Financial liberalization, exchange rates and stock prices: Long-run relationships and short-run dynamics in four Latin America countries

by

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Abstract

This paper provides an analysis of the long-run relationships and short-run dynamics between stock prices and exchange rates as well as the channels through which exogenous shocks influence these markets. We use monthly data for the period 1980-2005 for four Latin America, namely, Argentina, Brazil, Chile and Mexico. We conduct our analysis by means of cointegration analysis and multivariate Granger causality tests. The main finding of our analysis suggests that stock and foreign exchange markets in these economies are positively related and that the US stock market acts as a channel for these links. Moreover, it is shown that these links are independent of foreign exchange restrictions. Finally, stability tests proposed by Hansen and Johansen (1993) are applied and it is shown that the dimension of the cointegration space is sample dependent while the estimated coefficients do not exhibit instability in recursive estimations.

Keywords: Stock markets, foreign exchange market, capital market integration, financial crises, Latin America stock markets

JEL classification: F21, F31, F36

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1. Introduction

During the last fifteen years we have witnessed a substantial development in the structure of both mature and emerging financial markets. The emerging markets have been in the centre of interest of private and institutional investors as well portfolio managers. More specifically this growing interest has come from the recognition of the important positive link between the process of financial liberalization which these economies have undergone and economic development.[1] The flow of portfolio investments to emerging financial has increased from a mere \$6.2 billion in 1987 to \$37.2 billion in 1992 to a total of \$211.6 for the period 2000-2006 (BIS, 2006). The flow of these funds has been mainly directed to bonds, certificates of deposit and commercial papers although during the recent period there is a major shift towards investing on stocks.

Financial liberalization and especially abolishing capital controls directed to an emerging country's stock market is the result of the decision by the government of this country to allow foreigners to invest on equity in the respective stock market. Over the recent period several researchers have studied the possible benefits and costs of such liberalization processes. A major argument in favour of such move is that the opening of financial markets in the emerging economies by removing existing capital controls will help them to attract foreign capital to finance economic growth. Furthermore, this increased capital flows will speed up the development of stock markets which may lead to long-run economic growth. The argument for such a positive relationship is based on the prediction that the liberalization of stock markets reduces the aggregate cost of equity capital. An additional implication is that

[1]Bekaert and Harvey (1997, 2000, 2002); Bekaert *et al*. (2002a, 2002b); Bekaert *et al*. (2006); De Santis and Imrohoroglu (1997); Edwards and Susmel (2003); Edwards *et al*. (2003); Henry (2000); Huang and Yang (1999); Kaminsky and Schumkler (2003) and Kim and Singal (2000), are among the numerous studies that have studied the effects o financial liberalization on emerging economies.

following the opening of financial markets we should observe a rise in physical investment as a result of the decline cost of equity capital. Finally, there will be an urgent need for increased transparency and accountability by the firms' management and this will also lead to result to improved allocation of resources, reduction of the risk to hold stocks and again to a reduction of the cost of capital, (Henry, 2000; Kim and Singal, 2000).

The most important concern by the governments and policy practitioners of emerging economies is that such abolition of capital markets and the subsequent rise in capital flows may give rise to some unpleasant effects. These uncertainties are the result of the fact that international capital flows are very sensitive to changes in interest rates as well as expectations about future economic growth and expected returns from holding stocks. Such changes of even a small magnitude may result to negative effects in the domestic economy. Moreover, the financial liberalization will result to an exposure to foreign influence making domestic stock prices more volatile as a result of external impact from global financial markets. A final important concern deals with the prediction that capital inflows will cause an appreciation of the domestic currency and this will cause exports to fall for those countries which are export-oriented whereas in the case of not enough available investment plans to absorb the money inflow will result to rising inflation.

During the 1980s and early 1990s several Latin America and Asian countries have undergone a number of structural reforms, abolition of capital controls and global integration processes. But these processes were accompanied by financial crises such the 1994 Mexican currency crisis and the recent 1997-1998 turmoil in East Asian financial markets. A substantial number of papers which have scrutinized the operation of the emerging markets of this period have shown that the most common feature of these markets is the high volatility observed in the returns of financial securities. Given this stylized fact of the emerging markets instability, many authors have attempted to evaluate the effect of financial reform on several characteristics of emerging markets, (Bekaert and Harvey, 2000; Henry 2000; and Bekaert *et al*., 2000a,b). Furthermore, Bekaert and Harvey (1995, 1997), De Santis and Imrohoroglu (1997), Huang and Yang (1999), Aggarwal *et al*. (1999), Kaminsky and Schmukler (2003), Edwards and Susmel (2003), Bekaert *et al*. (2006) and Cunado *et al*. (2006) are examples of studies that have examined the issue of increased volatility in emerging markets following financial liberalization. The evidence from these studies is mixed implying that there is no definite conclusion that the abolishment of capital controls and the opening of the markets has contributed to increased degree of uncertainty.

This paper focuses on the effects of financial liberalization that four Latin America emerging economies, namely Argentina, Brazil, Chile and Mexico, have implemented in late 1980s and early 1990s from a rather different perspective. More specifically we do not examine the link between market openness and stock market volatility which ahs been studied extensively but we focus our attention on the question whether financial liberalization has any significant effect on the link between the stock market and foreign exchange markets for these four economies.

Alternative theories of exchange rate determination have for long documented the positive relationship between the stock market performance and the exchange rate changes. Thus, models of current account dynamics (Dornbusch and Fisher, 1980) show that movements in exchange rates have an impact in the international competitiveness and the position of the trade balance. As a result we expect a change in the output of the economy and finally a change the firms' performance and their stock prices. However, it is also possible to track the reverse links, implying that changes in stock prices may also affect exchange rates. This is accomplished through the change in demand for money caused by economic agents' wealth given that domestic stocks are part of their portfolio. This positive relationship has also been shown within the portfolio-balance models developed by Branson (1983) and Frankel (1983).[2]

The present analysis examines the link between stock and foreign exchange markets for the four Latin America countries. The analysis brings on the surface several important features of this link. First, we provide the theoretical framework within we which we should investigate the existence of such a relationship. Furthermore, we explain the two different channels through which exogenous disturbances may affect the stock and foreign exchange markets. Second, we provide a thorough analysis of the long-run relationships between the two markets by applying a complete set of statistical tests based on the Johansen (1988, 1991) multivariate cointegration analysis. Third we also examine the short run dynamics between the two markets using the multivariate Granger causality developed by Dolado and Lutkepohl (1996). These are appropriate causality tests in the presence of cointegration since the analysis is made in level and not in returns. Following Granger (1981) it is appropriate to analyze the relationship between two or more variables un both levels and first differences abstracting from the problem of nonstationary data within the context of cointegration analysis. Third, we consider the case of omitted variables. Therefore, in case where there is no evidence of cointegration between the domestic

[2] The relationship between stock and foreign exchange markets has been analyzed empirically by many authors mainly for the US case as well as other industrialized countries. For example, Aggarwal, (1981), Roll (1992) and Chow *et al*. (1997) among others have been directly examined this link. Studies that they have investigated the link between stock returns and an extensive set of macroeconomics variables have also examined the relationship between the two markets indirectly, (e.g Schwert, 1990). A relevant study for the case of Latin America economies is undertaken by Adrangi et al. (1999)

stock and foreign exchange markets we take the position that this may be the outcome that an important variable is missing which otherwise will provide the channel of transmissions from one market to the other. We consider the US stock market as a potential candidate for such a missing channel. Given the high degree of financial markets integration in the last decade changes in the US stock market (that may be taken as representing the world markets) may influence substantially domestic stock markets and at the same time making the link between the stock and foreign exchange markets stronger. Following Caporale and Pittis (1997) we adopt the appropriate testing procedures for cointegration and causality in the presence of omitted variables. Finally, given that at least one statistically significant cointegrating vector has been found we examine the stability of the long-run relationships through time, since evidence that the stock and foreign exchange markets are cointegrated, might be exploitable by the investors only if this evidence is sample independent and stable overtime. To this end, we apply the Hansen and Johansen (1993) recursive stability tests.

Within this analytical framework we attempt to answer several important issues for the four Latin America emerging markets. First, whether a long-relationship between the stock and foreign exchange markets exists and second, in the case that such a relationship exists the use of stability tests will help us to examine whether this relationship is stable over time given the financial liberalization process and additionally to evaluate the potential effect of the 1994 Mexican crisis and the 1997-1998 Asian financial crisis on this relationship. Third, we look into the causal patterns between these two markets and analyze how different causality direction may have an impact of the transmission channel of disturbances. A final question analyzes the effect of world markets influence to the link between the domestic stock and foreign exchange markets. The results of our analysis have important policy implications with respect to the appropriate monetary and exchange rate policy that these emerging economies should adopt. Moreover, these results should be of interest to investors and portfolio managers who are active in financial markets and to the multinational companies which need to consider hedging techniques to limit foreign exchange exposure.

The structure of the rest of the paper is as follows. Section 2 discusses the theoretical framework of the relationship between stock and foreign exchange markets. In section 3 we present the cointegration methodology and the associated Hansen and Johansen (1993) recursive stability test as well as the Dolado and Lutkepohl (1996) approach to multivariate Granger causality. Section 4 presents the data and discusses the empirical results and in section 5 we give our summary and concluding remarks.

2. Models of stock and foreign exchange markets relationships

For the analysis of the relationship between the stock and foreign exchange markets in each Latin America country the starting point is the following specification

$$I_t = b_0 + b_1 E_t + u_t \qquad (1)$$

where I_t is the domestic stock price index, E_t is the real exchange rate defined as foreign prices adjusted by the nominal exchange rate (defined as units of domestic currency per unit of foreign currency) relative to domestic prices and u_t is a white noise error term. The choice of the real exchange rate is based on the fact that it measures a country's competitiveness in international goods and services markets as opposed to the nominal exchange rate. The real exchange rate is defined as follows:

$$E_t = P_t / s_t P_t^* \qquad (2)$$

where s_t is the nominal exchange rate, P_t is the domestic price and P_t^* is the corresponding foreign price level.

Economic theory suggests that the sign of the coefficient b_i can either be positive or negative. There are several explanations put forward to justify the positive relationship between foreign exchange and stock markets. The first one looks on the impact that exchange rate movements have on economic activity. As Cornell (1983) and Wolff (1988) among others have shown, a fall in the real exchange rate will cause an increase in the competitiveness of the domestic economy leading an improvement of the balance of trade, an increase in domestic aggregate demand as well as the level of output. The second explanation is based on the well documented fact that economic activity as is proxied by macroeconomic variables such as inflation, money supply, industrial production, economic growth, employment rate or corporate profits affects the stock market.[3] Specifically, the stock price of a company is considered to embody all available information related to the expected future cash-flows which are influenced by the expected future performance of the firm and the expected future changes in economic activity. This long-run positive relationship between stock prices and economic activity has been empirically justified in several relevant works. The implication of such a relationship is that a decline in real exchange rate will result to an increase in stock prices and hence, $b_i < 0$. Such a transmission mechanism is

[3] See for example, Fama (1981), Smith and Sims (1993) and Flannery and Protopapadakis (2002).

consistent with the *flow* approach to exchange rate determination, (Dornbusch and Fisher, 1980)

An alternative channel link between the two markets is offered by the portfolio-balance approach to the exchange rate determination (Branson, 1983; Frankel 1983). The key element of this class of models is that economic agents allocate their wealth among domestic money as well as domestic and foreign bonds. Equilibrium exchange rate is the result of the interaction between demand and supply of assets. The model predicts that, in case of an increase in domestic stock prices, wealth and the demand for money rise and this will result to an increase in domestic interest rates. Consequently, we will observe increasing capital inflows and an appreciation of the domestic currency and a fall in the real exchange rate. We call this transmission mechanism the *stock* approach.[4]

Equation (1) has been tested with the use of data in first difference in a number of studies but the evidence is mixed since they found no clear-cut results in favour of the *flow* or the *stock* channels of links between the stock and foreign exchange markets. This outcome was attributed by some authors to the fact that such a specification formalized as in eq. (1) may actually be an incomplete system. Incomplete systems arise in the case when an important variable is missing. Under these circumstances Lutkepohl (1982) and Caporale and Pittis (1997) have shown that statistical inference about the relationship among a set of variables as well as causality tests are invalid. Additionally, Caporale and Pittis (1997) have shown that the inclusion of the omitted variable in the particular structure will make inference valid. To take account, for the significant role of the omitted variable to reveal its impact on revealing the channels of links among variables, Caporale and Pittis (1997) consider

[4] For a comprehensive analysis of alternative models of exchange rate determination see MacDonald and Hallwood (2000) and Copeland (2005). See Garvin (1989) for a complete analysis of the *flow* and *stock* approaches

the case of a first order bivariate VAR models under the assumption of no cointegration and no causality between the two variables of the system. They prove that statistical inference are influenced in case where this bivariate system appears to be subset of a trivariate system and at the same time the omitted variable either causes none of the variables of the bivariate system, or causes on the variables or both variables under consideration.[5]

In the present analysis a good candidate for a complete trivariate system is considered to be the US stock market, which can be taken to represent the world stock markets. Thus, we assume that the US stock market provides an important link between the foreign exchange market and the domestic stock market. Under the *flow* approach this implies that any change in the US stock market is taken to reflect a change in the foreign economy given our discussion about the relationship between the stock market and economic activity. Furthermore, within the standard Mundell-Fleming framework an increase in the foreign imports will result to a rise in exports of the Latin America countries under consideration give the strong trade links of the two areas. As a result agents will observe an appreciation of the corresponding domestic currency and a subsequent improvement of their terms of trade (a rise in the real exchange rate), an increase in the domestic aggregate demand and domestic output. This result will finally cause the domestic capital market to rise.[6] When we consider the *stock* approach, it is clear that increased integration of world capital markets will imply that an increase in the US stock market will result to a rise of the domestic stock market, as well and the implied increase in domestic wealth will lead to an increase in the real exchange rate as discussed above.

[5] The analysis by Caporale and Pittis (1997) implies that evidence of no cointegration between two variables may be due to the omission of an important omitted variable tat would change the causality pattern.

[6] Canova and DeNicolo (1995) have shown that the relationship between stock returns and economic activity is magnified when foreign influences are taken into consideration.

Based on these considerations the complete system can be specified as follows:

$$I_t = b_0 + b_1 E_t + b_2 I_t^* \quad (3)$$

where I_t^* denotes the US stock price index. The sign of coefficients b_1 and b_2 is expected to be positive under both the *flow* and the *stock* approaches.

3. Econometric methodology

3.1. Cointegration analysis

In order to evaluate whether a long run relationship exists between the stock market of each Latin America country and the foreign exchange market we apply the Johansen (1988, 1991) and Johansen and Juselius (1990, 1992), multivariate cointegration methodology with well behaved (Gaussian) errors that need to be estimated in the form of unrestricted VARs.

The Johansen (1988, 1991) framework involves estimating the following vector-error correction model (VECM):

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \ldots.. + \Gamma_{k-1} z_{t-k+1} + \Pi z_{t-k} + \gamma D_t + \mu + \varepsilon_t \quad (4)$$

where z_t is a column vector of stochastic variables, $\varepsilon_t \sim Nid_p(0,\Sigma)$. The parameters $(\Gamma_1, \ldots\ldots\ldots, \Gamma_{k-1}, \gamma)$ define the short-run adjustment to the changes of the process, whereas $\Pi = \alpha\beta'$ defines the short-run adjustment, α, to the cointegrating relationships, β. If the short-run effects are basically different from the long-run effects, due for instance, to costly arbitrage and/or imperfect information, the explicit specification of the short-run effects is probably crucial for a successful estimation of the steady-state relations of interest. D_t is a vector of nonstochastic variables, such as centered seasonal dummies which sum to zero over a full year by construction and are necessary to account for short-run effects which could otherwise violate the Gaussian assumption, and/or intervention dummies; μ is a drift and T is the sample size.

Johansen (1991) shows that if $Z_t \sim I(1)$, the following restrictions on model (3) have to be satisfied:

$$\Pi = \alpha\beta' \tag{5}$$

where Π has reduced rank, r, α and β are $(p\text{x}r)$ matrices, and

$$\Psi = \alpha_{\perp}(-I + \Gamma_1)\beta_{\perp} = \varphi\eta' \tag{6}$$

where Ψ is a $(p-r)\text{x}(p-r)$ matrix of full rank, φ and η are $(p-r)\text{x}(p-r)$ matrices, and $\alpha_{\perp}$ and $\beta_{\perp}$ are $p\text{x}(p-r)$ matrices orthogonal to α and β, respectively. The parameterization in (4) and (5) facilitates the investigation of, on the one hand, the r linearly-independent stationary relations between the levels of the variables and, on the other hand, the $p-r$ linearly-independent non-stationary relations. This duality between the stationary relations and the non-stationary common trends is very useful for a full understanding of the generating mechanisms behind the chosen data. While the AR representation of the model is useful for the analysis of the long-run relations in the data, the MA representation is useful for the analysis of the common stochastic and deterministic trends that have generated the data.

The usefulness of this methodology for the present analysis is essentially related to the determination of the rank of the matrix, Π. Specifically, the number of cointegrating vectors provides evidence for the degree of integration across stock markets. First, if $n-r=0$ $(r=n)$, (full rank), then there is absence of any stochastic trends with all elements in z_t being stationary, i.e. $I(0)$, and cointegration is not defined. Second, if $n-r=n$ $(r=0)$, there are no stationary long-run relationships among the elements of z_t. Therefore, both short-term and long-term gains may occur from international portfolio diversification. Finally, if $n-r>1$ (reduced rank) then there are more than one common stochastic trends that drive the system of the stock markets. This case would imply that while long-run integration is not complete, the convergence process has started whereas the number of independent stochastic trends reflects the degree of this convergence as well as any gains from portfolio diversification and institutional adjustments arising from this process. Based on the reduced rank definition if $n-r=1$ $(r=n-1)$, there exists a single common stochastic trend which drives the system and creates the non-stationary property of the data. This finding suggests that although stock markets may exhibit transitory deviations from their long-run value, stock prices are gradually moved towards this single, common permanent component. Hence, there are no long-run portfolio diversification gains to be made. However, even in this case there exist short-run diversification gains and their effect will depend on the size and persistence of any transitory deviations from the trend as well as the time horizon of the investors.

An equally important issue, along with the existence of at least one cointegration vector, is the issue of the stability of such a relationship through time as well as the stability of the estimated coefficients of such a relationship. Strictly speaking in case that the rank of the cointegration space is sample dependent then the evidence in favour of cointegration is weak even though the likelihood ratio tests provide evidence of at least one cointegrating vector. Thus, Septhon and Larsen (1991) have shown that Johansen's test may be characterised by sample dependency. Hansen and Johansen (1993) have suggested methods for the evaluation of parameter constancy in cointegrated VAR models, formally using estimates obtained from the Johansen FIML technique. Three tests have been constructed under the two VAR representations. In the "Z-representation" all the parameters of model (4) are re-estimated during the recursions while under the "R-representation" the short-run parameters Γ_i, $i = 1...k$, are fixed to their full sample values and only the long-run parameters α and β are re-estimated.

The first test is called the *Rank* test and is used to examine the null hypothesis of sample independency of the cointegration rank of the system. This is accomplished by first estimating the model over the full sample, and the residuals corresponding to each recursive subsample are used to form the standard sample moments associated with Johansen's reduced rank. The eigenvalue problem is then solved directly from these subsample moment matrices. The obtained sequence of trace statistics is scaled by the corresponding critical values, and we accept the null hypothesis that the chosen rank is maintained regardless of the subperiod for which it has been estimated if it takes values greater than one.

A second test deals with the null hypothesis of constancy of the cointegration space for a given cointegration rank. Hansen and Johansen propose a likelihood ratio test that is constructed by comparing the likelihood function from each recursive subsample with the likelihood function computed under the restriction that the cointegrating vector estimated from the full sample falls within the space spanned by the estimated vectors of each individual sample. The test statistic is a χ^2 distributed with $(p-r)r$ degrees of freedom.

The third test examines the constancy of the individual elements of the cointegrating vectors β. However, when the cointegration rank is greater than one, the elements of those vectors can not be identified, except under restrictions. Fortunately, one can exploit the fact that there is a unique relationship between the eigenvalues and the cointegrating vectors. Therefore, when the cointegrating vectors have undergone a structural change, this will be reflected in the estimated eigenvalues. Hansen and Johansen (1993) have derived the asymptotic distribution as well as the asymptotic variance of the estimated eigenvalues.

3.2. Granger causality tests

The second stage of our analysis involves the examination of the short-run dynamics between stock and foreign exchange markets by employing standard Granger causality tests for cointegrating systems. This will enable us to shed light to the potential interrelationships among the variables of the system under examination and to reveal whether the transmission mechanism is either of *flow* and/or *stock* type.

For this analysis we adopt the methodology developed by Dolado and Lutkepohl (1996), which amounts for the construction of Wald tests that have standard asymptotic χ^2 - distributions. This approach has the advantage of avoiding the well known problem of pretest bias that may be present when estimating a first-order differenced VAR if variables are known to be I(1) with no cointegration, and an error correction model if they are known to be cointegrated.

The implementation of the methodology by Dolado and Lutkepohl (1996) is applied directly on the least squares estimators of the coefficients of the estimated VAR process specified in levels of the variables. These OLS estimates are consistent and asymptotically efficient given that each equation has the same lag length.[7] This approach is applied in two steps. First, we select the appropriate lag structure by testing a $\mathrm{VAR}(k)$ against a $\mathrm{VAR}(K+1)$, $k \geq 1$ and we then construct a Wald test.

[7] See for example Enders (1995).

Second, if the true data generating process is a VAR(k), a VAR$(K+1)$ model is fitted and we then apply standard Wald tests on the first k VAR coefficient matrix.

For the present study the Dolado and Lutkepohl (1996) procedure amounts to the estimation of the VAR in levels given by the VECM representation in eq. (4),

$$Y_t = \mu + A_1 Y_{t-1} + ... + A_p Y_{t-k} + \varepsilon_t \tag{7}$$

where A_i are pxp coefficient matrices.

For the bivariate case the appropriate VAR model for each Latin America country is given by:

$$\begin{bmatrix} I \\ E \end{bmatrix} = \begin{bmatrix} A_{10} \\ A_{20} \end{bmatrix} + \begin{bmatrix} A_{11}(L) & A_{21}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix} \begin{bmatrix} I_{t-i} \\ E_{t-i} \end{bmatrix} + \begin{bmatrix} \varepsilon_I \\ \varepsilon_E \end{bmatrix} \tag{8}$$

whereas for the trivariate case the appropriate VAR model for each Latin America country is given by:

$$\begin{bmatrix} I \\ E \\ I^* \end{bmatrix} = \begin{bmatrix} A_{10} \\ A_{20} \\ A_{30} \end{bmatrix} + \begin{bmatrix} A_{11}(L) & A_{12}(L) & A_{13}(L) \\ A_{21}(L) & A_{22}(L) & A_{23}(L) \\ A_{31}(L) & A_{32}(L) & A_{33}(L) \end{bmatrix} \begin{bmatrix} I_{t-i} \\ E_{t-i} \\ I^*_{t-i} \end{bmatrix} + \begin{bmatrix} \varepsilon_I \\ \varepsilon_E \\ \varepsilon^*_I \end{bmatrix} \tag{9}$$

where A_{i0} are the parameters representing intercept terms and A_{ij} the polynomials in the lag operator.

Following Phylaktis and Ravazzolo (2002) we test several alternative hypotheses related to the two channels of link between stock and foreign exchange markets which were described in the previous section. The hypotheses of interest are given as follows:

(a) $A_{13}(L) \neq 0, A_{23}(L), A_{12}(L) \neq 0$.

(b) $A_{13}(L) \neq 0, A_{21}(L) \neq 0$.

(c) $A_{13}(L) \neq 0, A_{23}(L) \neq 0, A_{12}(L) \neq 0, A_{21}(L) \neq 0$.

Hypothesis (a) describes the *flow* channel; hypothesis (b) reflects the *stock* channel; and hypothesis (c) tests for the combined transmission mechanism of *flow* and *stock* channels.

4. Empirical results

4.1. Data and preliminary results

In this paper we study four Latin America equity markets, namely those of Argentina, Brazil, Chile and Mexico. We use monthly data for the period January, 1980 to December 2005. The stock price indices are expressed in local currency and they are the *Standard and Poors'* indices. These are end-of-month value-weighted indices of a large sample of firms in each market and they are taken from *Datastream*. The *Standard and Poors'* indices correspond quite closely to the standard published indices, such as the Burcap index for Argentina, the Bovespa index for Brazil, the General index for Chile, the Bolsa for Mexico, but they have the advantages of being constructed on a consistent basis across countries and of netting out cross-listed securities.[8] For the US stock market we use the S&P500 Composition Index. Nominal

[8] Data availability was also an important factor for adopting the Standard and Poors' indices. Some local indices were reported for shorter periods and given the fact that we need to use as as a series is possible we decided not to use them.

exchange rates are defined as units of domestic currency per U.S. dollar, and the consumer price index is used as proxy for the respective domestic price level. All the data consists of end-of-month observations obtained from the IFS data base of the International Monetary Fund in *Datastream*. All the series are in logarithms.

As we have already discussed a central feature of our analysis is to investigate the impact of stock markets liberalization process. Table 1 (Panel A) presents the dates of the beginning of opening of the markets of the four Latin America emerging markets. These dates are identical across authors or very close, so we can consider them as the appropriate dates that can be taken for the purpose of our analysis. Panel B provides alternative signals of liberalization which include the Official Liberalization Date and the dates of introduction of the First Country Fund and First ADR. Panel C reports various indicators of direct and indirect for institutional investors which are used to assess the extent of liberalization in these economies. Therefore, in order to evaluate the effect of financial liberalization on the stock and foreign exchange markets we consider the pre and post-liberalization process with the addition of a dummy, for the specific date the liberalization began, in equation (1) for the bivariate model and in equation (3) for the trivariate model.[9]

For the case of Argentina the starting date of liberalization is November 1989 with the implementation of the New Foreign Investment Regime which lifted all legal limits on the type and nature of foreign investments. Furthermore, the market was fully opened in October 1991 with the adoption of the Deregulation Decree which

[9] Financial liberalization can take different forms: relaxation of foreign exchange restrictions, reduction of foreign ownership restriction and allowance of capital and dividends to be repatriated. In additional, in several emerging economies there have been indirect means of liberalization such as the creation of Country Funds and American Depository Receipts (ADRs) which allow foreigners to participate in local markets. These different approaches make the choice of liberalization date difficult in some cases but overall there is a broad agreement on this matter. For a comprehensive and detailed analysis about the liberalization dates and various indicators of direct and indirect barriers for international investors in various emerging economies see DeSantis and Imrohoroglu (1997, table A1), Kim and Singal (2000, Appendix) and Henry (2000, Tables I and III).

eliminated most restrictions on foreign investment, including taxes on capital gains. In addition in 1991 we have the implementation of the Convertibility Plan which established the one-to-one link between the peso and the U.S. dollar and the creation of the currency board. In Brazil, the liberalization process took place in May 1991 when foreign investors were allowed to hold up to 49% of voting common stocks and 100% of non-voting preferred stocks. Chile implemented a liberalization scheme with different stages with a gradual lift of all barriers to capital movements. The admission of the first country fund in October 1989 is taken as the date for the initial opening of the market although according to the International Finance Corporation (1995) the market was fully opened in January 1995. Finally, Mexico implemented a stabilization policy programme in 1988 when the peso was pegged to the US dollar with a fluctuation band. In May 1989 the market was fully open when all restrictions in foreign investments were abolished. Foreign investment in now permitted up to 100% in the majority of Mexico's firms in all economic sectors.[10]

To examine whether the series under consideration are stationary we apply the Elliot *et al*. (1996) GLS augmented Dickey-Fuller test (DF-GLS$_u$) and Ng and Perron (2001) GLS versions of the modified Phillips-Perron (1988) tests (MZ_a^{GLS} and MZ_t^{GLS}). The null hypothesis is that of a unit root against the alternative that the initial observation is drawn from its unconditional distribution and uses GLS-detrending as proposed by Elliott *et al*. (1996) and extended by Elliott (1999), to maximize power, and a modified selection criterion to select the lag truncation parameter in order to minimize size distortion. In the GLS procedure of Elliot *et al.* (1996), the standard unit root tests (without trend) are applied after the series are first

[10] The International Finance Corporation has a complete data set with all opening dates and legal arrangements since 1974, (IFC, various issues). See also Bekaert (1995), DeSantis and Imrohoroglu (1997, table A1), Kim and Singal (2000, Appendix) and Henry (2000, Tables I and III).

detrended under the local alternative $\rho = 1 + \alpha / T$. This was found to provide substantial power gains for the DF-GLS$_u$ test resulting to power functions that lie just under the asymptotic power envelope. Ng and Perron (2001) find similar gains for the MZ_a^{GLS} and MZ_t^{GLS} tests. They also found that a modification of the AIC criterion (MIC), give rise to substantial size improvements over alternative selection rules such as BIC. For robustness, we then apply the Kwiatkowski *et al*. (1992) KPSS test for the null hypothesis of level or trend stationarity against the alternative of non-stationarity. The results are reported in Table 2 and they show that we are unable to reject the null hypothesis of non-stationarity with the DF-GLS$_u$ and MZ_a^{GLS} and MZ_t^{GLS} tests and we reject the null hypothesis of stationarity with the KPSS test for the levels of all stock prices, the exchange rates and consumer price indices. The results are reversed when we take the first difference of all series which leads us to the conclusion that all variables are realizations of $I(1)$ processes.

4.2. Cointegration analysis

We apply Johansen's (1988, 1991) multivariate cointegration analysis on two distinct sets of variables. We first analyze the bivariate case for each emerging economy, which includes the domestic stock price index and the corresponding real exchange rate. We then move to the trivariate case which includes in addition the US stock price index.[11]

The first stage of our analysis is the determination of the cointegration rank index, r, for the bivariate systems. Since we are interested to investigate the impact of stock markets on foreign exchange and stock markets we include a dummy variable

[11] To check for the statistical adequacy of model (1) we apply several univariate misspecification tests in order to investigate whether the estimated residuals do not deviate from being Gaussian white noise errors. We estimate a structure of two lags for each case which was chosen based on these misspecification tests and the Sims (1980) test for the presence of serial correlation in each equation. These tests are available upon request.

on the date of full opening of the corresponding Latin America stock market.[12] The results of the trace test are presented in Table 3[13]. Our results indicate that at the five percent level of significance we are unable to reject the null hypothesis of no cointegration between the domestic stock index and the real exchange for each bivariate case.

The finding of no cointegration between the stock market index and the real exchange in Argentina, Brazil, Chile and Mexico may suggest that a significant variable that causes the other two is missing, as we already explained in section 2. We repeat our cointegration analysis for the case of the trivariate system between the domestic stock market, the real exchange rate and the US stock market which can been thought as representing the rest of the world.[14] Table 4 provides a full account of the estimated trace statistic for each emerging market along with several multivariate misspecification tests. The overall evidence is that there exists one statistically significant cointegration vector among the three variables. This finding implies that the US stock market has significant influence on the Latin America stock markets. Furthermore, the misspecification tests show that there is no evidence of serial correlation although there is presence of nonnormality.[15]

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[12] An alternative method to examine the influence of stock market openness to the existence of long-run relationships between the variables under consideration is to split the sample period into pre and post liberalization sub-periods.

[13] In the presence of a dummy the critical values reported by MacKinnon *et al*. (1999) are not valid. We use the program Disco (Johansen and Nielsen, 1993) to conduct a Monte Carlo simulation which was repeated 10,000 times of 400 observations. Cheung and Lai (1993) have shown than the trace test is more robust than the maximum eigenvalue tests to the presence of non-normal errors. The estimated trace and maximum eigenvalue statistics have been corrected for small sample adjustment, (Reimers, 1992).

[14] In all cases the dummy for the period of financial liberalization period is statistically significant.

[15] Gonzalo (1994) shows (a) that the performance of the maximum likelihood estimator of the cointegrating vectors is little affected by non-normal errors and (b) nonnormality is not an appropriate criteria upon which to select a lag length. Thus, choosing a higher order lag structure to eliminate non-normality may result in obtaining misleading results regarding the number of cointegrating vectors due to the decrease of power and the loss of inference. The crucial criterion for selecting lag length in the VAR specification is to select that lag length which will eliminate serial correlation in the residuals. Lee and Tse (1996) have shown similar results when conditional heteroskedasticity is present.

In Table 5 we perform three tests which are important for checking the statistical adequacy of our model (Panel A). The first test examines the null hypothesis of exclusion from the cointegration vector. The results are statistically significant implying that for each economy none of the variables is excluded from the long-run relationship. We then test the null hypothesis of stationarity given a specific cointegration rank. The results show that all variables are nonstationary in each Latin America country. Finally, we test for weak exogeneity of each variable to the long run parameters. The hypothesis is strongly rejected in all cases. Panel B reports a set of multivariate misspecification tests which show that there is only presence of non-normality which is expected given the distributional properties of the stock returns.

Table 6 reports the long-run cointegrating vectors for the trivariate model. In all cases the coefficient is positively related to the domestic stock markets. We also note that there is a positive relationship between the US stock market and the respective domestic equity market. These results provide support for the increasing capital market integration of the four Latin America economies with the US market.

These findings allow us to assess the effects of financial liberalization on the link between the foreign exchange and stock markets. First, given that no evidence of cointegration was found between the two markets as well as that the dummy which captures the effect of financial liberalization was statistically insignificant implies that there may be additional factors beyond market openness that amount for the existence of a link between the two markets. Home bias effect of investors and lack of information on company stocks may be significant factors that make investors to restrict investment in global markets.[16] Moreover, access to market information is equally important to portfolio managers with the access to the stock market itself.[17]

4.3. Recursive tests for parameter stability

As we explained in section 3 the finding of cointegration is necessary but not sufficient condition for the existence of a long-run relationship between the foreign exchange and the domestic and the US stock markets. We also need to examine the parameter stability of our results. Figures 1(a)-(d) to 3(a)-(d) present the Hansen-Johansen (1993) recursive analysis on the parameter stability of the cointegrated-VAR models. The first set of figures shows that the rank of the cointegration space is independent on the sample size from which it has been estimated, since we are unable to reject the null hypothesis of a constant rank except for the case of Brazil. In the Brazilian case we observe that the first cointegration vector is not above one which the critical value for the whole sample. The second set of figures indicates that we reject the null hypothesis for parameter constancy since the path of the test statistics is above 1 in several cases. This evidence is confirmed by the third set of figures which traces the path of the first eigenvalue with 95% confidence bands.

We note that there is instability around the Mexican financial crisis in the end of 1994 and early 1995. There is also some evidence of instability in 1998 in the aftermath of the Asian crisis which is more evident for the case of Brazil which faced significant financial problems in 1998 and 1999 which have led to a substantial depreciation of its currency. We also note that in the case of Argentina there is an

[16] See also Bekaert (1995), Bekaert and Harvey (2000) and Levine and Zervos (1996)

[17] Indicators an market information accession are available on annual basis for most of the emerging markets by the International Finance Corporation. The technological advances in the 1990s have also improved substantially the degree of accession of market investors.

additional shift in the time path of the eigenvalue which reflects the episode of the economic events of 2001-2002 and the subsequent abandonment of the dolarization exchange rate policy. However, the time path of these test statistics remains below one for the period after 1998. It is evident from these results that our model captures quite well the events of on these financial markets in the late 1980s and throughout the 1990s. Furthermore, we note that the effects of the peso problem had much larger negative effects on the four Latin America financial markets than the Asian crisis and in addition we observe that in all cases the effects of both crises did not have long lasting effects. Finally, these results may be considered as evidence for financial contagion during the period of crises with the link between foreign and stock markets being increased.[18]

4.4. Multivariate Granger causality tests

The last stage of our analysis deals with the examination of the short-run dynamics between the four Latin America foreign exchange and stock markets. We apply the methodology suggested by Dolado and Lutkepohl (1996) presented in section 3 and we examine the alternative channels of transmission between foreign exchange and stock markets.

The results for the trivariate model are given in Table 7. These test statistics indicate that in Argentina and Brazil, the markets are connected through the "flow" channel since the restrictions $A_{12}(L)=0,\ A_{13}(L)=0$ and $A_{23}(L)=0$ are rejected, and in Mexico the markets are connected through the "stock" channel since the restrictions $A_{13}(L)=0$ and $A_{21}(L)=0$ are rejected. Finally, in Chile the markets are connected through both channels since all the above restrictions are rejected.

[18] Forbes and Rigobon 92002) and Corsetti et al. (2005) are recent studies on the issues of contagion and interdependence. Pericoli and Sbracia (2003) provide a comprehensive survey on these issues.

A final interesting point is derived from the evidence that the restriction $A_{13}(L)=0$ is strongly rejected in all four countries while the restriction $A_{31}(L)=0$ is accepted in all four economies. Therefore, there is clear evidence that the US stock market is driving the four Latin America economies, a result which is not surprising given the economic relations of the economies with the US economy. However, there is no feedback from the Latin America stock markets to the US stock market.

5. Summary and concluding remarks

This paper examined the long-run relationships and the short-run dynamics between stock prices and exchange rates for a group of Latin America countries. The main purpose of the analysis was to investigate whether these links were influenced by the financial liberalization process implemented in the Latin America emerging markets in late 1980s and early 1990. In addition we examined whether these relationships were affected by the Mexican peso crisis of 1994 and the Asian crisis of 1997.

The first stage of the analysis involved the use of Johansen (1998, 1991) multivariate cointegration methodology in order to reveal the existence of statistically significant long-run relationships between the domestic stock market, the respective real exchange and the US stock market. Furthermore, we examined the temporal stability of the estimated cointegrated VAR systems. The second stage involved the examination of the short-run dynamics with the application of multivariate Granger causality tests.

There are several interesting finding stem from our analysis. First, we were unable to find a statistically significant cointegration vector in each Latin America country between each domestic stock market and its respective real exchange rate for either the 1980s or the 1990s. Second, when our VAR system was completed with the inclusion of the US stock market which was taken as a proxy for the world equity markets our results were in favour of the existence of a statistically significant long-run relationship for the three variables in each case. This finding provides evidence that the US stock market operates a transmission mechanism through which the domestic stock and foreign exchange markets are linked. In addition, for each emerging market the estimated coefficients of the real exchange rate and the US stock market are positive suggesting a positive relationship with the respective domestic stock market. Finally, with the use of the Hansen and Johansen (1993) recursive test for parameter constancy we showed that the identified long-run relationship is stable over time except for the case of Brazil whereas the estimated parameters for each economy exhibit instability during the period of the Mexican peso crisis but not during the Asian crisis. It also evident that around the Mexican peso crisis these linkages between the two markets were stronger confirming earlier studies on contagion effects but at the same time these events were short lived and since in mid 1995 the long-run relationships appeared to have regained its normal evolution.

Third, the results of the multivariate Granger causality tests indicate that the US stock market drives the system of the four emerging economies which confirms the influence that the US economy has on these Latin America economies. Furthermore, the causality results show that in Argentina and Brazil the link between the two markets is done through the "flow channel", in Mexico through the "stock channel" and in , in Chile the markets are connected through both channels.

Finally, our results indicate that financial liberalization has not been a key factor in the link either between the domestic stock and foreign exchange markets or between the domestic and the US stock markets. Besides to free capital movement, access to market information is considered to be important for increasing foreign investment. The fact that Argentina, Brazil, Chile and Mexico have a substantial volume of trade with the US maybe a reason that independently of the degree of financial liberalization of these countries there is a significant degree of financial integration as well.

Our overall results are useful for the understanding of the interrelationships between the foreign exchange and domestic and foreign stock markets. Thus, these findings maybe of interest to portfolio managers, private and institutional investors as well as hedge funds that are active in emerging markets.

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Table 1.A: Different liberalization dates across authors

	Bekaert and Harvey (2000)	DeSantis and Imrohoroglou (1997)	Kim and Singal (2000)
Argentina	1989.11	1991.10	1989.11
Brazil	1991.05	1991.05	1991.05
Chile	1992.01	1988.12	1987.09
Mexico	1989.05	1989.05	1989.11

Table 1.B. Alternative signals of liberalization

Country	Official Liberalization Date	First Country Fund	First ADR
Argentina	1991.10	1991.10	1991.08
Brazil	1991.05	1987.10	1992.01
Chile	1988.12	1989.09	1990.06
Mexico	1989.05	1989.06	1989.01

Notes: These dates are taken from Bekaert and Harvey (1998) and Bekaert *et al.* (2002). These dates coincide with those of the International Financial Corporation.

Table 1.C. Emerging stock markets: Direct and indirect barriers for institutional indicators-end 1995

Country	Foreign ownership Limit	Dividends Repatriation	Capital Repatriation	Withholding Taxes on Dividends	Taxes on capital Gains
Argentina	100%	Free	Free	0.0%	0.0%
Brazil	100%[a] 49%	Free	6 years	15%	0.0%
Chile	25%	Free	1 year	0.0%	0.0%
Mexico	30%[b] 100%	Free	Free	0.0%	0.0%

Notes: The table is based on the information given in the International Financial Corporation Fact Book and De Santis and Imrohoroglou (1997), Bekaert and Harvey (1998) and Bekaert *et al.* (2002).
(a) 100% of non-voting preferred stocks and 49% of voting common stocks
(b) 30% for banks and 100% for other stocks.

Table 2. Unit root and stationarity tests

Variables	DF-GLS$_u$ t_μ	t_τ	MZ_a^{GLS}	MZ_t^{GLS}	KPSS η_μ	η_τ
$\ln \arg se$	-0.44 [0]	-2.30 [0]	-0.77 [0]	-0.68 [0]	1.340*	0.285*
$\Delta \ln \arg se$	-16.96* [0]	-16.77* [0]	-153.57* [0]	-8,76* [0]	0.092	0.031
$\ln brse$	-0.63 [0]	-1.30 [0]	-1.06 [0]	-0.62 [0]	0.538*	0.353*
$\Delta \ln brse$	-14.72* [0]	-14.74* [0]	-116.38* [0]	-7.62* [0]	0.169	0.066
$\ln chse$	1.74 [2]	-0.51 [2]	0.86 [0]	1.81 [0]	2.100*	0.269*
$\Delta \ln chse$	-9.37* [1]	-9.49* [1]	-86.50* [0]	-6.57* [0]	0.238	0.138
$\ln mxse$	2.06 [1]	-1.09 [1]	1.05 [1]	2.12 [1]	2.124*	0.334*
$\Delta \ln mxse$	-13.43* [0]	-13.55* [0]	-155.24* [0]	-8.51* [0]	0.233	0.133
$\ln usse$	1.97 [0]	-1.85 [0]	1.11 [0]	2.01 [0]	1.149*	0.304*
$\Delta \ln usse$	-18.82 [0]	-16.58 [0]	-205.27 [0]	-10.61* [0]	0.106	0.085
$\ln E \arg$	-1.21 [0]	-1.53 [0]	-4.17 [0]	-1.20 [0]	0.912*	0.358*
$\Delta \ln Ear$	-17.37* [0]	-17.42* [0]	-153.51* [0]	-8.16* [0]	0.161	0.086
$\ln Ebr$	-1.434 [0]	-1.565 [0]	-4.17 [0]	1.51 [0]	0.912*	0.216*
$\Delta \ln Ebr$	-15.30* [0]	-18.42* [0]	-116.77* [0]	-7.64* [0]	0.114	0.090
$\ln Ech$	-0.11 [0]	-0.89 [0]	-0.11 [0]	-0.10 [0]	0.707*	0.441*
$\Delta \ln Ech$	-18.26* [0]	-18.29* [0]	-187.80* [0]	-9.68* [0]	0.181	0.090
$\ln Emx$	1.34 [0]	-1.11 [0]	0.91 [0]	1.49 [0]	2.138*	0.413*
$\Delta \ln Emx$	-4.14 [0]	-4.71 [0]	-11.57 [5]	-2.40 [5]	0.127	0.077

Notes:

- The DF-GLS$_u$ is due to Elliot et al. (1996) and Elliott (1999) is a test with an unconditional alternative hypothesis. The standard Dickey-Fuller tests are detrended (with constant or constant and trend). The critical values for the DF-GLS$_u$ test at the 5% significance level are:-2.73 (with constant) and -3.17 (with constant and trend), respectively (Elliott,1999).
- MZ_a and MZ_t are the Ng and Perron (2001) GLS versions of the Phillips-Perron tests. The critical values at 5% significance level are: -8.10 and -1.98 (with constant), respectively (Ng and Perron, 2001, Table 1).
- η_μ and η_τ are the KPSS test statistics for level and trend stationarity respectively (Kwiatkowski *et al*. 1992). For the computation of theses statistics a Newey and West (1994) robust kernel estimate of the "long-run" variance is used. The kernel estimator is constructed using a quadratic spectral kernel with VAR(l) pre-whitening and automatic data-dependent bandwidth selection [see, Newey and West, 1994 for details]. The 5% critical values for level and trend stationarity are 0.461 and 0.148 respectiveley, and they are taken from Sephton (1995, Table 2).

(*) indicates significance at the 95% confidence level.

Table 3. Johansen - Juselius bivariate cointegration trace tests

$$I_t = b_0 + b_1 E_t + u_t$$

	$r = 0$	$r \leq 1$
Market		
Argentina	19.16	6.72
Brazil	18.19	6.59
Chile	18.33	3.66
Mexico	17.56	8.54

Notes: r denotes the number of eigenvectors. Trace is the Johansen likelihood ratio statistics. The 5% critical values are 22.13 and 8.21, respectively and they have been simulated using the Johansen and Nielsen (1993) DisCo programme for a model with one intervention dummy. A structure of four lags was chosen according to a likelihood ratio test, corrected for the degrees of freedom (Sims, 1980) and the Ljung-Box Q statistic for detecting serial correlation in the residuals of the equations of the VAR. A model with a constant restricted in the cointegrating vector is chosen according the Johansen (1992 a, b, 1994) testing strategy. (*) denotes statistical significance at the five percent critical level.
A small sample adjustment has been made in all the likelihood ratio statistics, equal to

$$-2\ln Q = -(T - kp)\sum_{i=r_{0+1}}^{k} \ln(1 - \hat{\lambda}_i)$$

as suggested by Reimers (1992).

Table 4. Johansen - Juselius Trivariate cointegration trace tests

$$I_t = b_0 + b_1 E_t + b_2 I_t^*$$

	$r=0$	$r \leq 1$	$r \leq 2$
Market			
Argentina	37.61*	17.15	5.43
Brazil	38.50*	18.76	5.87
Chile	63.54*	15.79	2.62
Mexico	71.91*	13.92	2.57

Notes: r denotes the number of eigenvectors. Trace is the Johansen likelihood ratio statistic. The 5% critical values are 34.22, 22.13 and 8.21, respectively and they have been simulated using the Johansen and Nielsen (1993) DisCo programme for a model with one intervention dummy. A structure of two lags was chosen according to a likelihood ratio test, corrected for the degrees of freedom (Sims, 1980) and the Ljung-Box Q statistic for detecting serial correlation in the residuals of the equations of the VAR. A model with an unrestricted linear trend in the VAR equation and a constant restricted in the cointegrating vector is chosen according the Johansen (1992 a, b) testing strategy. (*) denotes statistical significance at the five percent critical level.

A small sample adjustment has been made in all the likelihood ratio statistics, equal to

$$-2\ln Q = -(T - kp)\sum_{i=r_{0+1}}^{k} \ln(1 - \hat{\lambda}_i)$$

as suggested by Reimers (1992).

Table 5. Statistical Properties and Misspecification Tests of the Model

(a) Tests for long-run exclusion, stationarity, and weak exogeneity

	Long-run exclusion	Stationarity	Weak exogeneity
$\ln \arg se$	8.15*	13.62*	8.08*
$\ln E\arg$	10.09*	12.43*	9.72*
$\ln usse$	11.93*	14.11*	8.17*
$\ln brse$	6.84*	13.98*	6.55*
$\ln Ebr$	4.14*	10.01*	9.38*
$\ln usse$	8.96*	12.24*	6.25*
$\ln chse$	13.00*	33.61*	33.48*
$\ln Ech$	11.11*	24.80*	5.98*
$\ln usse$	5.06*	28.05*	9.03*
$\ln mxse$	13.22*	37.26*	18.88*
$\ln Emx$	14.02*	33.86*	18.07*
$\ln usse$	9.01*	26.29*	12.58*

Notes: The long-run exclusion restriction and the long-run weak exogeneity tests are χ^2 distributed with one degree of freedom and the 5% critical level is 3.84, and the stationarity test is a χ^2 distributed with three degrees of freedom and the 5% critical level is 7.81. An (*) denotes statistical significance at the 5 percent critical level.

(b) Multivariate Residuals Diagnostics

Argentina

L-B(333)	LM(1)	LM(4)	χ^2 (6)
304.10[0.87]	11.58[0.24]	1.09[1.00]	7795.66[0.00]

Brazil

L-B(504)	LM(1)	LM(4)	χ^2 (6)
517.30[0.33]	5.89[0.75]	6.42[0.70]	29355[0.00]

Chile

L-B(738)	LM(1)	LM(4)	χ^2 (6)
689.40[0.90]	14.61[0.19]	4.27[0.89]	36927[0.00]

Mexico

L-B(738)	LM(1)	LM(4)	χ^2 (6)
582.61[0.31]	4.35[0.89]	6.16[0.72]	491.59[0.00]

Notes: L-B is the multivariate version of the Ljung-Box test for autocorrelation based on the estimated auto- and cross- correlations of the first [T/4] lags distributed as a χ^2 with the degrees of freedom given in parentheses. LM(1) and LM(4) are the tests for first- and fourth-order autocorrelation distributed as χ^2 with 9 degrees of freedom and χ^2 is a normality test which is a multivariate version of the Shenton-Bowman (1977) test modified in Doornik and Hansen (1994). Numbers in brackets refer to marginal significance levels.

Table 6. Estimated Coefficients

$$I_t = b_0 + b_1 E_t + b_2 I_t^* + u_t$$

	b_0	b_1	b_2
Argentina	8.25 (1.05)	2.09 (0.56)	1.33 (0.34)
Brazil	9.33 (2.01)	1.97 (0.98)	1.46 (0.28)
Chile	8.69 (1.67)	3.56 (0.55)	2.89 (1.01)
Mexico	10.23 (2.02)	3.09 (1.05)	1.09 (0.11)

Notes: Numbers in parentheses report standard errors. The eigenvector is normalized with respect to the domestic price index.

Table 7. Multivariate Granger causality tests

	$A_{12}(L)=0$	$A_{13}(L)=0$	$A_{21}(L)=0$	$A_{23}(L)=0$	$A_{31}(L)=0$
Argentina	2.090* [0.03]	0.410* [0.02]	0.359 [0.83]	4.754* [0.00]	1.531 [0.83]
Brazil	2.986* [0.02]	0.245** [0.06]	3.954 [0.55]	8.713* [0.01]	5.493 [0.12]
Chile	7.612** [0.02]	6.030* [0.04]	43.544* [0.00]	8.393* [0.02]	8.465 [0.41]
Mexico	1.983 [0.37]	3.247* [0.03]	26.443* [0.00]	12.664* [0.00]	8.166 [0.38]

Notes: The restrictions are explained in equation (6). The Wald tests are distributed as χ^2 with two degrees of freedom. Figures in brackets are p-values and (*) and (**) denote statistical significance at the 5% and 10% level respectively.