Modelling the linkages between US and Latin American stock markets

by

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ABSTRACT

This paper examines the linkages between US and Latin American the stock markets during the 1995-2002 period using recently-developed cointegration techniques that allow for structural shifts in the long-run relationship. Our results suggest that, if we apply conventional cointegration tests, we only find a long-run relationship in the cases of Brazil and Mexico for the Dow Jones (DJ) index and in the case of Brazil for the Standard and Poor's 500 (SP500) index. In contrast, if we introduce the possibility of structural breaks, we find strong evidence in favour of such relationship between the Argentine, Chilean and Venezuelan indices and the DJ index after the 1998 financial turmoil, and between the Brazilian and Mexican indices and the DJ index before such turbulence, while some marginal cointegration is detected between the Mexican and DJ indices from February 1998. Additionally, we find evidence of a cointegrating relationship between the Argentine, Chilean and Mexican indices and the SP500 index from August 1998, April 1999 and October 1999, respectively, and between the Brazilian and the SP500 indices before November 1997, as well as some marginal cointegration between the Mexican and SP500 indices before October 1999. The results suggest that the gains from international diversification for investors with long holding periods is limited.

JEL classification numbers: C22, F36, G15

KEY WORDS: Stock market, Cointegration, Structural change

1. Introduction

Since the stock market crash in October 1987, there have been an increasing interest in empirical and theoretical investigations of the linkages between asset markets. More recently, the financial crisis in Asian markets has renewed this interest.

This issue is an important concern for investors since in recent years global market have became more integrated as a result of a broad tendency toward liberalisation and deregulation in the money and capital markets of developed as well as developing countries, therefore reducing the opportunities for international diversification. Furthermore, market comovements can also lead to market contagion as investors incorporate into their trading decisions information about price changes in other markets in an attempt to form a complete information set, carrying the risk that errors in one market may be transmitted elsewhere [see, e. g., King and Wadhwani (1990)].

While there exists a extensive research examining the international stock market linkages in United States, Europe, Japan and even the Pacific-Basin stock markets, there is a scant research on the linkages between US and Latin American stock markets in contrast with the growing economic importance of Latin America in the emerging markets.

The purpose of this paper is to investigate the long-run relationships between six major Latin American stock markets and the United States during the 1995-2002 period. Our study differs from the previously published papers in several ways. First, while most of the studies examine mainly short-run market relationships through correlation tests [see, e. g., DeFusco *et al.* (1996) and Aggarwall *et al.* (1999)], we explore whether there are linkages and long-run comovements between the US and the major Latin American stock markets. Evidence of such long-run comovement would suggest greatly overstated benefits for US investors with longer-term investment horizons who diversify in these emerging markets [see, e. g., Kasa (1992)]. Second, we make use of recently-developed cointegration techniques that allow for structural shifts in the long-run relationship. These shifts could be due to changes experienced in those countries during the sample period, such as the movement from being relatively isolated from outside influences to their stock markets being opened up and exchange rates floated, or the 1997-98 global financial crisis. Finally, indices from the major stock exchanges are employed as well as the use of daily data.

The remained of the paper is organised as follows. The econometric methodology is presented in Section 2, while Section 3 describes the data set and reports the empirical results. Finally, some concluding remarks are provided in Section 4.

2. Econometric methodology

The econometric methodology used in this paper has the following main objectives: (i) to analyse the order of integration of the variables and the stability of the stochastic trend, (ii) to examine the long run relationship between the variables using cointegration techniques, and (iii) to test for parametric instability in the estimated cointegration relations[1].

To test for unit roots, in addition to the traditional augmented Dickey-Fuller (ADF) test, we use sequential ADF tests in order to detect structural changes in the stochastic trend as well as changes in the degree of integration [Fernández-Serrano and Peruga (1999a y 1999b)]. As it is well known, to test the null hypothesis of a unit root in a time series Y_t, the standard ADF test computes the pseudo t-ratio (t_δ) in the following regression:

$$\Delta Y_t = \mu + \beta t + \delta Y_{t-1} + \sum_{i=1}^{q} \gamma_i \Delta Y_{t-i} + \varepsilon_t, \tag{1}$$

The sequential ADF test usually employed in the literature [Banerjee, Lumsdaine and Stock (1992), Zivot and Andrews (1992), Perron and Vogelsang (1992) and Montañés (1996)] involves estimating the following regressions:

$$\Delta Y_t = \mu + \mu' D_{\lambda t} + \delta Y_{t-1} + \sum_{i=1}^{q} \gamma_i \Delta Y_{t-i} + \varepsilon_t, \tag{2}$$

where

$$D_{\lambda t} = \begin{cases} 0 & t < [\lambda T] \\ 1 & t \geq [\lambda T] \end{cases}, \quad \lambda \in (\tau,\ 1-\tau) \tag{3}$$

is a dummy variable that selects the last $(1-\tau)\%$ observations of the sample, and [.] indicates integer part. For each possible break date in the sample, $[\lambda T]$, two statistics are computed from regression (2): t_{δ} and $|t_{\mu'}|$. t is the standard

[1]It should be noted that we are primarily interested in the shift of mean only but not the shift in variance. However, Hamori and Tokihisa (1997) have shown that the limiting distribution of the standard unit root is not invariant to changes in variance (heteroskedasticity). Therefore, our results should be taken with caution.

pseudo t-ratio for testing the null hypothesis of a unit root $\delta = 0$, while $|t_{\mu'}|$ is the absolute value of the t statistic for testing the null hypothesis $\mu' = 0$ (i.e., it is a test of stability in the stochastic trend). If we impose the existence of a unit root in (2), we have the following restricted regression:

$$\Delta Y_t = \mu + \mu' D_{\lambda t} + \sum_{i=1}^{q} \gamma_i \Delta Y_{t-i} + \varepsilon_t, \tag{4}$$

and, based on it we can compute $|t_{(\mu')}|$. Regarding the break fraction parameter τ, we follow the convention of using $\tau \in [0.15, 0.85]$.

From regressions (2) and (4), we obtain a sequence of estimated values for each statistic. From this sequence, we take two summary values: the supreme and the mean. Therefore we compute the following six statistics: Inf t_δ, Mean t_δ, Sup $|t_{\mu'}|$, Mean $|t_{\mu'}|$, Sup $|t_{(\mu')}|$ and Mean $|t_{(\mu')}|$.

Following Zivot and Andrews (1992), we consider as a breakpoint the observation associated with the corresponding supreme: Ninf t_δ, Nsup $|t_{\mu'}|$ and Nsup $|t_{(\mu')}|$.

Until now, we have assumed that the time series follows a stochastic process that has always the same degree of integration [i. e., changes in the parameter δ in regression (1) are not allowed]. To consider the possibility that such parameter may not be constant in all the sample (and, therefore, the order of integration of the stochastic process could change depending on the subsample examined), we can study the following set of regressions:

$$\Delta Y_t = \mu + \gamma_1[1-D_{\lambda t}]Y_{t-1} + \gamma_2 D_{\lambda t} Y_{t-1} + \sum_{i=1}^{k} \delta_i \Delta Y_{t-i} + u_t \tag{5}$$

$$\Delta Y_t = \mu + \alpha_1[1-D_{\lambda t}]Y_{t-1} + \sum_{i=1}^{k} \delta_i \Delta Y_{t-i} + u_t \tag{6}$$

$$\Delta Y_t = \mu + \alpha_2 D_{\lambda t} Y_{t-1} + \sum_{i=1}^{k} \delta_i \Delta Y_{t-i} + u_t \tag{7}$$

Regression (5) simultaneously considers both subsamples resulting from the division of the sample using a dummy variable. Since none restriction is imposed in any of the subsamples, this regression tries to test simultaneously the null hypothesis of a unit root against the alternative hypothesis of stationarity in both subsamples. In this way, for example, the time series could be integrated of order one in one subsample and integrated of order zero in the other. In the other two regressions, the existence of a unit root is imposed in one subsample (in the second subsample in regression (6) and in the first subsample in regression (7)), allowing the possibility of the variable being stationarity in the not restricted subsample.

For each possible break point in the sample, the statistics $t_{\gamma 1}$, $t_{\gamma 2}$, $t_{\alpha 1}$, and $t_{\alpha 2}$ are then computed. The first two statistics ($t_{\gamma 1}$ and $t_{\gamma 2}$) test for a unit root in the first or second subsamples, respectively. The statistic $t_{\alpha 1}$ tests separately for a unit root in the first subsample, while the statistic $t_{\alpha 2}$ does the same in the second subsample. As in the previous case, after finishing this testing procedure, we will have four sequences of estimated statistics and from these sequences we compute the summary statistics: $\text{Sup } t_{\gamma 1}$, $\text{Mean } t_{\gamma 1}$, $\text{Sup } t_{\gamma 2}$, $\text{Mean } t_{\gamma 2}$, $\text{Sup } t_{\alpha 1}$, $\text{Mean } t_{\alpha 1}$, $\text{Sup } t_{\alpha 2}$ and $\text{Mean } t_{\alpha 2}$. In this case, the estimators of the break points are $\text{Nsupt}_{\gamma 1}$, $\text{Nsupt}_{\gamma 2}$, $\text{Nsupt}_{\alpha 1}$ and $\text{Nsupt}_{\alpha 2}$, respectively.

In order to estimate the cointegrating vector and analyse its stability, we have made use of Gregory and Hansen (1996)´s generalization of the usual residual based cointegration tests, that allows for a broader view of cointegration by considering an alternative hypothesis in which the cointegration vector suffers shift at an unknown time. Therefore, we test for a unit root in the residuals of the following cointegrating regression:

$$Y_t = \mu_1 + \mu_2 D_{\lambda t} + \alpha_1 X_t + \alpha_2 D_{\lambda t} X_t + \varepsilon_t, \quad t = 1, \ldots, T \tag{8}$$

where X_t is a vector of I(1) regressors and ε_t is I(0). Once we have estimated (7) by OLS, we apply the ADF test to the cointegrating residual $\hat{\varepsilon}_t$. For each possible break point we compute the ADF(t) test. As before, we will have a sequence of estimated statistics and from this sequence we compute the summary statistics: InfADF and MeanADF, being NinfADF the estimator of the break point.

3. Empirical results

In this paper we have used daily closing prices of the six major Latin American markets (Argentina, Brazil, Chile, Mexico, Peru and Venezuela) and the United States. Specifically, the indices under study are the Merval Index of the Buenos Aires Stock Exchange (Argentina), the Bovespa Index of the São Paulo Stock Exchange (Brazil), the General Stock Price Index of the Santiago de Chile Stock Exchange (Chile), the Price and Quotations Index of the Mexican Stock Exchange (Mexico), General Index of Lima Stock Exchange (Peru), the General Index of the Caracas Stock Exchange (Venezuela), the Dow Jones (DJ) and the Standard and Poor's 500 (SP500). The total capitalisation of this six markets constitutes more than 90% of the market capitalisation of all Caribean and latin American equity markets during the period [see Standard & Poor´s (2000)]. The data are obtained from Ecowin and cover the period 2 January 1995 through 14 February 2002 (1859 observations). We have normalised the data by subtracting the mean from the series and dividing the results by the standard deviation of the original series. It should be notice that all the markets examined trade simultaneously during the day, and therefore the market linkages can be analysed in a more appropriate setting. Regarding the missing observations associated with the possibility of holidays differing across countries, we have followed the convention of using the previous day closing price when there is no trading in a country under study.

Given the importance of the United States stock market in the region, we take it as the reference in order to study the linkages between US and Latin American stock markets. In Figure 1 we present the graphs of the levels of the Latin American indices together with the DJ index, while Figure 2 shows them with the SP500 index.

Figure 1. Levels of the indices together with USA Dow Jones Index

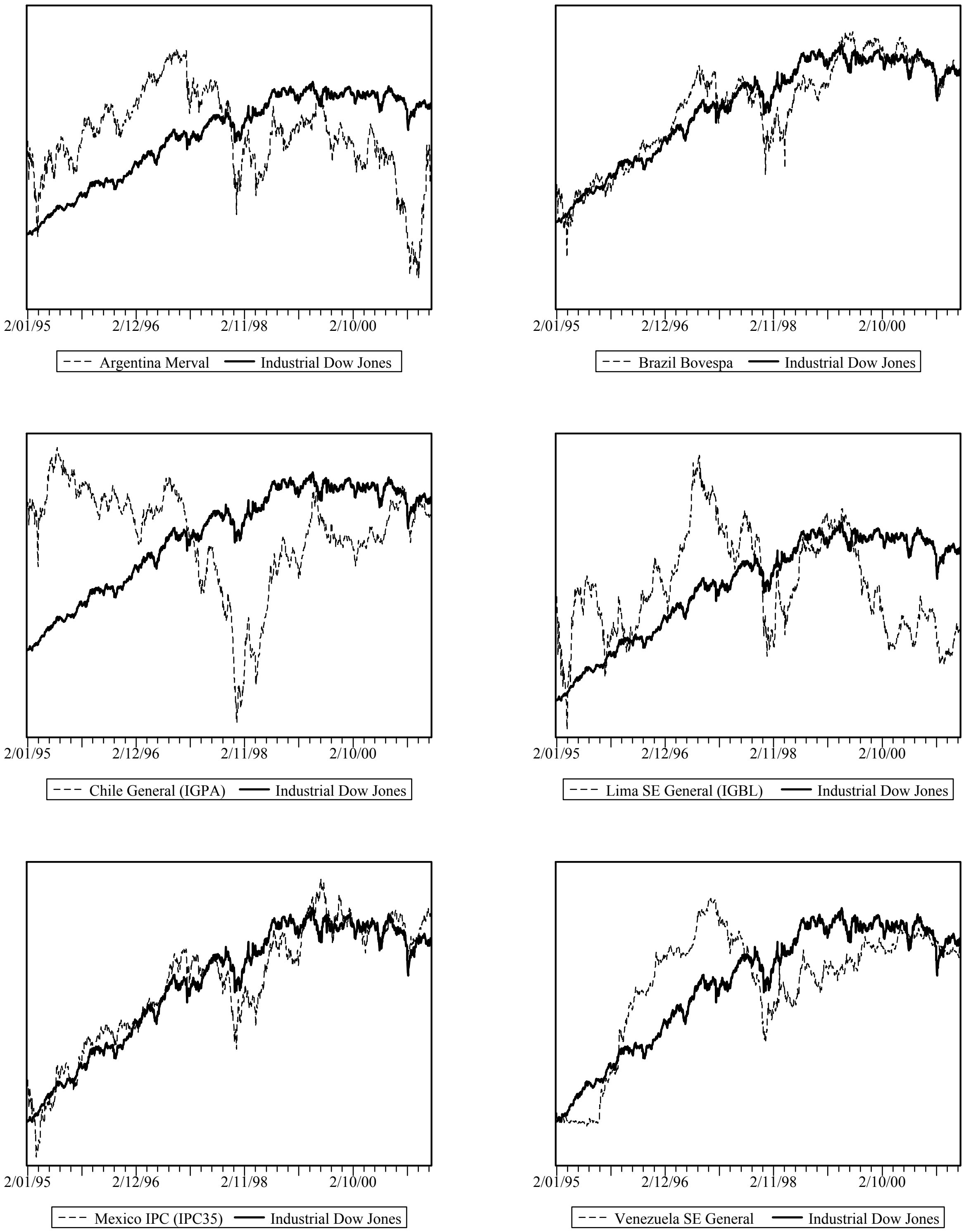

Figure 2. Levels of the indices together with USA SP500 Index

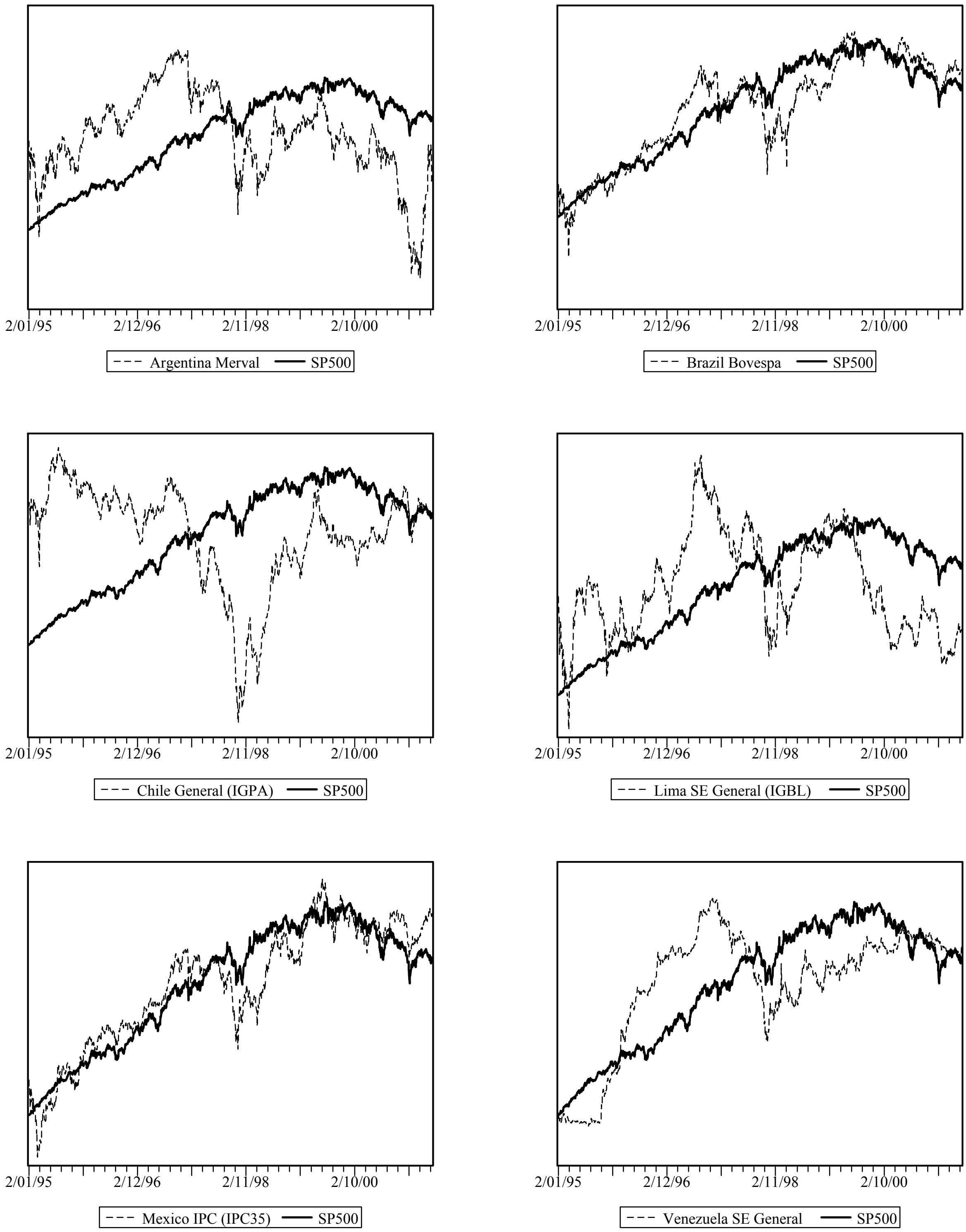

Table 1 provides summary statistics of the returns series (defined as the log-difference of the price), while Table 2 gives the correlation coefficients between the stock indices. As can be seen in Table 1, the Jarque-Bera (1980) test for joint normal kurtosis and skewness rejects the normality hypothesis. As Table 2 shows, the correlation among stock markets returns are positively and generally significantly different from zero in all cases. In addition, such correlation hardly changes when computed for Latin American and US stock markets.

Table 1
Summary statistics daily returns

	Argentina	Brazil	Chile	Lima	Mexico	USA DowJones	USA SP500	Venezuela
Mean	-0.000	0.001	0.000	0.000	0.001	0.001	0.000	0.001
Median	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Maximun	0.161	0.288	0.052	0.075	0.121	0.058	0.058	0.201
Minimun	-0.1476	-0.172	-0.038	-0.088	-0.143	-0.075	-0.071	-0.108
Std.Dev.	0.025	0.027	0.008	0.013	0.019	0.011	0.011	0.019
Skewnexx	-0.011	0.708	0.339	-0.007	0.064	-0.467	-0.249	1.097
Kurtosis	8.167	16.575	8.126	10.208	8.652	8.596	7.504	17.270
Jarque-Bera	2066.975	14421.50	2070.033	4022.475	2474.051	2491.963	1589.841	16136.91
(probability)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Autocorrelations								
1	0.106	0.046	0.297	0.221	0.110	-0.016	-0.027	0.243
2	-0.021	-0.006	0.121	-0.015	-0.042	-0.053	-0.040	-0.001
3	-0.037	-0.046	0.054	0.032	-0.007	-0.014	-0.041	-0.051
4	-0.014	-0.034	0.041	0.034	0.039	0.005	0.005	0.033
5	0.008	-0.065	0.052	0.016	-0.004	-0.029	-0.038	0.107

Table 2
Correlation coefficients between daily market returns

	Brazil	**Chile**	**Lima**	**Mexico**	**USA Dow Jones**	**USA SP500**	**Venezuela**
Argentina	0.590 (0.000)	0.398 (0.000)	0.301 (0.000)	0.477 (0.000)	0.374 (0.000)	0.377 (0.000)	0.209 (0.000)
Brazil	1	0.472 (0.000)	0.372 (0.000)	0.519 (0.000)	0.413 (0.000)	0.433 (0.000)	0.222 (0.000)
Chile		1	0.372 (0.000)	0.356 (0.000)	0.284 (0.000)	0.277 (0.000)	0.220 (0.000)
Lima			1	0.320 (0.000)	0.201 (0.000)	0.195 (0.000)	0.212 (0.000)
Mexico				1	0.479 (0.000)	0.516 (0.000)	0.219 (0.000)
USA Dow Jones					1	0.896 (0.000)	0.130 (0.000)
USA SP-500						1	0.130 (0.000)
Venezuela							1

Panel A of Table 3 reports the results for stability in the stochastic trend, while Panel B presents the results for the degree of partial integration. As can be seen, for the level variables the ADF test cannot reject the null hypothesis of a unit root in any stock market at the 5% significance level. In contrast, in the case of the first differences the ADF test always rejects the null hypothesis of a unit root. These results are corroborated by the sequential statistics Inf t_{δ} and Mean t_{δ} that are more robust than the ADF test.

Regarding the results in Panel B of Table 3, for the first differences there is not any sign of change in the order of integration. However, for the level variables, the results are no so conclusive. The results from the $\text{Sup}\, t_{\gamma 2}$, $\text{Mean}\, t_{\gamma 2}$, $\text{Sup}\, t_{\alpha 2}$ and $\text{Mean}\, t_{\alpha 2}$ statistics suggest the DJ index presents a stationary behaviour from 8 October 1998 after a global market turmoil. For the Brazil and Mexico indices, the $\text{Sup}\, t_{\gamma 2}$ y $\text{Sup}\, t_{\alpha 2}$ tests (with higher power than those based in the means) detect a change in the order of integration from 15 January 1999, showing the spillover effects from the severe currency crisis in Brazil. In the case of Venezuela, only the statistic $\text{Sup}\, t_{\gamma 2}$ suggests change in the order of integration from 10 October 1995 perhaps reflecting the effects of the Mexican crisis.

Table 3
Individual stock Indices

	Argentina		Brazil		Chile		Lima		Mexico		USA – DJ		USA – SP500		Venezuela	
	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t	Y_t	ΔY_t
Panel A: stability analysis of stochastic trend																
ADF	-2.04	-29.80*	-1.68	-30.00*	-1.68	-24.60*	-1.93	-19.82*	-1.47	-24.69*	-2.29	-43.747*	-2.23	-44.24*	-2.30	-14.94*
Inf t_δ	-2.90	-29.84*	-3.18	-30.04*	-2.23	-24.76*	-2.57	-19.90*	-2.89	-24.73*	-3.26	-43.834*	-3.43	-44.39*	-3.35	-15.26*
Ninf t_δ	19/1/0	18/8/97	13/1/9	4/7/97	3/7/	8/9/98	1/5/00	8/7/97	27/8/97	6/3/00	10/4/97	11/5/99	10/4/97	22/3/00	1/2/96	19/9/97
Mean t_δ	-2.35	-29.81*	-1.89	-30.01*	-1.71	-24.62*	-2.06	-19.83*	-1.69	-24.69*	-2.03	-43.784*	-1.83	-44.30*	-1.70	-15.03*
Sup $\|t_{\mu'}\|$	2.18	1.32	2.80	1.40	2.19	2.55	2.03	1.72	2.62	1.34	2.49	2.057	2.67	2.59	2.57	2.96
Nsup$\|t_{\mu'}\|$	19/1/0	18/8/97	13/1/9	4/7/97	11/9/9	8/9/ 98	2/3/00	8/7/97	9/9/98	6/3/00	10/4/97	11/5/99	10/4/97	22/3/00	1/2/96	19/9/97
Mean$\|t_{\mu'}\|$	1.21	0.61	0.97	0.63	0.82	0.55	0.93	0.64	0.93	0.49	0.89	1.423	1.08	1.72	0.68	1.38
Supt $_{(\mu)'}\|$	1.22	0.23	1.36	0.35	2.65	0.27	1.70	0.47	1.35	0.38	2.02	0.219	2.52	0.20	3.04	0.75
Nsup$\|t_{(\mu)'}\|$	19/8/9	11/9/98	7/7/97	13/1/99	9/9/98	21/9/ 98	9/7/97	2/2/96	7/3/00	23/10/9	12/5/99	23/10/97	22/3/00	23/10/97	22/9/97	8/2/96
Mean$\|t_{(\mu)'}\|$	0.55	0.07	0.60	0.05	0.59	0.04	0.63	0.07	0.49	0.04	1.39**	0.03	1.67	0.03	1.43**	0.09
Panel B: order of partial integration																
Sup $t_{\gamma 1}$	-2.17	-29.80*	-2.15	-30.02*	-2.69	-24.59*	-2.90	-19.82*	-1.69	-24.69*	-2.39	-43.75*	-2.32	-44.24*	-2.33	-14.94*
Nsup $t_{\gamma 1}$	13/9/9	4/1/95	22/01/	5/1/95	28/12/	4/1/95	15/9/0	12/1/95	2/6/95	6/1/95	26/5/95	2/1/95	9/1/00	2/1/91	25/1/95	19/1/95
Mean $t_{\gamma 1}$	-1.57	-16.01*	-1.18	-18.90*	-1.60	-17.44*	-1.90	-12.50*	-0.95	-14.20*	-1.15	-24.09*	-1.03	-23.72*	-1.12	-6.78*
Sup $t_{\gamma 2}$	-2.72	-30.00*	-4.49*	-30.23*	-3.88	-24.949*	-2.79	-20.78*	-3.92**	-25.03*	-4.60*	-43.75*	-3.40	-44.24*	-4.33*	-14.94*
Nsup $t_{\gamma 2}$	23/10/	10/1/95	15/1/9	10/1/95	15/1/9	10/1/ 95	14/3/9	12/1/95	15/1/99	11/1/95	8/10/98	2/1/95	11/4/97	2/1/91	10/10/95	19/01/95
Mean $t_{\gamma 2}$	-1.83	-18.82*	-1.85	-16.14*	-1.51	-13.39*	-1.32	-9.12*	-2.03	-12.97*	-2.78*	-31.81*	-2.17	-32.74*	-2.28	-8.91*
Sup $t_{\alpha 1}$	-2.14	-27.46*	-2.16	-29.31*	-2.69	-24.04*	-2.91	-19.60*	-1.57	-24.41*	-1.90	-39.72*	-1.83	-40.71*	-2.15	-14.80*
Nsup $t_{\alpha 1}$	19/1/0	22/1/01	30/1/9	17/1/01	22/2/9	8/1/01	13/6/9	19/1/01	22/12/0	18/1/01	18/10/00	18/1/01	20/12/00	18/1/01	8/12/00	19/1/01
Mean $t_{\alpha 1}$	-1.54	-21.88*	-1.10	-24.03*	-1.59	-20.12*	-1.89	-17.66*	-0.85	-20.44*	-0.95	-25.21*	-0.80	-24.51*	-0.90	-12.32*
Sup $t_{\alpha 2}$	-2.72	-28.25*	-4.49*	-28.10*	-3.88	-21.69*	-2.21	-18.88*	-3.90	-23.67*	-4.60*	-43.02*	-3.40	-43.61*	-3.67	-15.53*
Nsup $t_{\alpha 2}$	23/10/	5/2/96	15/1/9	2/2/96	15/1/9	1/2/96	5/5/00	8/2/96	15/1/99	8/2/96	8/10/98	29/1/96	11/4/97	7/2/96	12/2/96	19/2/96
Mean $t_{\alpha 2}$	-1.82	-24.43*	-1.81	-21.90*	-1.50	-16.579*	-1.26	-15.22*	-2.05	-19.83*	-2.89*	-33.50*	-2.17	-34.41*	-2.07	-13.57*

^{m} Denotes marginal significance. * Denotes significance at the 95% level. ** Denotes significance at 90% Level

Tables 4 and 5 show the results of our analysis on the existence and stability of a cointegration relationship between each Latin American index and both US indices. Table 4 reports the results corresponding to the DJ index, while Table 5 presents those obtained when using the SP500 index as representative of the US stock market. To that end, we make use of the test for cointegration proposed by Engle and Granger (1987) and the stability test in cointegrating relations suggested by Gregory and Hansen (1996). Panel A of Tables 4 and 5 report the results of these test for the whole sample, as well as the estimated correlation coefficients**.** As can be seen, in three of the six cases considered the correlation coefficients are positive and statistically significant, while for Chile we find a negative and significant correlation coefficient. In addition, the correlation coefficients obtained using the DJ index are lightly higher than those from the SP500 index. Regarding the ADF tests, the results for the DJ index suggest that we can only reject the null hypothesis of no-cointegration in the cases of Brazil and Mexico. Furthermore, the stability tests InfADF and MeanADF detect evidence of instability for the cointegrating regression between the Brazilian and DJ indices, being April 1998 the break point as suggested by the NinfADF statistics. In the case of Mexico, we only find marginal evidence of instability from the MeanADF statistic, being February 1998 the break point. When we use the SP500 index, the ADF statistic only detects cointegration in the case of Brazil and some evidence of instability from August 1998 in the case of Argentina as shown by the InfADF statistic. Therefore, there is evidence of a profound change in the second half of 1998, when Latin America started to suffer a deterioration of economic performance due to a decline of capital inflows, the resulting tightening of domestic credit conditions and a sharp deterioration in the terms of trade, coupled with the Russian financial crisis and the extreme difficulties experienced in some countries (specially in Brazil) to reduce macroeconomic unbalances.

Panel B of Tables 4 and 5 reports the results of the cointegration tests for the different subsamples. These subsamples has been obtained by dividing the sample in two taking as a breaking points those suggested by the NinfADF statistic found for each cointegrating relationship (see Panel A). For the case of the DJ index (Panel B of Table 4), there is a substantial increase in the positive correlations with respect to the results in Panel A. In the case of Argentina and Venezuela there is now evidence of a cointegrating relationship in the second subsample as indicated by all test statistics. The same is truth for Brazil and Mexico in the first subsample, while for Mexico we only marginally reject the null of no-cointegration in the second subsample with the ADF statistic. Regarding Chile, we obtain now a positive and statistically significant correlation coefficient in the second subsample, and both InfADF and MeanADF statistics also indicate cointegration between the Chilean and the DJ indices. Finally, for the case of Peru, we still fail to obtain evidence of cointegration

Table 4.
Co-integration tests (national index & Dow Jones)

		ADF	**InfADF**	**NinfADF**	**MeanADF**
Panel A: whole sample					
Argentina	-0.038	-2.068	-4.222	15/01/99 (T = 1055)	-3.263
Brazil	0.928*	-3.195*	-4.677**	27/04/98 (T = 866)	-3.604**
Chile	-0.425*	-1.888	-3.369	1/09/98 (T = 957)	-2.386
Lima	0.265*	-1.867	-3.444	25/09/98 (T = 975)	-2.699
México	0.943*	-3.043*	-4.141	24/02/98 (T = 822)	-3.398^{m}
Venezuela	0.757	-1.516	-3.770	19/03/98 (T = 839)	-2.298
Panel B: subsamples					
Argentina					
2/01/95 - 15/01/99 (T= 1055)	0.566	-0.816	-3.578		-2.269
18/01/99 - 14/02/02 (T= 804)	0.668*	-2.965*	-3.648*		-3.300
Brazil					
2/01/95 - 27/04/98 (T= 866)	0.965*	-3.851*	-4.748**		-4.076*
28/04/98 - 14/02/02 (T= 993)	0.780*	-2.940*	-4.660**		-3.411*
Chile					
2/01/95 - 1/09/98 (T= 957)	-0.644	0.133	-2.843		-0.432
2/09/98 – 14/02/02 (T= 902)	0.678*	-1.864	-4.871*		-3.834*
Lima					
2/01/95 - 25/09/98 (T= 975)	0.755*	-1.491	-3.478		-1.816
28/09/98 -14/02/02 (T= 884)	0.471*	-1.568	-4.019		-2.514
México					
2/01/95 - 24/02/98 (T= 822)	0.970*	-3.922*	-4.722**		-4.234*
25/02/98 - 14/02/02 (T= 1037)	0.787	-2.560^{m}	-3.787		-3.124
Venezuela					
2/01/95 - 19/03/98 (T= 839)	0.949	-1.198	-3.106		-2.322
20/03/98 - 14/02/02 (T= 1020)	0.496*	-2.751**	-4.483^{m}		-3.808**

^{m} Denotes marginal significance.
*Denotes significance at 95% level.
**Denotes significance at 90% level.

As for the SP500 index (Panel B of Table 5), there is a full agreement in all the statistics ADF, InfADF and MeanADF regarding the existence of a cointegrating relationship between the Argentine, Chilean and Mexican indices and the SP500 index during the second subsample and between the Brazilian and the SP500 indices during the first subsample. The MeanADF statistic also indicate marginal evidence of cointegration during the first subsample in the case of the Mexican index.

Table 5
Co-integration tests (national index & SP-500)

	A	ADF	InfADF	NinfADF	MeanADF
Panel A: whole sample					
Argentina	-0.041	-2.065	-4.633**	28/08/98 (T = 955)	-3.057
Brazil	0.908*	-2.620**	-3.665	31/10/97 (T = 740)	-3.031
Chile	-0.477*	-1.983	-3.299	23/04/99 (T = 1125)	-2.234
Lima	0.261*	-1.894	-3.617	1/09/98 (T = 957)	-2.606
México	0.921*	-2.163	-3.043	13/10/99 (T = 1248)	-2.704
Venezuela	0.703	-1.403	-3.245	4/12/97 (T = 764)	-2.213
Panel B: subsamples					
Argentina					
2/01/95 - 28/08/98 (T= 955)	0.753*	-0.092	-2.979		-1.541
31/08/98 - 14/02/02 (T= 904)	0.694*	-2.863*	-3.679		-3.188
Brazil					
2/01/95 – 31/10/97 (T= 740)	0.965*	-4.023*	-4.513[m]		-3.991*
3/11/97 – 14/02/02 (T= 1119)	0.653*	-1.914	-3.456		-2.695
Chile					
2/01/95 – 23/04/99 (T= 1125)	-0.747*	-2.042	-3.800		-2.289
26/04/99 - 14/02/02 (T= 734)	-0.373*	-3.061*	-4.313		-3.411
Lima					
2/01/95 – 01/09/98 (T= 957)	0.765*	-1.160	-2.763		-1.658
2/09/98 – 14/02/02 (T= 902)	0.596	-1.883	-3.857		-2.304
México					
2/01/95 – 13/10/99 (T= 1248)	0.892*	-2.173	-4.397		-3.496[m]
14/10/99 - 14/02/02 (T= 611)	0.510	-3.078*	-3.256		-2.805
Venezuela					
2/01/95 – 04/12/97 (T= 764)	0.946*	-1.777	-3.958		-2.671
5/12/97 – 14/02/02 (T= 1095)	0.041	-1.952	-2.931		-2.604

[m] Denotes marginal significance.
*Denotes significance at 95%level.
**Denotes significance at 90%level.

Since we have applied residual-based cointegration tests, it could be of interest to compare the results to those obtained from applying the Johansen´s VAR approach to cointegration (see Johansen, 1988, and Johansen and Juselious, 1990). As it is well known, in this approach two key issues must be addressed.

The first concerns the choice of the lag length in the VAR and the second concerns the inclusion or otherwise of deterministic trends and constants. Regarding the lag length (lg in the tables), it is found using the minimum lag that eliminates serial correlation in the residuals. With respect to the deterministic components, three cases were first considered: a constant in the short-run dynamics and a constant in the cointegrating vector (model 1), a constant in the cointegration vector and neither a constant nor a trend in the short-run dynamics (model 2), and a constant only in the short-run dynamics and neither a constant nor a trend in the cointegration vector (model 3). In each case, the rank (Rd in the tables was determined using the Johansen trace test. As can be seen in Tables 6 and 7, for all bivariate relationships under study there is evidence of lack of cointegration, except for the case of Chile with model 3. In order to explore the possibility of structural change, we included a dummy variable taking value one in the subsamples where we have previously found cointegration. This dummy variable interacts with the US index and is included as an intercept shifter. Even in this case, the results (Tables 6 and 7) fail to reject the null of no cointegration.

Finally, we also used the bounds testing procedure to the analysis of level relationships within an autoregressive distributed lag (ARDL) framework proposed by Pesaran and Shin (1999) and Pesaran *et al.* (2001), that allows the underlying variables to be either I(1) or I(0). This approach could be relevant in our case given the results obtained in Table 3. To investigate if there exists a long-run relationship between the independent variable Y_t and a regressor X_t the following conditional equilibrium correction model (ECM) is estimated:

$$\Delta Y_t = c_0 + c_1 t + \pi_y Y_{t-1} + \pi_x X_{t-1} + \sum_{i=1}^{p} \beta_i \Delta Y_{t-i} + \sum_{j=0}^{q} \delta_i \Delta X_{t-i} + \varphi w_t + u_t \quad (9)$$

where c_0 is the drift, t is a time trend and w_t is a vector of exogenous variables (e. g. dummy variables). The null hypothesis in this test is the absence of the long-run relationship between the variables defined by $H_0: \pi_y = \pi_x = 0$, against the alternative that each of the is not zero using an F-statistic [$F(\pi_y, \pi_x)$ in the tables]. However, the asymptotic distribution of this F-statistic in non-standard. The appropriate critical values can be found in Pesaran *et al.* (2001) and depend on the specification of the deterministic components. Pesaran *et al.* (2001) also provide an additional test for cointegration based on the *t*-test for $H_0: \pi_y = 0$ [$t(\pi_y)$ in the tables] suggested by Banerjee *et al.* (1998).

Table 6
Johansen Co-integration analysis (national index & Dow Jones)

	Model 1			Model 2			Model 3		
	Lg	**Trace**	**Rd**	**Lg**	**Trace**	**Rd**	**Lg**	**Trace**	**Rd**
Panel A: whole period, no dummy									
Argentina	(10)	10.350	0	(10)	15.995	0	(10)	10.350	0
Brazil	(7)	12.435	0	(7)	17.362	0	(7)	12.435	0
Chile	(3)	13.318	0	(3)	18.611	0	(3)	13.318	1
Lima	(7)	9.941	0	(7)	15.063	0	(7)	9.441	0
México	(7)	14.991	0	(7)	19.510	0	(7)	14.592	1
Venezuela	(5)	9.757	0	(5)	15.559	0	(5)	9.757	0
Panel B: whole period, with dummy									
Argentina	(10)	7.138	0	(10)	7.332	0	(10)	6.249	0
Brazil	(7)	14.938	0	(7)	14.469	0	(7)	9.394	0
Chile	(3)	14.319	0	(3)	15.79	0	(3)	7.829	0
Lima	(7)	12364	0	(7)	11.253	0	(7)	6.437	0
México	(7)	15.275	0	(7)	16.107	0	(7)	11.259	0
Venezuela	(5)	14.739	0	(5)	14.508	0	(5)	5.377	0

Table 7
Johansen Co-integration analysis (national index & Dow Jones)

	Model 1			Model 2			Model 3		
	Lg	**Trace**	**Rd**	**Lg**	**Trace**	**Rd**	**Lg**	**Trace**	**Rd**
Panel A: whole period, no dummy									
Argentina	(10)	11.999	0	(10)	17.556	0	(10)	11.999	0
Brazil	(7)	11.523	0	(7)	16.314	0	(7)	11.523	0
Chile	(3)	14.359	0	(3)	19.661	0	(3)	14.898	1
Lima	(7)	11.922	0	(7)	16.800	0	(7)	11.922	0
México	(7)	12.407	0	(7)	17.225	0	(7)	12.407	0
Venezuela	(5)	9.511	0	(5)	14.656	0	(5)	9.512	0
Panel B: whole period, with dummy									
Argentina	(10)	8.344	0	(10)	9.178	0	(10)	5.959	0
Brazil	(7)	13.232	0	(7)	13.183	0	(7)	7.723	0
Chile	(3)	10.643	0	(3)	12.271	0	(3)	7.191	0
Lima	(7)	13.455	0	(7)	11.889	0	(7)	6.311	0
México	(7)	10.184	0	(7)	12.121	0	(7)	7.605	0
Venezuela	(5)	11.439	0	(5)	12.549	0	(5)	4.752	0

As can be seen in Tables 8 and 9, for most bivariate relationships under study the computed F-statistics are found to fall below the critical values, indicating that we cannot reject the null hypothesis of no cointegration. The only exceptions are Brazil [where the $t_I(\pi_y)$ statistic suggest cointegration for the whole sample, while the $F_I(\pi_y, \pi_x)$ statistic is inconclusive) and Mexico (where the $t_I(\pi_y)$ statistic is inconclusive]. It is interesting to note that these cases roughly correspond to those found in Tables 4 and 5. To investigate the possibility of structural change, we included a dummy variable taking value one in the subsamples where we have previously found cointegration. In this case of the SP500, all the statistics [except $F_I(\pi_y, \pi_x)$] indicate the existence of a long-run relationship between Venezuelan and the DJ indices when we include a dummy variable. In addition, the $t_I(\pi_y)$ statistic suggest some evidence of cointegration for Brazil, being inconclusive for Mexico (Table 8). For the DJ index, we find evidence of cointegration for Brazil [as indicated by the $t_I(\pi_y)$ statistic] and Mexico [as indicated by the $t_{III}(\pi_y)$ statistic] (Table 9).

The results from Tables 4 and 9 suggest that conventional cointegration tests that do not allow for changes in regime might lead to biases when testing for the null hypothesis of no cointegration in favour of acceptance. This in turns illustrates how the formal consideration (through adequate statistical procedures) of eventual structural breaks may be useful for a more correct specification of an econometric model.

Table 8
Bounds testing for cointegration analysis (National index & Dow Jones)

	No intercept and no trend				Unrestricted intercept and no trend				Unrestricted intercept and no trend			
	p	q	$F_I(\pi_y, \pi_x)$	$t_I(\pi_y)$	p	q	$F_{III}(\pi_y, \pi_x)$	$t_{III}(\pi_y)$	p	q	$F_V(\pi_y, \pi_x)$	$t_V(\pi_y)$
Panel A: whole period, no dummy												
Argentina	1	2	2.3904[a]	-1.9338[a]	1	2	2.3891[a]	-1.9332[a]	1	2	2.5634[a]	-2.1385[a]
Brazil	1	4	3.5224[b]	-2.6153[c]	1	4	3.5289[a]	-2.6173[a]	1	4	3.5202[c]	-2.6498[a]
Chile	1	1	1.4069[a]	-1.5078[a]	1	1	1.4060[a]	-1.5071[a]	1	1	1.2278[a]	-1.5180[a]
Lima	3	1	1.5593[a]	-1.7279[a]	3	1	1.5578[a]	-1.7272[a]	3	1	3.8369[a]	-2.7385[a]
México	2	1	2.9934[a]	-2.4013[b]	2	1	2.9882[a]	-2.3987[a]	2	1	4.4884[a]	-2.8382[a]
Venezuela	3	3	1.5220[a]	-1.5571[a]	3	3	1.5263[a]	-1.5585[a]	3	3	1.1843[a]	-1.5382[a]
Panel B: whole period, with dummy												
Argentina	1	2	1.8035[a]	-1.6230[a]	1	2	2.4304[a]	-1.2978[a]	1	2	1.8035[a]	-1.6230[a]
Brazil	1	4	4.0027[b]	-2.8278[c]	1	4	4.3574[a]	-2.9478[b]	1	4	5.0870[a]	-3.1887[a]
Chile	1	1	1.0807[a]	-1.4617[a]	1	1	2.2218[a]	-1.3281[a]	1	1	0.4885[a]	-0.1789[a]
Lima	3	1	1.5292[a]	-1.6357[a]	3	1	1.6504[a]	-1.5990[a]	3	1	3.1242[a]	-2.4945[a]
México	2	1	3.3251[b]	-2.5763[b]	2	1	3.0721[a]	-2.3810[a]	2	1	5.4875[a]	-3.1587[a]
Venezuela	3	3	2.3067[a]	-2.1303[c]	3	3	6.0799[a]	-3.4141[c]	3	3	7.6014[c]	-3.8728[c]

Notes: $F_I(\pi_y, \pi_x)$ and $t_I(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9) when $c_0 = c_1 = 0$.

$F_{III}(\pi_y, \pi_x)$ and $t_{III}(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9) when $c_1 = 0$.

$F_V(\pi_y, \pi_x)$ and $t_V(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9).

[a] indicates that the statistic lies below the 0.05 bound.

[b] indicates that the statistic lies within the 0.05 bound.

[c] indicates that the statistic lies above the 0.05 bound.

Table 9
Bounds testing for cointegration analysis (National index & SP500)

	No intercept and no trend				Unrestricted intercept and no trend				Unrestricted intercept and no trend			
	p	q	$F_I(\pi_y, \pi_x)$	$t_I(\pi_y)$	p	q	$F_{III}(\pi_y, \pi_x)$	$t_{III}(\pi_y)$	p	q	$F_V(\pi_y, \pi_x)$	$t_V(\pi_y)$
Panel A: whole period, no dummy												
Argentina	1	2	2.4460[a]	-1.8877[a]	1	2	2.4446[a]	-1.8872[a]	1	2	2.3027[a]	-1.9613[a]
Brazil	1	4	3.5677[b]	-2.6369[c]	1	2	3.5697[a]	-2.6375[a]	1	2	3.6417[a]	-2.6971[a]
Chile	1	1	1.3970[a]	-1.5324[a]	1	1	1.3962[a]	-1.5318[a]	1	1	2.3332[a]	-1.6260[a]
Lima	3	1	1.7918[a]	-1.7521[a]	3	1	1.7903[a]	-1.7515[a]	3	1	1.3332[a]	-1.8971[a]
México	2	1	2.1775[a]	-2.0154[a]	2	1	2.1706[a]	-2.0116[a]	2	1	4.4468[a]	-2.8365[a]
Venezuela	3	3	1.4874[a]	-1.4995[a]	3	3	1.4916[a]	-1.5006	3	3	1.2428[a]	-1.5414[a]
Panel B: whole period, with dummy												
Argentina	1	2	2.4650[a]	-1.6784[a]	1	2	2.6423[a]	-0.6770[a]	1	2	1.3899[a]	-0.8118[a]
Brazil	1	4	3.9377[a]	-2.8031[c]	1	2	3.8434[a]	-2.7445[a]	1	2	4.4964[a]	-2.9964[a]
Chile	1	1	1.6201[a]	-1.7683[a]	1	1	3.5180[a]	-2.6477[a]	1	1	3.8501[a]	-2.7733[a]
Lima	3	1	1.8021[a]	-1.5538[a]	3	1	1.8473[a]	-1.0078[a]	3	1	1.2115[a]	-1.1855[a]
México	2	1	1.8227[a]	-1.8936[a]	2	1	4.7194[a]	-2.9688[c]	2	1	5.2940[a]	-3.1355[a]
Venezuela	3	3	1.9968[a]	-1.8532[a]	3	3	4.8046[a]	-2.4799[a]	3	3	5.8905[a]	-3.2216[a]

Notes: $F_I(\pi_y, \pi_x)$ and $t_I(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9) when $c_0 = c_1 = 0$.

$F_{III}(\pi_y, \pi_x)$ and $t_{III}(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9) when $c_1 = 0$.

$F_V(\pi_y, \pi_x)$ and $t_V(\pi_y)$ are, respectively, the F-statistic for testing $\pi_y = \pi_x = 0$ and $\pi_y = 0$ in equation (9).

[a] indicates that the statistic lies below the 0.05 bound.

[b] indicates that the statistic lies within the 0.05 bound.

[c] indicates that the statistic lies above the 0.05 bound

4. Concluding remarks

In this paper we have provided some new evidence on the relationship between the US and Latin American stock markets, using daily data covering 2 January 1995- 14 February 2002 period. We use both the Dow Jones (DJ) and the Standard and Poor's 500 (SP500). We depart from previously published papers by making use of recently-developed cointegration techniques that allow for structural shifts in the cointegration vector.

Our results suggest that, if we apply cointegration tests without structural breaks, we only find cointegration in the cases of Brazil and Mexico for the DJ index and Brazil for the SP500 index. In contrast, if we introduce the possibility of structural breaks, we find strong evidence in favour of such relationship between the Argentine, Chilean and Venezuelan indices and the DJ index after the 1998 financial turmoil, and between the Brazilian and Mexican indices and the DJ index before such turbulence, while some marginal cointegration is detected between the Mexican and DJ indices in from February 1998. Additionally, we find evidence of a cointegrating relationship between the Argentine, Chilean and Mexican indices and the SP500 index from August 1998, April 1999 and October 1999, respectively and between the Brazilian and the SP500 indices before November 1997, as well as some marginal cointegration between the Mexican and SP500 indices before October 1999.

Therefore, the analysis carried out in this paper has provided some evidence in favour of modelling the long-run relationship between US and Latin American stock markets using an evolving formulation that formally considers eventual structural breaks rather than the conventional specification. In particular, the empirical results obtained in this paper suggest that the 1997-98 global financial crisis has had a profound impact in the long-run common trends existing between these stock markets. This crisis started in the Southeast Asian countries in 1997, led to the Russian default of August 17, 1998, culminating with the floatation of the Brazilian real on January 15, 1999, after a series of speculative attacks in 1998

The evidence of cointegration between the Latin American and US stock markets provided in this paper would indicate the presence of common factors reducing the amount of independent variation among them. This in turn would imply that, although it is still possible in this region to derive portfolio diversification in the short run. As a result, the gains from international diversification for investors with long holding periods may be limited.

On the other hand, although Granger (1986) suggests a certain link between cointegration and inefficiency in asset markets, Caporale and Pittis (1998) argue that cointegration links among stock markets should be interpreted as evidence of predictability, without referring to the question of market efficiency. Indeed, Dwyer and Wallace (1992), Lien (1996) and Masih and Masih (1999) contend that the existence of cointegration does not necessarily contradict the notion of market efficiency, once the latter is defined as the lack of arbitrage opportunities (Fama, 1991).

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