim N(0, \sigma^2)$$
$$\sigma^2_{t,i} = k_{i,t} + \beta_{1i}\sigma^2_{t-1,i} + \beta_{2i}\sigma^2_{t-2,i} + \beta_{3i}\mu^2_{t-1,i} + \beta_{4i}\mu^2_{t-2,i} + \beta_{5i}\chi_i\mu^2_{t-1,i}$$
$$where: \chi_i = \begin{cases} 1, to: \mu_{i,t-1} < 0 \\ 0, to: \mu_{i,t-1} \geq 0 \end{cases}$$
$$\sigma^{i,j}_t = \rho_{i,j}\sqrt{\sigma^2_{i,t}\sigma^2_{j,t}}$$
$$q_{i,j} = \varpi + Aq_{I,J,t-1} + B\varepsilon_{i,j;t-1}\varepsilon^{/}_{i,j;t-1} + Cq_{I,J,t-2} + D\varepsilon_{i,j;t-2}\varepsilon^{/}_{i,j;t-2}$$
$$where: \varepsilon_{i,t} = \mu_{i,t} / \sigma_{i,t}$$
$$to: i, j = SP, BOV, MER, IGPA, IPC$$
$$and: \varpi = 1 - A - B - C - D$$

Results confirm some stylized facts about stochastic processes applied to the conditional variance. Every market presents high internal variance persistence, measured by the sum of the variance equation's coefficients. However, LR tests, not only applied for each equation but also for the whole system, reject the hypothesis that the sum of the coefficients equals to one[9]. The same test applied to the conditional correlation estimator also rejects that hypothesis, despite its high persistence. The efficiency hypothesis of the Latin-American stock markets cannot be accepted, since the estimated autoregressive parameters are statistically significant. This result is not confirmed in US market.

[9] LR statistics for the whole system is 319,51, for a chi-square distribution with 5 degrees of freedom.

TABLE I – Multivariate GARCH-GJR Model – V1-GJR-G(2,2)-D(2,2)

S&P 500		IBOVESPA		MERVAL		IGPA		IPC	
C	0.0004271*	C	0.0002725	C	0.0000474	C	-0.0001630	C	0.0004186
	(0.0001957)		(0.0004474)		(0.0004091)		(0.0001586)		(0.0003372)
$S\&P_{t-1}$	0.0209000	$S\&P_{t-1}$	0.0605000	$S\&P_{t-1}$	-0.0384000	$S\&P_{t-1}$	0.0339000*	$S\&P_{t-1}$	0.0524000
	(0.0243000)		(0.0455000)		(0.0375000)		(0.0152000)		(0.0333000)
$IBOV_{t-1}$	0.0017698	$IBOV_{t-1}$	0.0693000*	$IBOV_{t-1}$	-0.0254000	$IBOV_{t-1}$	0.0092124	$IBOV_{t-1}$	0.0076934
	(0.0088331)		(0.0230000)		(0.0178000)		(0.0070964)		(0.0149000)
MER_{t-1}	0.0038730	MER_{t-1}	-0.0134000	MER_{t-1}	0.0700000*	MER_{t-1}	0.0058312	MER_{t-1}	0.0178000
	(0.0097973)		(0.0214000)		(0.0240000)		(0.0063019)		(0.0131000)
$IGPA_{t-1}$	0.0165000	$IGPA_{t-1}$	0.0830000	$IGPA_{t-1}$	0.0365000	$IGPA_{t-1}$	0.2207000*	$IGPA_{t-1}$	0.0648000
	(0.0237000)		(0.0577000)		(0.0491000)		(0.0214000)		(0.0385000)
IPC_{t-1}	0.0048168	IPC_{t-1}	0.1216000*	IPC_{t-1}	0.1466000*	IPC_{t-1}	0.0348000*	IPC_{t-1}	0.1096000*
	(0.0110000)		(0.0274000)		(0.0233000)		(0.0088892)		(0.0238000)
K_1	0.0000017*	k_2	0.0000133*	k_3	0.0000060*	k_4	0.0000009*	k_5	0.0000113*
	(0.0000002)		(0.0000020)		(0.0000007)		(0.0000002)		(0.0000017)
$\sigma^2_{t-1,1}$	0.9310000*	$\sigma^2_{t-1.2}$	0.9039000*	$\sigma^2_{t-1.3}$	0.9145000*	$\sigma^2_{t-1.4}$	0.9175000*	$\sigma^2_{t-1.5}$	0.8628000*
	(0.0075325)		(0.0086655)		(0.0046645)		(0.0077073)		(0.0107000)
$\mu^2_{t-1,1}$	-0.0124000	$\mu^2_{t-1.2}$	0.0229000*	$\mu^2_{t-1.3}$	0.0562000*	$\mu^2_{t-1.4}$	0.0625000*	$\mu^2_{t-1.5}$	0.0405000*
	(0.0076073)		(0.0095044)		(0.0065106)		(0.0092560)		(0.0106000)
χ_1	0.1365000*	χ_2	0.1046000*	χ_3	0.0461000*	χ_4	0.0245000*	χ_5	0.1405000*
	(0.0123000)		(0.0134000)		(0.0066697)		(0.0096635)		(0.0138000)

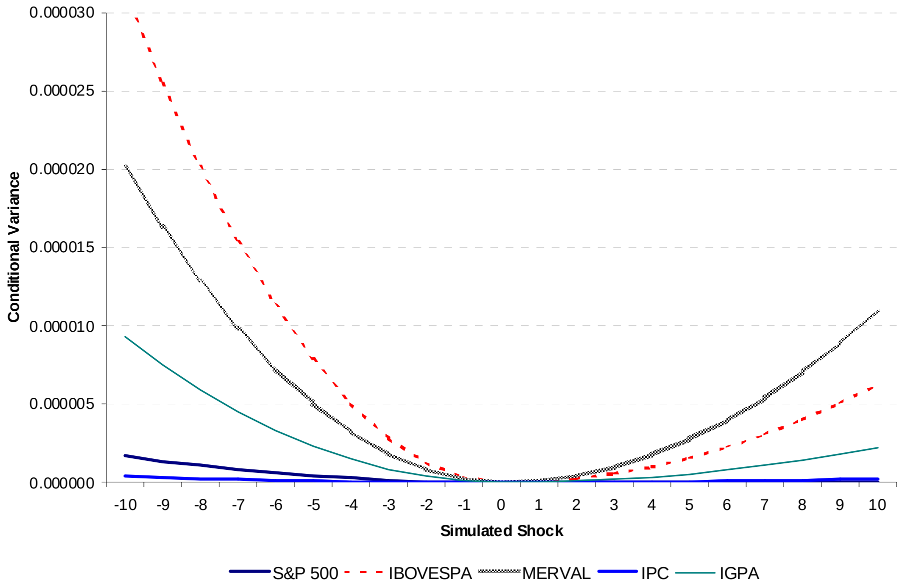
Conditional Correlation Structure	
$q_{i,j;t-1}$	0.9583000
	(0.0037571)
$\varepsilon_{i,j;t-1}\varepsilon'_{i,j;t-1}$	0.0262000
	(0.0022223)

Note: (*) Indicate significance at 5%. Standard Deviations are in the brackets.

The model allows identifying some common channels of information transmission over the returns. While none of the Latin-American stock markets affect the US market, the US stock market only affects through mean returns the Chilean market. The performance of the Mexican market has influence over returns of other Latin markets. The last result is the opposite as those in EDWARDS (1998), who did not find any effects on the interest rates' market between Mexico and Chile. However, it is possible that, for other asset markets, there are different relations of dependence among countries.

Another interesting feature is the market asymmetric response when faced by different types of shocks. All markets report higher volatility levels in the presence of bad news, when compared to positive deviations from the mean. The LR statistics accepts the joint significance of the asymmetry terms (150.43 in a chi-squared distribution with one degree of freedom). Graph 2 presents the estimated reaction curves for each market. The deviations in relation to the median of the conditional volatility are estimated in the vertical axis, while the horizontal one presents the shocks. The result may be considered a long-run effect, since the dynamics between periods is not considered[10].

GRAPH 2 - News Impact Curves



An interesting result is that shock's effects over the Mexican stock market are equivalent to the S&P 500's response, despite a large differential in the average level of conditional variance. It also draws attention the higher asymmetry value estimated to the American market in comparison with that of the Mexican IPC. The Chilean stock market's reaction is less disproportional when compared to the volatilities estimated in the IBOVESPA and MERVAL indexes. A higher asymmetry in the Brazilian asset's response can be verified by a larger reaction estimated to the Brazilian market, in face of negative surprises, than the one of the Argentinean stock market.

4. Testing for Contagion Using Conditional Correlation

The objective of this section is to test for the existence of contagion among markets, according to those concepts postulated in FORBES and RIGOBON (2000, p.13), presented in the previous section. In our framework, as long as the multivariate GARCH model is correctly specified, the estimated markets' conditional correlations provide the measure of the

[10] The concave form of the curve is a consequence of a standardized residual simulation (namely, $\varepsilon = \mu / \sigma$) applied over the conditional variance. To eliminate dynamics, the long-run response was calculated just like the conditional mean under AR structures, taking the form:

$$\sigma^2 = \frac{(\beta_3 + \beta_4)}{(1 - \beta_1 - \beta_2)} \varepsilon$$

relationship necessary for testing our procedures. Table J compares the sample correlation between the returns and the median of the conditional variance estimated by the model, presented in the upper diagonal block. One can notice that the results do not differ much when comparing the whole sample.

TABLE J – Returns' Median Conditional Correlation – August, 94 to June, 03

	S&P 500 *USA*	*IBOVESPA* *BRA*	*IPC* *MEX*	*MERVAL* *ARG*	*IGPA* *CHI*
S&P 500 - USA	**1**	0.417134	0.464627	0.331940	0.330201
IBOVESPA - BRA	0.413124	**1**	0.385868	0.492353	0.440519
IPC – MEX	0.442572	0.388471	**1**	0.343400	0.330427
MERVAL – ARG	0.294935	0.477175	0.341101	**1**	0.285871
IGPA – CHI	0.312176	0.482160	0.304869	0.338996	**1**

Five events have their conditional median correlation compared to that estimated by the model. Three of them are considered to be endogenous, since that the stock markets' returns reflect the shocks originated in countries included in the equation system: the depreciation of the Mexican peso (12/19/94 to 12/31/94); the crisis with the Argentinean peso (12/01/01 to 12/31/02); and the change in the Brazilian exchange rate regime (01/15/99 to 03/31/99). The other two are considered to be exogenous once they occurred in countries that are not included in the system, such as the Asian countries' crisis (10/17/97 to 10/31/97) and the Russian one (08/01/98 to 12/31/98)[11].

One can point out that this test compares a specific period of crisis to the whole correlation series estimated by the model. This objection, based on FORBES and RIGOBON (2002), states that structural changes may affect the estimation of the mean correlation, inserting a bias in the parameter. However, the solution proposed there –to compare the crisis period with the month that immediately precedes its start – is not necessarily the most adequate, since it is also arbitrary establishing a normality period. BAIG and GOLDFAJN (2000) illustrate this point by the strong rise in the Russian interest rate at the end of 1997 as a mean to avoid the deterioration of the country's international reserves. In a sense, this policy action signalized that there were severe problems with the Russian economy, marking the period, if not as a sharp crisis, as a period of high volatility. Hence, it is not adequate to establish September 1998 as a normality month that preceded the crisis.

[11] The choice of the periods was as follows: Mexican and Asian crises – FORBES and RIBOBON (2002), p. 2238; Russian crisis – BAIG and GOLDFAJN (2000), p. 5, with the restructuring of rublo-denominated debt and the change in the exchange rate band; Argentinean collapse – CHUI, HALL and TAYLOR (2004), p. 34, with the introduction of capital controls over the external flows. In the Brazilian crisis, the choice of period reflects the sharp increase in exchange rates returns after the change in regime in January 15th.

Table K presents the median correlations estimated for the high volatility's periods. The lower diagonal reports the median of the correlation estimated for the whole sample. For those events defined as endogenous, bold values highlight the correlations of the countries in a crisis.

TABLE K – Returns' Median Conditional Correlation – Crises Periods

	S&P 500 USA	*IBOVESPA BRA*	*IPC MEX*	*MERVAL ARG*	*IGPA CHI*	
S&P 500 – USA	**1**	0.2134	**0.0657**	0.2173	0.2247	Mexican Crisis (12/19/1994 to
IBOVESPA – BRA	0.4171	**1**	**0.1841**	0.5405	0.3507	
IPC – MEX	**0.4643**	**0.3855**	**1**	**0.2229**	**0.0876**	
MERVAL – ARG	0.3310	0.4928	**0.3429**	**1**	0.3514	
IGPA – CHI	0.3309	0.4411	**0.2855**	0.3303	**1**	
S&P 500 – USA	**1**	**0.5091**	0.5231	0.2861	0.3908	Exchange Rate Regime – Brazil – (01/15/1999 to
IBOVESPA – BRA	**0.4171**	**1**	**0.3859**	**0.3721**	**0.4934**	
IPC – MEX	0.4643	**0.3855**	**1**	0.2946	0.3020	
MERVAL – ARG	0.3310	**0.4928**	0.3429	**1**	0.2279	
IGPA – CHI	0.3309	**0.4411**	0.2855	0.3303	**1**	
S&P 500 – USA	**1**	0.4078	0.4972	**0.1506**	0.3658	Argentinean Crisis (12/01/2001 to
IBOVESPA – BRA	0.4171	**1**	0.3887	**0.2332**	0.4888	
IPC – MEX	0.4643	0.3855	**1**	**0.2075**	0.2755	
MERVAL – ARG	**0.3310**	**0.4928**	**0.3429**	**1**	**0.1619**	
IGPA – CHI	0.3309	0.4411	0.2855	**0.3303**	**1**	
S&P 500 – USA	**1**	0.5414	0.5261	0.4701	0.3597	Asian Crisis (10/17/1997 to
IBOVESPA – BRA	0.4171	**1**	0.5105	0.5957	0.4472	
IPC – MEX	0.4643	0.3855	**1**	0.4656	0.3297	
MERVAL – ARG	0.3310	0.4928	0.3429	**1**	0.3961	
IGPA – CHI	0.3309	0.4411	0.2855	0.3303	**1**	
S&P 500 – USA	**1**	0.5518	0.5529	0.5479	0.4170	Russian Crisis (01/08/1998 a
IBOVESPA – BRA	0.4171	**1**	0.6267	0.7067	0.5796	
IPC – MEX	0.4643	0.3855	**1**	0.5870	0.5147	
MERVAL – ARG	0.3310	0.4928	0.3429	**1**	0.5686	
IGPA – CHI	0.3309	0.4411	0.2855	0.3303	**1**	

There are some patterns in the exogenous events responses. For the Asian and Russian crises, each estimated correlation rises sharply, in a kind of level-shift change. This is the ultimate signal of contagion, as described in FORBES and RIGOBON (2000). The possibility of a structural break, based in FORBES and RIBOBON (2002), is ruled out by the data presented in table L, since an eventual bias in the sample – a bad selection of crisis dates – can not be sustained, as long as median correlations in normal periods are lower before 1999. In this sense, as long as these events occurred before 1999, the verified volatility increase is not a consequence of a structural break during the period.

TABLE L – Returns' Median Conditional Correlation – Partial Samples – Excludes Crises Periods

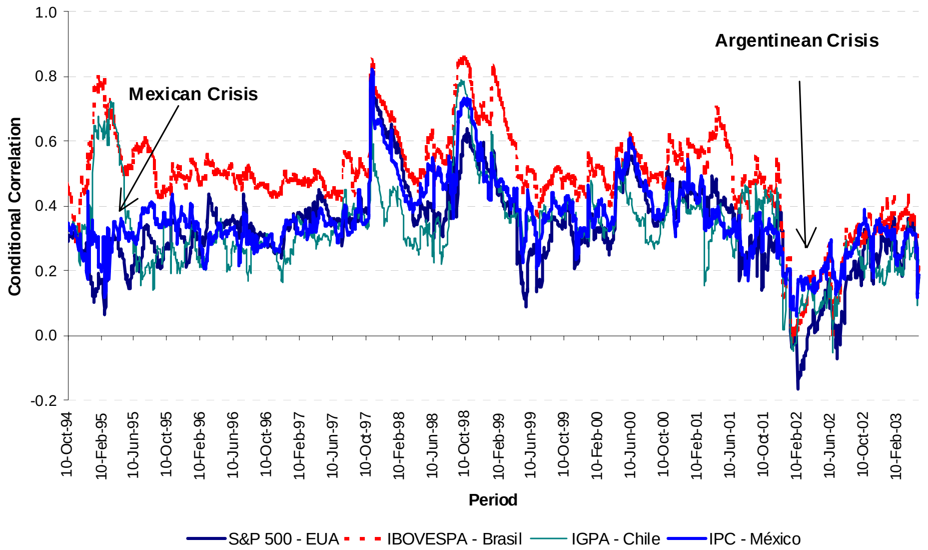
	S&P 500 USA	*IBOVESPA BRA*	*IPC MEX*	*MERVAL ARG*	*IGPA CHI*	
S&P 500 – USA	**1**	0.3847	0.3969	0.3436	0.2942	1994-1998
IBOVESPA – BRA	0.4592	**1**	0.3559	0.5044	0.3975	
IPC – MEX	0.5108	0.4210	**1**	0.3441	0.2672	
MERVAL – ARG	0.3432	0.4932	0.3650	**1**	0.3233	
IGPA – CHI	0.3630	0.4578	0.3167	0.3625	**1**	
	1999-2003					

The endogenous events show some peculiar features. The change in the Brazilian exchange rate regime, in 1999, was an atypical event of information transmission in the stock markets, since only in this episode the estimated conditional correlation does not point to a unique direction. During stress events, the Chilean and the Mexican stock markets increase their correlation with the US market. Coincidently, these are the markets with a lower asymmetric response under negative news. The correlations of IBOVESPA and MERVAL do not rise under the same circumstances.

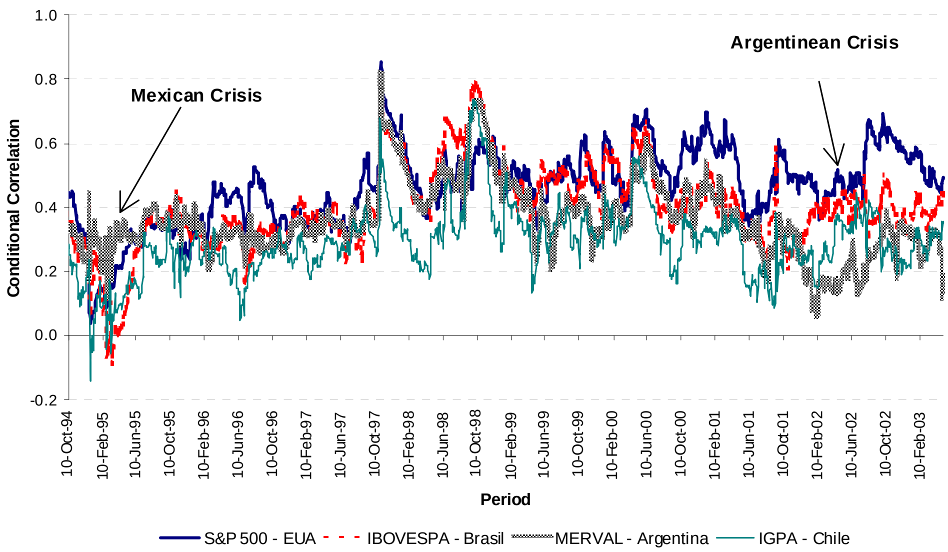
5. Contagion or Interdependence: Sufficient Concepts? Analyzing Mexico and Argentina

The results in the analysis of endogenous events raise a few questions about the sufficiency of the concepts proposed in FORBES and RIGOBON (2000) concerning contagion and interdependence. The decline of the conditional correlation among markets was a common feature of the Argentinean and Mexican crises. Therefore, it is reasonable to believe that their effects over other markets were insignificant. Graphs 3 and 4 present the estimated correlations between Mexican and Argentinean stock markets with the other markets.

GRAPH 3 - Estimated Conditional Correlation - Argentina



GRAPH 4 - Estimated Conditional Correlation - Mexico



The concepts presented in FORBES and RIGOBON (2000) are based on the variation of some measure of market's relation: if the common shocks in markets do not alter this measure of relationship, then, it is a matter of interdependence; if the measure rises in stress events, then, it is a matter of contagion. However, in the Argentinean and Mexican cases, the effects of the stock markets' returns over the others seem to have been small, despite the crises' intensity. All markets had a decrease in the estimated correlation, which allows us to consider the hypothesis that, under certain circumstances, the market experiencing the crisis is isolated.

Table M reports the results of simple dynamic models where the endogenous variables are the deviations from the mean correlation among Latin American markets and the US[12]. Dummies variables, whose values equal the unity in crisis periods, were introduced in order to check the existence of a significant deviation of the conditional correlation from the sample mean. Coherently with table K's results, the crisis periods' dummies are significant only in contagion (Asian and Russian Crisis, with positive sign) and isolation cases (Mexican and Argentinean Crisis, with negative sign). The coefficient associated with the Brazilian event was never significant. An alternative set of regressions includes the nominal exchange rate variation of the local currency against US dollars, in order to check if nominal variations in the asset prices can be important to explain these deviations. Only in the IGPA and MERVAL equations these movements are significant at 5%. The negative sign implies that local currency devaluations increase the isolation's probability.

TABLE M – Regression Results – Deviation from Sample Correlations – Latin America and US Markets

	IBOVESPA BRA		*IPC MEX*		*MERVAL ARG*		*IGPA CHI*	
	Without ER	With ER	Without ER	With ER	Without ER	With ER	Without ER	With ER
C	-0.001553	0.002874	-0.003182	-0.003657	0.003108	-0.000311	-0.013642	0.002958
	(0.010579)	(0.010363)	(0.012628)	(0.013046)	(0.015256)	(0.012664)	(0.015712)	(0.016409)
AR(1)	0.501967**	0.539139**	0.616022**	0.614423**	0.387877**	0.530193**	0.484771**	0.602673**
	(0.114714)	(0.111236)	(0.102646)	(0.104774)	(0.138628)	(0.121177)	(0.151809)	(0.150977)
Mexican Crisis	-0.118575**	-0.116079**	-0.175367	-0.191063**	-0.085859	-0.076271	-0.063456	-0.084638
	(0.038478)	(0.036756)	(0.045752)	(0.086678)	(0.054847)	(0.045483)	(0.054660)	(0.051837)
Brazilian Exchange Rate Regime	0.005905	0.084967	-0.043976	-0.041981	0.024896	-0.005911	-0.035040	-0.048482
	(0.055460)	(0.067281)	(0.064066)	(0.065887)	(0.081137)	(0.067698)	(0.078479)	(0.073428)
Argentinean Crisis	-0.000947	0.022381	0.042227	0.041559	-0.141339*	-0.032682	0.044964	0.043289
	(0.028388)	(0.029741)	(0.033605)	(0.034356)	(0.051626)	(0.052040)	(0.040510)	(0.037784)
Asian Crisis	0.227883**	0.227741**	0.197434	0.196576**	0.258942**	0.257889**	0.128680	0.150997**
	(0.052869)	(0.050471)	(0.062881)	(0.064146)	(0.075131)	(0.062201)	(0.076551)	(0.072073)
Russian Crisis	0.096549**	0.095078**	0.069653	0.068341	0.136273**	0.130259**	0.087388	0.072405
	(0.038391)	(0.036658)	(0.045335)	(0.046560)	(0.054557)	(0.045196)	(0.055701)	(0.052369)
Exchange Rate Variation	-	-0.209842*	-	0.038012	-	-0.140795**	-	-0.746552**
		(0.110189)		(0.177032)		(0.038471)		(0.332251)

Note: (**) Indicate significance at 5%. (*) Indicate significance at 10%. Standard Deviations are in the brackets.

[12] The choice of US market as a benchmark in this analysis was mainly due to the GARCH results, where the US market characteristics were closer to the definition of an "efficient" market. However, similar results can be achieved comparing different sets of correlations. Data frequency was changed to quarterly basis, as long as other tests were performed with different macroeconomic variables.

There is an expressive part of the literature about crises in Latin America justifying the events based on the economies' fundamentals. In this sense, BAIG and GOLDFAJN (2000, p. 12) draws attention to the change in the composition of international capital flows to Brazil, where foreign direct investments (FDI) turned to be a considerable part of the Capital and Financial account. Indeed, as table N shows – by comparing the FDI inflows and the outcome of the Financial and Capital account for the three countries – Brazil did not incur on deficits in those accounts. Argentina and Mexico, on the other hand, had some problems with their external position, largely due to the decrease in the long-run capital inflows, such as FDI. The Argentinean case had a dramatic aspect, since the capital inflows were in a sharp decrease since the Brazilian crisis.

CHUI, HALL and TAYLOR (2004) also highlight the lower intensity in the answers of a group of countries in face of the Argentinean crisis when compared to the Asian one. The authors point to the improvement in these countries' fundamentals, for instance, the adoption of a flexible exchange rate system in Brazil in 1999. Hence, not only the capital account composition, emphasized in BAIG and GOLDFAJN (2000), but also the economic policy mix adopted by those countries, under trade and financial aspects, may have contributed to the smaller effects of those crises throughout the time.

TABLE N – Capital Flows to Brazil, Mexico and Argentina – 1993-2001 – In US$ Billions

		1993	1994	1995	1996	1997	1998	1999	2000	2001
Foreign Direct Investment (a)	Argentina	2,793	3,635	5,609	6,949	9,160	7,291	23,988	10,418	2,166
	Brasil	1,292	3,072	4,859	11,200	19,650	31,913	28,576	32,779	22,457
	Mexico	4,389	10,973	9,526	9,186	12,831	12,285	13,166	16,449	26,569
Capital and Financial Account (b)	Argentina	20,344	11,377	5,003	11,764	16,850	19,009	14,494	7,824	-13,268
	Brasil	7,685	8,193	29,658	33,922	25,400	20,438	8,395	29,649	20,295
	Mexico	33,760	15,787	-10,487	4,248	25,745	19,747	17,563	22,611	25,403
(a) / (b)	Argentina	0.137	0.319	1.121	0.591	0.544	0.384	1.655	1.332	-0.163
	Brasil	0.168	0.375	0.164	0.330	0.774	1.561	3.404	1.106	1.107
	Mexico	0.130	0.695	-0.908	2.162	0.498	0.622	0.750	0.727	1.046

Source: International Finance Statistics – International Monetary Fund.

Anyway, it should be highlighted that, as in BAIG and GOLDFAJN (2000), higher contagion effects were found in the markets of external debt bonds. This result is consistent with EDWARDS (1998), who finds higher connections among Latin-American markets by comparing the yield of the debt bonds. In other words, in spite of the evidences about the effects of countries' fundamentals on crises propagation, one should consider the specific information transmission among different markets assets.

6. Conclusions:

The results derived from the estimation of GARCH models allow us to draw some conclusions about the information transmission mechanism among the Latin America's stock markets. The main conclusions are:

a) All the evidences point to the occurrence of contagion, as defined in FORBES and RIGOBON (2000), over Latin America only in the Russian and Asian crises. The Mexican and Argentinean crises had a specific propagation dynamics that is not, necessarily, an event of contagion, while the effects of Brazilian exchange rate developments over Latin America should be seen as a phenomenon of interdependence among the markets;

b) The small direct influence of the average US market returns over the Latin-America's stock markets. Although this is a counter-intuitive fact, the correlation between the returns is increasing over time, as table L shows. This is a signal of a change in the structure of the relation among markets, considering the financial opening of the Latin economies. The increasing correlation between Latin and US market in crises periods, implying that the US market works as reference in times of high volatility, also draws attention.

c) The different pattern in the markets' responses in face of international crises according to their origins. On one hand, crises that were originated outside Latin America had contagion effects by increasing the correlation between the returns in all markets. On the other hand, crises originating in Latin American countries do not present a uniform response, having the "isolation" of the troubled country (Mexico, 1994 and Argentina, 2001) or the propagation throughout close countries (Brazil, 1999) as possible outcomes.

d) There are evidences that the propagation of the so-called internal crises is related to the countries' macroeconomic fundamentals. The more solid Brazilian position, in comparison with those of the other Latin–American countries, may have been a determinant factor in the propagation of the country's volatility to the other markets. Anyhow, if this is a specific feature of stock markets or if other assets have the same pattern of behaviour remains to be analyzed.

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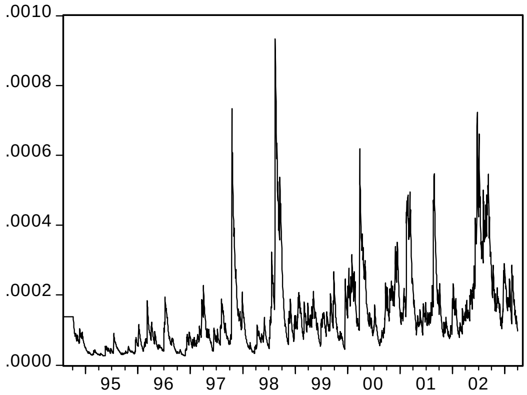
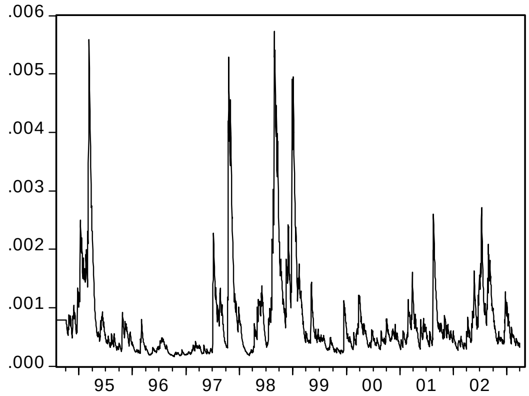
I. *Appendix:*

Estimated Conditional Correlations and Volatilities – V(1)-GJR-GARCH(1,1)-D(1,1)

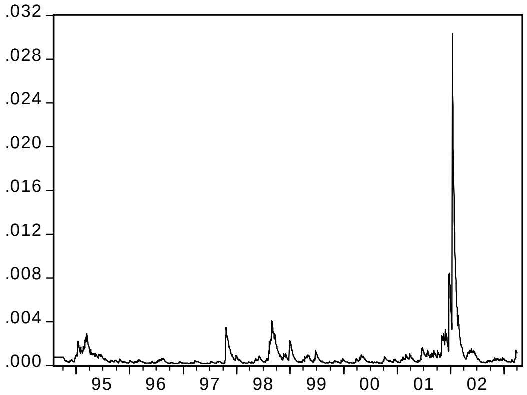
Estimated Volatility - S&P 500



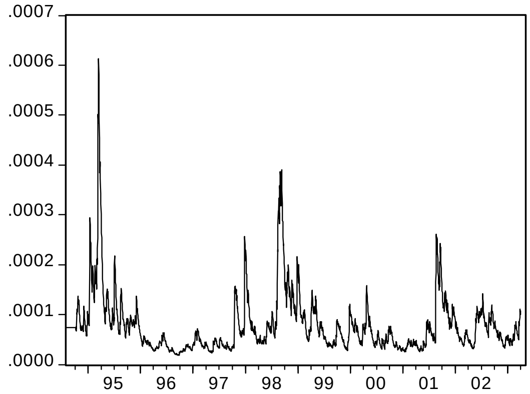
Estimated Volatility - IBOVESPA



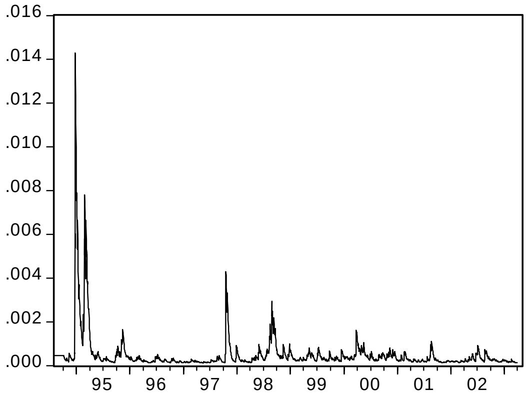
Estimated Volatility - MERVAL



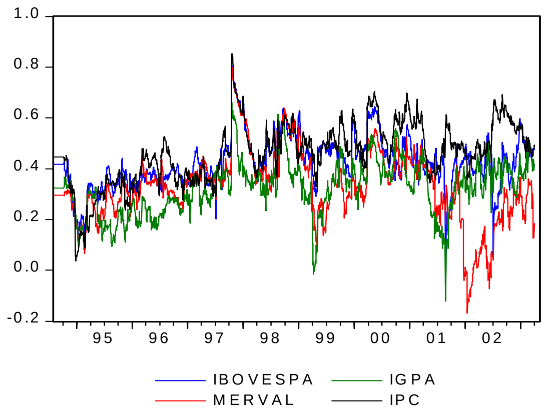
Estimated Volatility - IGPA

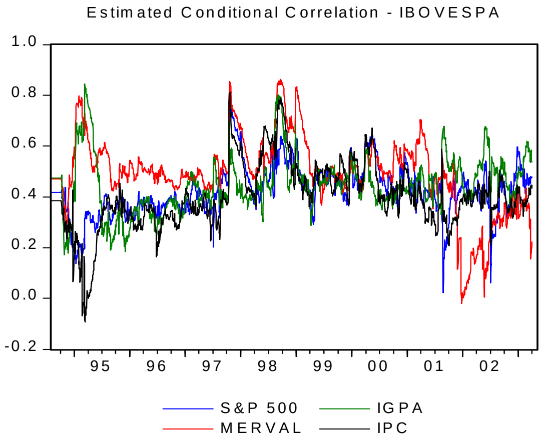


Estimated Volatility - IPC

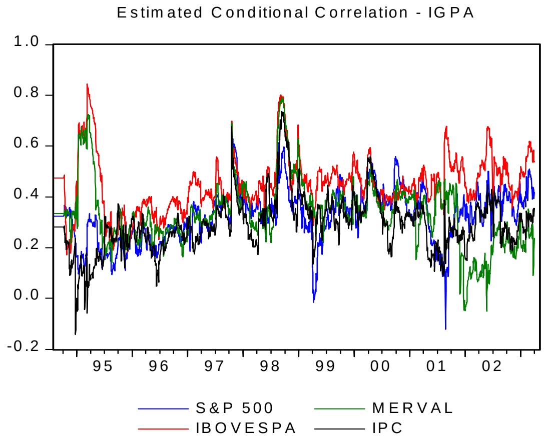


Estimated Conditional Correlation - S&P 500





Estimated Conditional Correlation - IBOVESPA
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
95 96 97 98 99 00 01 02
S&P 500
IGPA
MERVAL
IPC



Estimated Conditional Correlation - IGPA
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
95 96 97 98 99 00 01 02
S&P 500
MERVAL
IBOVESPA
IPC

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